DO SHORT-SELLING REGULATIONS MATTER?

Evidence from Europe

GIANNOULA KARAMICHAILIDOU

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Abstract

The Lehman Brothers bankruptcy on September 15, 2008 led to turmoil in financial markets, sharp declines in the stock prices of financial institutions, and fears for the stability of the global financial system. This prompted financial regulatory authorities in many countries to enforce bans on the short-selling of stocks, mainly of financial institutions. The aim of the short-selling bans was to halt further stock price declines and restore the confidence in the financial system. This study examines the impact of the 2008 short-selling regulations on (i) market liquidity, (ii) the components of the bid/ask spread, (iii) the speed of price adjustment to new information, and (iv) the spot-futures informational dynamics in Europe over the period January 2008–December 2009. The main purpose of this study is to provide new empirical evidence on the consequences of the short-selling measures in order to assist financial authorities to design more efficient and effective short-selling policies in the future. Overall, the results show that (i) liquidity deteriorates, (ii) both informational and non-informational components of the bid/ask spread increase, and (iii) the speed of information transmission is faster for the European financial stocks subject to total (naked and covered) short-selling bans with listed derivatives. In contrast, there is no evidence that short-selling bans affected the informational linkage between the spot and futures markets in Europe at the Banks super-sector level.

For Milada and Sotiria

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TABLE OF CONTENTS

CHAPTER 4 Short-selling regulations and the components of the bid/ask spread

CHAPTER 5 Short-selling regulations and the speed of price adjustment to new information

CHAPTER 6 Short-selling regulations and the spot-futures informational dynamics

List of Tables

List of Figures

CHAPTER 1

Introduction

1.1 Overview

Short-selling is an indispensable part of trading in financial markets. Nevertheless, attempts to restrict this trading technique go as far back as the establishment of the first stock exchanges (Bris, Goetzmann, and Zhu (2007)). Since short sellers profit from price declines, most short-selling regulations have emerged after financial crises and have aimed at stabilizing prices and calming the markets (FSA (2009)). This study examines the impact of the short-selling regulations that were introduced in Europe in the autumn of 2008 on four dimensions of market microstructure: (i) liquidity, (ii) components of the bid/ask spread, (iii) speed of price adjustment to new information, and (iv) spot-futures informational dynamics.

In general, empirical evidence shows that short-selling provides markets with liquidity (Bris, Goetzmann, and Zhu (2007)), increases the speed with which new information is incorporated into prices (Chen and Rhee (2010)), enhances the price discovery process (Fung and Jiang (1999); Jiang, Fung, and Cheng (2001)), and consequently improves pricing efficiency (Boehmer and Wu (2013)). Nevertheless, there is a strand of literature arguing that aggressive short-selling can drive stock prices below their fundamental values, thus supporting government intervention (Shkilko, Van Ness, and Van Ness (2012); Blau, Van Ness, Van Ness, and Wood (2010)).

The bankruptcy of Lehman Brothers in mid-September of 2008 triggered sharp declines in the stock prices of financial institutions and prompted the regulatory bodies to introduce short-selling regulations in many markets. The regulators were worried that the tumbling share prices of already troubled financial institutions were sending the wrong signals to investors. The rationale behind the intervention of the regulators was to restrict the sharp downward movement in share prices and restore confidence in the financial system.

The main purpose of this study is to provide empirical evidence that illustrates the effects of the 2008 short-selling regulations on European markets in order to assist regulators to design more efficient and effective short-selling policies in the future. Because short-selling regulations were applied mainly to the stocks of financial institutions, this study focuses on investigating the impact of these short-selling regulations on the stocks of major European financial institutions. However, a control sample of stocks of non-financial firms that were not subject to the new short-selling rules is also used in the analysis to control for general market movements.

Overall, the results show that short-selling regulations were associated with (i) a deterioration in the liquidity of the stocks subject to these regulations and (ii) an increase in order processing, inventory holding, and information asymmetry costs for the market makers who traded these stocks. In particular, the results show that the impact of the short-selling regulations was greater on the order processing and inventory holding costs (non-informational component of the bid/ask spread) relative to the information asymmetry costs (informational component of the spread). Moreover, empirical evidence shows that the trading of derivatives that were written on the financial stocks subject to the short-selling regulations worsened the impact of these regulations on the liquidity of these financial stocks.

With regard to information efficiency, the results show that information is incorporated more slowly into prices of the stocks subject to the short-selling regulations. However, these results hold only for stocks without listed derivatives. In contrast, it is found that information is incorporated faster into the prices of stocks subject to the short-selling regulations with listed derivatives. Finally, at an index level, there is no evidence showing that tight (relaxed) short-selling regulations are associated with a smaller (greater) contribution from the spot market to the price discovery process in the European Banks super-sector. These findings should be interpreted under the caveat that there may be other forces in the market driving the spot-futures informational dynamics in Europe during the sample period.

The remainder of chapter 1 is organized as follows. Section 1.2 discusses the motivation and contribution of this study while section 1.3 summarizes the contents of the following six chapters.

1.2 Motivation and Contribution

Prior to September 2008, most European countries did not have specific regulations on the short-selling of stocks. However, after the collapse of Lehman Brothers, many European financial regulatory authorities adopted various forms of short-selling rules. For example, some countries banned naked short-selling of either financial stocks (e.g., Belgium France, and Portugal) or all stocks (e.g., Austria). Other countries prohibited both naked and covered short-selling of either financial stocks only (e.g., Denmark, Ireland, the UK, and Norway) or all listed stocks (e.g., Italy).[1] This study is motivated by the range of the various short-selling rules and the consequent range of the potential effects of these rules on the aspects of liquidity, information efficiency, and price discovery in European markets that have yet to be fully explored. Understanding the consequences of the new short-selling measures would contribute to the design of more efficient and effective short-selling policies in the future, particularly since European authorities are already showing a willingness to enforce a Pan-European short-selling regime (CESR (2010)).

Liquidity is an essential factor in financial markets. Prior studies find that short-selling regulations are associated with higher bid/ask spreads on stocks subject to these regulations (see for example, Bris (2008), Boulton and Braga-Aves (2010), Boehmer, Jones, and Zhang

[1] See Clifford Chance (2009c).

(2013), and Beber and Pagano (2013)). However, liquidity has two dimensions: a price dimension measured by the bid/ask spread and a quantity dimension measured by depth (Lee, Mucklow, and Ready (1993)). By considering both dimensions, this study provides a more complete picture of the impact of the short-selling regulations on liquidity. Because liquidity is only considered to be unambiguously reduced when an increase in bid/ask spreads is accompanied by a decrease in depth (Lee, Mucklow, and Ready (1993)). In addition, this study extends prior research by incorporating two measures of depth into the analysis along with the bid/ask spreads on the stocks of major European financial institutions.

The main hypothesis is that the liquidity of financial stocks that were subject to the short-selling regulations deteriorates after the introduction of these regulations. Overall, the results support this hypothesis. It is found that the short-selling regulations lead to an increase in the bid/ask spreads and a decrease in both measures of depth. A second testable hypothesis is that the impact of the short-selling regulations is stronger the more severe these regulations are. Again, the results are in line with this hypothesis by showing that liquidity deteriorates more for the financial stocks that were subject to more stringent short-selling regulations. A third hypothesis is that the trading of derivatives written on financial stocks subject to the short-selling regulations has no effect on the impact of short-selling regulations on the liquidity of the financial stocks. In this case, empirical evidence does not support this hypothesis. The results show that the presence of derivatives is associated with greater deterioration in the liquidity of the financial stocks that were subject to the short-selling regulations relative to the liquidity of the financial stocks subject to the short-selling regulations absent listed derivatives.

The price dimension of liquidity or the bid/ask spread is comprised of two components – a non-informational component and an informational component (Bessembinder and Kaufman

(1997)). Prior studies find that bid/ask spreads on stocks that were subject to the short-selling regulations increase significantly following the enactment of these regulations. This study contributes to the literature by further decomposing the bid/ask spread into its non-informational and informational components. The main objective of this decomposition is to identify the channel through which bid/ask spreads increased during the period when short-selling regulations were in force in Europe. In a related study, Boehmer, Jones, and Zhang (2013) find that both bid/ask spread components increase significantly during the short-selling ban in the US. The main testable hypothesis here is that the non-informational and informational components of the bid/ask spread increase during the short-selling bans in Europe. Overall, the results provide supporting evidence for this hypothesis. More specifically, empirical evidence shows that the impact of the short-selling regulations is greater on the non-informational component relative to the informational component.

Another crucial factor in financial markets is information. Theory predicts that short-selling constraints will reduce the speed of price adjustment to new information, especially if the news is bad (Diamond and Verrecchia (1987)). Prior studies find that relaxing short-selling regulations increases the speed with which new information is incorporated into prices thereby improving the information efficiency of the price mechanism (see for example, Bris, Goetzmann, and Zhu (2007), Chen and Rhee (2010), and Boehmer and Wu (2013)). This study adds to the literature by using a new dataset associated with the introduction of the 2008 short-selling regulations to test Diamond and Verrecchia’s (1987) theoretical prediction in a European context. Contrary to prior research, this study also analyses the specific effect of a tightening in short-selling regulations on the speed of price adjustment to new information. Hence, the two hypotheses here are (i) the introduction of the short-selling regulations reduces the speed of price adjustment to new information for the stocks subject to these regulations and (ii) the speed of price adjustment to new information is slower for stocks subject to more severe short-selling regulations. Overall, the results show that the speed of price adjustment to new information increases for the financial stocks subject to the short-selling regulations following the enforcement of these regulations. These findings contradict the theoretical predictions of Diamond and Verrecchia (1987). Also, the results do not support the hypothesis that the speed of price adjustment to new information is slower for stocks subject to more severe short-selling regulations. Empirical evidence shows that the speed of price adjustment to new information is faster for stocks that were subject to more stringent short-selling regulations. Furthermore, the results show that the speed of price adjustment to new information is faster for stocks subject to the short-selling regulations with listed derivatives relative to stocks subject to the short-selling regulations without listed derivatives. These results, again, are not in line with prior expectations.

Another theoretical prediction of the Diamond and Verrecchia's (1987) model is that short-selling constraints in the spot market will reduce the contribution from that market to the price discovery process. In line with this, prior studies find that relaxing short-selling regulations increases the contribution from the spot market to price discovery and strengthens the contemporaneous informational relationship between the spot and futures markets (see for example, Fung and Jiang (1999) and Jiang, Fung, and Cheng (2001)). This study extends prior research by investigating the impact of tightening and relaxing short-selling regulations on the spot-futures informational dynamics in Europe at a super-sector level. In particular, this study tests the hypothesis that the contribution from the European Banks super-sector index to price discovery will decrease (increase) after the introduction (relaxation) of the short-selling regulations that affected the components of that index. No empirical evidence is found showing that the information share of the Banks super-sector index decreased following the introduction of the short-selling regulations. This result too contradicts Diamond and Verrecchia (1987) theoretical predictions.

An examination of the spot-futures informational dynamics for other European super-sectors, the components of which were not affected by the short-selling regulations, reveals that there were other factors driving the spot-futures informational dynamics during the sample period.

In summary, this study contributes to the literature by providing additional empirical evidence on the impact of the 2008 short-selling regulations on (i) both dimensions of market liquidity, (ii) the components of the bid/ask spread, (iii) the speed of price adjustment to new information, and (iii) the price discovery between spot and futures markets in a European context.

1.3 Thesis outline

After describing the institutional background of short-selling regulations in Europe (*Chapter 2*), this study proceeds with four empirical essays (*Chapters 3–6*). The sample for all empirical essays spans the period from mid-January 2008 to mid-December 2009. The stocks of 78 financial institutions from 10 European countries comprise the sample for the first three empirical essays.[2] To control for market movements during the sample period, this study constructs a control sample of 78 non-financial stocks that were not subject to the short-selling regulations. This control sample is used in the analysis of the first three empirical essays. In contrast, the last essay conducts an analysis at an index level rather than an individual stock level as described below.

Chapter 2 provides the institutional background for this study. Firstly, it explains the mechanics of short-selling and, secondly, it discusses the rationale behind the introduction of

[2] These countries are Belgium, Denmark, France, Germany, Ireland, the Netherlands, Portugal, Spain, Switzerland, and the UK.

the short-selling regulations. Then *Chapter 2* defines and analyses the various short-selling measures that were implemented in each of the 10 European countries in the sample.[3]

Chapter 3 (Essay 1) examines the impact of the short-selling regulations on the liquidity of the European financial stocks that were subject to these regulations. The analysis is conducted at both the country and aggregate level conditional on the stringency of the short-selling regulations as well as the presence of listed derivatives. The univariate analysis is conducted by using the measures of bid/ask spreads that were proposed by Stoll (2000) and measures of depth that were proposed by Heflin, Shaw, and Wild (2005). Boehmer, Jones, and Zhang's (2013) approach is used to conduct the multivariate analysis, which incorporates control variables that have been shown in the literature to affect liquidity.

Chapter 4 (Essay 2) investigates the effects of the short-selling regulations on the components of the bid/ask spread. In particular, the bid/ask spread is decomposed into its two components – the non-informational component, which is measured by the realized spread, and the informational component, which is measured by the price impact – following Bessembinder and Kaufman (1997). Again, the multivariate regression framework of Boehmer, Jones, and Zhang (2013) is employed to control for other factors that have been found in the literature to affect the bid/ask spread components. The analysis is conducted at both the country and aggregate level.

Chapter 5 (Essay 3) examines the impact of the short-selling regulations on the speed of price adjustment to new information for the European financial stocks that were subject to these regulations. The bivariate VAR model, which was proposed by Hasbrouck (1991), is applied to estimate the speed of price adjustment to new information. As in *Chapter 3*, the analysis is

[3] Additionally, Chapter 2 provides details of the short-selling rules that were introduced in two more countries – Italy and Norway – that are analysed in the Appendix 5A of Chapter 5.

conducted at country and aggregate levels conditional on the severity of the short-selling regulations and the trading of derivatives.

Chapter 6 (Essay 4) looks at the impact of the 2008 short-selling regulations on the informational dynamics between the Dow Jones STOXX 600 super-sector indexes and their associated futures contracts, focusing mainly on the European Banks super-sector index. The approach of Hasbrouck (1995) is used to estimate the information share of the spot market in the price discovery process under tight and relaxed short-selling regulations.

Lastly, *Chapter 7* discusses the findings and the limitations of this study. *Chapter 7* also concludes with proposals for future research on the topics examined in this study.

CHAPTER 2

Short-selling regulations: Institutional Background

2.1 Introduction

The collapse of Lehman Brothers on September 15, 2008 led to a massive loss of confidence in the financial system, particularly in the banking system, and a rapid decline in the price of bank stocks. One immediate consequence was the introduction of short-selling regulations in many developed economies. The aim of this chapter is to (i) explain the mechanics of short-selling, (ii) discuss the rationale behind the adoption of the 2008 short-selling rules, (iii) define the various short-selling measures, and (iv) illustrate the specific features of the short-selling regulations that were applied in 12 European countries.[4] These countries comprise the sample employed in the subsequent empirical chapters.[5]

Chapter 2 is organized as follows. Section 2.2 defines short-selling and describes how it works. Section 2.3 reviews the arguments in the academic literature for and against restrictions on short-selling. Section 2.4 introduces the various short-selling rules that were implemented in the autumn of 2008 and describes the specific short-selling rules that were applied in the sample countries. Section 2.5 summarises the main dimensions of the new short-selling rules and concludes.

2.2 How does short-selling work?

Short-selling is defined as a sale of a security that the seller does not own. There are two types of short-selling: “covered” and “naked”. The main difference between covered and naked short-selling involves the stock lending process. More specifically, in covered short-selling, the seller borrows the shares from a long-term shareholder, such as an insurance company or a pension fund (through her broker) paying a borrowing fee in order to deliver the stock to the buyer on the settlement date. On delivery of the stock, the seller receives the

[4] These countries are Belgium, Denmark, France, Germany, Ireland, Italy, the Netherlands, Norway, Portugal, Spain, Switzerland, and the UK.
[5] Italy and Norway are only analyzed in chapter 5. Their results are presented in Appendix 5A.

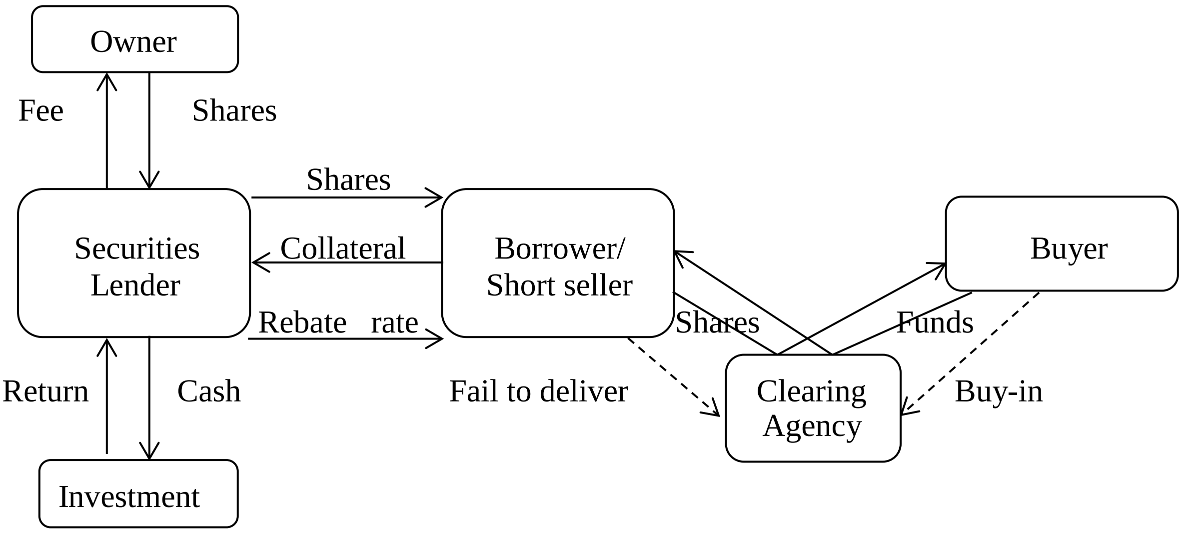
sale proceeds from the buyer. Later, before returning the shares to the original lender, the seller must repurchase the borrowed shares on the market. Naked short sellers, on the other hand, do not arrange to borrow any shares from the stock lending market to settle the transaction (Gruenewald, Wagner, and Weber (2009b)). Frequently, naked short-selling is used by market makers in intra-day trading when they do not have enough supply to meet their clients' demand. However, naked short-selling can be manipulative, and this is the type of short-selling regulators feared the most and attempted to curtail by introducing short-selling rules in late 2008. The problem associated with manipulative naked short-selling is that it injects additional shares of the firm targeted by the short-sellers/speculators to the market. This causes an artificial increase in the supply of the particular shares since the sale of these shares is not backed by borrowed shares from another shareholder. The artificial increase in the shares' supply in turn leads to a subsequent drop in the price of the shares (FSA (2009)).

Figure 2.1 depicts the mechanics of the stock lending and short-selling process.[6] The parties involved in the stock lending process are: (i) the securities lender that acts on behalf of the shares' owner, (ii) the borrower or short seller, and (iii) the buyer. Once the lender and the buyer have been located, the short seller settles the short sale by delivering the borrowed shares to a clearing agency. If the short seller fails to deliver the shares, the buyer may instruct the clearing agency to locate and buy back the shares. This practice is known as a "buy-in" (Gruenewald, Wagner, and Weber (2009a)). In order to borrow the shares, the short seller pays the lender cash collateral. Consequently, if the short seller fails to return the shares, the lender can cover the loan by using the collateral to buy the shares on the open market. Also, while the stock is on loan, the lender invests the collateral and receives interest on the investment. Part of the interest received, called the rebate rate, is returned to the short

[6] See Gruenewald, Wagner, and Weber (2009a: p.5).

seller. When the demand to sell the shares of a stock short exceeds the number of shares available to borrow, the stock is said to be "on special". This might result in a negative rebate rate. On such an occasion, the short seller actually pays a premium to the lender in order to borrow the shares. In the end, the lender earns the spread between the current market interest rate and the rebate rate that she pays the short seller on the collateral.[7]

Figure 2.1 Stock lending and short-selling process



Overall, there are important risks associated with both covered and naked short-selling practices that may result in significant losses. Firstly, in the case of covered short-selling, the borrowed shares could be recalled by the lender at any time unless there is an agreement between the two parties excluding this possibility (Fabozzi (2004: Chapter 2). Secondly, when trading in a stock that is illiquid, or many short sellers close their positions at the same time, short sellers might not be able to find shares to buy back and close out their position. This is known as a short squeeze. When a short squeeze takes place, short sellers might eventually be forced to pay much higher prices than the price they received when they sold the shares short in the first place (Taulli (2004: Chapter 4). On the other hand, naked short-

[7] See Fabozzi, F. (2004), Short-selling strategies, risks, and rewards, Chapter 2.

selling bears settlement risk; i.e., failure to deliver the promised shares to the buyer when delivery is due. This can potentially affect the proper functioning of the market and lead to greater transaction costs and lower levels of trading. Finally, both covered and naked short-selling are subject to the risk of an unexpected rise in the price of the stock sold short (FSA (2009)).

2.3 The rationale behind the introduction of short-selling regulations

2.3.1 Introduction

> "While we still regard short-selling as a legitimate investment technique in normal market conditions, the current extreme circumstances have given rise to disorderly markets. As a result, we have taken this decisive action, after careful consideration, to protect the fundamental integrity and quality of markets and to guard against further instability in the financial sector" (Hector Sants – chief executive of the Financial Services Authority in the UK – September 19, 2008).[8]

Financial regulators feared that short-selling would drive, or more importantly, induce further share price declines of already troubled major financial institutions, mostly banks, during the Global Financial Crisis. This is because the market was using bank stock prices as a proxy measure of credit risk. Consequently, in an environment of great uncertainty about which banks held the most valueless assets, a tumbling stock price was taken as a signal that the bank concerned had a lot of valueless assets on its books. Once this perception prevailed, other banks would be reluctant to lend to a bank involved in the interbank market since this form of lending is typically uncollateralised. This led to a "freeze" in the interbank market. If nothing had been done, this would have impaired the functioning of the financial system and could have had a damaging effect on economies with the supply of short-term credit drying up (Brunnermeier (2009)). Therefore, financial regulators attempted to limit the downward

[8] http://www.telegraph.co.uk/finance/newsbysector/banksandfinance/2987827/Market-turmoil-FSA-bans-speculators-from-short-selling.html

pressure caused by short-selling and the consequent downside speculation with the main objective of restoring confidence in the financial system.

In general, short-selling is considered to be a legitimate trading practice that (i) enhances the price discovery process, (ii) increases market liquidity, (iii) facilitates hedging and risk management activities, and (iv) mitigates market bubbles. However, according to Financial Regulatory and Supervisory Authorities, in very volatile market conditions, short-selling can (i) be abusive, pushing down the prices of financial instruments to a level that is not justified by the fundamentals, (ii) contribute to disorderly markets, leading to a massive long selling as a result of the overreaction of panicked investors to the short-sellers' signals or due to settlement failures, and (iii) damage financial stability and confidence in the financial system.[9] Under these conditions, regulatory interventions are justified in order to calm the markets and restore confidence in the financial system. As Sir Callum McCarthy – the Financial Services Authority's outgoing chairman in the UK – stated on the short-selling ban:

> "This is a measure which reflects the present turbulence in markets. It is designed to have a calming effect – something which the equity markets for financial firms badly need." [10]

2.3.2 Review of academic literature for and against short-selling restrictions

Opponents of short-selling argue that short sellers manipulate prices. Brunnermeier and Pedersen (2005) develop a model to investigate the dynamics of predatory trading. They show that when large investors are forced to unwind their position, other strategic investors trade in the same direction, which increases the costs of liquidation and leads to prices overshooting. The authors argue that in such cases, coordinated action by regulators is justified. Shkilko, Van Ness, and Van Ness (2012) provide empirical evidence which

[9] See ISLA report (2009) and FSA (2009).
[10] http://www.ft.com/cms/s/0/16102460-85a0-11dd-a1ac-0000779fd18c.html#axzz2Nvffzlto

supports this model. The authors find that aggressive short-selling can occasionally be accompanied by more aggressive long-selling, and both lead to price declines. Moreover, Blau, Van Ness, Van Ness, and Wood (2010) find that in volatile markets, short sellers are momentum traders who tend to follow negative returns on down days and positive returns on up days, thus exaggerating price movements. These findings also suggest that in some cases regulations on short-selling are needed.

Other theoretical papers show that short-selling constraints lead to bubble formation and subsequent stock price crashes. Scheinkman and Xiong (2003), for example, develop a model with short-selling constraints in which overconfidence generates disagreements among agents concerning asset fundamentals. The authors argue that short-selling constraints contribute to bubble creation. Hong and Stein (2003) present a theoretical framework of market crashes based on differences of opinion and the presence of short-selling constraints. In their model, there are two investors that are subject to short-selling constraints and risk-neutral rational arbitrageurs that are not subject to short-selling constraints. Both investors believe that only their own signal is informative. The authors argue that due to short-selling constraints an investor with a pessimistic signal will stay out of the market so her information will not be reflected in the market price. However, as time passes and when both investors get a bad signal, any new information as well as past hidden information is revealed when the market is falling. The price movements can be severe enough to lead to a crash. Bai, Chang, and Wang (2006) demonstrate that short-selling restrictions limit both risk-sharing and trades based on private information. Limiting trades based on private information reduces the informativeness of market prices and increases the risk perceived by less informed investors. Thus, constraining short-selling may lead to market crashes as prices can fall sharply.

Experimental studies also show that permitting short-selling reduces the probability of a bubble formation. Ackert, Charupat, Church, and Deaves (2002), for example, analyze the results of experimental asset markets. The authors find that crashes are moderated when short-selling is allowed as it provides an equilibrating force in the market. Haruvy and Noussair (2006) use an experimental environment to examine the effect of allowing short-selling on the incidence and magnitude of market bubbles. They find that short-selling reduces market prices; however, allowing large short-selling capacity reduces prices to levels below their fundamental values.

2.3.4 Summary

The collapse of Lehman Brothers in the autumn of 2008 revealed the extent of counterparty credit risk problems. More precisely, major financial institutions were not aware of the extent of exposure of their counterparties to toxic assets. Due to this opaqueness, banks were reluctant to lend to each other, which reduced the much needed supply of liquidity during the financial crisis. Since the market was using the changes in stock prices of financial institutions to infer their level of credit risk, financial authorities introduced short-selling bans. This was because they feared that short-selling could exert downward pressure on the stock prices of these financial institutions and send the wrong signal to the markets (Brunnermeier (2009)).

The academic literature has mixed views on the wisdom of intervention by financial authorities to restrict short-selling. On the one hand, in extreme market conditions, short sellers can exert downward pressure on prices and drive these prices to levels that are not justified by the fundamentals (see for example, Shkilko, Van Ness, and Van Ness (2012), and Blau, Van Ness, Van Ness, and Wood (2010)). On the other hand, short-selling constraints can create price bubbles and lead to a crash (see for example, Hong and Stein (2003) and Scheinkman and Xiong (2003)).

2.4 Short-selling rules in Europe

2.4.1 Introduction

This section reviews the short-selling rules introduced in Belgium, Denmark, France, Germany, Ireland, Italy, Netherlands, Norway, Portugal, Spain, Switzerland, and the UK. Table 2.1 describes the possible short-selling measures that a financial authority can implement. Additionally, Appendix 2A summarizes the short-selling rules that were applied in the sample countries.

Table 2.1 Short-selling measures [11]

Short-selling rule	Description
Ban on naked short-selling	The seller is not allowed to sell the shares of a stock unless she has borrowed these shares before the transaction.
Ban on naked and covered short-selling	The seller is not allowed to sell the shares of a stock unless she owns them.[12]
Circuit breakers	Circuit breakers that are intended to halt the trading on a stock for the rest of the day when the volatility of the price of that stock exceeds a particular threshold set by the stock exchange or the regulator.
Disclosure obligations	The seller is obliged to disclose to the regulator and/or the market her short position in the specified stocks if a particular threshold is met e.g., 0.25% of the issued share capital.
Flagging	The seller or her broker is obliged to indicate that the transaction taking place is a short sale. IOSCO (2009), p.15, defines flagging of short-selling as "…a system that requires putting a marker on each individual short sale order that a broker sends to the exchange or alternative trading facility for execution".
Lending restrictions	Local regulatory authorities encourage the stock lenders to abstain from lending the shares of specified stocks to short-sellers or recommend that lenders receive a consent form from their clients permitting them to lend out their clients' stocks.
Penalty for fails	Stock exchanges impose a penalty to short sellers when they fail to deliver the shares promised to the buyer on the settlement day.
Up-tick rule	The seller is not allowed to sell short the stock if the current price is lower than the previous sale price. The purpose of this rule is to prevent short sellers from inducing prices declines.

2.4.2 Belgium

Under Royal Decree, the King in Belgium has the authority to intervene when good functioning, integrity, and transparency of financial markets are threatened due to market abuse and manipulation.[13] Non-compliance may result in an administrative fine, which may reach €2,500,000.[14]

Prior to the Lehman Brothers bankruptcy in mid-September, 2008, there were no specific short-selling regulations in place in Belgium. However, following the Lehman Brothers

[11] See Clifford Chance (2009c).
[12] "The seller owns the stock when (i) the seller has purchased or entered into an unconditional contract to purchase the stock but has not yet received delivery; (ii) the seller has a title to other securities, which are convertible into or exchangeable for the stock to which the order relates and has tendered the application to convert or exchange; (iii) the seller has an option and has exercised such an option to acquire the stock to which the order relates; (iv) the seller has rights or warrants and has exercised such rights or warrants to subscribe to and to receive the stock to which the order relates; (v) the seller is making a sale of a stock that trades on a "when issued" basis and has entered into a binding contract to purchase such security, subject only to the condition of issuance of the security; (vi) the seller has bought the stock in one market and then sells the same stock in another market regardless of whether it is an overseas market." (See IOSCO (2009)).
[13] See CESR (2009: p.2).
[14] See Clifford Chance (2009c: p.10).

collapse, the Comission Bancaire, Financiere et des Assurances (CBFA) introduced temporary short-selling regulations on September 22, 2008. The new short-selling rules included a naked short-selling ban and a net short position disclosure obligation.[15] More precisely, a naked short-selling ban applied to five financial institutions listed on Euronext Brussels.[16] Transactions that created short positions in stocks of the specified financial institutions or derivatives written on these stocks, on-exchange or in the over the counter (OTC) market, in Belgium or abroad, were subject to the naked short-selling ban. An investor who wished to sell short the shares of the five specified financial institutions either directly through the spot market or synthetically through the derivatives market had the following choices: (i) possess the shares before selling them on the market, (ii) borrow the shares on the stock lending market before the transaction, and (iii) have proof that the shares had been purchased and would be delivered on the third day at the latest following the short sale.[17]

The CBFA also requested market participants to abstain from lending the shares of the specified financial institutions if the lending process led to the creation of an economic short position.[18] Moreover, financial intermediaries had to ascertain that their clients were able to deliver the shares they sold short on time.[19]

With respect to disclosure requirements, anyone who held a net short position with an economic interest of more than 0.25% of the issued share capital by the five specified financial institutions had to report this position within one day of the transaction to the CBFA by fax or email and to the markets by internationally distributed press release.[20]

[15] See Clifford Chance (2009c: p.10).
[16] The five financial institutions were Dexia SA, Fortis NV/SA, KBC Groep NV, KBC Ancora CVA, and ING Groep NV.
[17] See Clifford Chance (2008a: p.8).
[18] See Clifford Chance (2008a: p.8).
[19] See Clifford Chance (2009c: p.10).
[20] See Clifford Chance (2008a: p.8).

While short-selling rules were applicable to most market participants, there were some important exemptions. More precisely, any existing positions prior to September 22, 2008 were exempt from these rules. Also, liquidity providers as defined by the Rulebook of Euronext, market makers on the cash and derivatives markets, and block trade counterparties were also exempt from the short-selling rules.[21,22]

On August 12, 2011, the Financial Services Market Authority (FSMA), which replaced the CBFA in April 2011, modified the short-selling rules by introducing a covered short-selling ban on the same financial institutions.[23] The covered short-selling ban was lifted on February 13, 2012 while disclosure requirements remained in force. Additionally, the FSMA introduced the "locate rule", which instructed investors who sold the shares of the specified financial institutions without possessing or having borrowed them to make all the necessary arrangements so that they could deliver these shares on time.

National temporary short-selling rules were replaced with the permanent regulation on short-selling of the European Union No. 236/2012, which came into force on November 1, 2012.[24]

2.4.3 Denmark

There were no short-selling regulations in Denmark before the autumn of 2008.[25] On October 13, 2008, the Danish Financial Supervisory Authority (Finanstilsynet) banned the creation of short positions, or an increase of an existing short position, in the stocks of all Danish banks that were traded on a regulated market. According to this measure, a person that entered into a short position or increased her existing short position was allowed to sell under two circumstances. The person had to either own the number of shares she intended to sell or

[21] See Clifford Chance (2008a: p.8).
[22] https://europeanequities.nyx.com/members/market-makers-and-liquidity-providers
[23] http://www.fsma.be/en/consumers/press/archwarndiv/article/press/div/2011-08-11_shortselling.aspx
[24] http://www.fsma.be/en/RSS/Article/press/div/2012/2012-02-13_shortselling.aspx
[25] See Clifford Chance (2009c: p.15).

alternatively, enter into an agreement which specified that she would own the number of the shares she intended to sell and would have them delivered to the buyer on time.[26]

Transactions in other financial instruments that led to capital gains from price declines in the shares of specified banks were also prohibited unless these agreements were entered into to hedge risk on shares covered by the short-selling ban. Violators of the new short-selling rules would be liable to a fine.[27] However, market makers were exempt from all the newly introduced short-selling rules. A market maker was defined as a securities dealer willing to trade on a continuous basis in financial markets on her own account by buying and selling securities from her own holdings at prices defined by her. Additionally, these rules were not applied to trading in own shares in buy-back programs or to financial instruments as part of the stabilization of a security's price.[28]

With regards to disclosure requirements, no rules on disclosing short positions were introduced. Also, financial intermediaries were not required to obtain confirmation statements from their clients that the latter did not engage in short-selling.[29]

The short-selling ban was lifted on November 1, 2012 and replaced with the European Union regulation on short-selling No. 236/2012.[30]

2.4.4 France

Prior to September 2008, there were rules restricting short-selling in France. For example, delivery had to take place three days after the execution of the short-selling order. If delivery

[26] See Clifford Chance (2008d: p.1).
[27] See CESR (2009: pp.3–4).
[28] See CESR (2009: p.4).
[29] See Clifford Chance (2008d: p.1).
[30] http://www.investmenteurope.net/investment-europe/news/2221558/eu-regulation-lifts-danish-short-selling-ban

of securities required more than three days after the execution of a short-selling order, collateral would be needed.[31]

On September 22, 2008, the French Financial Authority (Autorité des Marchés Financiers or AMF) banned naked short-selling of specified financial institutions and introduced disclosure requirements.[32] These short-selling rules applied to all investors and service providers, both on-exchange and in the OTC market, in France and abroad. More specifically, the short-selling ban covered spot, forward, and option transactions made on own account or on behalf of a third party. Intra-day short positions were also subject to the naked short-selling ban. A seller had to cover her position within three days of the short sale transaction taking place. If a seller did not own the shares she intended to sell, she could cover the short sale position by borrowing the shares from the securities lending market. However, a convertible bond or a depository receipt was not adequate unless the bond or the depository receipt had been converted into shares prior to the delivery date, which was three days after the short sale transaction. Furthermore, hedging for trades in index derivatives or baskets required the seller to own the securities if the latter were subject to the short-selling ban.[33]

Regarding disclosure requirements, any person that held a net position representing an economic interest of 0.25% or more of the issued share capital of the stocks subject to the short-selling ban had to disclose her short positions to the AMF and to the market within one day of the short sale. Short positions had to be calculated daily for the entire position, including derivatives on a delta-adjusted basis.[34] Also, financial institutions were requested to abstain from lending the shares of stocks of the specified financial institutions, while

[31] See Clifford Chance (2009c: p.16).
[32] The specified financial institutions were Allianz, April Group, AXA, BNP Paribas, CIC, CNP Assuranses, Crédit Agricole, Dexia, Euler Hermes, HSBC Holdings, Natixis, NYSE Euronext, Paris Re, Scor,Société Générale.
[33] See Clifford Chance (2008a: p.10).
[34] See Clifford Chance (2008a: p.10).

financial intermediaries were required to obtain a formal statement from their clients that they possessed the shares they intended to sell short.[35]

Commitments made before September 22, 2008, or lending unrelated to creating an economic net short position were exempt from the short-selling regulations. Market makers, liquidity providers, and block trade counterparties were also exempt from the short-selling rules.[36]

On August 12, 2011, the AMF introduced a covered short-selling ban on the stocks of the same financial institutions, which was lifted on February 13, 2012.[37,38] On November 1, 2012, the AMF replaced national temporary short-selling rules with the permanent European Union regulation on short-selling, No. 236/2012.[39]

[35] See Clifford Chance (2009c: p.16).
[36] See Clifford Chance (2008a: p.10).
[37] http://www.amf-france.org/documents/general/10109_1.pdf
[38] http://www.amf-france.org/documents/general/10310_1.pdf
[39] http://www.amf-france.org/documents/general/10623_1.pdf

2.4.5 Germany

Before the introduction of the 2008 short-selling ban in Germany, investment funds were not allowed to sell short. Hedge funds were the exemption. However, the German Financial Supervisory Authority (Bandesanstalt fur Finanzdienstleisunsaufsicht or BaFin) had the legal power to ban short selling with respect to hedge funds as well. Generally, short selling was not used by German banks.[40]

On September 20, 2008, BaFin banned the naked short-selling of stocks of 11 financial institutions on an intra-day basis.[41] Nevertheless, single stock and index linked derivatives, along with warrants, were exempt from the short-selling ban. Although the seller had to own the shares or have a legal claim for the transfer of the title in the shares or arrange to borrow the shares prior to or at the same time as the short sale transaction, the securities did not need to be booked into the account of the seller.[42] No disclosure requirements were introduced.[43] The ban on naked short-selling of stocks of financial institutions was lifted on January 31, 2010.[44]

On December 15, 2008, the Frankfurt Stock Exchange also banned the naked short-selling but of stocks of all listed firms. Any violations of the ban were treated as market manipulation.[45]

Market makers were exempt from the short-selling ban. Also, although financial intermediaries were not obliged to monitor their customers' compliance with the short-selling ban, they were required to report suspicious transactions.[46]

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[40] See Clifford Chance (2009c: p.16).
[41] The 11 financial institutions were Aareal Bank AG, Allianz SE, AMB Generali Holding AG, Commerzbank AG, Deutsche Bank AG, Deutsche Börse AG, Deutsche Postbank AG, Hannover Rückversicherung AG, Hypo Real Estate Holding AG, MLP AG, Münchener Rückversicherungs-Gesellschaft AG.
[42] See Clifford Chance (2008a: p.11).
[43] See Clifford Chance (2008a: p.11).
[44] See Clifford Chance (2009c: p.16).
[45] See Clifford Chance (2008a: p.11).

As of November 1, 2012, the European Union regulation on short-selling No. 236/2012 came into force in Germany.[47]

2.4.6 Ireland

Prior to September 2008, there were no specific short-selling regulations in Ireland. After the introduction of the short-selling ban, any pre-ban short positions were allowed to continue, be decreased or be closed out.[48] On September 19, 2008, Ireland's Financial Regulator banned transactions generating a net economic benefit from a price decline in the shares of four financial institutions.[49] Intra-day short positions were also prohibited.[50]

The short-selling ban applied to new short positions as well as increases in existing short positions, on-exchange and in the OTC market, covering any economic interest, direct or indirect, in the price of stocks of the specified financial institutions, such as spread bets, contracts-for-differences (CFDs), and baskets.[51]

With regards to disclosure obligations, any person that held on a particular day an economic interest of 0.25% or more of the issued share capital of the specified financial institutions had to disclose her position to the Regulatory Information Service by 15:30 of the same day. The information disclosed referred to the name of the person holding the short position, the name of the company whose shares were sold short, and the amount of the short position. Furthermore, disclosure was required even if the size of the net exposure had not changed since the last disclosure.[52]

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[46] See Clifford Chance (2008a: p.14).
[47] http://www.bafin.de/EN/Supervision/StockExchangesMarkets/ShortSelling/shortselling_node.html
[48] See Clifford Chance (2009c: p.20).
[49] The four financial institutions were AIB Group, Bank of Ireland, Anglo Irish Bank, Irish Life and Permanent.
[50] See Clifford Chance (2009c: p.20).
[51] See Clifford Chance (2008a: p.14).
[52] See Clifford Chance (2008a: p.14).

Market makers trading in good faith as a principal to fulfil clients' orders or hedge positions arising from clients' orders were exempt from the short-selling ban and disclosure obligations.[53]

On December 30, 2011, the Central Bank of Ireland ended the short-selling ban on financial instruments.[54] The European Union regulation on short-selling No. 236/2012 was applicable in Ireland as of November 1, 2012.

2.4.7 Italy

There were no specific short-selling regulations in Italy prior to September 2008. After Lehman Brothers collapsed on September 15, 2008, the Italian Financial Authority (Commissione Nazionale per le Societa e la Borsa or CONSOB) introduced a ban on the naked short-selling of stocks of banks and insurance firms listed and traded on the Italian regulated markets.[55] On October 1, 2008, a covered short-selling ban was implemented on the same stocks. Under the new short-selling ban, the seller could cover her position only if she owned the shares she intended to sell short. On October 10, 2008, the ban was extended to all listed firms, financial and non-financial.[56]

On January 1, 2009, the covered short-selling ban lapsed for the non-financial firms, which now were subject only to a naked short-selling ban until July 31, 2009.[57] In contrast, financial institutions continued to be subject to both naked and covered short-selling bans until May 31, 2009.[58,59]

[53] See Clifford Chance (2008a: p.14).
[54] http://www.centralbank.ie/press-area/press-releases/Pages/CentralBankofIrelandannouncestheendofthebanonshortsellingofcertainfinancialinstruments.aspx
[55] See Clifford Chance (2009c: p.23).
[56] See Clifford Chance (2009c: p.23).
[57] See Gruenewald, Wagner, and Weber (2009b: Appendix).
[58] See Clifford Chance (2009a: p.2).
[59] See the list of financial institutions at http://www.consob.it/mainen/documenti/english/resolutions/res16813.htm

From May 31, 2009 to July 31, 2009, both financial and non-financial stocks listed and traded on Italian regulated markets were subject to a naked short-selling ban only. Since June 1, 2009, only firms increasing their capital remained subject to the covered short-selling ban where ownership of the shares was required.[60] All short-selling bans covered physical interests. Establishing short positions was allowed in convertible bonds, warrants, derivatives and indexes.[61] CONSOB did not introduce any net short position disclosure requirements.[62]

Originally, all market participants had to comply with the rules whether they traded on-exchange or in OTC market, in Italy or abroad.[63] After the amendment of the rules on October 10, 2008, however, short-selling rules did not apply to OTC trades.[64]

Market makers were exempt from the short-selling regulations. Specialists and liquidity providers' activities recognized as such by the Italian Stock Exchange and carried out in regulated markets were also exempt from the short-selling rules. Lastly, sanctions for breach of the short-selling rules were the same as those for any other breach of the market abuse regime.[65]

On August 12, 2011, CONSOB banned the creation of, or increase in, net short positions of financial stocks.[66] All existing short-selling regulations were replaced with the European Union regulation on short-selling No. 236/2012 as of November 1, 2012.[67]

2.4.8 Netherlands

Prior to 2008, there were no specific short-selling regulations in the Netherlands. Nevertheless, the Dutch clearing house charged a fee to clearing members for providing

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[60] See Gruenewald, Wagner, and Weber (2009b: Appendix).
[61] See Clifford Chance (2008a: p.15).
[62] See Clifford Chance (2008a: p.15).
[63] See Clifford Chance (2008a: p.15).
[64] See Clifford Chance (2009a: p.2).
[65] See Clifford Chance (2009a: p.2).
[66] http://www.consob.it/mainen/press_release/comunicato_20110825.htm
[67] http://www.consob.it/mainen/markets/short_selling/index.html

automatic borrowing or "buy-in" facilities when clearing members failed to deliver the sold short securities at settlement.[68]

On September 22, 2008, the Financial Supervisory Authority of the Netherlands (Autoriteit Financiele Markten or AFM) introduced a ban on the naked short-selling of all listed firms trading on domestic stock exchanges and the OTC market. Intra-day positions were also covered by the naked short-selling ban; however, all other instruments were excluded from the ban.[69]

On October 5, 2008, a covered short-selling ban was introduced on eight specified financial institutions.[70] More precisely, the ban covered the creation of, or an increase in, a net position giving rise to an economic exposure to shares of specified financial institutions and insurers. Consequently, the ban covered both the shares of specified financial institutions and insurers and the related instruments where the price or value depended on the price or value of these shares (e.g., CFDs, spread bets, options, futures, and depository receipts). Any economic interest held as part of a basket, index, or ETF, where the predominant components were the specified financial institutions, also had to be included in the calculation of the net short positions.[71]

The AFM clarified that a short position in the equity of a relevant Dutch financial firm could be offset by a long position in convertible bonds in the same firm, provided this did not result in a new or increased net short position.[72] Moreover, if a client wished to establish a short

[68] See Clifford Chance (2009c: p.26).
[69] See Clifford Chance (2008a: p.18).
[70] The eight financial institutions were Aegon N.V., ING Groep N.V., Fortis N.V., Binckbank N.V., Kas Bank N.V., SNS Reaal N.V., Van Der Moolen Holding N.V., Van Lanschot N.V.
[71] See Clifford Chance (2008b: p.3).
[72] See Clifford Chance (2008b: p.3).

position with the intermediary, the intermediary needed to ascertain that the client's order was fully covered and report suspicious transactions.[73]

With regard to disclosure requirements, a net economic short position of more than 0.25% of the issued share capital of a specified financial institution had to be reported to the AFM within one day of the position being established. Disclosure obligations included all instruments such as CFDs, spread bets, and options giving rise to a direct or indirect exposure in the equity capital of a firm. Disclosure was required even if the position had not changed since the previous disclosure. If the net short position fell below 0.25%, one final disclosure was required. Additionally, the person making the disclosure had to provide details of (i) short and long positions making up the net position and (ii) the component instruments making up the position.[74]

All market participants had to comply. However, market makers on the derivatives markets, liquidity providers on the cash markets, and block trade counterparties were exempt from the short-selling rules.[75]

The prohibition expired on June 1, 2009. While the disclosure regime remained in force until December 31, 2009, it was amended slightly so that an increase or decrease of a short interest past additional threshold over 0.25% was now disclosed.[76,77]

On November 1, 2012, AFM adopted the European Union regulation on short-selling No. 236/2012.[78]

[73] See Clifford Chance (2008a: p.18).
[74] See Clifford Chance (2008a: p.18).
[75] See Clifford Chance (2009c: p.26).
[76] http://www.afm.nl/en/professionals/afm-actueel/nieuws/2009/mei/verbod_short_selling_vervalt.aspx
[77] See Clifford Chance (2009c: p.26).
[78] http://www.afm.nl/en/professionals/afm-actueel/nieuws/2012/okt/meldingsregeling-shortposities.aspx

2.4.9 Norway

Generally, domestic and foreign investment firms in Norway are not allowed to establish naked short positions either on their own account or on behalf of their clients. If an investment firm trades on its own account, it must have access to the financial instruments to deliver them on time. Furthermore, if an investment firm sells financial instruments on behalf of their client who does not own these financial instruments, it can only do so if the client is able to access the financial instruments. Lastly, if investment firms mediate securities lending agreements, they should ensure that satisfactory collateral is provided.[79]

On October 8, 2008, the Financial Supervisory Authority of Norway (Kredittilsynet) banned both naked and covered short-selling of stocks of five specified financial institutions.[80] However, no obligations were imposed to disclose short positions. Market makers in derivatives and warrants listed on the Oslo Stock Exchange were exempt from the short-selling ban.[81] This short-selling ban was lifted on October 9, 2009.[82]

[79] See Clifford Chance (2009c: p.28).
[80] The five financial institutions were DnB NOR ASA, Voss Veksel og Landmandsbank ASA, Storebrand ASA, Bluewater ASA, Protector ASA.
[81] See Clifford Chance (2008c: p.1).
[82] http://www.finanstilsynet.no/en/Secondary-menu/Documents/Press-releases/Market-conditions-no-longer-an-obstacle-to-short-selling-in-Kredittilsynets-view/

2.4.10 Portugal

There were no short-selling regulations in Portugal prior to September 2008. According to the Portuguese Securities Code, however, financial intermediaries were required to ensure that the seller had legal title to the securities before executing a sale order. If the seller could not prove possession of the financial instruments she intended to sell, the financial intermediary was entitled, but not forced, to refuse to execute such an order.[83]

On September 24, 2008, the Comissao do Mercado de Valores Mobiliarios (CMVM) banned the naked short-selling of the stocks of eight specified financial institutions traded on Euronext Lisbon but not in the OTC market.[84] The ban covered equities, convertible bonds, depository receipts, and baskets. Additionally, sell orders for stocks subject to the short-selling ban could not be accepted if the seller did not hold or was not in a position to prove that she would hold the stocks at the time of the transaction.[85]

A seller could cover her position either by owning the shares or through borrowing the shares in the securities lending market. Thus, a short sale would be adequately covered if there was a stock loan agreement in place at the time of the short sale. Provided that the seller could ensure that the shares would be available for settlement, a short sale could be executed even if the shares had not yet been delivered into the seller's account. On the other hand, a convertible stock or depository receipt would not be sufficient if the shares were not available on the settlement day.[86]

On September 29, 2008, the CMVM also implemented two disclosure regimes: (i) an obligation to disclose short positions above 0.25% of the issued share capital of the specified

[83] See Clifford Chance (2009c: p.31).
[84]The eight financial institutions were Banco Commercial Português, Banco Espìrito Santo, Banco BPI, Banif SGPS, Finibanco Holding SGPS, Banco Santander, Banco Popular Español, Espìrito Santo Financial Group.
[85] See Clifford Chance (2008a: p.19).
[86] See Clifford Chance (2008a: p.19).

financial institutions, and (ii) an obligation for members of Euronext Lisbon and PEX to make a daily disclosure of any short position acquired or reduced in shares traded on Euronext Lisbon and PEX.[87] On January 9, 2009, the CMVM lifted the daily reporting requirement. Market makers were exempt from both short-selling the ban and the two disclosure obligations.[88]

In July 2011, CMVM extended the net short position disclosure requirements to all the equities listed on Portuguese regulated markets. The new threshold to disclose the short positions to the regulator decreased to 0.20% of the issued share capital and to the market set to 0.50%.[89] On November 1, 2012, Portugal directly implemented the European Union regulation on naked short-selling and disclosure requirements of net short positions No. 236/2012.[90]

2.4.11 Spain

Naked short-selling has been prohibited in Spain since 1967. On September 24, 2008, the Comision Nacional del Mercado de Valores (CNMV) introduced disclosure requirements for transactions taking place on stock exchanges and in the OTC market. According to these requirements, any natural or legal person holding net short positions above 0.25% of the issued share capital in any of the 20 specified financial stocks had to disclose these positions.[91] A short position was defined as the net result of all positions in financial instruments – derivatives and underlying assets – that gave any natural or legal person a positive exposure or profit due to decreases in prices of stocks of the specified financial

[87] See Clifford Chance (2009c: p.31).
[88] See CliffordChance (2009b: p.2).
[89] http://www.esma.europa.eu/system/files/2011-39.pdf
[90] http://www.esma.europa.eu/system/files/2012-666.pdf
[91] The 20 financial institutions were Banco de Andalucía, Banco de Castilla, Banco de Crédito Baleur, Banco de Galicia, Banco Pastor, Banco Guipuzcoano, Banco Popular Español, Banco Sabadell, Banco Santander, Banco de Valencia, Banco de Vasconia, Banco Español de Crédito, Bankinter, BBVA, Caja de Ahorros del Mediterráneo, Grupo Catalana Occidente, Mapfre, Inverfiatc, Bolsas y Mercados, Españoles, and Renta 4.
[91] See Clifford Chance (2008a: p.22).

institutions. Changes in short positions had to be disclosed to the public before 19:00 on the day following these changes.[92]

A short sale would be considered as covered if the seller possessed the shares by either (i) a purchase or (ii) a securities loan agreed before the sale, or (iii) an irrevocable exercise of convertible security, option, or any kind of derivative instrument. If the stock loan was registered after the sale date, the sale would be considered as a naked short sale unless the seller demonstrated she held a sufficient number of the shares she intended to sell before the order was placed.[93] Stock lending was also subject to disclosure requirements. Penalties were imposed by the CNMV and separately by IBERCLEAR whenever a trade could not be settled due to insufficient securities.[94] Lastly, market makers were exempt from the short-selling rules.[95]

On August 12, 2011, the CNMV banned the creation of net short positions on Spanish financial stocks. The ban was lifted on January 31, 2013. The European Union regulation on naked short-selling and disclosure requirements of net short positions No. 236/2012 was also in force in Spain as of November 1, 2012.[96]

2.4.12 Switzerland

There were no specific short-selling regulations or short position disclosure requirements in Switzerland prior to September 2008. On September 19, 2008, the Swiss Federal Banking Commission (SFBC) banned naked short-selling in all securities traded on stock exchanges and in the OTC market and Six Swiss Exchange banned the creation of net short positions in

[92] See Clifford Chance (2008a: p.22).
[93] See Clifford Chance (2008a: p.22).
[94] See Clifford Chance (2009c: p.37).
[95] See CESR (2009: p.11).
[96] http://www.cnmv.es/loultimo/prorroga%201%20nov_en.pdf,
http://www.cnmv.es/DocPortal/Aldia/FAQ_ENGLISH_v4_20_septiembre.pdf

specified financial institutions on an intra-day basis.[97] The Six Swiss Exchange's short-selling rules did not cover the net short positions created using derivatives.[98]

At the time of the trade, the seller had to have arranged in advance to borrow the sold short shares to ensure delivery of these shares on the settlement day. Thus, the SFBC allowed banks to lend securities to their clients in order to facilitate the coverage of short-selling transactions. Although the Swiss-regulated entities had to ensure that their clients were able to deliver the securities at settlement, they were not expected to systematically ensure that their clients held the sold short shares at the time of the trade. Nevertheless, regulated entities were recommended to monitor clients more closely if they suspected that their clients were unable to deliver the sold short shares on the settlement date.[99] There were no net short position disclosure requirements introduced. The ban introduced by the Six Swiss Exchange lapsed on January 16, 2009, but the naked short-selling ban on all stocks remained in force.[100] Market makers were exempt from both short-selling bans.

2.4.13 UK

Prior to the new short-selling rules introduced in September 2008, rules were already in place that covered settlement obligations in the UK. More specifically, according to the London Stock Exchange Rule 1400, member firms had to ensure that when they sold short on a substantial scale, either on their own account or on behalf of their customers, they were in a position to settle the short-selling transaction. If member firms believed that they would be unable to deliver the shares sold short on the settlement day they could not undertake their short-selling strategy.[101]

[97] These financial institutions were UBS AG, Credit Suisse, Julius Bär, Swiss Life, Swiss Re, Zurich, Bâloise.
[98] See Clifford Chance (2008a: p.23).
[99] See Clifford Chance (2008a: p.23).
[100] See Clifford Chance (2008b: p.5).
[101] See Clifford Chance (2009c: p.6).

On January 16, 2008, the Financial Services Authority (FSA) introduced short-selling regulations affecting transactions on organised stock exchanges and OTC markets. These regulations banned the creation of, or increase in, a net short position that gave rise to an economic exposure to shares of specified financial institutions and insurers. Short intra-day positions were also covered by the ban.[102] Issued share capital included ordinary and preferred shares, and both naked and covered short positions were taken into account in estimating the net short position. Additionally, any related instruments where the price or value depended on the price or value of the share of the specified financial institutions such as CFDs, spread bets, options, futures, and depository receipts were also covered by the short-selling rules.[103] Any economic interest held as part of a basket, index, or exchange traded fund (ETF), where the predominant components were the specified financial institutions, had to be included in the calculation.[104] A short position could be covered by taking a long position in the shares of the specified financial institutions or in an instrument where the shares were the underlying asset.[105]

Regarding disclosure obligations, if a person held a net short position of 0.25% or more of the issued share capital of the specified financial institutions, she was required to disclose her position by 15:30 at the latest on the business day following each day on which the short position was held. Even if the size of the short position had not changed since the previous disclosure, disclosure was still required. Finally, if the net short position fell below 0.25%, one final disclosure had to be made.[106] In November 2008, this rule was amended. Once disclosure of a short position was made, additional disclosures were only required when that

[102] See the list of specified financial stocks on http://www.fsa.gov.uk/pubs/other/Shortselling_list.pdf
[103] These financial instruments should be accounted for on a delta basis.
[104] See Clifford Chance (2008a: p.3).
[105] See Clifford Chance (2008a: p.3).
[106] See Clifford Chance (2008a: p.3).

short position changed.[107] After the short-selling ban was lifted in January 16, 2009, the disclosure requirements were amended for a second time. The disclosure of net short positions had to be made by no later than 15:30 on the business day following the day on which a short position in stocks of the specified financial institutions reached, exceeded or fell below 0.25%, 0.35%, 0.45%, and 0.55%, respectively, of their issued share capital and each 0.1% threshold thereafter. The disclosure obligations were lifted on December 31, 2009.[108]

Short-selling rules did not apply to short positions created before their introduction. Market makers were also exempt from the short-selling rules when acting in their role. A market maker is considered by the FSA as an entity that, ordinarily as a part of her business, deals as principal in equities, options or derivatives exchange-traded and OTC to fulfil clients' orders by trading on their behalf or hedge positions arising from such trading.[109]

Although no additional restrictions on stock lending were introduced, the FSA asked firms to be vigilant and notify the authorities if they suspected that stock was borrowed for the purpose of banned short-selling.[110]

On November 1, 2012, the European Union regulation on naked short-selling and disclosure requirements of net short positions No. 236/2012 was directly applied in the UK.[111]

2.4.14 The five dimensions of short-selling regulations

The preceding survey of short-selling rules in Europe shows that there are five dimensions of these short-selling rules: (i) type of the short-selling ban, (ii) coverage of the short-selling

[107] See Clifford Chance (2008e: p.3).
[108] See Clifford Chance (2009b: p.3).
[109] See Clifford Chance (2008a: p.3).
[110] See Clifford Chance (2008a: p.3).
[111] http://www.fsa.gov.uk/about/what/international/short-selling

ban, (iii) disclosure requirements of short positions, (iv) stock lending restrictions, and (v) exemptions from the short-selling regulations. Table 2.2 summarizes the various short-selling rules in each country, based on these five dimensions.

Two main types of short-selling bans were introduced in European markets: a naked short-selling ban and a covered short-selling ban. The main difference between the two types of bans is that in a naked short-selling ban borrowing the shares sold short is considered adequate to cover a short position while in a covered short-selling ban only the ownership of the shares sold is sufficient to cover a short position. Belgium, France, Germany, Portugal, and Spain introduced a naked short-selling ban while Denmark, Ireland, the Netherlands, Norway, and UK adopted a covered short-selling ban. It is important to note that borrowing the shares sold short under a naked short-selling ban regime would be adequate only if the seller borrowed the shares before or at the same time as placing the sale order, but not had the shares booked in their account prior to the settlement date. The shares would also have to be obtained and delivered to the buyer when delivery was due.

With the exception of Germany, Italy, Norway, and Switzerland, the short-selling rules applied to short positions in specified financial institutions created in both spot and derivatives markets. This implied that one could not sell short the shares of financial institutions subject to short-selling regulations synthetically using derivatives such as put options without first borrowing the stocks in the case of the naked short-selling ban or owning the shares in the case of the covered short-selling ban.

Seven out of 12 jurisdictions introduced disclosure requirements along with their short-selling bans. Denmark, Germany, Italy, Norway, and Switzerland did not apply any obligations to disclose short positions of stocks subject to their short-selling bans. Stock lending was another activity to be restricted by some European countries such as Belgium, France, and the Netherlands. In these countries the regulatory authorities requested market participants to refrain from lending the shares of securities that were subject to the short-selling rules. Finally, market makers were exempt from all types of short-selling rules in all the sample countries.

Table 2.2 Summary of the 2008 short-selling rules introduced in 12 European countries

This table summarises the short-selling rules introduced in the sample countries grouped by five dimensions: type of short-selling ban, coverage of the short-selling ban, disclosure requirements of short positions, stock lending restrictions, and exemptions from the short-selling rules.

Country	Type of the ban	Coverage of the ban	Disclosure requirements	Lending restrictions	Exemptions
Belgium	Naked	Spot and derivatives markets	Yes	Yes	Market makers
Denmark	Naked and covered (prohibition of a creation of, or an increase in, a short position)	Spot and derivatives markets	No	No	Market makers
France	Naked	Spot and derivatives markets	Yes	Yes	Market makers
Germany	Naked	Spot market	No	No	Market makers
Ireland	Naked and covered (prohibition of a creation of, or an increase in, a short position)	Spot and derivatives markets	Yes	No	Market makers
Italy	See Appendix 2A	Spot market	No	No	Market makers
Netherlands	Naked and covered (prohibition of a creation of, or an increase in, a short position)	Spot and derivatives markets	Yes	Yes	Market makers
Norway	Naked and covered	Spot market	No	No	Market makers
Portugal	Naked	Spot and derivatives markets	Yes	No	Market makers
Spain	Naked	Spot and derivatives markets	Yes	No	Market makers
Switzerland	Naked and covered (prohibition of a creation of, or an increase in, a short position)	Spot market	No	No	Market makers
UK	Naked and covered (prohibition of a creation of, or an increase in, a short position)	Spot and derivatives markets	Yes	No	Market makers

2.5 Summary

This chapter provided the institutional background on short-selling regulations in Europe. Section 2.2 described the mechanics of short-selling. Section 2.3 reviewed the academic arguments for and against the practice of short-selling. Section 2.4 described the regulatory framework governing short-selling in 12 European countries, both before and after the collapse of Lehman Brothers in mid-September 2008.

The next four chapters attempt to provide empirical evidence on the consequences of various short-selling regulations in order to (i) understand the implications of these short-selling policies on the European stock markets and (ii) assist regulators in improving their decisions related to short-selling in the future.

APPENDIX 2A[112,113]

Table 2A.1 Description of the short-selling rules in the sample countries

Country	Authority	From	Until	Type of restriction	Exemptions	Type of targeted firms	List of firms
Belgium	Comission Bancaire, Financiere et des Assurances (CBFA) Now FSMA	22/09/2008	**11/08/2011**	Naked short selling ban (incl. transactions in derivatives without complete coverage)	Market makers on derivatives and OTC market, liquidity providers on cash market, block trade counterparties	Financial institutions	Dexia SA, Fortis NV/SA, KBC Groep NV, KBC Ancora CVA, ING Groep NV
		22/09/2008	**01/11/2012**	Disclosure regime for net short positions with economic interest of > 0.25% of capital	As specified above	As specified above	As specified above
		12/08/2011	**12/02/2012**	**Creation or increase in net short position is prohibited**	**As specified above**	**As specified above**	**As specified above**
		01/11/2012		**EU short-selling regulation**			
Denmark	Finanstilsynet	13/10/2008		Short selling ban (incl. financial instruments if not used for hedging purposes)	Market makers, trading in own shares in buy-backs, financial instruments as part of stabilization of stock's price	All Danish banks which are licensed under the Financial Business Act and traded on a regulated market	
		01/11/2012		**EU short-selling regulation**			

112 See Gruenewald, Wagner, and Weber (2009b: Appendix).
113 Updated information in bold (except for the country column) is added by the author of this thesis.

Table 2A.1 Description of the short-selling rules in the sample countries

Table continued from previous page

Country	Authority	From	Until	Type of restriction	Exemptions	Type of targeted firms	List of firms
France	Authorité des Marchés Financiers (AMF)	22/09/2008	**11/08/2011**	Naked short selling ban (incl. Spot, forward and option transactions)	Market makers, liquidity providers, block trade counterparties	Financial institutions	Allianz, April Group, AXA, BNP Paribas, CIC, CNP Assuranses, Crédit Agricole, Dexia, Euler Hermes, HSBC Holdings, Natixis, NYSE Euronext, Paris Re, Scor,Société Générale
		22/09/2008	**31/01/2011**	Disclosure regime for net short positions with economic interest of > 0.25% of capital (to the AMF and the market)	As specified above	As specified above	
		12/08/2011	**12/02/2012**	**Creation or increase in net short position is prohibited**	**As specified above**	**As specified above**	**As specified above**
		01/11/2012		**EU short-selling regulation**			

Table 2A.1 Description of the short-selling rules in the sample countries

Table continued from previous page

Country	Authority	From	Until	Type of restriction	Exemptions	Type of targeted firms	List of firms
Germany	Bundesanstalt für Finanzdienst-leistungsaufsicht (BaFin)	20/09/2008	31/01/2010	Naked short-selling ban	Fixed-price transactions, market makers	Financial institutions	Aareal Bank AG, Allianz SE, AMB Generali Holding AG, Commerzbank AG, Deutsche Bank AG, Deutsche Börse AG, Deutsche Postbank AG, Hannover Rückversicherung AG, Hypo Real Estate Holding AG, MLP AG, Münchener Rückversicherungs-Gesellschaft AG
		19/05/2010	**31/03/2011**	Naked short-selling ban		Banks, credit default swaps in euro-area	
	Frankfurt Stock	15/12/2008	**31/01/2010**	Naked short selling ban		All listed firms	
		01/11/2012		**EU short-selling regulation**			
Ireland	Irish Financial Services Regulatory Authority (IFSRA)	19/09/2008	**30/12/2011**	Short selling ban (incl. spread-bets and CFDs)	Market makers	Financial institutions	AIB Group, Bank of Ireland, Anglo Irish Bank, Irish Life and Permanent
		23/09/2008	**30/12/2011**	Disclosure regime for net short positions with economic interest of > 0.25% of capital	As specified above	As specified above	As specified above
		01/11/2012		**EU short-selling regulation**			

Table 2A.1 Description of the short-selling rules in the sample countries

Table continued from previous page

Country	Authority	From	Until	Type of restriction	Exemptions	Type of targeted firms	List of firms
Italy	Commissione Nazionale per le Società e la Borsa (CONSOB)	23/09/2008	30/09/2008	Naked short selling ban	Market makers, liquidity providers, specialists	Banks and insurance companies listed and traded on the Italian regulated markets	Commissione Nazionale per le Società e la Borsa (CONSOB)
		1/10/2008	9/10/2008	Short-selling ban	As specified above	As specified above	As specified above
		10/10/2008	31/12/2008	Short-selling ban	As specified above	All listed firms	
		1/01/2009	31/01/2009	Short-selling ban	As specified above	Financial institutions and firms increasing their capital	
		1/01/2009	31/07/2009	Naked short selling ban		All listed firms	
		1/02/2009	31/05/2009	Short-selling ban	As specified above	Financial institutions and firms increasing their capital	See list on http://www.consob.it/mainen/documenti/english/resolutions/res16813.htm
		1/06/2009	**11/08/2011**	Short-selling ban	As specified above	Firms increasing their capital	See list on http://www.consob.it/mainen/intermediares/faq_short_selling.htm#list
		12/08/2011	**11/02/2012**	**Creation or increase in net short position is prohibited**	**As specified above**	**Financial institutions**	
		13/02/2012	**31/10/2012**	**Disclosure requirements**	**As specified above**	**Financial institutions**	
		01/11/2012		**EU short-selling regulation**			

Table 2A.1 Description of the short-selling rules in the sample countries

Table continued from previous page

Country	Authority	From	Until	Type of restriction	Exemptions	Type of targeted firms	List of firms
Netherlands	Autoriteit Financiële Markten (AFM)	22/09/2008	4/10/2008	Naked short-selling ban		All listed firms	
		5/10/2008	31/05/2009	Short-selling ban	Market makers	Financial institutions	Aegon N.V., ING Groep N.V., Fortis N.V., Binckbank N.V., Kas Bank N.V., SNS Reaal N.V., Van Der Moolen Holding N.V., Van Lanschot N.V.
		5/10/2008	31/05/2009	Disclosure regime for net short positions with economic interest of > 0.25% of capital (to AFM)	Market makers	Financial institutions	As specified above
		1/06/2009	31/12/2009	Disclosure regime for net short positions with economic interest of > 0.25% of capital and at additional thresholds of 0.1% (to AFM)		Financial institutions	As specified above
		01/11/2012		**EU short-selling regulation**			
Norway	Kredittilsynet	8/10/2008	9/10/2009	Short selling ban (incl. primary capital certificates)	Market makers	Financial institutions	DnB NOR ASA, Voss Veksel og Landmandsbank ASA, Storebrand ASA, Bluewater ASA, Protector ASA

Table 2A.1 Description of the short-selling rules in the sample countries

Table continued from previous page

Country	Authority	From	Until	Type of restriction	Exemptions	Type of targeted firms	List of firms
Portugal	Comissão do Mercado de Valores Mobiliários (CMVM)	24/09/2008	**31/10/2012**	Naked short-selling ban	Market makers, liquidity providers	Financial institutions	Banco Commercial Português, Banco Espìrito Santo, Banco BPI, Banif SGPS, Finibanco Holding SGPS, Banco Santander, Banco Popular Español, Espìrito Santo Financial Group
		29/09/2008	9/01/2009	Daily notification obligation for stock exchange members of all short selling activities		All listed firms	
		29/09/2008	**31/10/2012**	Disclosure for net short positions with economic interest of > 0.25% of capital	Market makers, liquidity providers	Financial institutions	As specified above
		01/11/2012		**EU short-selling regulation**			

Table 2A.1 Description of the short-selling rules in the sample countries

Table continued from previous page

Country	Authority	From	Until	Type of restriction	Exemptions	Type of targeted firms	List of firms
Spain	Comisión National de Mercado de Valores (CNMV)	24/09/2008	16/02/2012	Disclosure regime for net short positions with economic interest of > 0.25% of capital and any increases or decreases of such short positions at latest on the day after each change (to CNMV)		Financial institutions	Banco de Andalucía, Banco de Castilla, Banco de Crédito Baleur, Banco de Galicia, Banco Pastor, Banco Guipuzcoano, Banco Popular Español, Banco Sabadell, Banco Santander, Banco de Valencia, Banco de Vasconia, Banco Español de Crédito, Bankinter, BBVA, Caja de Ahorros del Mediterráneo, Grupo Catalana Occidente, Mapfre, Inverfiatc, Bolsas y Mercados, Españoles, and Renta 4
		12/08/2011	**31/01/2013**	**Creation or increase in net short position is prohibited**		**Financial institutions**	
		16/02/2012	**31/01/2013**	**Disclosure requirements for net short positions above 0.2% of capital on issue and for changes in these positions of magnitude of 0.1%**		**All listed firms**	
		01/11/2012		**EU short-selling regulation**			

Table 2A.1 Description of the short-selling rules in the sample countries

Table continued from previous page

Country	Authority	From	Until	Type of restriction	Exemptions	Type of targeted firms	List of firms
Switzerland	SIX Swiss Exchange	19/09/2008	16/01/2008	Short-selling ban		Financial institutions	UBS AG, Credit Suisse, Julius Bär, Swiss Life, Swiss Re, Zurich, Bâloise
UK	Financial Services Authority (FSA)	19/09/2008	16/01/2008	Short-selling ban	Market makers	Financial institutions	See list on http://www.fsa.gov.uk/pubs/other/Shortselling_list.pdf
		23/09/2008	16/01/2009	Disclosure regime for net short positions with economic interest of > 0.25% of capital	Market makers	Financial institutions	As specified above
		16/01/1009	31/12/2009	Disclosure regime for net short positions with economic interest of > 0.25% of capital and above 0.25% reaching or falling below bands placed every 0.1%	Market makers	Financial institutions	As specified above
		01/11/2012		**EU short-selling regulation**			

CHAPTER 3

Essay 1: Short-selling regulations and market liquidity

3.1 Introduction

In the autumn of 2008, regulators around the world introduced short-selling rules to halt stock price declines of major financial institutions and restore confidence in the financial system. This essay examines the impact of these short-selling rules on the market liquidity of stocks of European financial institutions. Prior studies, which mainly focus on the price dimension of liquidity as measured by the bid/ask spread, find that market liquidity significantly deteriorates under the new short-selling regulations (see for example, Clifton and Snape (2008) and Beber and Pagano (2013)). However, Lee, Mucklow, and Ready (1993) argue that liquidity also has a quantity dimension, depth, which is measured by the average number of shares quoted at the best bid and ask prices respectively. The authors make the point that a deeper understanding of liquidity requires the study of both price and quantity dimensions.

There are studies that either examine one dimension of liquidity or focus on both dimensions of liquidity but study only one market. For example, Beber and Pagano (2013) study the impact of 2008 short-selling rules but only on the price dimension of liquidity for the affected stocks across 30 countries. Marsh and Payne (2012), on the other hand, examine the effects of the 2008 short-selling regulations on both dimensions of liquidity but only in the UK market. This essay contributes to the literature by examining the impact of the 2008 short-selling regulations on both dimensions of liquidity across 10 European countries.[114] From an investor's perspective, depth that is expressed in currency terms is more important than simply the number of shares available for trading at the best bid and ask prices (Chiu, Chung, Ho, and Wang (2012)). This essay also adds to the literature by incorporating depth in currency terms into the analysis. The aim of this research is to provide empirical evidence on the consequences of financial regulators' policies on both dimensions of liquidity. This

[114] These countries are: Belgium, Denmark, France, Germany, Ireland, the Netherlands, Portugal, Spain, Switzerland, and the UK.

should give us a better understanding of the effects of short-selling restrictions on broad aspects of market liquidity and help regulators in their policy-making.

This chapter is organized as follows. Section 3.2 reviews prior studies that examine the impact of the introduction of 2008 short-selling rules on market liquidity. Section 3.3 develops the hypotheses. Section 3.4 describes the sample and the data. Section 3.5 discusses the methodology. Section 3.6 presents and discusses the empirical results. Section 3.7 summarizes the main findings and concludes.

3.2 Literature Review

3.2.1 Short-selling regulations and market liquidity in the US

There were two instances during the financial crisis where short-selling regulations were tightened in the US. On July 15, 2008, the Securities Exchange Commission (SEC) in the US introduced an Emergency Order (EO), which required that anyone engaging in short-selling of 19 specified financial stocks should borrow the shares before the sale took place and deliver them at the settlement. The EO became effective on July 21, 2008 and was lifted on August 12, 2008. Fundamentally, this new rule prohibited naked short-selling. On September 18, 2008, SEC reintroduced a short-selling ban, which covered nearly 800 US financial stocks. This ban was lifted on October 8, 2008.

Bris (2008) and Boulton and Braga-Aves (2010) examined the impact of the EO on the bid/ask spreads of the financial stocks covered by this EO. The EO allowed the construction of a control sample of financial stocks that were not affected by the new short-selling rules. Both papers construct a control sample of financial stocks with similar characteristics, such as market capitalization, closing price, and volatility of the daily returns, in order to isolate the potential effects of the EO on the liquidity of the financial stocks subject to the short-selling rules.

Bris (2008) identifies two control samples of financial stocks that were not covered under the EO but had similar market value, closing price, and volatility of daily returns to those of the banned financial stocks. The first control sample is comprised of US financial stocks while the second sample is comprised of non-US financial stocks listed on the US stock exchanges. The univariate analysis shows that both quoted and percentage quoted bid/ask spreads are greater for the stocks affected by the EO rule compared to those of the US financial stocks in the control sample in both pre-EO and EO periods. Although the difference is statistically significant in both periods, it is greater in the EO period. For example, the difference between the percentage quoted spread of the US banned financial stocks and the US non-banned financial stocks in the pre-EO period is 6.31%. In the EO period this difference increases to 19.90%. Both results are significant at the 1% level. Moreover, in the pre-EO period, the average quoted and percentage bid/ask spreads of the non-US financial stocks in the control sample are greater than those of the 19 specified financial stocks in the banned sample. However, in the EO period, the average quoted and percentage spreads of the financial stocks in the banned sample increase to the level of the respective spreads of the non-US financial stocks in the control sample. For example, the difference between the percentage quoted spread of the US financial banned stocks and the non-US financial non-banned stocks is -23.01% in the pre-EO period, significant at the 1% level, implying that the percentage quoted spread of the US financial banned stocks is much narrower than that of the non-US financial non-banned stocks during that period. In the EO period, this difference increases to 2.42% but it is not statistically significant.

Boulton and Braga-Aves (2010) also study the change in bid/ask spreads around the EO. Each financial stock subject to the EO short-selling rules is matched with a financial stock not affected by these short-selling rules by mean market value, closing price, volatility of daily returns, and daily turnover from January 1, 2007 to December 31, 2007. The authors find that the quoted spread increases for both samples of stocks during the EO. However, the increase is more pronounced in the financial stocks of the banned sample (33.6% versus 14.2%). Both results are statistically significant at the 1% level. Furthermore, results from the multivariate analysis show that the EO dummy variable is approximately 0.19, implying that in the EO period the quoted spread on financial stocks in the banned sample is greater by 19% than the quoted spread on financial stocks in the control sample. This estimate is statistically significant at the 5% level.

To sum up, both studies find that liquidity as measured by the quoted and percentage quoted bid/ask spreads on financial stocks covered under the EO worsens after the introduction of the EO.

Boehmer, Jones, and Zhang (2013) examine how the short-selling ban introduced by SEC on September 18, 2008 affected the liquidity of the financial stocks subject to this ban. The authors use percentage quoted and effective spreads to measure liquidity. The banned sample comprises 727 NYSE and NASDAQ financial stocks subject to the short-selling ban. Each stock in the banned sample is matched with a stock not subject to the short-selling ban to create a control sample of non-banned stocks. The difference between the control sample in this study and the control samples in previous studies is that non-banned stocks in the present study are not financial stocks. The new short-selling ban applied to stocks of most financial institutions, which made it nearly impossible to match each financial banned stock to a financial non-banned stock.[115] Consequently, the authors match the banned stocks from the

[115]The short-selling ban initially covered approximately 800 US financial stocks and was then extended to nearly 1000 US financial stocks (Boehmer, Jones, and Zhang (2013)).

financial industry with non-banned stocks from other industries by listing exchange, listed options status, market value, and dollar trading volume.

Univariate analysis reveals that, in the pre-ban period, the percentage effective spread is 2.78% for the banned stocks and 2.56% for the non-banned stocks. In the ban period, the percentage effective spread increases to 4.26% for the banned stocks and 3.62% for the non-banned stocks, suggesting that effective spreads increase more for the banned stocks during the ban period. Results from the multivariate analysis show that the coefficient estimate of the ban dummy variable for both the percentage quoted spread and percentage effective spread is 0.0035, implying that, in the ban period, the percentage quoted and effective spreads of the banned stocks are greater by 35 basis points than the respective percentage quoted and effective spreads of the non-banned stocks. This coefficient estimate is statistically significant at the 1% level. Overall, the findings of the paper suggest deterioration in liquidity as measured by percentage quoted and effective spreads after the introduction of the short-selling ban.

3.2.2 Short-selling regulations and market liquidity in the UK

How short-selling regulations affected liquidity has also been investigated in the UK market, where FSA followed SEC in the US and introduced a short-selling ban on financial stocks on September 19, 2008. Clifton and Snape (2008) and Hansson and Fors (2009) focus on the impact of the short-selling ban on the price dimension of liquidity, or bid/ask spread, while Marsh and Payne (2012) incorporate into their analysis both price and quantity dimensions of liquidity as measured by the bid/ask spread and depth respectively.

All three studies construct a control sample of non-financial stocks that were not subject to the short-selling ban. However, the matching procedure of the non-banned stocks in the control sample to the banned stocks in the banned sample in each paper differs. For example, Clifton and Snape (2008) use the banned financial stocks that are constituents of the FTSE 100 as their banned sample and the remaining non-banned constituents of the FTSE 100 as their control sample. In the Hansson and Fors (2009) paper, the banned sample comprises 26 banned stocks that are components of the FTSE 350, while the control sample comprises the remaining constituents of the FTSE 350. On the other hand, Marsh and Payne (2012) match ten non-financial non-banned stocks to each of the 23 financial banned stocks by average market values computed over the first half of 2008.

The findings of the papers that examined the UK market can be summarised as follows. Clifton and Snape (2008) show that after the short-selling ban is enforced, the percentage quoted bid/ask spread on banned stocks increases by 140% from 15 basis points to 36 basis points. In contrast, there is a 66% increase in spreads on the control sample stocks from 12 basis points to 20 basis points. Hansson and Fors (2009) find that bid/ask spreads increase by 131% for the banned stocks and by 55% for the non-banned stocks. In the post-ban period, bid/ask spreads drop by 29% for the banned sample and by 9% for the control sample. Their multivariate analysis shows that the coefficient estimate of the ban dummy variable is 0.248, significant at the 1% level, implying that bid/ask spreads on the banned stocks are greater by 24.8% than the bid/ask spreads on the non-banned stocks during the ban period. Finally, Marsh and Payne (2012) show that quoted bid/ask spread on both financial banned stocks and non-financial non-banned stocks is similar and average around 19 basis points in the pre-ban period. However, during the ban period, quoted bid/ask spread on the banned stocks rises to 52 basis points, while quoted bid/ask spread on the non-banned stocks rises to only 35 basis points. Additionally, the authors examine the impact of the short-selling ban on the limit order book liquidity provision. They find that depth collapses at all prices in the order book for the banned stocks and that both sides of the book suffer from liquidity drains during the short-selling ban period. Although depth also deteriorates for the control stocks during the short-selling ban, the change in the book depth for banned stocks is much greater.

Overall, empirical evidence shows that liquidity deteriorates for both banned and control samples after the introduction of the short-selling ban in the UK; nevertheless, the liquidity deterioration is far more pronounced for the banned stocks.

3.2.3 Short-selling regulations and market liquidity: International evidence

Beber and Pagano (2013) examine the effects of the 2008 short-selling regulations on market liquidity as measured by the percentage bid/ask spread across 30 countries. Univariate analysis shows that after the introduction of short-selling bans, bid/ask spreads in countries such as the UK and Ireland are more than twice as large as their pre-ban levels. Regression analysis also shows that across all the countries bid/ask spreads are higher in the ban period and subsequently lower in the post-ban period. Total short-selling bans, which prohibit both naked and covered short-selling, are found to have a greater impact on liquidity than the naked short-selling bans alone. For example, the coefficient estimate of the naked ban dummy variable is 1.28 while the coefficient estimate of the total (naked and covered) ban dummy variable is 1.98. Both results are statistically significant at the 1% level and imply that bid/ask spreads increase by 1.28 and 1.98 percentage points under naked and total short-selling bans respectively. These results also suggest that bid/ask spreads increase more in countries where total short-selling bans were introduced. In contrast, disclosure requirements of short positions during the ban period are found to reduce bid/ask spreads. Furthermore, the authors find that bid/ask spreads increase more for smaller, more volatile stocks, and for stocks without listed options.

In order to control for funding liquidity problems and increased aggregate risk during the financial crisis, the authors use the TED spread (i.e., the spread between the US dollar

LIBOR rate and the US Treasury Bill rate) and the VIX (i.e., the CBOE implied volatility index for S&P 500 options) in their regressions. Moreover, the authors use the lagged values of the country-level credit default swap spreads for financial stocks and a financial stress index to account for endogeneity of the short-selling bans enactment. Again, the coefficient estimate of the short-selling ban dummy variable is found to be positive and significant, supporting previous results that short-selling bans are indeed associated with greater bid/ask spreads.

3.2.4 Summary

Bid/ask spreads are commonly used to estimate liquidity in a stock market. The greater the bid/ask spread on a stock, the less liquid this stock is considered to be. Empirical evidence presented in this section suggests that the introduction of the 2008 short-selling regulations is associated with an increase in bid/ask spreads on the stocks affected by these regulations (Bris (2008); Boulton and Braga-Aves (2010); Clifton and Snape (2008); Hansson and Fors, (2009); Marsh and Payne (2012); Beber and Pagano (2013)). This finding implies that liquidity is reduced for the banned stocks. However, liquidity has more than one dimension (Lee, Mucklow, and Ready (1993)). Consequently, investigating depth along with the bid/ask spread is important for drawing a more complete conclusion on the impact of short-selling regulations on the liquidity of the affected stocks. As discussed in this section, Marsh and Payne (2012) also incorporate depth in their analysis of the UK market and find that short-selling regulations are associated with both greater bid/ask spreads and lower depth. This result suggests that liquidity unambiguously deteriorates in the UK after the introduction of the short-selling ban.

The next section develops the hypotheses on the expected effects of the 2008 short-selling regulations on bid/ask spreads and two measures of depth.

3.3 Hypotheses Development

One consequence of the short-selling regulations introduced in the autumn of 2008 was the increase in bid/ask spreads of stocks subject to these regulations (see for example, Boulton and Braga-Aves (2010) and Marsh and Payne (2012)). Although market makers were exempt from the short-selling rules, other liquidity providers who supplied liquidity through their limit orders were not permitted to sell short. As a result, market makers had to manage a higher number of orders and thus increase the bid/ask spread to compensate for the greater risks involved in their transactions that related to inventory management (Beber and Pagano (2013)). Market makers also faced less competition from other liquidity providers, which increased their market power and subsequently the prices they charged for their services (Grundy, Lim, and Verwijmeren (2012)). According to both arguments, bid/ask spreads on stocks affected by short-selling regulations would be expected to increase following the introduction of these short-selling regulations.

Furthermore, theoretical models predict a positive relationship between bid/ask spreads and information asymmetry as market makers tend to increase bid/ask spreads in the presence of information asymmetry to protect themselves from better informed traders (see for example, Glosten and Milgrom (1985) and Kyle (1985)). During the 2008 short-selling bans, many short sellers, who are considered to be sophisticated and informed traders, were left out of the market. This increased uncertainty about the true value of the stocks subject to these bans and information asymmetry in the markets.[116] Based on the findings of earlier studies and the theoretical predictions, this essay hypothesizes that:

[116] For example, Aitken, Frino, McCorry, and Swan (1998) examine whether short-selling is bad news in an Australian setting where information is publicly available immediately after the execution of a short trade over the period 1994–1996. The authors find that short-selling conveys information which leads to the reassessment of the stock value. Dechow, Hutton, Meulbroek, and Sloan (2001) investigate whether short sellers target stocks of companies that are priced high relative to their fundamentals in the US over the period 1976–1993. The authors show that short sellers can indeed predict future returns and choose stocks that are temporarily

Hypothesis 1. The bid/ask spreads of European financial stocks subject to the 2008 short-selling bans increase (decrease) during the ban (post-ban) period.

Lee, Mucklow, and Ready (1993) argue that liquidity has two dimensions. Hence, it is essential to examine both dimensions (i.e., bid/ask spread and depth) in order to draw conclusions about liquidity in the markets. More precisely, liquidity unambiguously improves (deteriorates) if bid/ask spreads narrow (widen) and depth increases (decreases). Models that include both bid/ask spread and depth in their analysis predict that market makers quote wider spreads and narrower depths to protect themselves from traders with superior information in the presence of high information asymmetry (for example, Mann and Ramanlal (1996) and DuPont (2000)). Consequently, this essay hypothesizes that:

Hypothesis 2A. The depth of European financial stocks subject to the 2008 short-selling bans decreases (increases) during the ban (post-ban) period.

Depth is defined as the number of shares available for trading at the best bid and ask price. However, from an investor's perspective, the amount of currency that changes hands is more important (Chiu, Chung, Ho, and Wang (2012)). Therefore, this essay also examines the impact of the 2008 short-selling regulations on depth in currency terms (or euro depth) and hypothesizes that:

Hypothesis 2B. The euro depth of European financial stocks subject to the 2008 short-selling bans decreases (increases) during the ban (post-ban) period.

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overpriced. Desai, Krishnamurthy, and Venkataraman (2006), whose study examines US stocks for the period 1997–2002, also show that short sellers use accounting information, particularly conveyed by accruals, to target stocks of mispriced firms. Their results reveal that short sellers anticipate restatement announcements of previously reported financials and accordingly take positions to exploit the mispricing.

Short-selling bans that were introduced in Europe in 2008 took two forms: a naked short-selling ban and a total (naked and covered) short-selling ban.[117] A total short-selling ban is more stringent than a naked short-selling ban as in the latter case short-selling is still permitted if the short seller is able to borrow the stocks to be shorted. Consequently, this essay hypothesizes that:

Hypothesis 3. The liquidity of European financial stocks subject to the 2008 short-selling bans deteriorates more during the ban period if these bans prohibit both naked and covered short-selling.

Assuming that no new regulations are introduced in the derivatives markets, one would expect that derivatives trading would mitigate the deterioration in liquidity due to short-selling bans. Beber and Pagano (2013) find that the listing of options weakens the effects of short-selling bans on market liquidity across 30 countries. The authors argue that, under short-selling regulations, investors could still establish short positions on the underlying stocks via the option market, which should mitigate the impact of short-selling bans on market liquidity.[118] However, in eight out of the 10 European countries examined in this essay short-selling bans also covered the derivatives markets.[119] Consequently, this essay hypothesizes that:

[117] See Chapter 2 for a detailed description of short-selling rules in Europe.
[118] Latest empirical evidence shows that transaction costs in derivatives markets substantially increase after the introduction of the 2008 short-selling rules. For example, Grundy, Lim, and Verwijmeren (2012) examine whether bearish option strategies were used by short sellers to avoid the short-selling ban in the US. The authors find that during the ban there are increased spreads and less trading in the option market, implying that investors did not substitute short-selling in the spot market with short-selling in the option market. Battalio and Schultz (2011) also show that bid/ask spreads of options on stocks subject to the short-selling ban in the US increase more than those on stocks not subject to the short-selling ban. The authors argue that options market makers must have found it more difficult to hedge customers' long positions in puts or short positions in calls, so they increased ask prices of puts and lowered the bid prices of calls to discourage their customers from creating synthetic short positions.
[119] See Chapter 2 for more details.

Hypothesis 4. The liquidity of European financial stocks subject to the 2008 short-selling bans is not affected by the absence or presence of listed derivatives (options and/or futures) during the ban period.

3.4 Sample and Data

3.4.1 Introduction

Intraday quote and trade data on 78 financial stocks from 10 European countries that were subject to the short-selling bans are obtained from Thomson Reuters Tick History (TRTH) database of Securities Industry Research Centre of Asia-Pacific (SIRCA).[120] Data on market value is obtained from Datastream. The sample period, which spans from mid-January 2008 to mid-December 2009, is divided into three sub-periods: a pre-ban period covering the period before the introduction of the new short-selling regulations, a ban period covering the period when the short-selling regulations are in force, and a post-ban period covering the period where the short-selling regulations are lifted. Table 3.1 presents a summary of the type of short-selling bans introduced in each country, the number of financial stocks affected by these short-selling bans that are included in the sample, and the duration of each sample sub-period.[121] Additionally, Table 3.2 reports the trading hours of the stock exchanges in which the sample stocks primarily trade.[122]

[120] These countries are Belgium, Denmark, France, Germany, Ireland, the Netherlands, Portugal, Spain, Switzerland, and the UK.

[121] Several stocks of European financial institutions were excluded from the analysis for the following main reasons: (i) data insufficiency (Nordea AB – traded on Copenhagen Stock Exchange, AMB General Holdings, Julius Bar, Banco de Andalucia, Banco de Castilla, Banco de Credito Balear, Banco de Galicia, Banco de Vasconia, Caja de Ahorros del Mediterraneo, Inverfiate, Alliance & Leicester, Bradford & Bingley, Arbuthnot Banking Group, European Islamic Investment Bank, Friends Provident, HBOS, Highway Insurance Group, Islamic Bank of Britain, London Scottish Bank, and Tawa), (ii) unavailability of a good matching stock (HSBC Holdings – traded on Euronext Paris, Banco Popular Espanol and Banco Santander-traded on Euronext Lisbon), and (iii) bankruptcy of the financial institution (Just Retirement Holdings, British Insurance Holdings, Finibanco, Banco Guipuzcoano, Banco Pastor, Hypo Real Estate, Irish Life and Permanent, Van der Moolen Holdings, and Swiss Reinsurance).

[122] Information on trading hours was obtained from the respective stock exchanges' websites.

Table 3.1 Short-selling regulations in Europe

This table describes the type of short-selling bans that were introduced, the number of financial stocks affected by short-selling bans that are included in the sample, and the duration of each sample sub-period for each European country that is used in the analysis.

Country	**Type of the short-selling ban**	**Number of banned stocks in the sample**	**Pre-ban period**	**Ban period**	**Post ban period**
Belgium	Naked	4	15/01/2008-19/09/2008	22/09/2008-15/12/2009	N/A
France	Naked	10	15/01/2008-19/09/2008	22/09/2008-15/12/2009	N/A
Germany	Naked	9	15/01/2008-19/09/2008	22/09/2008-15/12/2008	N/A
Portugal	Naked	5	15/01/2008-23/09/2008	24/09/2008-15/12/2009	N/A
Spain	Naked	11	15/01/2008-23/09/2008	24/09/2008-15/12/2009	N/A
Denmark	Total	3	15/01/2008-10/10/2008	13/10/2008-15/12/2009	N/A
Ireland	Total	2	15/01/2008-18/09/2008	19/09/2008-15/12/2009	N/A
Netherlands	Total	6	15/01/2008-03/10/2008	06/10/2008-29/05/2009	01/06/2009-15/12/2009
Switzerland	Total	5	15/01/2008-18/09/2008	19/09/2008-16/01/2009	19/01/2009-15/12/2009
UK	Total	23	15/01/2008-18/09/2008	19/09/2008-16/01/2009	19/01/2009-15/12/2009
		Total: 78			

Table 3.2 Trading hours in European markets

Country	**Exchange**	**Trading hours (Local time)**
Belgium	Euronext Brussels	9am-5:30pm
France	Euronext Paris	9am-5:30pm
Germany	Frankfurt SE	9am-5:30pm
Portugal	Euronext Lisbon	9am-5:30pm
Spain	Madrid Stock Exchange	9am-5:30pm
Denmark	Copenhagen Stock Exchange	9am-5:00pm
Ireland	Irish Stock Exchange	8am-4:30pm
Netherlands	Euronext Amsterdam	9am-5:30pm
Switzerland	SIX Swiss Exchange	9am-5:30pm
UK	London Stock Exchange	8am-5:00pm

3.4.2 Data filtering

Raw trade and quote data obtained from the TRTH database of SIRCA undergo the following filtering process. Firstly, the first and last 15 minutes of trading are removed to control for volatility spikes that occur during the opening and closing trading sessions (Bessembinder and Kaufman (1997)). Secondly, data are winsorised at 1 and 99 percentiles, respectively, to ameliorate the impact of outliers on the results. Thirdly, to minimize errors, data with the following characteristics are excluded from the analysis (Huang and Stoll (1996)):

(i) Bid/ask quotes if the spread is greater than €3 or negative.[123]

(ii) Trade price p_t when $(p_t-p_{t-1})/p_{t-1} > 0.10$

(iii) Ask quote a_t when $(a_t-a_{t-1})/a_{t-1} > 0.10$

(iv) Bid quote b_t when $(b_t-b_{t-1})/b_{t-1} > 0.10$

The daily data that are used in this analysis are derived from the filtered intraday data.

3.4.3 Control sample

Short-selling regulations were introduced on financial stocks in a period of extreme market conditions. Consequently, the construction of a control sample would be appropriate. Each financial stock subject to a short-selling ban is matched to a non-financial stock from the same market not subject to a short-selling ban. The advantage of a control sample is that it removes market movements that could potentially drive the results rather than the short-selling bans per se. The disadvantage is that the control sample is comprised of stocks from other sectors and there might be other factors that affect the observed differences between the two samples. Nevertheless, if the matching is robust, the latter problem can be ameliorated.

[123] For stocks traded in the UK, Switzerland, and Denmark, bid/ask quotes are removed if the spread is greater than 2.5 GBP, 4 CHF, and 23 DKK, respectively.

This essay follows the matching procedure of Boehmer, Jones, and Zhang (2013). Each financial stock is matched to a non-financial stock that (i) is traded on the same stock exchange, and (ii) has similar characteristics and derivatives listings with the financial stock.[124] The matching score is computed as:

$$Score = \left(\frac{P_{banned} - P_{non-banned}}{\frac{P_{banned} + P_{non-banned}}{2}}\right)^2 + \left(\frac{EV_{banned} - EV_{non-banned}}{\frac{EV_{banned} + EV_{non-banned}}{2}}\right)^2 + \left(\frac{MV_{banned} - MV_{non-banned}}{\frac{MV_{banned} + MV_{non-banned}}{2}}\right)^2$$

where P_{banned} is the average price of each financial stock in the sample that is subject to the short-selling rules (banned stock) and $P_{non-banned}$ is the average price of a non-financial stock that is not subject to the short-selling rules (non-banned stock) over the period from January, 2008 to July, 2008. EV_{banned} is the average trading volume in euros (or turnover) for each banned stock, and $EV_{non-banned}$ is the average trading volume in euros of a non-banned stock over the period from January, 2008 to July, 2008.[125] Finally, MV_{banned} is the average market value of each banned stock, and $MV_{non-banned}$ is the average market value of a non-banned stock over the same period.

Table 3.3 reports the pairs of matched stocks. Column (1) reports the names of the banned firms while column (2) reports the names of the non-banned firms. Column (3) describes the derivatives status, i.e., whether the pair of matched firms has listed options and/or futures. Lastly, column (4) reports the matching score for each pair of matched firms.

[124] More specifically, the matched sample was identified as follows. Firstly, daily data on price, trading volume, and market value of all stocks traded on stock exchanges specified in Table 3.2 were downloaded from Datastream. Secondly, the daily data were averaged over the period January 2008–July 2008 for each variable and each stock. Thirdly, two groups were created: a group comprised of financial stocks subject to the short-selling bans and a group of stocks that were not affected by the short-selling bans. Fourthly, a matching code was written and used applying the SAS software to calculate a matching score for each financial stock subject to the short-selling regulations with each remaining stock not subject to the short-selling regulations. Fifthly, each financial stock subject to a short-selling ban was matched to the non-financial stock not subject to a short-selling ban with the lowest score. Each pair of matched stocks was also required to have similar derivatives status.

[125] Although the variable is named as EV (i.e., Euro Volume), turnover for British, Danish, and Swiss stocks is expressed in their domestic currencies (i.e., British pound, Danish krone, and Swiss franc, respectively).

Overall, as can be seen in column (4) of Table 3.3, the matching score is below 1 for most financial banned stocks. The ideal matching score would be 0, which would indicate that, on average, the matched stocks have identical characteristics such as price, trading volume, and market value. Matching scores below or close to 1 still indicate that the matched stocks have very similar characteristics. The average matching score across all stocks is 0.64. Notably, the matching scores for three out of four Belgian financial stocks and one Irish financial stock are much greater than 1, implying that there were no stocks traded on Euronext Brussels and Irish Stock Exchange, respectively, which matched closely the large stocks of the banned Belgian and Irish banks.[126]

[126] The average matching scores per country are: Belgium (2.28), Denmark (0.30), France (0.53), Germany (0.48), Ireland (2.73), the Netherlands (0.51), Portugal (0.87), Spain (0.71), Switzerland (0.61), and the UK (0.27).

Table 3.3 Pairs of matched stocks

This table reports the names of the banned firms that were subject to short-selling regulations (column 1) and their matched non-banned firms that were not subject to short-selling regulations (column 2). Column (3) reports the derivatives status of each pair of matched firms while column (4) reports the best lowest matching score achieved which is calculated as

$$Score = \left(\frac{P_{banned} - P_{non-banned}}{\frac{P_{banned} + P_{non-banned}}{2}}\right)^2 + \left(\frac{EV_{banned} - EV_{non-banned}}{\frac{EV_{banned} + EV_{non-banned}}{2}}\right)^2 + \left(\frac{MV_{banned} - MV_{non-banned}}{\frac{MV_{banned} + MV_{non-banned}}{2}}\right)^2$$

(1) Banned firm	(2) Non-banned firm	(3) Derivatives status	(4) Matching score
	Belgium		
Ageas (ex- Fortis)	GBL New	Options	2.11
KBC Group	Solvay	Options and futures	2.15
Dexia	Delhaize Group	Options and futures	3.62
KBC Ancora	CMB	No derivatives listed	1.24
	Denmark		
Danske Bank	Novo Nordisk B	Options and futures	0.64
Jyske Bank	William Demant Holdings	No derivatives listed	0.10
Sydbank	Torm	No derivatives listed	0.17
	France		
April Group	Fimalac	No derivatives listed	0.21
AXA	France Telecom	Options and futures	0.04
BNP Paribas	Sanofi	Options and futures	0.24
Credit Agricole	Vivendi	Options and futures	0.17
Natixis	Stmicroelectronics	Options and futures	0.66
CNP Assurances	Safran	Options and futures	0.67
Euler Hermes	Ipsen	No derivatives listed	0.27
NYSE Euronext	Biomerieux	Options	1.68
Scor Se	Jcdecaux	Options and futures	0.29
Societe Generale	Danone	Options and futures	1.09
	Germany		
Allianz	Siemens	Options and futures	0.26
Aareal Bank	Software	Futures	0.44
Commerzbank	Deutsche Telekom	Options and futures	1.48
Deutsche Boerse	Man	Futures	0.24
Deutsche Bank	Daimler	Options and futures	0.16
Deutsche Postbank	Continental	Futures	0.29
Hannover Ruck.	Gea Group	Options and futures	0.17
MLP	Comdirect bank	No derivatives listed	0.57
Münchener Ruck. G.	Rwe	Options and futures	0.71
	Ireland		
Allied Irish Banks	Independent News & Media	Futures	4.36
Bank of Ireland	Ryanair Holdings	Options and futures	1.09
	Netherlands		
Aegon	KPN Kon	Options and futures	0.19
Binckbank	Wessanen Kon. Certs.	Options	0.04
ING Groep	Royal Dutch Shell A	Options and futures	0.45
KAS Bank	Macintosh Retail	No derivatives listed	0.50
SNS Reaal	Vopak	Options	0.37
Van Lanschot	Hal Trust	No derivatives listed	1.53
	Portugal		
Millenium BCP	Sonae SGPS	Futures	1.90
Banco Espirito Santo	Brisa-Autsds. de Portugal	Futures	0.63
Banco BPI	Jeronimo Martins	Futures	0.36
Espirito Santo Financial Group	Martifer	No derivatives listed	0.82
Banif SGPS	Altri SGPS	No derivatives listed	0.63

Table 3.3 Pairs of matched stocks

Table continued from previous page

(1) Banned firm	(2) Non-banned firm	(3) Derivatives status	(4) Matching score
Spain			
Banco Santander	Telefonica	Options and futures	0.35
BBVA	Iberdrola	Options and futures	0.24
Banco Popular Espanol	Abertis Infraestructuras	Options and futures	1.89
Banco Espanol de Credito	Sacyr Vallehermoso	Options and futures	0.34
Banco de Sabadell	Ferrovial	Options and futures	0.92
Mapfre	Enagas	Options and futures	2.46
Bankinter	Acerinox R	Options and futures	0.35
Banco de Valencia	Zardoya Otis	Futures	0.36
Bolsas y Mercados Espanoles	Fomento Constr. Y Cntr.	Options and futures	0.54
Grupo Catalana Occidente	Almirall	No derivatives listed	0.25
Renta 4	Natra	No derivatives listed	0.14
Switzerland			
UBS	Novartis R	Options and futures	0.82
Credit Suisse	ABB R	Options and futures	0.51
Zurich Financial Services	Syngenta	Options and futures	0.54
Swiss Life Holding	Swisscom R	Options and futures	0.85
Baloise	Geberit R	Options and futures	0.32
UK			
Aberdeen Asset Management	Aegis Group	No derivatives listed	0.17
Admiral Group	Weir group	No derivatives listed	0.10
Alliance Trust	Ferrexpo	No derivatives listed	0.01
Aviva	National Grid	Options and futures	0.04
Barclays	Tesco	Options and futures	0.77
Chesnara	Diploma	No derivatives listed	0.02
Close Brothers Group	SIG	No derivatives listed	0.10
F&C Asset Management	BBA Aviation	No derivatives listed	0.06
HSBC Holdings	BP	Options and futures	0.10
Investec	William Hill	Options and Futures	0.13
Legal & General Group	Morrison(WM)Spmkts.	Options and futures	0.66
Lloyds TSB Group	BT Group	Options and futures	0.58
Novae Group	Dechra Pharmaceutical	No derivatives listed	0.16
Old Mutual	Kingfisher	Options and futures	0.56
Provident Financial	Travis Perkins	Futures	0.21
Prudential	BAE Systems	Options and futures	0.19
Rathbone Brothers	Synergy Health	No derivatives listed	0.14
Royal Bank of Scotland Group	Xstrata	Options and futures	0.16
RSA Insurance group	Cairn Energy	Options and futures	0.87
Schroders	Segro	Futures	0.06
St James' Place	Bwin Party Digital En	No derivatives listed	0.15
Standard Chartered	Diageo	Options and futures	0.16
Standard Life	Invensys	Options and futures	0.71

3.5 Methodology

3.5.1 Measurement of the dependent variables

This essay employs two measures of the bid/ask spread – the quoted half-spread and the effective half-spread – expressed in percentage terms. The quoted spread measures the cost of completing a buy and a sell trade if trades are executed at the quoted prices. In this essay, half-spreads are used to standardize on the cost associated with one trade because the spread is the cost of two trades (see Stoll (2000)). Consequently, execution costs for a single trade can be calculated as

$$\%\,Quoted\ half-spread = \frac{1}{2} \times \left[\frac{Ask_{it} - Bid_{it}}{M_{it}}\right] \times 100 \tag{3.1}$$

where Ask_{it} and Bid_{it} are the posted ask price and bid price, respectively, for stock i at time t and the quote midpoint $M_{it} = \frac{Ask_{it}+Bid_{it}}{2}$ is a proxy for the stock's true value.

The effective half-spread is a better measure of trading costs when transactions take place inside or outside the quoted spread. It is computed as

$$\%\,Effective\ half-spread = D_{it} \times \left[\frac{P_{it} - M_{it}}{M_{it}}\right] \times 100 \tag{3.2}$$

where P_{it} is the transaction price for security i at time t; D_{it} is a trade indicator variable that equals 1 for buyer-initiated orders (i.e., when the transaction price is greater than the quote midpoint) and -1 for seller-initiated orders (i.e., when the transaction price is less than the quote midpoint).[127] M_{it} is the quote midpoint.

[127] See Lee and Ready (1991).

The volume of liquidity is measured by depth and euro depth. Depth is measured as the average number of shares that can be traded at the best bid and ask quotes (Heflin, Shaw, and Wild (2005)) and it is calculated as

$$Depth = \frac{(Bid\ Size + Ask\ Size)}{2} \tag{3.3}$$

Euro depth is calculated as the average of the sum of the number of shares quoted at the best ask price plus the number of shares quoted at the best bid price, times their respective quoted prices (Heflin, Shaw, and Wild (2005)). It is calculated as

$$Euro\ depth = \frac{(Bid\ Size \times Bid\ Price + Ask\ Size \times Ask\ Price)}{2} \tag{3.4}$$

3.5.2 Regression modeling

Financial regulators enforced restrictive short-selling regimes, particularly on financial stocks, as a response to the global financial meltdown. This essay conducts a difference-in-difference multivariate analysis to ameliorate the impact of severe market movements during the sample period, using daily data. To test Hypotheses 1, 2A, and 2B, firstly, the regression analysis is run at a country level and has the following form as shown in Boehmer, Jones, and Zhang (2013)[128]:

$$Diff\_Y_{it} = \alpha_i + \beta D_{it}^{Ban} + \theta X_{it} + e_{it} \tag{3.5}$$

where $Diff\_Y_{it}$ is the difference between the liquidity measure of the banned stocks (B) and the liquidity measure of the control or non-banned stocks (NB). Y_{it} is either the percentage quoted half-spread (PQHS), percentage effective half-spread (PEHS), natural logarithm of depth (depth), or natural logarithm of depth in currency terms (euro depth). X_{it} is a vector of

[128] Belgium, Denmark, and Ireland are excluded from the regression analysis at a country level due to the small sample size.

the differences in the values of the control variables of the matched pairs of stocks. The control variables that are included in the analysis are:

(i) Market value

(ii) Trading volume in euros (in their respective domestic currencies for the UK, Switzerland, and Denmark)

(iii) Proportional daily range of transaction prices

(iv) Daily value-weighted average price (VWAP)

D_{it}^{Ban} is a dummy variable that takes the value of 1 when short-selling bans are in force and 0 otherwise. This variable is of most interest to the analysis as it measures the impact of the short-selling bans on the liquidity of the financial stocks subject to these bans after controlling for other factors that could drive the results.

The four control variables have been documented in the literature as determinants of the bid/ask spread. For example, a negative relationship between the market value of a firm and the bid/ask spread on that firm's stock is found in Stoll (2000) and Boulton and Braga-Aves (2010). Trading volume is also inversely related to the bid/ask spread (see Tinic and West (1972), McInish and Wood (1992), and Manyah and Paudyal (1996) among others). Barnea (1974), Stoll (1978a, b), and Tripathy and Peterson (1991) report a negative relationship between dollar trading volume and the bid/ask spread. In contrast, information asymmetry models predict a positive relationship between price volatility and the bid/ask spread (see for example, Copeland and Galai (1983) and Easley and O'Hara (1987)). Lastly, an inverse relationship between the percentage bid/ask spread and the stock's price is documented in Stoll (1978a, b), Aitken and Frino (1996), and Menyah and Paudyal (1996). With regards to depth, models predict that market makers quote wider spreads and narrower depths to protect themselves from traders with superior information in the presence of high information asymmetry or inventory management considerations (see for example, Mann and Ramanlal (1996) and DuPont (2000)). Consequently, one would expect the opposite relationship between depth and the control variables from that documented for the relationship between the bid/ask spread and the control variables.

The same analysis is conducted separately for countries that have a post-ban period available, as specified in Table 3.1 of section 3.4.1.[129] A post-ban period dummy variable – $D_{it}^{Post\text{-}ban}$ – is added to equation (3.5), which now takes the following form:

$$Diff\_Y_{it} = \alpha_i + \beta D_{it}^{Ban} + \gamma D_{it}^{Post\_ban} + \theta X_{it} + e_{it} \qquad (3.6)$$

where $D_{it}^{Post\text{-}ban}$ is equal to 1 when short-selling bans are lifted and 0 otherwise.

Secondly, matched pairs of stocks from all countries are pooled together.[130] To test Hypotheses 1, 2A, and 2B at an aggregate level, again equation (3.5) is used. To test Hypothesis 3, two dummy variables are added to equation (3.5), which creates the specification below:

$$Diff\_Y_{it} = \alpha_i + \beta D_{it}^{N} + \gamma D_{it}^{T} + \theta X_{it} + e_{it} \qquad (3.7)$$

where D_{it}^{N} takes the value of 1 if the short-selling ban is naked and 0 otherwise, and D_{it}^{T} equals 1 if a total (naked and covered) short-selling ban is introduced and 0 otherwise. Additionally, equation (3.7) is run separately for stocks with listed derivatives and stocks without listed derivatives to test Hypothesis 4.

Fixed effects are used to account for any differences between the matched pairs of stocks in the pre-ban period. Furthermore, the control variables undergo logarithmic transformation to reduce any skewness and heteroskedasticity problems that might occur in the regression

[129] These countries are the Netherlands, Switzerland, and the UK.
[130] Because only three countries out of 10 have a post-ban period available, the regression analysis at an aggregate level excludes the post-ban period.

analysis, as in Benston and Hagerman (1974). Lastly, the Newy-West procedure is used to generate estimates robust to autocorrelation and heteroskedasticity in residuals.

3.6 Empirical Results

3.6.1 Univariate analysis

This section reports the results of the univariate analysis for four liquidity measures, which are the percentage quoted half-spread (PQHS), percentage effective half-spread (PEHS), depth, and euro depth as defined in section 3.5.1. Because mean values are sensitive to outliers, this section focuses on the median values of all the liquidity measures. However, mean values are provided in Appendix 3A.

Table 3.4 shows the average median values of PQHS for both banned (B) and non-banned (NB) stocks, respectively, along with the difference in PQHS between the two samples, across 10 European countries. PQHS (B) – PQHS (NB) is the median of the difference between the PQHS of the banned stocks and the PQHS of the non-banned stocks (or relative PQHS of the banned stocks).[131] Variable Diff is the change in the median PQHS between the pre-ban and ban periods, and ban and post-ban periods. Diff-in-Diff is the change in the relative PQHS of the banned stocks between the different sub-periods.

[131] Although a robust matching process is used as described in section 3.4.3, stocks in each sample belong to different industries and thus might be affected by different factors, which could lead to discrepancies in their difference variable (i.e., PQHS (B) – PQHS (NB)). Consequently, the average median value of the variable PQHS (B) – PQHS (NB) is different from the difference simply between the average median value of PQHS (B) and the average median value of PQHS (NB). The same holds for the other three liquidity variables.

Table 3.4 Median percentage quoted half-spread (PQHS) for banned (B) and non-banned (NB) stocks

This table reports the average median values of the percentage quoted half-spread (PQHS) for the banned (B) and non-banned (NB) stocks, respectively, across 10 European countries. The sample period spans from mid-January 2008 to mid-December 2009 and it is divided into three sub-periods: the pre-ban period, the ban period, and the post-ban period. The results for countries that have two sub-periods, the pre-ban period and the ban period (i.e., Belgium, Denmark, France, Germany, Ireland, Portugal, and Spain) are shown in Panel A. Only Switzerland, the Netherlands, and the UK have a post-ban period available. Their results are reported in Panel B. The sub-periods differ for each country and are specified in Table 3.1. Table 3.4 also shows the statistical significance of the test for equality of medians of the PQHS for both samples as well as their difference (i.e., PQHS (B) – PQHS (NB)).Variable Diff is the change in the median PQHS between the pre-ban (P1) and ban (P2) periods (P2–P1), and ban and post-ban (P3) periods (P3–P2), respectively. Diff-in-Diff is the change in the difference between the PQHS of the banned stocks and the PQHS of the non-banned stocks from the pre-ban period to the ban period (P2–P1), and from the ban period to the post-ban period (P3–P2), respectively. The Wilcoxon test is used and the p-values of the Z-statistic are given in the parentheses. * indicates statistical significance at the 10% level. ** indicates statistical significance at the 5% level. *** indicates statistical significance at the 1% level.

Country	Sub-periods	(1) PQHS(B)	(2) PQHS(NB)	(3) PQHS(B) – PQHS(NB)		(4) Diff PQHS(B)	(5) Diff PQHS(NB)	(6) Diff-in-Diff PQHS(B) – PQHS(NB)
Panel A: Countries with two sub-periods								
Belgium	Pre-ban	0.0452	0.0410	0.0047	P2-P1	0.0441***	0.007***	0.0306***
	Ban	0.0893	0.0480	0.0353	(p-value)	(0.0001)	(0.0001)	(0.0001)
Denmark	Pre-ban	0.1069	0.1022	-0.0023	P2-P1	0.0843***	0.0283***	0.0569***
	Ban	0.1912	0.1305	0.0546	(p-value)	(0.0001)	(0.0001)	(0.0001)
France	Pre-ban	0.0417	0.0344	0.0028	P2-P1	0.0046***	-0.0001	0.0108***
	Ban	0.0463	0.0343	0.0136	(p-value)	(0.0001)	(0.1318)	(0.0001)
Germany	Pre-ban	0.0550	0.0547	0.0001	P2-P1	0.0448***	0.0391***	0.0048***
	Ban	0.0998	0.0938	0.0049	(p-value)	(0.0001)	(0.0001)	(0.0001)
Ireland	Pre-ban	0.1543	0.2715	-0.1194	P2-P1	0.3548***	0.1387***	0.1354***
	Ban	0.5091	0.4102	0.0160	(p-value)	(0.0001)	(0.0001)	(0.0010)
Portugal	Pre-ban	0.1377	0.1047	0.0048	P2-P1	-0.0464***	-0.0224***	0.0083***
	Ban	0.0913	0.0823	0.0131	(p-value)	(0.0002)	(0.0001)	(0.0001)
Spain	Pre-ban	0.0714	0.0575	0.0132	P2-P1	0.0156***	0.0134***	-0.0009*
	Ban	0.0870	0.0709	0.0123	(p-value)	(0.0001)	(0.0001)	(0.0756)
Panel B: Countries with three sub-periods								
Netherlands	Pre-ban	0.0573	0.0467	0.0049	P2-P1	0.0277***	-0.0019	0.0284***
	Ban	0.0850	0.0448	0.0333	(p-value)	(0.0001)	(0.4969)	(0.0001)
	Post-ban	0.0557	0.0337	0.0139	P3-P2	-0.0293***	-0.0111***	-0.0194***
					(p-value)	(0.0001)	(0.0002)	(0.0001)
Switzerland	Pre-ban	0.0494	0.0471	0.0014	P2-P1	0.0182***	0.0095***	0.0089***
	Ban	0.0676	0.0566	0.0103	(p-value)	(0.0001)	(0.0001)	(0.0001)
	Post-ban	0.0662	0.0496	0.0145	P3-P2	-0.0014	-0.007***	0.0042***
					(p-value)	(0.3341)	(0.0001)	(0.0001)
UK	Pre-ban	0.0628	0.0589	0.0070	P2-P1	0.0427***	0.0076***	0.0247***
	Ban	0.1055	0.0665	0.0317	(p-value)	(0.0001)	(0.0001)	(0.0001)
	Post-ban	0.0687	0.0553	0.0064	P3-P2	-0.0368***	-0.0112***	-0.0253***
					(p-value)	(0.0001)	(0.0001)	(0.0001)

The data in column (1) of Panel A of Table 3.4, show that the PQHS on the banned stocks increases for all the countries from the pre-ban to the ban period with the exception of Portugal. For example, the average median value of the PQHS (B) in Belgium is 0.0452 in the pre-ban period, which increases to 0.0893 in the ban period. For Portuguese financial stocks, the PQHS (B) decreases from 0.1377 in the pre-ban period to 0.0913 in the ban period. The data in column (4) show that all changes in the PQHS (B) are statistically significant at the 1% level. However, the data in column (2) show that PQHS (NB) also increases for the majority of the countries during the ban period. For example, the PQHS for the Danish non-banned stocks increases from 0.1022 in the pre-ban period to 0.1305 in the ban period. The data in column (5) show a significant change in the PQHS (NB) at the 1% level for all countries except for France, where the result is insignificant. Again in the case of Portugal there is a significant decrease in the PQHS (NB) in the ban period. This result is also significant at the 1% level.

The results reported in column (6) show that the relative PQHS of the banned stocks between the pre-ban (P1) period and the ban (P2) period increases for all countries except for Spain, and the increase is statistically significant at the 1% level. For example, the relative PQHS of the banned stocks in Germany increases by 0.0048 in the ban period. In Portugal, although both the PQHS (B) and PQHS (NB) decrease in the ban period (see columns (4) and (5)), the relative PQHS of the banned stocks increases by 0.0083 in the ban period, which is significant at the 1% level. In the case of Spain, both the PQHS (B) and PQHS (NB) increase during the ban period (see columns (4) and (5)) but there is no statistically significant change in the relative PQHS of the banned stocks during the same period.

Panel B of Table 3.4 reports the results for countries with a post-ban period available. For the Dutch financial stocks, the data in column (1) show that there is an increase in the PQHS of the banned stocks in the ban period and a subsequent decrease of the PQHS of the banned stocks in the post-ban period. The data in column (4) show that both changes are significant at the 1% level. On the other hand, the data in column (5) show that there is no statistically significant change in the PQHS of the non-banned stocks between the pre-ban period and the ban period, while there is a significant decrease of the PQHS (NB) in the post-ban period. Similar behavior is observed for the British stocks. The only difference is that in the case of the UK, the PQHS (NB) also increases significantly during the ban period. The data in column (6) show that both countries exhibit a statistically significant increase in the relative PQHS of the banned stocks during the ban period and a subsequent statistically significant decrease in the same variable during the post-ban period. Lastly, the data in columns (1) and (2), respectively, show that the PQHS for both Swiss samples increase in the ban period. However, in the post-ban period, the PQHS (B) remains nearly the same while the PQHS (NB) significantly decreases. Also, as can be seen in column (3), the relative PQHS of the Swiss banned stocks increases from 0.0014 in the pre-ban period to 0.0103 in the ban period and then further increases to 0.0145 in the post-ban period. The data in column (6) show that both results are statistically significant at the 1% level.

Figure 3.1 depicts the time series of the average PQHS for banned and non-banned stocks, respectively, covering a period of four months before the introduction of the short-selling bans and four months after the introduction of the short-selling bans. The UK was one of the first countries to apply short-selling regulations on September 19, 2008, and to subsequently lift these short-selling regulations on January 16, 2009. These two dates are chosen as the starting and ending points of the short-selling bans in the figure.[132] As can be seen in Figure 3.1, the PQHS of the banned stocks was significantly higher relative to the PQHS of the non-banned stocks after the introduction of the short-selling bans.

[132] Other European countries followed the UK and introduced short-selling regulations within the following two weeks.

Figure 3.1 Percentage quoted half-spread: Banned versus non-banned stocks

This figure shows the time series of the average percentage quoted half-spread (PQHS) aggregated across all banned and non-banned stocks, respectively. The time series for both groups of stocks covers four months before the introduction of the short-selling bans and four months during the short-selling bans. The starting and ending points of the short-selling bans in this figure are September 19, 2008, and January 16, 2009 – the dates of the introduction and end of the short-selling ban in the UK.

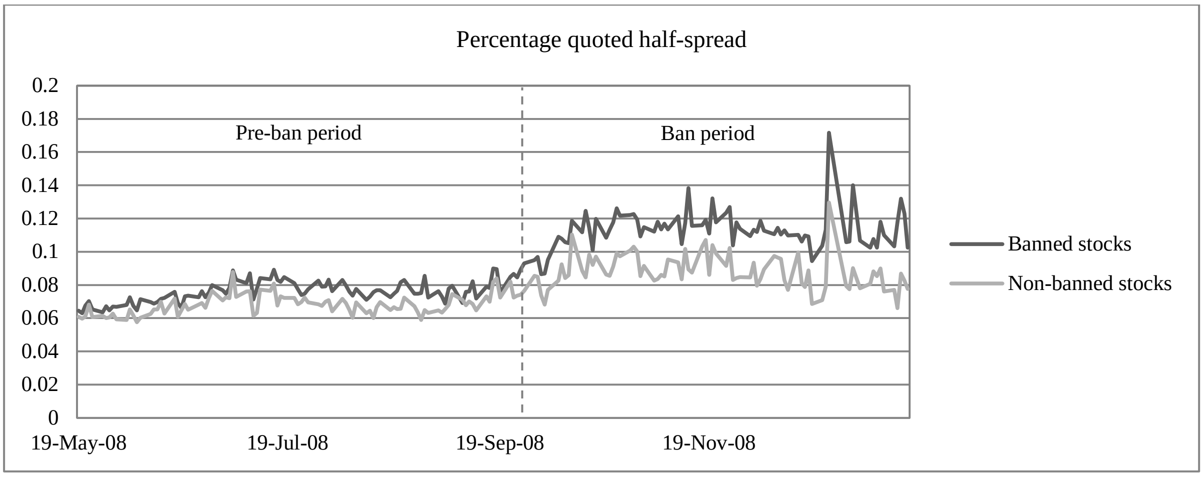


Table 3.5 reports the average median values of the PEHS for the banned (B) and non-banned (NB) stocks, respectively, as well as the difference in PEHS between the two samples. Changes in the PEHS are qualitatively similar to those in the PQHS reported in Table 3.4. In Panel A of Table 3.5, the relative PEHS of the banned stocks increases during the ban period in six out of seven countries (see column (6)). In Panel B of Table 3.5, the relative PEHS of banned stocks increases during the ban period and subsequently declines in the post-ban period for the Netherlands and the UK (see column (6)). However, there are some differences. For example, the data in column (6) of Panel B of Table 3.5 show that in the post-ban period, the relative PEHS of the Swiss banned stocks slightly falls, but the fall is not statistically significant. In contrast, column (6) of Panel B of Table 3.4 reports a statistically significant increase in the relative PQHS of the Swiss banned stocks in the same sub-period.

Table 3.5 Median percentage effective half-spread (PEHS) for banned (B) and non-banned (NB) stocks

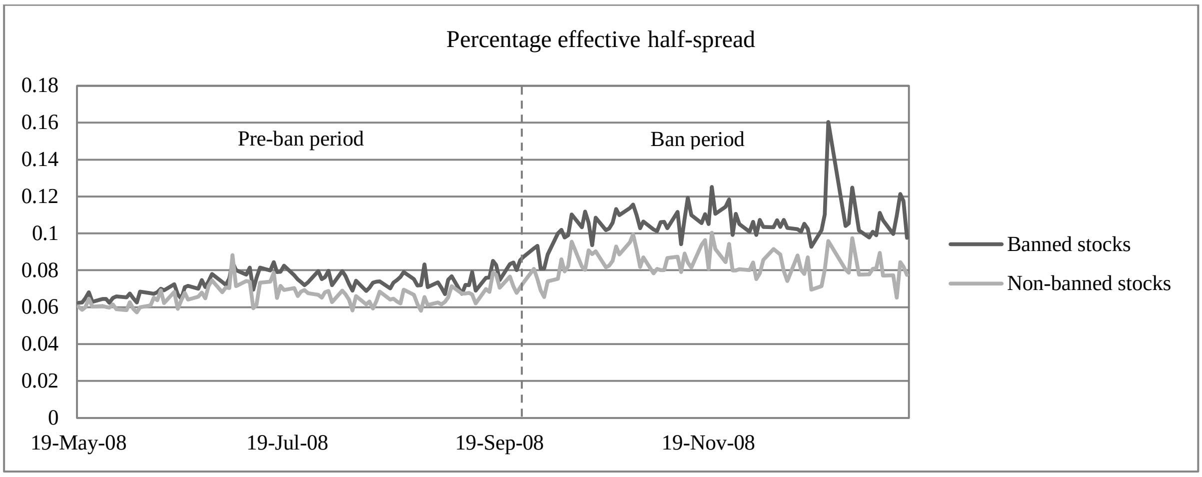
This table reports the average median values of the percentage effective half-spread (PEHS) for the banned (B) and non-banned (NB) stocks, respectively, across 10 European countries The sample period spans from mid-January 2008 to mid-December 2009 and it is divided into three sub-periods: the pre-ban period, the ban period, and the post-ban period. The results for countries that have two sub-periods, the pre-ban period and the ban period (i.e., Belgium, Denmark, France, Germany, Ireland, Portugal, and Spain) are shown in Panel A. Only Switzerland, the Netherlands, and the UK have a post-ban period available. Their results are reported in Panel B. The sub-periods differ for each country and are specified in Table 3.1. Table 3.5 also shows the statistical significance of the test for equality of medians of the PEHS for both samples as well as their difference (i.e., PEHS (B) – PEHS (NB)).Variable Diff is the change in the median PEHS between the pre-ban (P1) and ban (P2) periods (P2–P1), and ban and post-ban (P3) periods (P3–P2), respectively. Diff-in-Diff is the change in the difference between the PEHS of the banned stocks and the PEHS of the non-banned stocks from the pre-ban period to the ban period (P2–P1), and from the ban period to the post-ban (P3–P2) period, respectively. The Wilcoxon test is used and the p-values of the Z-statistic are given in the parentheses. * indicates statistical significance at the 10% level. ** indicates statistical significance at the 5% level. *** indicates statistical significance at the 1% level.

Panel A: Countries with two sub-periods								
		(1)	(2)	(3)		(4)	(5)	(6)
Country	Sub-periods	PEHS(B)	PEHS(NB)	PEHS(B) – PEHS(NB)		Diff PEHS(B)	Diff PEHS(NB)	Diff-in-Diff PEHS(B) – PEHS(NB)
Belgium	Pre-ban	0.0479	0.0440	0.0044	P2-P1	0.0527***	0.0069***	0.0381***
	Ban	0.1006	0.0509	0.0425	(p-value)	(0.0001)	(0.0001)	(0.0001)
Denmark	Pre-ban	0.1047	0.0995	0.0012	P2-P1	0.0777***	0.0173***	0.0564***
	Ban	0.1824	0.1168	0.0576	(p-value)	(0.0001)	(0.0001)	(0.0001)
France	Pre-ban	0.0454	0.0374	0.0038	P2-P1	0.0053***	-0.0011**	0.0115***
	Ban	0.0507	0.0363	0.0153	(p-value)	(0.0001)	(0.0174)	(0.0001)
Germany	Pre-ban	0.0447	0.0430	0.0011	P2-P1	0.0307***	0.0243***	0.0055***
	Ban	0.0754	0.0673	0.0066	(p-value)	(0.0001)	(0.0001)	(0.0001)
Ireland	Pre-ban	0.1211	0.2099	-0.0802	P2-P1	0.2871***	0.1315***	0.1196***
	Ban	0.4082	0.3414	0.0394	(p-value)	(0.0001)	(0.0001)	(0.0017)
Portugal	Pre-ban	0.1404	0.1176	0.0036	P2-P1	-0.0379***	-0.0263***	0.0113***
	Ban	0.1025	0.0913	0.0149	(p-value)	(0.0012)	(0.0001)	(0.0001)
Spain	Pre-ban	0.0758	0.0608	0.0127	P2-P1	0.0146***	0.0136***	-0.0001**
	Ban	0.0904	0.0744	0.0126	(p-value)	(0.0001)	(0.0001)	(0.0309)
Panel B: Countries with three sub-periods								
Netherlands	Pre-ban	0.0598	0.0493	0.0061	P2-P1	0.0296***	-0.0015	0.0317***
	Ban	0.0894	0.0478	0.0378	(p-value)	(0.0001)	(0.5087)	(0.0001)
	Post-ban	0.0594	0.0365	0.0164	P3-P2	-0.03***	-0.0113***	-0.0214***
					(p-value)	(0.0001)	(0.0001)	(0.0001)
Switzerland	Pre-ban	0.0515	0.0482	0.0019	P2-P1	0.0192***	0.0107***	0.0101***
	Ban	0.0707	0.0589	0.0120	(p-value)	(0.0001)	(0.0001)	(0.0001)
	Post-ban	0.0573	0.0464	0.0119	P3-P2	-0.0134***	-0.0125***	-0.0001
					(p-value)	(0.0001)	(0.0001)	(0.5082)
UK	Pre-ban	0.0566	0.0577	0.0014	P2-P1	0.0327***	0.0057***	0.0183***
	Ban	0.0893	0.0634	0.0197	(p-value)	(0.0001)	(0.0001)	(0.0001)
	Post-ban	0.0659	0.0541	0.0093	P3-P2	-0.0234***	-0.0093***	-0.0104***
					(p-value)	(0.0001)	(0.0001)	(0.0001)

Figure 3.2 depicts the time series of the average PEHS for banned and non-banned stocks, respectively, four months before the introduction of the short-selling bans and four months after the introduction of the short-selling bans. Figure 3.2 shows that after the introduction of the short-selling bans, the PEHS of the banned stocks was significantly higher relative to the PEHS of the non-banned stocks.

Figure 3.2 Percentage effective half-spread: Banned versus non-banned stocks

This figure shows the average percentage effective half-spread (PEHS) aggregated across all banned and non-banned stocks, respectively. The time series for both groups of stocks covers four months before the introduction of the short-selling bans and four months during the short-selling bans. The starting and ending points of the short-selling bans in this figure are September 19, 2008, and January 16, 2009 – the dates of the introduction and end of the short-selling ban in the UK.



In summary, the univariate analysis of the price dimension of liquidity (or bid/ask spread) provides support for Hypothesis 1. The relative bid/ask spreads of the banned stocks increase significantly in the ban period for nine out of 10 countries and decrease significantly in the post-ban period for two out of three countries.

Table 3.6 reports the average median values of depth for the banned (B) and non-banned (NB) stocks respectively, as well as the difference in depth between the two samples. The data in column (1) of Panel A of Table 3.6 show that the depth (B) decreases for Belgium, France, Portugal, and Spain, and increases for Denmark, Germany, and Ireland in the ban period. The data in column (4) show that all the changes are statistically significant, with exception the changes in Portugal and Spain. On the other hand, as can be seen in column (2), the depth (NB) decreases for France, Germany, and Portugal and increases for the remaining countries during the ban period. The data in column (5) show that the change in the depth (NB) is statistically significant at the 1% level for all countries with the exception of Belgium and Germany where the results are insignificant.

The results in column (6) of Panel A of Table 3.6 show that the relative depth of the banned stocks exhibits a significant drop in Belgium and Spain while the opposite is found for Germany, Ireland, France, and Portugal during the ban period. There is marginally significant increase in the relative depth of the banned stocks during the ban period for Denmark. For example, the data in column (3) show that, in the pre-ban period, the relative depth of the Spanish banned stocks is 1,106, which implies that during that period, on average the number of shares traded for the banned sample is greater by that amount compared to the number of shares traded for the non-banned sample. In the ban period, this difference decreases by 228 (see column (6)) to 878 shares (see column (3)). The data in column (6) show that this difference is statistically significant at the 1% level. Alternatively, the data in column (3) show that the relative depth for Irish banned stocks, in the pre-ban period is -3,187, indicating that the non-banned stocks are more liquid than the banned stocks during that period. In the ban period, as can be seen in column (3), the difference decreases to -567 shares, suggesting an improvement in depth for the Irish banned stocks relative to the depth of the Irish non-banned stocks. The data in column (6) show that this difference is statistically significant at the 1% level.

Panel B of Table 3.6 reports the average median values of the depth for banned (B) and non-banned (NB) stocks, respectively, as well as the relative depth of the banned stocks for countries with a post-ban period available. Again, the results are mixed. The data in column (1) of Panel B of Table 3.6 show that the depth (B) decreases during the ban period for Switzerland and the UK while it increases for the Netherlands. All the changes are statistically significant at the 1% level, as reported in column (6). Moreover, the data in column (2) show that the depth (NB) exhibits the same behavior as the depth (B) for Switzerland and the UK (see columns (2) and (5)); however, there is no significant change in the depth (NB) for the Netherlands, as can be seen in column (6).

In the post-ban period, the data in column (4) of Panel B of Table 3.6 show that the depth (B) decreases for the Netherlands, increases for Switzerland, and remains nearly unchanged for the UK. On the other hand, the depth (NB) remains unchanged for the Netherlands, but increases for both Switzerland and the UK, as reported in column (5). The data in column (3) show that the relative depth of the Dutch banned stocks increases from 93 shares in the pre-ban period to 673 shares in the ban period and subsequently decreases to -54 in the post-ban period. The results reported in column (6) imply relatively better liquidity for the Dutch banned stocks in the ban period and relatively worse liquidity for the banned stocks in the post-ban period. For British and Swiss banned stocks, the relative depth significantly decreases during the ban period and then significantly decreases even more in the post-ban period, suggesting worse liquidity for the banned stocks in both ban and post-ban periods.

Table 3.6 Median depth for banned (B) and non-banned (NB) stocks

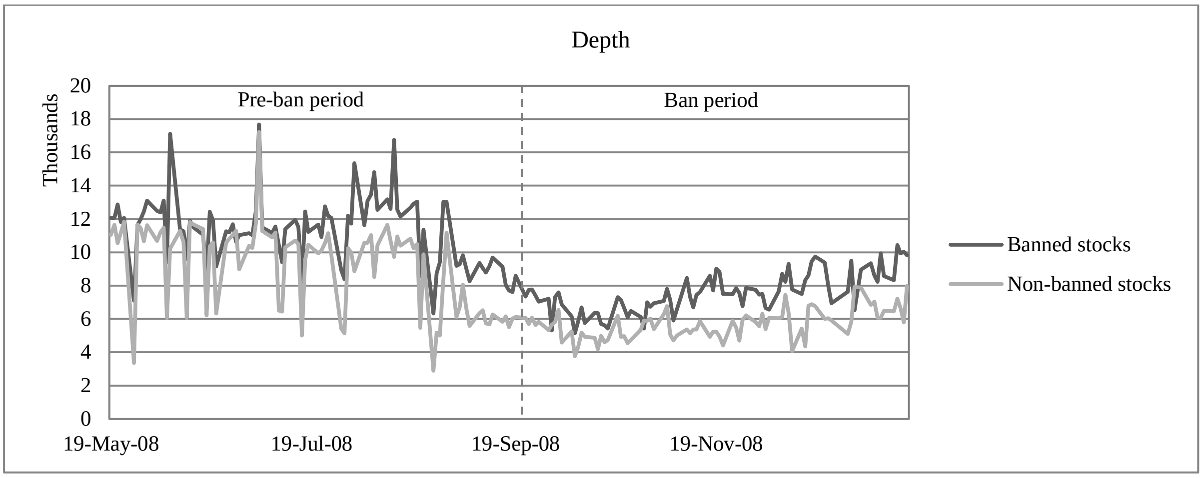
This table reports the average median values of the depth for the banned (B) and non-banned (NB) stocks, respectively, across 10 European countries. The sample period spans from mid-January 2008 to mid-December 2009 and it is divided into three sub-periods: the pre-ban period, the ban period, and the post-ban period. The results for countries that have two sub-periods, the pre-ban period and the ban period (i.e., Belgium, Denmark, France, Germany, Ireland, Portugal, and Spain) are shown in Panel A. Only Switzerland, the Netherlands, and the UK have a post-ban period available. Their results are reported in Panel B. The sub-periods differ for each country and are specified in Table 3.1. Table 3.6 also shows the statistical significance of the test for equality of medians of the depth for both samples as well as their difference (i.e., depth (B) – depth (NB)).Variable Diff is the change in the median depth between the pre-ban (P1) and ban (P2) periods (P2–P1), and ban and post-ban (P3) periods (P3–P2), respectively. Diff-in-Diff is the change in the difference between the depth of the banned stocks and the depth of the non-banned stocks from the pre-ban period to the ban period (P2–P1), and from the ban period to the post-ban period (P3–P2), respectively. The Wilcoxon test is used and the p-values of the Z-statistic are given in the parentheses. * indicates statistical significance at the 10% level. ** indicates statistical significance at the 5% level. *** indicates statistical significance at the 1% level.

Panel A: Countries with two sub-periods								
		(1)	(2)	(3)		(4)	(5)	(6)
Country	Sub-periods	Depth(B)	Depth(NB)	Depth(B) – depth(NB)		Diff Depth(B)	Diff Depth(NB)	Diff-in-Diff Depth(B) – depth (NB)
Belgium	Pre-ban	1,907	303	1,559	P2-P1	-367***	16	-458***
	Ban	1,540	319	1,101	(p-value)	(0.0001)	(0.3403)	(0.0001)
Denmark	Pre-ban	1,435	1,919	241	P2-P1	651***	791***	103*
	Ban	2,086	2,710	344	(p-value)	(0.0001)	(0.0018)	(0.0905)
France	Pre-ban	588	1,110	-90	P2-P1	-15***	-368***	98***
	Ban	573	742	8	(p-value)	(0.0001)	(0.0001)	(0.0001)
Germany	Pre-ban	772	1,105	-88	P2-P1	115***	-88	109***
	Ban	887	1,017	21	(p-value)	(0.0001)	(0.0929)	(0.0001)
Ireland	Pre-ban	3,205	6,606	-3,187	P2-P1	4,757***	937***	2,620***
	Ban	7,962	7,543	-567	(p-value)	(0.0001)	(0.0001)	(0.0001)
Portugal	Pre-ban	7,288	3,978	-127	P2-P1	-1,723	-1,334***	1,767***
	Ban	5,565	2,644	1,640	(p-value)	(0.6912)	(0.0001)	(0.0001)
Spain	Pre-ban	5,003	1,481	1,106	P2-P1	-703	426***	-228***
	Ban	4,300	1,907	878	(p-value)	(0.8661)	(0.0001)	(0.0003)
Panel B: Countries with three sub-periods								
Netherlands	Pre-ban	2,131	1,891	93	P2-P1	886***	700	580***
	Ban	3,017	2,591	673	(p-value)	(0.0010)	(0.9057)	(0.0001)
	Post-ban	2,437	2,372	-54	P3-P2	-580**	-219	-727***
					(p-value)	(0.0235)	(0.2481)	(0.0001)
Switzerland	Pre-ban	1,465	832	187	P2-P1	-779***	-387***	-69***
	Ban	686	445	118	(p-value)	(0.0001)	(0.0001)	(0.0017)
	Post-ban	825	621	113	P3-P2	139***	176***	-5**
					(p-value)	(0.0001)	(0.0001)	(0.0268)
UK	Pre-ban	10,906	8,767	-367	P2-P1	-4,599***	-2,534***	-620***
	Ban	6,307	6,233	-987	(p-value)	(0.0001)	(0.0001)	(0.0052)
	Post-ban	5,743	7,707	-1,415	P3-P2	-564	1,474***	-428***
					(p-value)	(0.5714)	(0.0001)	(0.0003)

Figure 3.3 depicts the time series of the average depth of the banned and non-banned stocks, respectively, four months before the introduction of the short-selling bans and four months after the introduction of the short-selling bans. Figure 3.3 shows that the depth of the banned stocks is generally higher relative to the depth of the non-banned stocks in the pre-ban period. The depth of banned stocks remains higher following the introduction of the short-selling bans, although the gap between the depths of the two samples narrows.

Figure 3.3 Depth: Banned versus non-banned stocks

This figure shows the average depth aggregated across all banned and non-banned stocks, respectively. The time series for both groups of stocks covers four months before the introduction of the short-selling bans and four months during the short-selling bans. The starting and ending points of the short-selling bans in this figure are September 19, 2008, and January 16, 2009 – the dates of the introduction and end of the short-selling ban in the UK.



Overall, the results of the univariate analysis do not provide strong evidence for Hypothesis 2A. The relative depth of the banned stocks significantly decreases only in four out of 10 countries in the ban period, while there is no significant increase in the relative depth of the banned stocks in the post-ban period in any of the three countries with a post-ban period available.

Table 3.7 presents the average median values of the euro depth (ED) for the banned (B) stocks and non-banned (NB) stocks, respectively, and the average median values of the relative ED of the banned stocks. The data in columns (1) and (2) of Panel A of Table 3.7 show that during the ban period, ED dramatically decreases for both samples. For example, the ED (B) in Ireland decreases from €25,480 in the pre-ban period to €12,751 in the ban period, while the respective figures for the ED (NB) are €17,392 and €13,247. The data in column (6) show that the relative ED of the banned stocks decreases in the ban period for all countries except for Germany, where the relative ED of the banned stocks increases during the ban period. All the changes are statistically significant at the 1% level for all countries except for Portugal, where the change in the relative ED of the banned stocks is insignificant (see column 6)). These results imply that the relative euro depth of the banned stocks deteriorates during the ban period with the exception of the German banned stocks.

The data in column (6) of Panel B of Table 3.7 show that the relative ED of the banned stocks decreases in all three countries during the ban period. The results are statistically significant at the 1% level. In the post-ban period, the relative ED of the banned stocks increases for the UK, decreases even more for Switzerland, and does not change significantly for the Netherlands. The results for the UK and Switzerland are statistically significant at the 1% level.

Table 3.7 Median euro depth (ED) for banned (B) and non-banned (NB) stocks

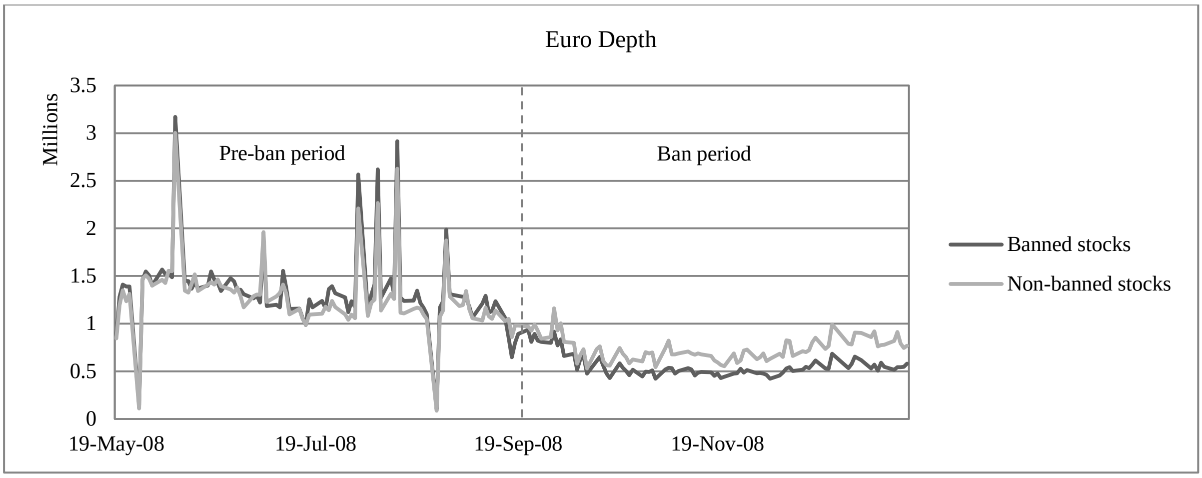
This table reports the average median values of the euro depth for the banned (B) and non-banned (NB) stocks, respectively, across 10 European countries. The sample period spans from mid-January 2008 to mid-December 2009 and it is divided into three sub-periods: the pre-ban period, the ban period, and the post-ban period. The results for countries that have two sub-periods, the pre-ban period and the ban period (i.e., Belgium, Denmark, France, Germany, Ireland, Portugal, and Spain) are shown in Panel A. Only Switzerland, the Netherlands, and the UK have a post-ban period available. Their results are reported in Panel B. The sub-periods differ for each country and are specified in Table 3.1. Table 3.7 also shows the statistical significance of the test for equality of medians of the ED for both samples as well as their difference (i.e., euro depth (B) – euro depth (NB)). Variable Diff is the change in the median ED between the pre-ban (P1) and ban (P2) periods (P2–P1), and ban and post-ban (P3) periods (P3–P2), respectively. Diff-in-Diff is the change in the difference between the ED of the banned stocks and the ED of the non-banned stocks from the pre-ban period to the ban period (P2–P1), and from the ban period to the post-ban period (P3–P2), respectively. The Wilcoxon test is used and the p-values of the Z-statistic are given in the parentheses. * indicates statistical significance at the 10% level. ** indicates statistical significance at the 5% level. *** indicates statistical significance at the 1% level.

Panel A: Countries with two sub-periods								
		(1)	(2)	(3)		(4)	(5)	(6)
Country	Sub-periods	ED(B)	ED(NB)	ED(B) – ED(NB)		Diff ED(B)	Diff ED(NB)	Diff-in-Diff ED(B) – ED(NB)
Belgium	Pre-ban	41,793	20,788	17,691	P2-P1	-28,835***	-5,212***	-18,366***
	Ban	12,958	15,576	-675	(p-value)	(0.0001)	(0.0001)	(0.0001)
Denmark	Pre-ban	329,192	389,858	-12,059	P2-P1	-81,152***	-91,762***	-37,218***
	Ban	248,040	298,096	-49,277	(p-value)	(0.0001)	(0.0001)	(0.0001)
France	Pre-ban	33,618	31,129	-321	P2-P1	-18,012***	-13,503***	-2,741***
	Ban	15,606	17,626	-3,062	(p-value)	(0.0001)	(0.0001)	(0.0001)
Germany	Pre-ban	29,695	31,713	-2,911	P2-P1	-12,501***	-13,741***	886***
	Ban	17,194	17,972	-2,025	(p-value)	(0.0001)	(0.0001)	(0.0001)
Ireland	Pre-ban	25,480	17,392	8,098	P2-P1	-12,729***	-4,145***	-4,887***
	Ban	12,751	13,247	3,211	(p-value)	(0.0001)	(0.0001)	(0.0001)
Portugal	Pre-ban	27,312	20,848	4,359	P2-P1	-12,451***	-9,110***	-2,181
	Ban	14,861	11,738	2,178	(p-value)	(0.0001)	(0.0001)	(0.4222)
Spain	Pre-ban	44,872	30,701	8,921	P2-P1	-15,273***	-7,033***	-6,702***
	Ban	29,599	23,668	2,219	(p-value)	(0.0001)	(0.0001)	(0.0001)
Panel B: Countries with three sub-periods								
Netherlands	Pre-ban	25,839	18,265	-1,150	P2-P1	-12,832***	-1,513***	-5,587***
	Ban	13,007	16,752	-6,737	(p-value)	(0.0001)	(0.0001)	(0.0001)
	Post-ban	16,826	18,311	-5,593	P3-P2	3,819***	1,559**	1,144
					(p-value)	(0.0001)	(0.0147)	(0.9869)
Switzerland	Pre-ban	232,851	242,512	16,962	P2-P1	-145,214***	-148,550***	-26,885***
	Ban	87,637	93,962	-9,923	(p-value)	(0.0001)	(0.0001)	(0.0001)
	Post-ban	107,234	139,592	-25,326	P3-P2	19,597***	45,630***	-15,403***
					(p-value)	(0.0001)	(0.0001)	(0.0003)
UK	Pre-ban	2,858,146	2,942,272	-354,944	P2-P1	-1,537,017***	-1,177,761***	-47,425***
	Ban	1,321,129	1,764,511	-402,369	(p-value)	(0.0001)	(0.0001)	(0.0066)
	Post-ban	1,399,744	1,897,434	-385,342	P3-P2	78,615**	132,923***	17,027***
					(p-value)	(0.0434)	(0.0001)	(0.0052)

Figure 3.4 plots the time series of the average euro depth of the banned and non-banned stocks, respectively, covering a period of four months before the introduction of the short-selling bans and four months after the introduction of the short-selling bans. The depth in currency terms for Denmark, Switzerland, and the UK – countries which use a different currency from the euro – is converted into euro depth. As shown in Figure 3.4, the euro depth of the banned stocks was noticeably lower relative to the euro depth of the non-banned stocks after the introduction of the short-selling bans.

Figure 3.4 Euro depth: Banned versus non-banned stocks

This figure shows the average euro depth aggregated across all banned and non-banned stocks, respectively. The time series for both groups of stocks covers four months before the introduction of the short-selling bans and four months during the short-selling bans. The starting and ending points of the short-selling bans in this figure are September 19, 2008, and January 16, 2009 – the dates of the introduction and end of the short-selling ban in the UK.



To sum up, the univariate analysis of the relative euro depth strongly supports Hypothesis 2B as the relative euro depth of the banned stocks decreases during the ban period for eight out of 10 countries. However, in the post-ban period, the relative depth of the banned stocks increases only for the UK, one of the three countries with a post-ban period available.

3.6.2 Summary of the results of the univariate analysis and discussion

Section 3.6.1 reported the results of the univariate analysis for four liquidity measures. PQHS and PEHS comprise the measures of price dimension of liquidity while depth and euro depth comprise the measures of quantity dimension of liquidity. The analysis focused on changes in the differences between the liquidity measures of the banned and non-banned stocks in an attempt to isolate the impact of short-selling regulations on the liquidity of the stocks affected by these regulations.

Hypothesis 1 is supported by the univariate analysis. The results in column (6) of Tables 3.4 and 3.5, respectively, show that the quoted and effective bid/ask spreads of the banned stocks widen more than the respective spreads of the non-banned stocks for nine out of 10 countries in the ban period, with the exception of Spain. Overall, these results support the findings of Beber and Pagano (2013), who document a significant increase in the percentage quoted spreads during the ban period for all countries. The main difference in conducting the univariate analysis between this essay and Beber and Pagano (2013) is that this essay focuses on the changes in the relative percentage bid/ask spreads of the banned stocks. More precisely, this essay incorporates a control sample of non-banned stocks into the analysis, while Beber and Pagano (2013) focus only on the changes in the percentage bid/ask spreads of the banned stocks.

Hypothesis 1 also asserts that bid/ask spreads decrease after the short-selling bans are lifted. This is supported in the case of the Netherlands and the UK but not Switzerland. In the case of the UK, the results support the findings of Hansson and Fors (2009). The authors also document a significant increase in the relative percentage quoted spreads of British banned stocks during the ban period and a significant subsequent decrease in the relative percentage quoted spreads of British banned stocks during the post-ban period.

Hypothesis 2B is also supported by the univariate analysis. The results in column (6) of Table 3.7 show that the relative euro depth of the banned stocks drops in eight out of 10 countries during the ban period, with the exception of Germany and Portugal. However, when the short-selling bans are lifted, the relative euro depth of the banned stocks increases only in the UK. On the other hand, there is no strong support for hypothesis 2A. The results in column (6) of Table 3.6 show that during the ban period, the relative depth of the banned stocks increases in six countries and decreases in only four countries. In the post-ban period, the relative depth of the banned stocks does not increase for any of the three countries with a post-ban period available.

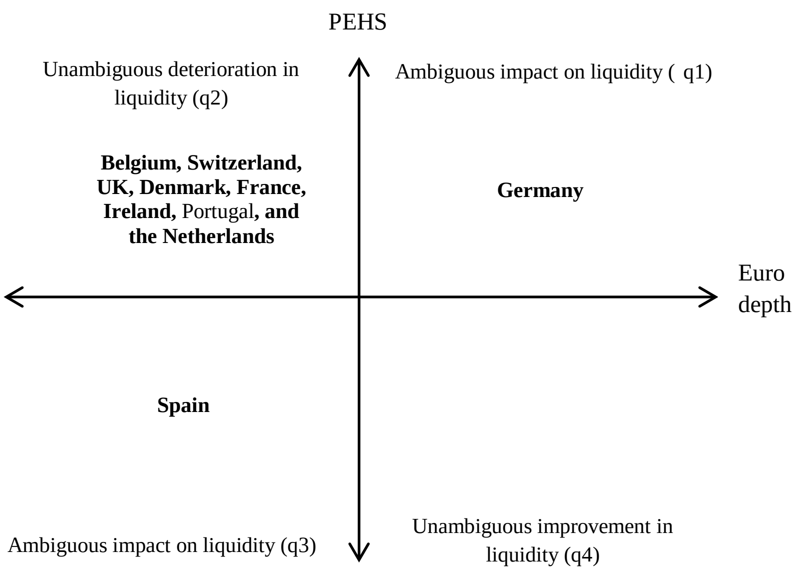
Euro depth is considered to be a more important measure of the quantity dimension of liquidity for investors (Chiu, Chung, Ho, and Wang (2012)). Examining the changes in the PEHS and euro depth across the sample countries reveals that liquidity unambiguously deteriorates for eight out of 10 countries in the ban period where the increase in the relative PEHS of the banned stocks is accompanied by a decrease in the relative euro depth of the banned stocks. In the case of Portugal, the increase in the relative bid/ask spread of the banned stocks is accompanied by a decrease in the relative euro depth of the same stocks although this decrease is not statistically significant. Thus, liquidity in Portugal deteriorates only in terms of higher bid/ask spreads. There is no clear impact on liquidity in Germany and Spain. In Germany, the deterioration in the price dimension of liquidity is accompanied by an improvement in the quantity dimension of liquidity. The opposite is observed for Spain. Figure 3.5 summarizes the results for both the price (PEHS) and quantity (euro depth) dimensions of liquidity between the pre-ban and ban periods (Panel A), when short-selling regulations are introduced, and between the ban and post-ban periods (Panel B), when short-selling regulations are lifted. Additionally, Appendix 3B provides charts that illustrate the changes in the four liquidity measures from one sub-period to the other for both banned and non-banned stocks at a country level.

The next section reports the results of the multivariate analysis that incorporates control variables such as differences in market value and turnover between the banned and non-banned stocks. These differences could potentially drive the univariate results that were reported in this section.

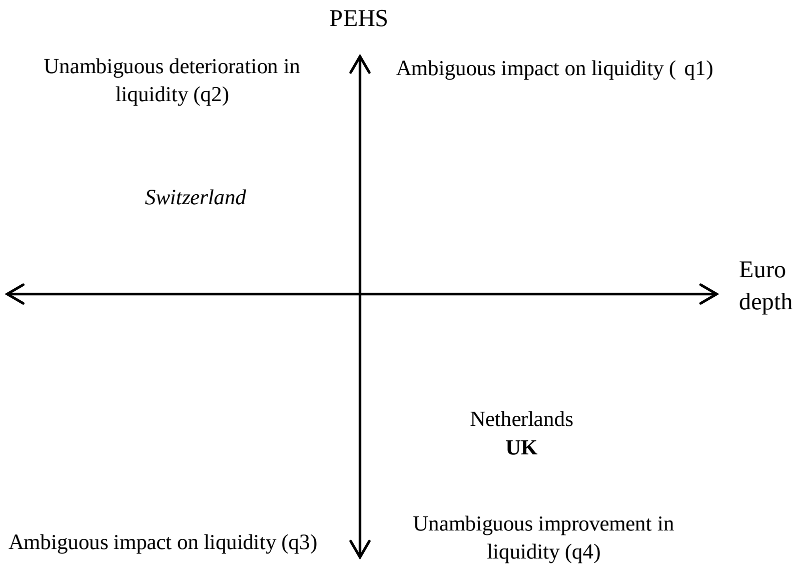
Figure 3.5 The impact of short-selling regulations on the liquidity of banned stocks

This figure summarizes the changes in both the relative price and quantity dimensions of liquidity of banned stocks as measured by the percentage effective-half spread (PEHS) and euro depth, respectively. Panel A illustrates the effects of introducing short-selling regulations on liquidity from the pre-ban period (P1) to the ban period (P2), while Panel B exhibits the impact of lifting short-selling regulations on liquidity from the ban period (P2) to the post-ban (P3) period. In quadrants 1 and 3 (q1 and q3) the impact on liquidity is ambiguous where an increase/decrease in the PEHS is accompanied by an increase/decrease in the euro depth. In quadrant 2 (q2), liquidity unambiguously deteriorates where an increase in the PEHS is associated with a decrease in the euro depth. Liquidity unambiguously improves in quadrant 4 (q4) where a decrease in the PEHS is accompanied by an increase in the euro depth. Countries in bold indicate statistically significant results for both the PEHS and euro depth changes. Countries in unbold indicate statistically significant results only for the PEHS. Countries in unbold and italics indicate statistically significant results only for the euro depth.

Panel A: Relative PEHS and euro depth of banned stocks (P1–P2)



Panel B: Relative PEHS and euro depth of banned stocks (P2–P3)



3.6.3 Multivariate analysis

3.6.3.1 Country-level liquidity effects

This section presents the results of the multivariate regression analysis. The dependent variables are the differences between the four liquidity measures of the banned stocks (B) and the respective liquidity measures of the non-banned stocks (NB). The purpose of this section is to incorporate control variables in difference form on the right-hand side of the equation to remove any potential effect on the differences of the dependent variables which are not related to short-selling regulations.

Tables 3.8 to 3.11 report the coefficient estimates from equations (3.5) and (3.6) as described in section 3.5.2 at a country level. Belgium, Denmark, and Ireland are excluded from the regression analysis at a country level due to the small sample size. These countries, however, are included in the regression analysis at an aggregate level where all the countries are pooled together (see Tables 3.12–3.14 in section 3.6.3.2).

Table 3.8 reports the results from two specifications. Specification (1) – i.e., equation (3.5) – is the equation used for the four countries with two sub-periods available: the pre-ban and the ban period, respectively. On the other hand, specification (2) – i.e., equation (3.6) – is run for the three countries with a post-ban period available. Data in Table 3.8 show that the coefficient estimate of the ban dummy variable is positive for all seven countries. This shows that the relative PQHS of the banned stocks – Diff_PQHS – widens for all countries during the ban period. However, the results are not statistically significant in the case of Germany and Switzerland. For example, the coefficient estimate of the ban dummy variable for France is 0.0094, significant at the 1% level, which implies that the relative PQHS of the French banned stocks increases by 94 basis points in the ban period.

In the post-ban period, the coefficient estimate of the post-ban dummy variable for the Dutch stocks is -0.0051, significant at the 5% level, suggesting that the relative PQHS of the Dutch banned stocks decreases by 51 basis points during the post-ban period. In contrast, the coefficient estimate for Switzerland is 0.0101, indicating that the relative PQHS of the Swiss banned stocks widens by 101 basis points in the post-ban period. In the case of the UK, the lifting of the short-selling ban is not found to have any significant effect on the Diff_PQHS.

The coefficient estimates of Diff_PRANGE, a proxy for volatility, are statistically significant for all countries except for France. In four out of seven countries, the coefficient estimates for Diff_VWAP are also statistically significant. These results suggest that changes in relative volatility and relative price levels help explain changes in the relative quoted spreads of the banned stocks.

Table 3.8 Country-level multivariate results: Percentage quoted-half spread (PQHS)

The following regression model has the same form as shown in Boehmer, Jones, and Zhang (2013):

$$Diff\_PQHS_{it} = \alpha_i + \beta D_{it}^{Ban} + \theta X_{it} + e_{it} \qquad (1)$$

where $Diff\_PQHS_{it}$ is the difference between the PQHS of the banned stocks (B) and the PQHS of the non-banned stocks (NB). X_{it} is a vector of differences between the control variables of the matched pair of stocks. These control variables are: market value – MV, trading volume in euros (or in their respective domestic currencies for the UK and Switzerland) – EUROVOL, proportional daily range of transaction prices – PRANGE, and daily volume-weighted average price – VWAP. D_{it}^{Ban} is a dummy variable that takes the value of 1 when short-selling bans are in force and 0 otherwise.

For countries with a post-ban period available (i.e., the Netherlands, Switzerland, and the UK), equation (1) takes the following form:

$$Diff\_PQHS_{it} = \alpha_i + \beta D_{it}^{Ban} + \gamma D_{it}^{Post\_ban} + \theta X_{it} + e_{it} \qquad (2)$$

where $D_{it}^{Post\text{-}ban}$ is equal to 1 when short-selling bans are lifted and 0 otherwise.

Fixed effects are used to control for any differences between the matched stocks in the pre-ban period. All the control variables are logarithmically transformed. The p-values of the coefficient estimates are given in the parentheses below these estimates. * indicates statistical significance at the 10% level. ** indicates statistical significance at the 5% level. *** indicates statistical significance at the 1% level.

Note: Belgium, Denmark, and Ireland are excluded from the regression analysis at a country level due to the small sample size.

		(1)				(2)		
	Predicted sign	France	Germany	Portugal	Spain	Netherlands	Switzerland	UK
Intercept		0.0027 (0.2443)	-0.0707*** (0.0001)	0.0302 (0.4987)	0.0122*** (0.0033)	-0.0200 (0.0978)	-0.0074*** (0.0001)	0.0062*** (0.0056)
Ban dummy variable	+	0.0094*** (0.0001)	0.0072 (0.4933)	0.0485** (0.0281)	0.0054** (0.0304)	0.0073*** (0.0001)	0.0033 (0.3332)	0.0176*** (0.0002)
Post-ban dummy variable	-					-0.0051** (0.0153)	0.0101*** (0.0001)	0.0002 (0.9514)
Diff_MV	-	-0.0313*** (0.0005)	-0.1358*** (0.0001)	-0.0271 (0.4325)	-0.0392 (0.0908)	-0.0159 (0.0893)	-0.0053 (0.0718)	-0.0082 (0.1186)
Diff_EUROVOL	-	0.0048*** (0.0063)	-0.0051 (0.4637)	-0.0011 (0.5980)	0.0052 (0.4030)	0.0015 (0.5495)	0.0042 (0.1114)	0.0002 (0.9451)
Diff_PRANGE	+	0.0003 (0.3427)	0.0011*** (0.0068)	0.0025*** (0.0062)	0.0010*** (0.0010)	0.0008*** (0.0009)	0.0005** (0.0276)	0.0017*** (0.0001)
Diff_VWAP	-	0.0109*** (0.0032)	0.0919*** (0.0001)	-0.0228 (0.5927)	0.0037 (0.5573)	-0.0082 (0.4228)	-0.0227*** (0.0001)	-0.016*** (0.0001)
R^2		0.51	0.44	0.73	0.68	0.57	0.62	0.40

Table 3.9 shows the results of the multivariate regression where the dependent variable is the relative PEHS of the banned (B) stocks. The results show that the coefficient estimate of the ban dummy variable is positive for all seven countries and statistically significant at the 1% level. Portugal and Spain are the exception, where the coefficient estimate of the ban dummy variable is significant at the 5% level. Portugal exhibits the greatest increase in Diff_PEHS during the ban period. This is wider by 480 basis points with the next being the UK, where the increase is 127 basis points. In the post-ban period, the coefficient estimate of the post-ban dummy variable is negative but not significant for the Netherlands and positive and significant at the 1% level for Switzerland and the UK. These results suggest that the relative PEHS of the banned stocks does not narrow as would be expected; on the contrary, it widens even more after the relaxation of the short-selling bans in these two countries.

The results reported in Tables 3.8 and 3.9 offer mixed support for Hypothesis 1. There is stronger support for the short-selling bans increasing bid/ask spreads; however, there is no support for the relaxation of the short-selling bans reducing bid/ask spreads. Overall, Diff_PQHS registers a significant increase for five out of seven countries while Diff_PEHS registers a significant increase for all seven countries during the ban period.

Table 3.9 Country-level multivariate results: Percentage quoted effective half-spread (PEHS)

The following regression model has the same form as shown in Boehmer, Jones, and Zhang (2013):

$$Diff\_PEHS_{it} = \alpha_i + \beta D_{it}^{Ban} + \theta X_{it} + e_{it} \qquad (1)$$

where $Diff\_PEHS_{it}$ is the difference between the PEHS of the banned stocks (B) and the PEHS of the non-banned stocks (NB). X_{it} is a vector of differences between the control variables of the matched pair of stocks. These control variables are: market value – MV, trading volume in euros (or in their respective domestic currencies for the UK and Switzerland) – EUROVOL, proportional daily range of transaction prices – PRANGE, and daily volume-weighted average price – VWAP. D_{it}^{Ban} is a dummy variable that takes the value of 1 when short-selling bans are in force and 0 otherwise.

For countries with a post-ban period available (i.e., the Netherlands, Switzerland, and the UK), equation (1) takes the following form:

$$Diff\_PEHS_{it} = \alpha_i + \beta D_{it}^{Ban} + \gamma D_{it}^{Post\_ban} + \theta X_{it} + e_{it} \qquad (2)$$

where $D_{it}^{Post\text{-}ban}$ is equal to 1 when short-selling bans are lifted and 0 otherwise.

Fixed effects are used to control for any differences between the matched stocks in the pre-ban period. All control variables are logarithmically transformed. The p-values of the coefficient estimates are given in the parentheses below these estimates. * indicates statistical significance at the 10% level. ** indicates statistical significance at the 5% level. *** indicates statistical significance at the 1% level.

Note: Belgium, Denmark, and Ireland are excluded from the regression analysis at a country level due to the small sample size.

		(1)				(2)		
	Predicted sign	France	Germany	Portugal	Spain	Netherlands	Switzerland	UK
Intercept		0.0045**	-0.0149***	0.0288	0.0115***	-0.0178	-0.0088***	0.0016
		(0.0255)	(0.0001)	(0.5359)	(0.0043)	(0.1369)	(0.0001)	(0.4021)
Ban dummy variable	+	0.0089***	0.0090***	0.0480**	0.0057**	0.0092***	0.0055***	0.0127***
		(0.0001)	(0.0011)	(0.0249)	(0.0197)	(0.0001)	(0.0009)	(0.0005)
Post-ban dummy variable	-					-0.0032	0.0058***	0.0079***
						(0.1169)	(0.0001)	(0.0001)
Diff_MV	-	-0.0296***	-0.0347***	-0.0290	-0.0384	-0.0226**	0.0030	-0.0070
		(0.0001)	(0.0001)	(0.3909)	(0.0862)	(0.0130)	(0.1262)	(0.0922)
Diff_EUROVOL	-	0.0035**	-0.0046	-0.0030	0.0048	-0.0007	0.0040	0.0010
		(0.0133)	(0.0599)	(0.2207)	(0.4250)	(0.7459)	(0.1306)	(0.6472)
Diff_PRANGE	+	0.0004	0.0006***	0.0028***	0.0011***	0.0009***	0.0007***	0.0012***
		(0.0764)	(0.0021)	(0.0017)	(0.0001)	(0.0001)	(0.0047)	(0.0001)
Diff_VWAP	-	0.0093***	0.0273***	-0.0267	0.0039	-0.0022	-0.0204***	-0.012***
		(0.0049)	(0.0006)	(0.5473)	(0.5131)	(0.8238)	(0.0001)	(0.0001)
R^2		0.52	0.44	0.72	0.68	0.59	0.53	0.41

The regression results of the relative depth of the banned stocks are presented in Table 3.10. The coefficient estimates of the ban dummy variable are positive for Germany and Portugal and negative for the other five countries. However, all the coefficients are not statistically significant, suggesting that the introduction of the short-selling bans had no significant effect on the relative depth of the banned stocks at a country level. In the post-ban period, all the coefficient estimates of the post-ban dummy variable are negative and significant at the 1% level for Switzerland and the UK and at the 5% level for the Netherlands. These results suggest that the relative depth of the banned stocks is lower in the post-ban period than in the pre-ban period. For example, the coefficient estimate of the post-ban dummy variable for Switzerland is -0.1159, implying that in the post-ban period the relative depth of the banned stocks was 11.59% lower than in the pre-ban period.

The coefficient estimates of Diff_EUROVOL are statistically significant for all countries. Also, the coefficient estimates of Diff_VWAP are statistically significant for all countries with the exception of Portugal and the Netherlands. These results imply that the changes in the relative euro volume and relative price levels are important in explaining the changes in the relative depth of the banned stocks.

In summary, the empirical evidence reported in Table 3.10 does not support hypothesis 2A in both ban and post-ban periods.

Table 3.10 Country-level multivariate results: Depth

The following regression model has the same form as shown in Boehmer, Jones, and Zhang (2013):

$$Diff\_depth_{it} = \alpha_i + \beta D_{it}^{Ban} + \theta X_{it} + e_{it} \qquad (1)$$

where $Diff\_depth_{it}$ is the difference between the logarithmically transformed depth of the banned stocks (B) and logarithmically transformed depth of the non-banned stocks (NB). X_{it} is a vector of differences between the control variables of the matched pair of stocks. These control variables are: market value – MV, trading volume in euros (or in their respective domestic currencies for the UK and Switzerland) – EUROVOL, proportional daily range of transaction prices – PRANGE, and daily volume-weighted average price – VWAP. D_{it}^{Ban} is a dummy variable that takes the value of 1 when short-selling bans are in force and 0 otherwise.

For countries with a post-ban period available (i.e., the Netherlands, Switzerland, and the UK), equation (1) takes the following form:

$$Diff\_depth_{it} = \alpha_i + \beta D_{it}^{Ban} + \gamma D_{it}^{Post\_ban} + \theta X_{it} + e_{it} \qquad (2)$$

where $D_{it}^{Post\text{-}ban}$ is equal to 1 when short-selling bans are lifted and 0 otherwise.

Fixed effects are used to control for any differences between the matched stocks in the pre-ban period. All control variables are logarithmically transformed. The p-values of the coefficient estimates are given in the parentheses below these estimates. * indicates statistical significance at the 10% level. ** indicates statistical significance at the 5% level. *** indicates statistical significance at the 1% level.

Note: Belgium, Denmark, and Ireland are excluded from the regression analysis at a country level due to the small sample size.

		(1)				(2)		
	Predicted sign	France	Germany	Portugal	Spain	Netherlands	Switzerland	UK
Intercept		-0.0060 (0.9449)	0.1434 (0.3093)	0.6741*** (0.0006)	0.6298*** (0.0001)	-0.2071 (0.7039)	-0.1371*** (0.0064)	-0.386*** (0.0001)
Ban dummy variable	-	-0.0808 (0.4675)	0.0683 (0.1342)	0.00004 (0.9998)	-0.0848 (0.1882)	-0.0933 (0.1430)	-0.0472 (0.4391)	-0.0739 (0.1287)
Post-ban dummy variable	+					-0.1439** (0.0178)	-0.1159*** (0.0050)	-0.185*** (0.0047)
Diff_MV	+	0.2290 (0.2553)	0.7001*** (0.0001)	0.0500 (0.8255)	-0.8941** (0.0257)	-0.5078 (0.2593)	0.5709*** (0.0001)	0.0730 (0.4304)
Diff_EUROVOL	+	0.3329*** (0.0011)	0.6156*** (0.0006)	0.2587*** (0.0002)	0.5212*** (0.0003)	0.2777*** (0.0001)	0.8105*** (0.0001)	0.385*** (0.0001)
Diff_PRANGE	-	0.0065 (0.0362)	-0.0140 (0.0594)	-0.0134 (0.1939)	-0.0158*** (0.0059)	0.0073*** (0.0071)	-0.0082 (0.0517)	-0.0006 (0.8501)
Diff_VWAP	+	-0.8942*** (0.0001)	-1.695*** (0.0001)	-1.415 (0.0001)	-0.5900*** (0.0024)	-0.0343 (0.9396)	-1.097*** (0.0001)	-0.787*** (0.0001)
R^2		0.88	0.87	0.76	0.93	0.90	0.82	0.91

Table 3.11 reports the results of the relative euro depth of the banned stocks. In the ban period, the coefficient estimates of the ban dummy variable are negative and statistically significant at the 1% level for Spain and the 5% level for Portugal, the Netherlands, and the UK, suggesting a significant decline in the relative euro depth of the banned stocks in these four countries. For example, the coefficient estimate of the ban dummy variable for the Netherlands is -0.3963, which implies that during the ban period the relative euro depth of the Dutch banned stocks was 39.63% less than in the pre-ban period. The coefficient estimates

are positive but not significant for France and Germany and negative but not significant for Switzerland.

The coefficient estimates of the post-ban dummy variable are negative and significant at the 1% level for the UK and negative and significant at the 10% for the Netherlands. The negative estimate suggests that the relative euro depth of the British and Dutch banned stocks decreased in the post-ban period. Although the coefficient estimate of the post-ban dummy variable is also negative for Switzerland, the result is insignificant. Changes in Diff_EUROVOL and Diff_VWAP appear to be significant in explaining the changes in the Diff_eurodepth.

Overall, the results provide some support for hypothesis 2B. The relative euro depth of the banned stocks decreases after the introduction of these bans for four out of the seven countries.

Table 3.11 Country-level multivariate results: Euro depth

The following regression model has the same form as shown in Boehmer, Jones, and Zhang (2013):

$$Diff\_eurodepth_{it} = \alpha_i + \beta D_{it}^{Ban} + \theta X_{it} + e_{it} \qquad (1)$$

where $Diff\_eurodepth_{it}$ is the difference between the logarithmically transformed euro depth of the banned stocks (B) and the logarithmically transformed euro depth of the non-banned stocks (NB). X_{it} is a vector of differences between the control variables of the matched pair of stocks. These control variables are: market value – MV, trading volume in euros (or in their respective domestic currencies for the UK and Switzerland) – EUROVOL, proportional daily range of transaction prices – PRANGE, and daily volume-weighted average price – VWAP. D_{it}^{Ban} is a dummy variable that takes the value of 1 when short-selling bans are in force and 0 otherwise.

For countries with a post-ban period available (i.e., the Netherlands, Switzerland, and the UK), equation (1) takes the following form:

$$Diff\_eurodepth_{it} = \alpha_i + \beta D_{it}^{Ban} + \gamma D_{it}^{Post\_ban} + \theta X_{it} + e_{it} \qquad (2)$$

where $D_{it}^{Post\text{-}ban}$ is equal to 1 when short-selling bans are lifted and 0 otherwise.

Fixed effects are used to control for any differences between the matched stocks in the pre-ban period. All control variables are logarithmically transformed. The p-values of the coefficient estimates are given in the parentheses below these estimates. * indicates statistical significance at the 10% level. ** indicates statistical significance at the 5% level. *** indicates statistical significance at the 1% level.

Note: Belgium, Denmark, and Ireland are excluded from the regression analysis at a country level due to the small sample size.

		(1)				(2)		
	Predicted sign	France	Germany	Portugal	Spain	Netherlands	Switzerland	UK
Intercept		-0.2735	0.8939	-0.5867	-0.6684	-0.1423	0.5808***	-3.12***
		(0.2284)	(0.1894)	(0.2026)	(0.0807)	(0.9062)	(0.0071)	(0.0001)
Ban dummy variable	-	0.1203	0.2978	-0.5683**	-0.9835***	-0.3963**	-0.0007	-0.739**
		(0.7040)	(0.1515)	(0.0395)	(0.0001)	(0.0151)	(0.9985)	(0.0479)
Post-ban dummy variable	+					-0.3598*	-0.3948	-1.59***
						(0.0636)	(0.3218)	(0.0054)
Diff_MV	+	-1.667***	3.851***	0.9226	3.437***	-1.792	0.3046	-0.0193
		(0.0001)	(0.0001)	(0.2217)	(0.0029)	(0.0723)	(0.6437)	(0.9676)
Diff_EUROVOL	+	1.077***	2.040**	0.3151***	1.043***	0.6245***	3.608***	2.16***
		(0.0003)	(0.0142)	(0.0001)	(0.0001)	(0.0001)	(0.0001)	(0.0001)
Diff_PRANGE	-	0.0293**	-0.0408	-0.0126	-0.0423***	0.0085	-0.0357	0.0145
		(0.0146)	(0.1802)	(0.3461)	(0.0007)	(0.0659)	(0.1051)	(0.6467)
Diff_VWAP	+	7.265***	0.1511	6.511***	2.680***	8.277***	4.182***	4.77***
		(0.0001)	(0.6724)	(0.0001)	(0.0053)	(0.0001)	(0.0001)	(0.0001)
R^2		0.93	0.94	0.97	0.96	0.98	0.95	0.75

3.6.3.2 Aggregate-level liquidity effects

This section conducts a multivariate analysis at an aggregate level by pooling all the countries together to test the four hypotheses developed in section 3.3. Table 3.12 reports the results for the four liquidity measures. The estimation results of the ban dummy variable show that the relative PQHS of the banned stocks is on average 139 basis points higher during the ban period (see column (1)). The results for the relative PEHS of the banned stocks are similar although the ban effect is slightly lower at 128 basis points (see column (2)). Both estimates are statistically significant at the 1% level. Furthermore, the results that are reported in columns (3) and (4) of Table 3.12 show that during the ban period the relative depth and euro depth of the banned stocks decline by 11.3% and 42.8%, respectively, from their pre-ban levels. Again, both results are statistically significant at the 1% level.

Overall, these results provide strong support for Hypotheses 1, 2A and 2B.

Table 3.12 Aggregate-level multivariate results

The following regression model has the same form as shown in Boehmer, Jones, and Zhang (2013):

$$Diff\_Y_{it} = \alpha_i + \beta D_{it}^{Ban} + \theta X_{it} + e_{it}$$

where $Diff\_Y_{it}$ is the difference between the four liquidity measures of the banned stocks (B) and the respective liquidity measures of the non-banned stocks (NB). The four liquidity measures are the PQHS, PEHS, logarithmically transformed depth, and logarithmically transformed euro depth. X_{it} is a vector of differences between the control variables of the matched pair of stocks. These control variables are: market value – MV, trading volume in euros (or in their respective domestic currencies for the UK, Switzerland, and Denmark) – EUROVOL, proportional daily range of transaction prices – PRANGE, and daily volume-weighted average price – VWAP. D_{it}^{Ban} is a dummy variable that takes the value of 1 when short-selling bans are in force and 0 otherwise. Fixed effects are used to control for any differences between the matched stocks in the pre-ban period. All control variables are logarithmically transformed. The p-values of the coefficient estimates are given in the parentheses below these estimates. * indicates statistical significance at the 10% level. ** indicates statistical significance at the 5% level. *** indicates statistical significance at the 1% level.

Note: The regression analysis at an aggregate level excludes the post-ban period as only three countries out of 10 have a post-ban period available.

		(1)	(2)		(3)	(4)
	Predicted sign	Diff_PQHS	Diff_PEHS	Predicted sign	Diff_depth	Diff_eurodepth
Intercept		0.0069 (0.3653)	0.0049 (0.4512)		0.2779*** (0.0001)	2.285*** (0.0001)
Ban dummy variable	+	0.0139*** (0.0002)	0.0128*** (0.0001)	-	-0.1130*** (0.0041)	-0.4275*** (0.0013)
Diff_MV	-	-0.0406** (0.0370)	-0.0340** (0.0363)	+	0.2176** (0.0284)	0.0785 (0.9077)
Diff_EUROVOL	-	0.0005 (0.8105)	0.0001 (0.9592)	+	0.4218*** (0.0001)	1.299*** (0.0001)
Diff_PRANGE	+	0.0014*** (0.0002)	0.0014*** (0.0001)	-	-0.0027 (0.2885)	-0.0025 (0.8084)
Diff_VWAP	-	-0.0082 (0.4655)	-0.0095 (0.3898)	+	-0.9890*** (0.0001)	5.426*** (0.0001)
R^2		0.57	0.62		0.91	0.91

3.6.3.3 Naked versus total short-selling bans

Table 3.13 reports the results of the four liquidity measures using equation (3.7). The ban dummy variable of equation (3.5) is now replaced with two new dummy variables – a naked dummy variable and a total dummy variable – to examine whether there are differences in liquidity effects between countries that applied naked short-selling bans and total short-selling bans, respectively. The estimation results reported in the first two columns of Table 3.13 show that the relative PQHS and PEHS are higher in countries with total short-selling bans. Although the coefficient estimates on both dummy variables are positive, only the coefficient estimate of the total dummy variable is statistically significant. For example, as shown in column (2) of Table 3.13, the coefficient estimate of the total dummy variable is 0.0213, implying an average increase of 213 basis points in the relative PEHS of the banned stocks during the ban period in countries that introduced total short-selling bans. This result supports Hypothesis 3, that total short-selling bans have a stronger impact on liquidity, in this case on the price dimension of liquidity.

In column (3) of Table 3.13, where the dependent variable is Diff_depth, the coefficient estimate of the naked ban dummy variable is -0.1274, significant at the 5% level. The coefficient estimate of the total ban dummy variable is -0.0896, significant at the 1% level. Although both results are statistically significant, the Wald test (p-value=0.5882) fails to reject the null hypothesis that the difference between the coefficient estimates of the naked and total ban dummy variables equals 0. This result implies that naked and total short-selling bans have a similar impact on the relative depth of the banned stocks.

The estimation results in column (4) of Table 3.13 show that the relative euro depth of the banned stocks decreases 66.25% in countries with total short-selling bans. This result is statistically significant at the 5% level. However, there is no statistically significant change in the relative euro depth of the banned stocks in countries that introduced naked short-selling bans.

In summary, the empirical evidence in Table 3.13 shows that the relative bid/ask spreads of banned stocks increase, and the relative depth and euro depth of the banned stocks decrease during the ban period in countries that introduced total short-selling bans. At the same time, there is little evidence that purely naked short-selling bans had much of an impact on liquidity. Overall, these results provide strong support for Hypothesis 3.

Table 3.13 Aggregate-level multivariate results: Naked versus total short-selling bans

The regression model has the following form:

$$Diff\_Y_{it} = \alpha_i + \beta D_{it}^N + \gamma D_{it}^T + \theta X_{it} + e_{it}$$

where $Diff\_Y_{it}$ is the difference between the four liquidity measures of the banned stocks (B) and the respective liquidity measures of the non-banned stocks (NB). The four liquidity measures are the PQHS, PEHS, logarithmically transformed depth, and logarithmically transformed euro depth. X_{it} is a vector of differences between the control variables of the matched pair of stocks. These control variables are: market value – MV, trading volume in euros (or in their respective domestic currencies for the UK, Switzerland, and Denmark) – EUROVOL, proportional daily range of transaction prices – PRANGE, and daily volume-weighted average price – VWAP. D_{it}^N is a dummy variable that takes the value of 1 if short-selling ban is naked and 0 otherwise and Dit^T is a dummy variable that takes the value of 1 if short-selling ban is total and 0 otherwise. Fixed effects are used to control for any differences between the matched stocks in the pre-ban period. All control variables are logarithmically transformed. The p-values of the coefficient estimates are given in the parentheses below these estimates. * indicates statistical significance at the 10% level. ** indicates statistical significance at the 5% level. *** indicates statistical significance at the 1% level.

Note: The regression analysis at an aggregate level excludes the post-ban period as only three countries out of 10 have a post-ban period available.

		(1)	(2)		(3)	(4)
	Predicted sign	Diff_PQHS	Diff_PEHS	Predicted sign	Diff_depth	Diff_eurodepth
Intercept		0.0002 (0.9657)	0.0007 (0.8851)		0.2637*** (0.0001)	2.4*** (0.0001)
Naked dummy variable	+	0.0061 (0.2506)	0.0078 (0.1061)	-	-0.1274** (0.0371)	-0.2989 (0.0993)
Total dummy variable	+	0.0271*** (0.0018)	0.0213*** (0.0021)	-	-0.0896*** (0.0029)	-0.6625** (0.0111)
Diff_MV	-	-0.0352** (0.0309)	-0.0305** (0.0289)	+	0.2339** (0.0157)	0.0192 (0.9776)
Diff_EUROVOL	-	0.0014 (0.5396)	0.0006 (0.7333)	+	0.4230*** (0.0001)	1.297*** (0.0001)
Diff_PRANGE	+	0.0013*** (0.0001)	0.0013*** (0.0001)	-	-0.0030 (0.2187)	-0.0004 (0.9659)
Diff_VWAP	-	-0.0138 (0.1879)	-0.0131 (0.2042)	+	-1.009*** (0.0001)	5.518*** (0.0001)
R^2		0.57	0.62		0.91	0.91

3.6.3.4 The role of listed derivatives in the impact of short-selling bans on market liquidity

Table 3.14 reports the results of the multivariate analysis for all four liquidity measures conditional on both the stringency of the short-selling rules and derivatives trading using equation (3.7). The results in Panel A show that for stocks with listed derivatives, the coefficient estimates of the naked ban dummy variable are not statistically significant for all but the relative depth liquidity measure. On the other hand, the coefficient estimates of the total dummy variable are statistically significant for all the four relative liquidity measures.[133] For example, as reported in column (2) of Table 3.14, the coefficient estimate of the total ban

[133] The Wald test (p-value=0.3317) again fails to reject the null hypothesis that the difference between the coefficient estimates of the naked and total ban dummy variables equals 0 for Diff_depth where both results are statistically significant. Consequently, there is no significant difference in the magnitude of the impact of the short-selling bans on the relative depth of the banned stocks with listed derivatives conditional on the stringency of these bans.

dummy variable for Diff_PEHS is 0.0234, significant at the 1 % level, suggesting that the relative PEHS of the banned stocks is 234 basis points wider during the ban period. The relative euro depth of the same banned stocks is 74% lower during the ban period, as can be seen in column (4) of Table 3.14. This result is statistically significant at the 1% level.

Panel B of Table 3.14 reports the results of the multivariate regression for stocks without listed derivatives, again using equation (3.7). The results show that none of the coefficient estimates of the naked and total dummy variables are statistically significant for any of the four liquidity measures.

The results in Panel A of Table 3.14 imply that there is a significant deterioration in liquidity for banned stocks with listed derivatives in terms of higher bid/ask spreads – both quoted and effective – and lower depth and euro depth in countries where total short-selling bans were introduced. In contrast, the results in Panel B of Table 3.14 for the stocks without listed derivatives show that neither naked short-selling ban nor total short-selling bans have any impact on the relative liquidity of the banned stocks. These results do not support Hypothesis 4.

Table 3.14 Aggregate-level multivariate results conditional on the stringency of short-selling bans and derivatives trading

The regression model has the following form:

$$Diff\_Y_{it} = \alpha_i + \beta D_{it}^{N} + \gamma D_{it}^{T} + \theta X_{it} + e_{it}$$

where $Diff\_Y_{it}$ is the difference between the four liquidity measures of the banned stocks (B) and the respective liquidity measures of the non-banned stocks (NB). The four liquidity measures are the PQHS, PEHS, logarithmically transformed depth, and logarithmically transformed euro depth. X_{it} is a vector of differences between the control variables of the matched pair of stocks. These control variables are: market value – MV, trading volume in euros (or in their respective domestic currencies for the UK, Switzerland, and Denmark) – EUROVOL, proportional daily range of transaction prices – PRANGE, and daily volume-weighted average price – VWAP. D_{it}^{N} is a dummy variable that takes the value of 1 if short-selling ban is naked and 0 otherwise and Dit^{T} is a dummy variable that takes the value of 1 if short-selling ban is total and 0 otherwise. Panel A presents the results for pairs of stocks that have listed derivatives (options and/or futures) while Panel B reports the results for pairs of stocks without listed derivatives. Fixed effects are used to control for any differences between the matched stocks in the pre-ban period. All control variables are logarithmically transformed. The p-values of the coefficient estimates are given in the parentheses below these estimates. * indicates statistical significance at the 10% level. ** indicates statistical significance at the 5% level. *** indicates statistical significance at the 1% level.

Note: The regression analysis at an aggregate level excludes the post-ban period as only three countries out of 10 have a post-ban period available.

Panel A: Stocks with derivatives traded (57 pairs of stocks)						
		(1)	(2)		(3)	(4)
	Predicted sign	Diff_PQHS	Diff_PEHS	Predicted sign	Diff_depth	Diff_eurodepth
Intercept		0.0014 (0.7401)	0.0021 (0.5555)		0.2141*** (0.0001)	2.189*** (0.0001)
Naked dummy variable	+	0.0032 (0.5367)	0.0041 (0.3628)	-	-0.15** (0.0240)	-0.3786 (0.0611)
Total dummy variable	+	0.0297*** (0.0013)	0.0234*** (0.0008)	-	-0.0758** (0.0360)	-0.74*** (0.0073)
Diff_MV	-	-0.0424** (0.0168)	-0.0367*** (0.0068)	+	0.2347** (0.0147)	0.3940 (0.6118)
Diff_EUROVOL	-	0.0025 (0.4993)	0.0009 (0.7672)	+	0.5205*** (0.0001)	1.436*** (0.0001)
Diff_PRANGE	+	0.0009*** (0.0019)	0.0012*** (0.0001)	-	-0.0020 (0.5172)	0.0082 (0.4326)
Diff_VWAP	-	-0.0061 (0.3190)	-0.0061 (0.2652)	+	-1.007*** (0.0001)	5.428*** (0.0001)
R^2		0.52	0.59		0.93	0.93
Panel B: Stocks without derivatives traded (21 pairs of stocks)						
Intercept		0.1093 (0.1732)	0.1047 (0.2670)		1.198 (0.0839)	0.6988 (0.6959)
Naked dummy	+	0.0093 (0.3728)	0.0146 (0.1269)	-	-0.0503 (0.5948)	0.0385 (0.8711)
Total dummy	+	0.0202 (0.1406)	0.0148 (0.3302)	-	-0.0076 (0.8903)	0.3061 (0.1211)
Diff_MV	-	0.1088 (0.0991)	0.1028 (0.1816)	+	0.7260 (0.2048)	-0.8516 (0.5734)
Diff_EUROVOL	-	-0.0005 (0.8854)	0.0002 (0.9552)	+	0.2524*** (0.0001)	0.8831*** (0.0001)
Diff_PRANGE	+	0.0021*** (0.0008)	0.0015*** (0.0056)	-	-0.0054 (0.2006)	-0.0104 (0.4688)
Diff_VWAP	-	-0.1720*** (0.0093)	-0.1601** (0.0401)	+	-1.728*** (0.0001)	4.689*** (0.0060)
R^2		0.65	0.67		0.78	0.79

3.6.4 Summary of the results of the multivariate analysis and discussion

At the country level of analysis, the multivariate results show that the relative liquidity of the banned stocks deteriorates in terms of higher effective spreads but not in terms of lower depth. More precisely, higher relative bid/ask spreads of the banned stocks are not found to be associated with lower relative depth of the same stocks during the ban period. The coefficient estimates of the ban dummy variable for Diff_depth, although negative for most countries, are not found to be statistically significant. On the other hand, the results for euro depth are mixed. The relative euro depth of the banned stocks decreases for Portugal, Spain, the Netherlands, and the UK but does not change significantly for France, Germany, and Switzerland.

In the post-ban period, the relative PEHS of the banned stocks increases even more for Switzerland and the UK but does not change significantly for the Netherlands. The relative depth of the banned stocks decreases significantly in all three countries. Finally, the relative euro depth of the banned stocks decreases even more for the UK and marginally for the Netherlands.

Multivariate analysis conducted by Hansson and Fors (2009) and Marsh and Payne (2012) for the UK market also shows that the percentage quoted spread of the British banned stocks is greater relative to the percentage quoted spread of the non-banned stocks during the short-selling ban period. Their results are qualitatively similar to the results reported in this essay. However, one important difference between the prior studies and this study is the construction of the control sample and the matching process. Hansson and Fors (2009) use the remaining components of FTSE 350 not subject to the short-selling ban as a control sample, while Marsh and Payne (2012) match 10 non-banned stocks to each banned stock by their market value averaged over the first half of 2008. In contrast, this essay matches each banned stock to a non-banned stock by its market value, trading volume, and price averaged over the first half of 2008. A non-banned stock which has the most similar characteristics as well as the most similar derivatives listings to the banned stock is chosen as a control stock.

At an aggregate level, there is a clear deterioration in the relative liquidity of the banned stocks in the ban period where higher relative bid/ask spreads of the banned stocks are accompanied by lower relative depth and euro depth of the same stocks (see Table 3.12). Also, conditioning on the stringency of the short-selling bans the results reveal that the relative liquidity of the banned stocks deteriorates in countries with total short-selling bans. These results support the findings of Beber and Pagano (2013). These authors find that the impact of short-selling bans on bid/ask spreads is more severe in countries that introduced total short-selling bans. However, they also find a significant increase in the percentage bid/ask spreads of the banned stocks under naked short-selling bans. In contrast, this essay fails to find any significant impact of the naked short-selling bans on the percentage bid/ask spreads. This could be attributed to the sample size. Beber and Pagano (2013) use 16,491 stocks from 30 countries in their analysis while this essay focuses on 78 European banned stocks, 39 of which were subject to naked short-selling bans.

Furthermore, the results reported in Table 3.14 show that the relative liquidity of banned stocks worsens under total short-selling ban but only in the presence of listed derivatives. This result contradicts the findings of Beber and Pagano (2013) who document higher percentage bid/ask spreads on banned stocks without listed options. One explanation could be that this essay focuses only on European countries where most financial regulators introduced short-selling bans in the derivatives markets as well. In contrast, Beber and Pagano (2013) use a large number of banned stocks traded in countries where derivatives markets were not affected by the short-selling bans. Consequently, the authors find that, on average, derivatives trading ameliorated the impact of the short-selling bans on the bid/ask spreads of the banned stocks with listed options.

Overall, the aggregate-level multivariate results show a deterioration in the liquidity of the banned stocks after the implementation of the short-selling bans, which supports

Hypotheses1, 2A, and 2B. Also, empirical evidence shows that total short-selling bans have a stronger impact on the liquidity of the banned stocks, thus supporting Hypothesis 3. In contrast, there is no supporting evidence for Hypothesis 4. The results show that, under total short-selling bans, liquidity deteriorates for banned stocks with listed derivatives.

3.7 Conclusions

This essay investigates the impact of the short-selling regulations that were introduced in the autumn of 2008 on the liquidity of the stocks of major European financial institutions across 10 countries. The main contribution of this research is the incorporation of depth and euro depth into the analysis. Prior studies mostly focus on the price dimension of liquidity, measured by the bid/ask spread. However, liquidity has two dimensions (Lee, Mucklow, and Ready (1993)). Consequently, this essay extends prior studies by providing empirical evidence of the effects of the 2008 short-selling bans on both dimensions of liquidity in a European context.

The country-level univariate results show that the relative liquidity of the banned stocks deteriorates significantly for eight out of 10 countries during the ban period. The impact of the short-selling bans on the relative liquidity of the banned stocks for the remaining two countries – Germany and Spain – is ambiguous. In the post-ban period, the relative liquidity of the banned stocks improves for two out of three countries with a post-ban period available. These countries are the Netherlands and the UK. In contrast, Swiss banned stocks experience a further deterioration in liquidity during the post-ban period.

The country-level multivariate results show that the relative liquidity of the banned stocks worsens during the ban period for Portugal, Spain, the Netherlands, and the UK in terms of both greater effective spreads and lower euro depth. The relative liquidity of the banned stocks also deteriorates for France, Germany, and Switzerland, however, only in terms of greater effective spreads. Contrary to what would be expected, the relative liquidity of banned stocks continues to worsen during the post-ban period for countries with a post ban period available.

The aggregate-level multivariate results reveal that the relative PQHS and PEHS of the banned stocks increase significantly, while the relative depth and euro depth decrease significantly during the ban period.

Overall, these results show that the liquidity of the banned stocks deteriorates significantly during the ban period, supporting the hypotheses developed in this essay and enhancing the findings of prior studies (see for example, Bris (2008), Clifton and Snape (2008), Hansson and Fors, (2009), Boulton and Braga-Aves (2010), Marsh and Payne (2012), and Beber and Pagano (2013)). However, there is no strong evidence showing that the relative liquidity of the banned stocks improves following the relaxation of the short-selling bans in countries with a post-ban period available.

Furthermore, the regression analysis shows that the impact of the short-selling bans on the liquidity of the banned stocks is stronger under total short-selling bans. This result again supports the hypothesis developed in this essay that the stringency of the short-selling bans plays an important role in their effects on liquidity and is consistent with the findings of Beber and Pagano (2013).

Finally, this essay documents a greater deterioration in the liquidity of the banned stocks if these stocks have listed derivatives. This result does not support the hypothesis that listed derivatives ameliorate the impact of the short-selling bans on the liquidity of the banned stocks, contradicting the finding of Beber and Pagano (2013) who report that liquidity deteriorates less for stocks with listed options.

APPENDIX 3A

Table 3A.1 Mean percentage quoted half-spread (PQHS) for banned (B) and non-banned (NB) stocks

This table reports the average mean values of the percentage quoted half-spread (PQHS) for the banned (B) and non-banned (NB) stocks, respectively, across 10 European countries. The sample period spans from mid-January 2008 to mid-December 2009 and it is divided into three sub-periods: the pre-ban period, the ban period, and the post-ban period. The results for countries that have two sub-periods, the pre-ban period and the ban period (i.e., Belgium, Denmark, France, Germany, Ireland, Portugal, and Spain) are shown in Panel A. Only Switzerland, the Netherlands, and the UK have a post-ban period available. Their results are reported in Panel B. The sub-periods differ for each country and are specified in Table 3.1. Table 3A.1 also shows the statistical significance of the test for equality of means of the PQHS for both samples as well as their difference (i.e., PQHS (B) – PQHS (NB)).Variable Diff is the change in the mean PQHS between the pre-ban (P1) and ban (P2) periods (P2–P1), and ban and post-ban (P3) periods (P3–P2), respectively. Diff-in-Diff is the change in the difference between the PQHS of the banned stocks and the PQHS of the non-banned stocks from the pre-ban period to the ban period (P2–P1), and from the ban period to the post-ban period (P3–P2), respectively. The t-test is used and the p-values of the t-statistic are given in the parentheses. * indicates statistical significance at the 10% level. ** indicates statistical significance at the 5% level. *** indicates statistical significance at the 1% level.

Country	Sub-periods	PQHS(B)	PQHS(NB)	PQHS(B) – PQHS(NB)		Diff PQHS(B)	Diff PQHS(NB)	Diff-in-Diff PQHS(B) – PQHS(NB)
Panel A: Countries with two sub-periods								
Belgium	Pre-ban	0.0519	0.0448	0.0071	P2-P1	0.0469***	0.0127***	0.0341***
	Ban	0.0988	0.0575	0.0412	(p-value)	(0.0001)	(0.0001)	(0.0001)
Denmark	Pre-ban	0.1125	0.1200	-0.0075	P2-P1	0.0822***	0.0790***	0.0031
	Ban	0.1947	0.1990	-0.0043	(p-value)	(0.0001)	(0.0001)	(0.5477)
France	Pre-ban	0.0482	0.0406	0.0075	P2-P1	0.0125***	0.0005	0.0120***
	Ban	0.0608	0.0412	0.0195	(p-value)	(0.0001)	(0.4021)	(0.0001)
Germany	Pre-ban	0.0636	0.0666	-0.0030	P2-P1	0.0426***	0.0363***	0.0063
	Ban	0.1062	0.1029	0.0033	(p-value)	(0.0001)	(0.0001)	(0.0629)
Ireland	Pre-ban	0.1607	0.3373	-0.1766	P2-P1	0.4299***	0.4189***	0.0110
	Ban	0.5906	0.7562	-0.1656	(p-value)	(0.0001)	(0.0001)	(0.7227)
Portugal	Pre-ban	0.1473	0.1365	0.0107	P2-P1	0.0116***	-0.0320***	0.0436***
	Ban	0.1589	0.1046	0.0542	(p-value)	(0.0092)	(0.0001)	(0.0001)
Spain	Pre-ban	0.0804	0.0624	0.0180	P2-P1	0.0112***	0.0107***	0.0005
	Ban	0.0916	0.0731	0.0185	(p-value)	(0.0001)	(0.0001)	(0.6605)
Panel B: Countries with three sub-periods								
Netherlands	Pre-ban	0.0667	0.0591	0.0076	P2-P1	0.0231***	0.0023	0.0207***
					(p-value)	(0.0001)	(0.4344)	(0.0001)
	Ban	0.0898	0.0615	0.0283	P3-P2	-0.0194***	-0.0010	-0.0183***
	Post-ban	0.0704	0.0604	0.0100	(p-value)	(0.0001)	(0.8019)	(0.0002)
Switzerland	Pre-ban	0.0491	0.0476	0.0014	P2-P1	0.0207***	0.0114***	0.0093***
					(p-value)	(0.0001)	(0.0001)	(0.0001)
	Ban	0.0699	0.0591	0.0107	P3-P2	-0.0008	-0.0067***	0.0075***
	Post-ban	0.0707	0.0523	0.0183	(p-value)	(0.4340)	(0.0001)	(0.0001)
UK	Pre-ban	0.0765	0.0686	0.0079	P2-P1	0.0417***	0.0224***	0.0192***
					(p-value)	(0.0001)	(0.0001)	(0.0001)
	Ban	0.1182	0.0910	0.0271	P3-P2	-0.0358***	-0.0169***	-0.0190***
	Post-ban	0.0824	0.0741	0.0082	(p-value)	(0.0001)	(0.0001)	(0.0001)

Table 3A.2 Mean percentage effective half-spread (PEHS) for banned (B) and non-banned (NB) stocks

This table reports the average mean values of the percentage effective half-spread (PEHS) for the banned (B) and non-banned (NB) stocks, respectively, across 10 European countries. The sample period spans from mid-January 2008 to mid-December 2009 and it is divided into three sub-periods: the pre-ban period, the ban period, and the post-ban period. The results for countries that have two sub-periods, the pre-ban period and the ban period (i.e., Belgium, Denmark, France, Germany, Ireland, Portugal, and Spain) are shown in Panel A. Only Switzerland, the Netherlands, and the UK have a post-ban period available. Their results are reported in Panel B. The sub-periods differ for each country and are specified in Table 3.1. Table 3A.2 also shows the statistical significance of the test for equality of means of the PEHS for both samples as well as their difference (i.e., PEHS (B) – PEHS (NB)).Variable Diff is the change in the mean PEHS between the pre-ban (P1) and ban (P2) periods (P2–P1), and ban and post-ban (P3) periods (P3–P2), respectively. Diff-in-Diff is the change in the difference between the PEHS of the banned stocks and the PEHS of the non-banned stocks from the pre-ban period to the ban period (P2–P1), and from the ban period to the post-ban (P3–P2) period, respectively. The t-test is used and the p-values of the t-statistic are given in the parentheses. * indicates statistical significance at the 10% level. ** indicates statistical significance at the 5% level. *** indicates statistical significance at the 1% level.

Country	Sub-periods	PEHS(B)	PEHS(NB)	PEHS(B) – PEHS(NB)		Diff PEHS(B)	Diff PEHS(NB)	Diff-in-Diff PEHS(B) – PEHS(NB)
Panel A: Countries with two sub-periods								
Belgium	Pre-ban	0.0556	0.0480	0.0075	P2-P1	0.0522***	0.0120***	0.0402***
	Ban	0.1078	0.0600	0.0477	(p-value)	(0.0001)	(0.0001)	(0.0001)
Denmark	Pre-ban	0.1089	0.1125	-0.0035	P2-P1	0.0776***	0.0766***	0.0010
	Ban	0.1866	0.1891	-0.0025	(p-value)	(0.0001)	(0.0001)	(0.8395)
France	Pre-ban	0.0510	0.0428	0.0082	P2-P1	0.0110***	0.0002	0.0112***
	Ban	0.0620	0.0426	0.0194	(p-value)	(0.0001)	(0.7221)	(0.0001)
Germany	Pre-ban	0.0492	0.0483	0.0009	P2-P1	0.0247***	0.0163***	0.0083***
	Ban	0.0739	0.0646	0.0092	(p-value)	(0.0001)	(0.0001)	(0.0001)
Ireland	Pre-ban	0.1288	0.2618	-0.1329	P2-P1	0.3389***	0.2973***	0.0415
	Ban	0.4677	0.5591	-0.0914	(p-value)	(0.0001)	(0.0001)	(0.0640)
Portugal	Pre-ban	0.1513	0.1444	0.0068	P2-P1	0.0124***	-0.0320***	0.0443***
	Ban	0.1636	0.1124	0.0512	(p-value)	(0.0040)	(0.0001)	(0.0001)
Spain	Pre-ban	0.0831	0.0656	0.0175	P2-P1	0.0109***	0.0099***	0.0008
	Ban	0.0940	0.0756	0.0184	(p-value)	(0.0001)	(0.0001)	(0.4682)
Panel B: Countries with three sub-periods								
Netherlands	Pre-ban	0.0695	0.0607	0.0087	P2-P1	0.0249***	0.0003	0.0253***
	Ban	0.0944	0.0604	0.0340	(p-value)	(0.0001)	(0.8497)	(0.0001)
	Post-ban	0.0679	0.0570	0.0108	P3-P2	-0.0266***	-0.0033	-0.0232***
					(p-value)	(0.0001)	(0.1438)	(0.0001)
Switzerland	Pre-ban	0.0513	0.0491	0.0021	P2-P1	0.0214***	0.0114***	0.0099***
	Ban	0.0726	0.0605	0.0121	(p-value)	(0.0001)	(0.0001)	(0.0001)
	Post-ban	0.0590	0.0457	0.0133	P3-P2	-0.0137***	-0.0149***	0.0012
					(p-value)	(0.0001)	(0.0001)	(0.1663)
UK	Pre-ban	0.0674	0.0663	0.0011	P2-P1	0.0297***	0.0168***	0.0129***
	Ban	0.0971	0.0830	0.0140	(p-value)	(0.0001)	(0.0001)	(0.0001)
	Post-ban	0.0785	0.0687	0.0098	P3-P2	-0.0186***	-0.0144***	-0.0042***
					(p-value)	(0.0001)	(0.0001)	(0.0001)

Table 3A.3 Mean depth for banned (B) and non-banned (NB) stocks

This table reports the average mean values of the depth for the banned (B) and non-banned (NB) stocks, respectively, across 10 European countries. The sample period spans from mid-January 2008 to mid-December 2009 and it is divided into three sub-periods: the pre-ban period, the ban period, and the post-ban period. The results for countries that have two sub-periods, the pre-ban period and the ban period (i.e., Belgium, Denmark, France, Germany, Ireland, Portugal, and Spain) are shown in Panel A. Only Switzerland, the Netherlands, and the UK have a post-ban period available. Their results are reported in Panel B. The sub-periods differ for each country and are specified in Table 3.1. Table 3A.3 also shows the statistical significance of the test for equality of means of the depth for both samples as well as their difference (i.e., depth (B) – depth (NB)).Variable Diff is the change in the mean depth between the pre-ban (P1) and ban (P2) periods (P2–P1), and ban and post-ban (P3) periods (P3–P2), respectively. Diff-in-Diff is the change in the difference between the depth of the banned stocks and the depth of the non-banned stocks from the pre-ban period to the ban period (P2–P1), and from the ban period to the post-ban period (P3–P2), respectively. The t-test is used and the p-values of the t-statistic are given in the parentheses. * indicates statistical significance at the 10% level. ** indicates statistical significance at the 5% level. *** indicates statistical significance at the 1% level.

Country	Sub-periods	Depth(B)	Depth(NB)	Depth(B) – depth(NB)		Diff Depth(B)	Diff Depth(NB)	Diff-in-Diff Depth(B) – depth(NB)
Panel A: Countries with two sub-periods								
Belgium	Pre-ban	2,983	323	2,659	P2-P1	-34	-2	-37
	Ban	2,948	326	2,622	(p-value)	(0.8097)	(0.4227)	(0.7939)
Denmark	Pre-ban	4,551	2,471	2,079	P2-P1	225	806***	-581***
	Ban	4,776	3,278	1,498	(p-value)	(0.4190)	(0.0001)	(0.0078)
France	Pre-ban	1,619	1,586	32	P2-P1	-228***	-415***	187***
	Ban	1,391	1,170	220	(p-value)	(0.0006)	(0.0001)	(0.0001)
Germany	Pre-ban	901	1,114	-212	P2-P1	101***	-108***	210***
	Ban	1,003	1,005	-2	(p-value)	(0.0002)	(0.0048)	(0.0001)
Ireland	Pre-ban	3,311	6,978	-3,667	P2-P1	5,037***	3,793***	1,243***
	Ban	8,348	10,772	-2,423	(p-value)	(0.0001)	(0.0001)	(0.0002)
Portugal	Pre-ban	31,118	51,395	-20,276	P2-P1	-12,615**	-35,465***	22,850***
	Ban	18,503	15,929	2,573	(p-value)	(0.0001)	(0.0001)	(0.0001)
Spain	Pre-ban	19,070	6,155	12,914	P2-P1	-3,634***	1,016***	-4,650***
	Ban	15,436	7,171	8,264	(p-value)	(0.0002)	(0.0094)	(0.0001)
Panel B: Countries with three sub-periods								
Netherlands					P2-P1	-173	-677***	503***
	Pre-ban	3,084	2,989	95	(p-value)	(0.0914)	(0.0001)	(0.0001)
	Ban	2,910	2,311	599	P3-P2	-126	376***	-502***
	Post-ban	2,784	2,687	96	(p-value)	(0.1346)	(0.0002)	(0.0001)
Switzerland					P2-P1	-1,288***	-2,021***	732***
	Pre-ban	3,690	5,945	-2,255	(p-value)	(0.0001)	(0.0001)	(0.0002)
	Ban	2,401	3,924	-1,522	P3-P2	894***	764**	-130
	Post-ban	3,296	4,689	-1,392	(p-value)	(0.0034)	(0.0118)	(0.4680)
UK					P2-P1	-3,946***	-3,555***	-391
	Pre-ban	13,388	12,010	1,378	(p-value)	(0.0001)	(0.0001)	(0.3315)
	Ban	9,442	8,455	987	P3-P2	2,316***	1,072***	1,244***
	Post-ban	11,759	9,527	2,231	(p-value)	(0.0001)	(0.0001)	(0.0027)

Table 3A.4 Mean euro depth (ED) for banned (B) and non-banned (NB) stocks

This table reports the average mean values of the euro depth for the banned (B) and non-banned (NB) stocks, respectively, across 10 European countries. The sample period spans from mid-January 2008 to mid-December 2009 and it is divided into three sub-periods: the pre-ban period, the ban period, and the post-ban period. The results for countries that have two sub-periods, the pre-ban period and the ban period (i.e., Belgium, Denmark, France, Germany, Ireland, Portugal, and Spain) are shown in Panel A. Only Switzerland, the Netherlands, and the UK have a post-ban period available. Their results are reported in Panel B. The sub-periods differ for each country and are specified in Table 3.1. Table 3A.4 also shows the statistical significance of the test for equality of means of the ED for both samples as well as their difference (i.e., euro depth (B) – euro depth (NB)). Variable Diff is the change in the mean ED between the pre-ban (P1) and ban (P2) periods (P2–P1), and ban and post-ban (P3) periods (P3–P2), respectively. Diff-in-Diff is the change in the difference between the ED of the banned stocks and the ED of the non-banned stocks from the pre-ban period to the ban period (P2–P1), and from the ban period to the post-ban period (P3–P2), respectively. The t-test is used and the p-values of the t-statistic are given in the parentheses. * indicates statistical significance at the 10% level. ** indicates statistical significance at the 5% level. *** indicates statistical significance at the 1% level.

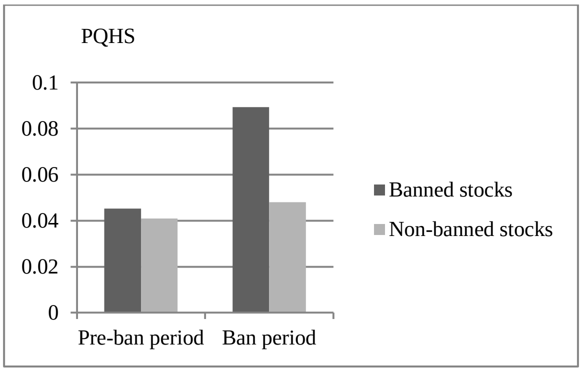
Country	Sub-periods	ED(B)	ED(NB)	ED(B) – ED(NB)		Diff ED(B)	Diff ED(NB)	Diff-in-Diff ED(B) – ED(NB)
Panel A: Countries with two sub-periods								
Belgium	Pre-ban	49,060	19,723	29,337	P2-P1	-35,424***	-5,166***	-30,257***
	Ban	13,636	14,556	-919	(p-value)	(0.0001)	(0.0001)	(0.0001)
Denmark	Pre-ban	763,845	733,246	30,598	P2-P1	-299,971***	-150,274***	-149,697***
	Ban	463,873	582,972	-119,098	(p-value)	(0.0001)	(0.0001)	(0.0001)
France	Pre-ban	34,393	33,408	985	P2-P1	-15,896***	-11,018***	-4,877
	Ban	18,497	22,389	-3,892	(p-value)	(0.0001)	(0.0001)	(0.0001)
Germany	Pre-ban	34,147	42,098	-7,950	P2-P1	-13,644***	-22,214***	8,569***
	Ban	20,503	19,884	618	(p-value)	(0.0001)	(0.0001)	(0.0001)
Ireland	Pre-ban	25,223	17,334	7,888	P2-P1	-12,179***	-5,292***	-6,887***
	Ban	13,043	12,041	1,001	(p-value)	(0.0001)	(0.0001)	(0.0001)
Portugal	Pre-ban	61,307	63,104	-1,797	P2-P1	-38,315***	-46,056***	7,741***
	Ban	22,991	17,047	5,944	(p-value)	(0.0001)	(0.0001)	(0.0001)
Spain	Pre-ban	140,899	83,042	57,857	P2-P1	-46,278***	-15,623***	-30,655***
	Ban	94,621	67,419	27,201	(p-value)	(0.0001)	(0.0001)	(0.0001)
Panel B: Countries with three sub-periods								
Netherlands	Pre-ban	42,065	47,070	-5,004	P2-P1	-26,697***	-19,544***	-7,153***
	Ban	15,368	27,525	-12,157	(p-value)	(0.0001)	(0.0001)	(0.0001)
	Post-ban	20,834	37,234	-16,399	P3-P2	5,466***	9,708***	-4,242***
					(p-value)	(0.0001)	(0.0001)	(0.0001)
Switzerland	Pre-ban	255,436	336,832	-81,396	P2-P1	-152,828***	-144,195***	-8,632
	Ban	102,608	192,637	-90,028	(p-value)	(0.0001)	(0.0001)	(0.5806)
	Post-ban	132,889	225,986	-93,096	P3-P2	30,281***	33,349**	-3,067
					(p-value)	(0.0001)	(0.0129)	(0.8154)
UK	Pre-ban	4,095,195	4,453,783	-358,587	P2-P1	-2,375,180***	-2,210,788***	-164,392***
	Ban	1,720,015	2,242,995	-522,980	(p-value)	(0.0001)	(0.0001)	(0.0006)
	Post-ban	1,832,757	2,730,153	-897,396	P3-P2	112,742***	487,158***	-374,416***
					(p-value)	(0.0058)	(0.0001)	(0.0001)

APPENDIX 3B

Figures 3B.1–3B.10 Changes in price dimension of liquidity: Banned (B) versus non-banned (NB) stocks

These figures illustrate the changes in the median values of the percentage quoted half-spread (PQHS) and percentage effective half-spread (PEHS) for banned (B) and non-banned (NB) stocks, respectively, from one sub-period to the other, across 10 European countries (Belgium, Denmark, France, Germany, Ireland, the Netherlands, Portugal, Spain, Switzerland, and the UK). There are three sub-periods: the pre-ban period, the ban period, and the post-ban period. Only Switzerland, the Netherlands, and the UK have a post-ban period available. The sample period covers the period from mid-January 2008 to mid-December 2009.

Figure 3B.1: Belgium



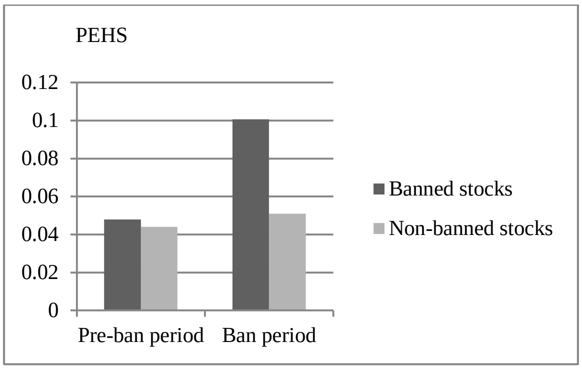
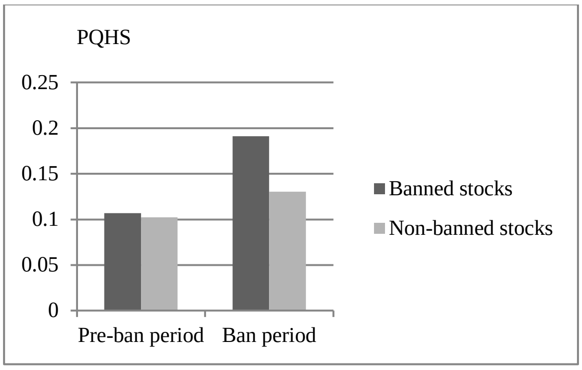


Figure 3B.2: Denmark



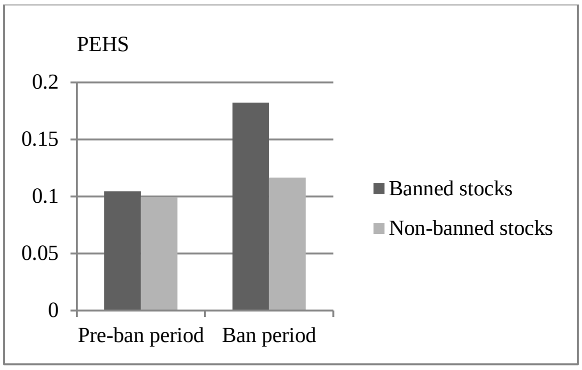
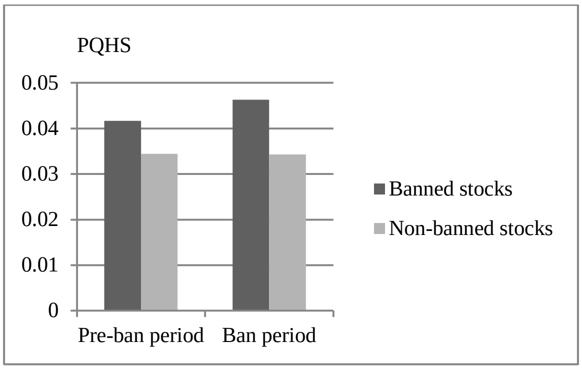
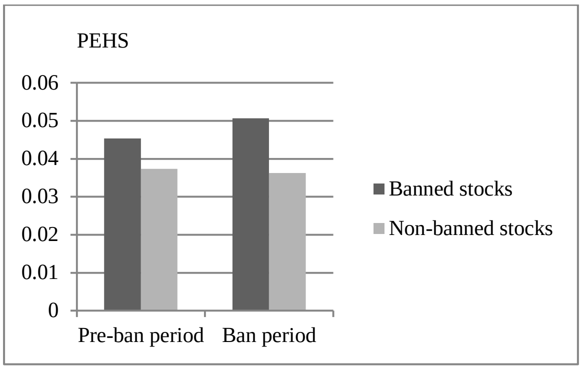


Figure 3B.3: France

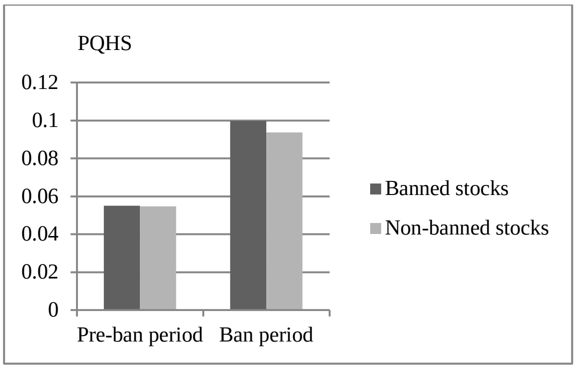




Figures 3B.1–3B.10 Changes in price dimension of liquidity: Banned (B) versus non-banned (NB) stocks

Figures continued from previous page

Figure 3B.4: Germany



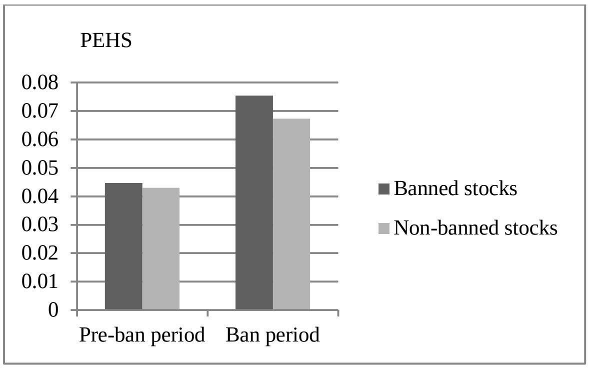
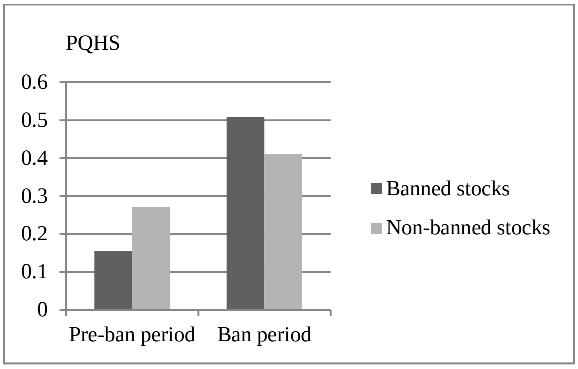


Figure 3B.5: Ireland



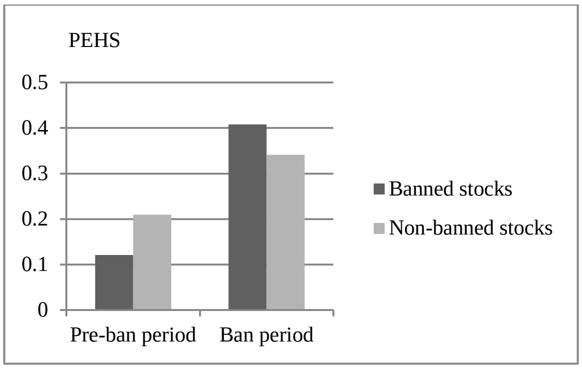
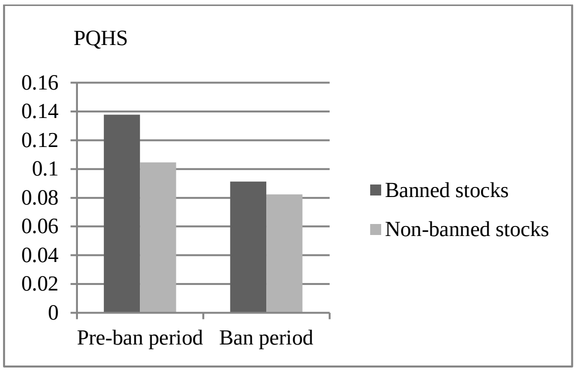
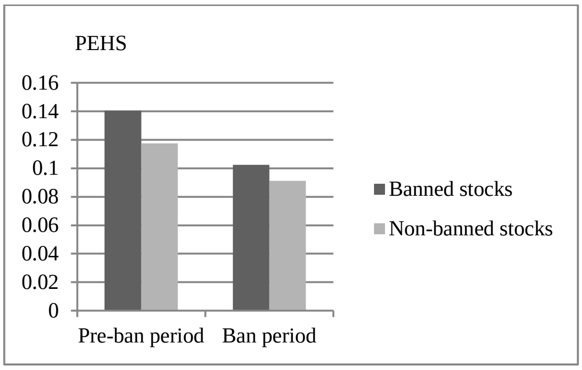


Figure 3B.6: Portugal





Figures 3B.1–3B.10 Changes in price dimension of liquidity: Banned (B) versus non-banned (NB) stocks

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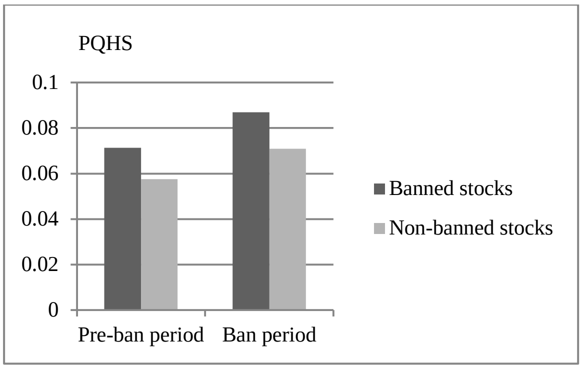
Figure 3B.7: Spain



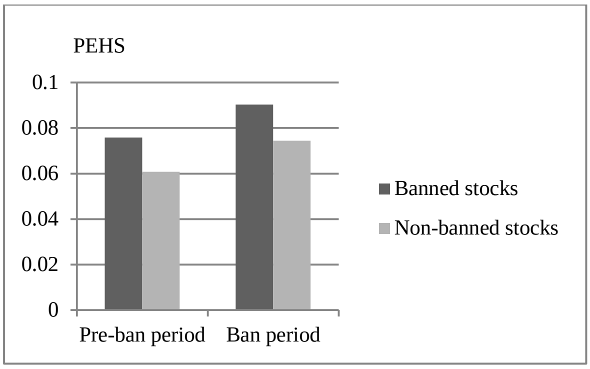
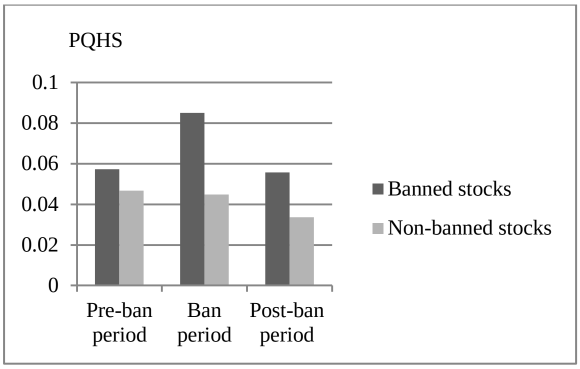


Figure 3B.8: Netherlands



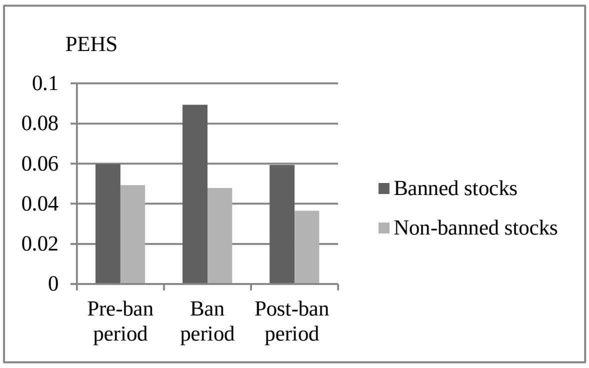
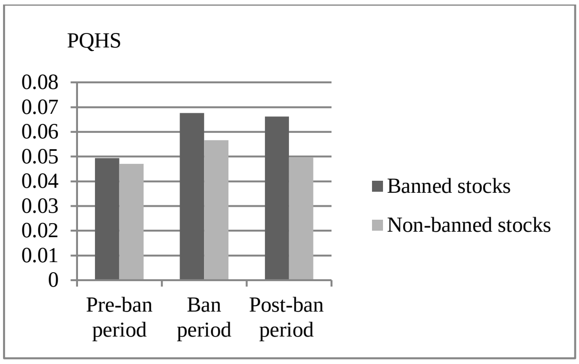
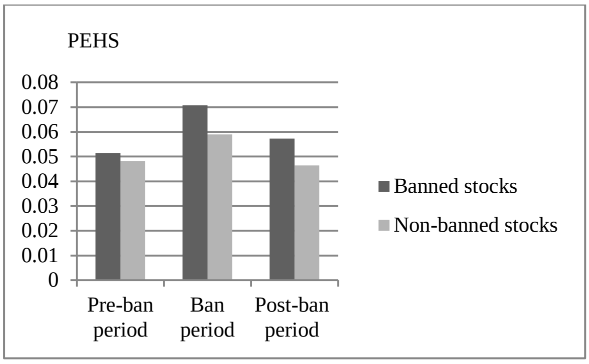


Figure 3B.9: Switzerland





Figures 3B.1–3B.10 Changes in price dimension of liquidity: Banned (B) versus non-banned (NB) stocks

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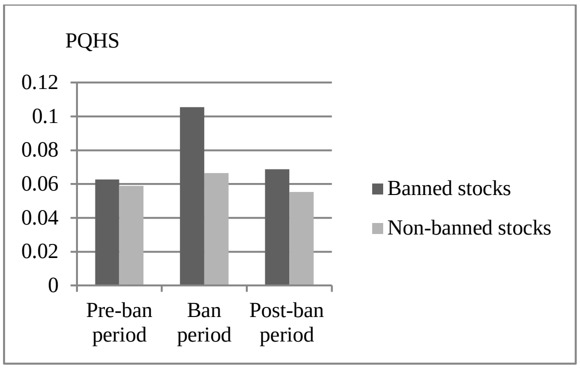
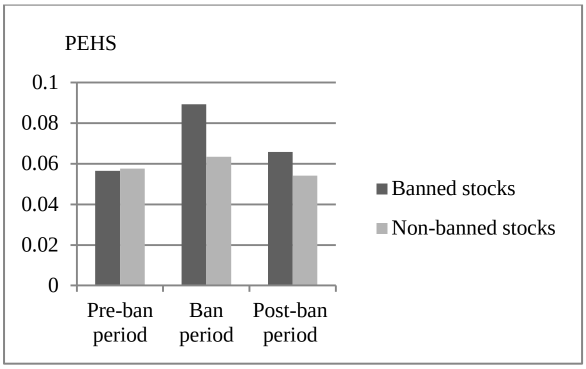
Figure 3B.10: UK

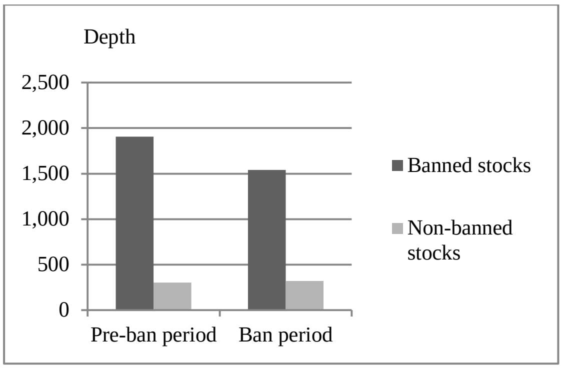




Figures 3B.11–3B.20 Changes in quantity dimension of liquidity: Banned (B) versus non-banned (NB) stocks

These figures illustrate the changes in the median values of the depth and euro depth for banned (B) and non-banned (NB) stocks, respectively, from one short-selling sub-period to the other, across 10 European countries (Belgium, Denmark, France, Germany, Ireland, the Netherlands, Portugal, Spain, Switzerland, and the UK). There are three sub-periods: the pre-ban period, the ban period, and the post-ban period. Only Switzerland, the Netherlands, and the UK have a post-ban period available. The sample period spans from mid-January 2008 to mid-December 2009.

Figure 3B.11: Belgium



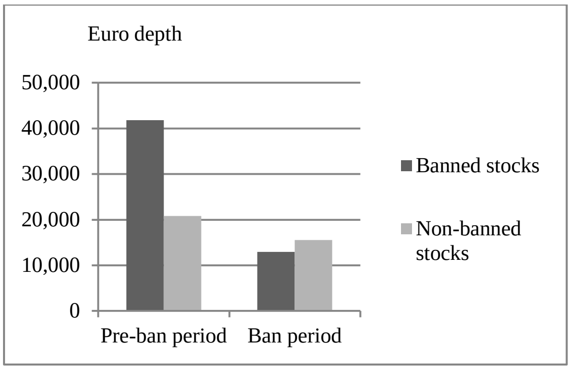
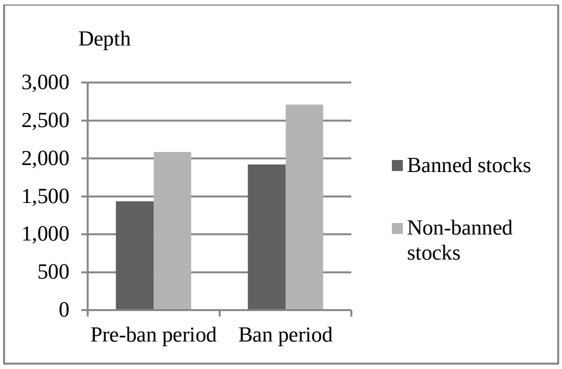


Figure 3B.12: Denmark



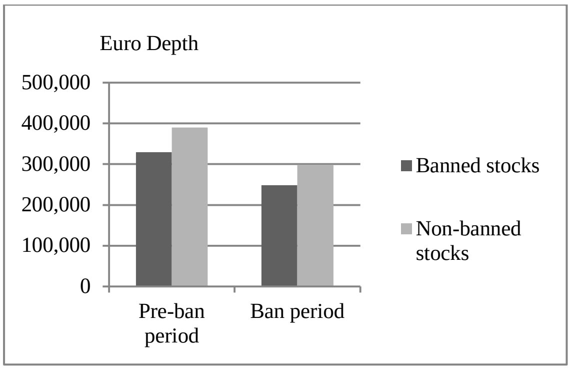
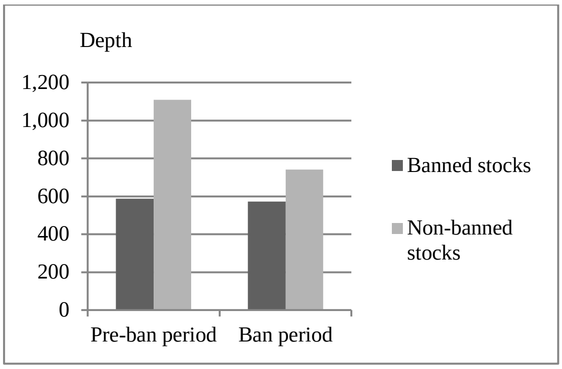
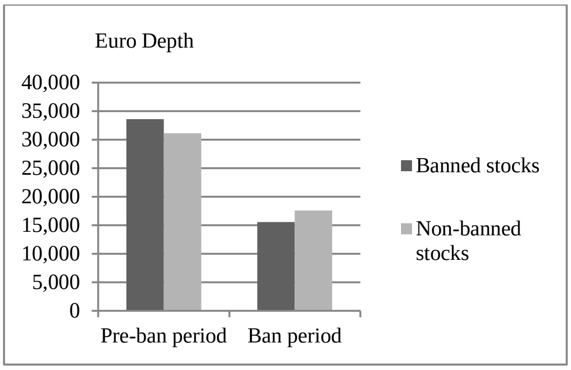


Figure 3B.13: France





Figures 3B.11–3B.20 Changes in quantity dimension of liquidity: Banned (B) versus non-banned (NB) stocks

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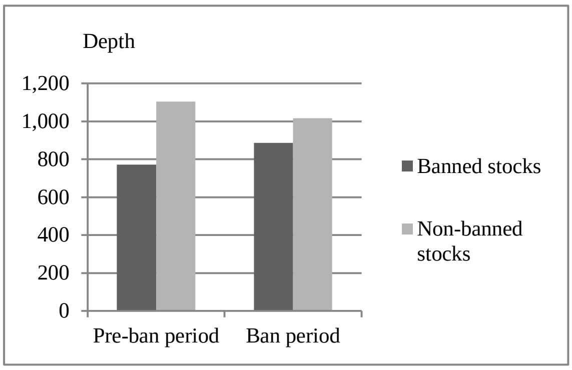
Figure 3B.14: Germany



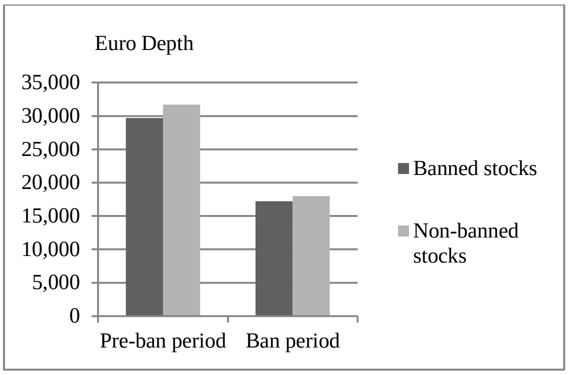
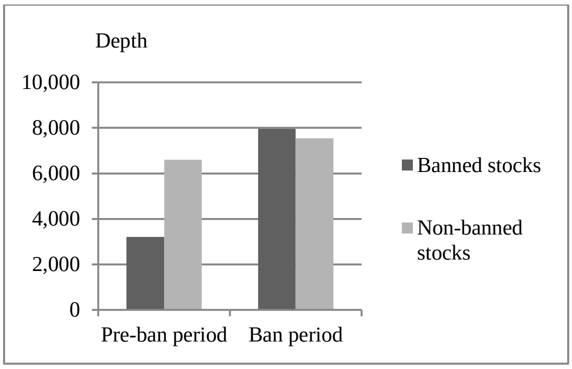
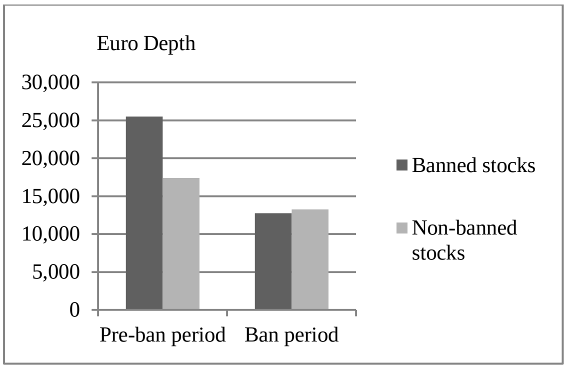


Figure 3B.15: Ireland





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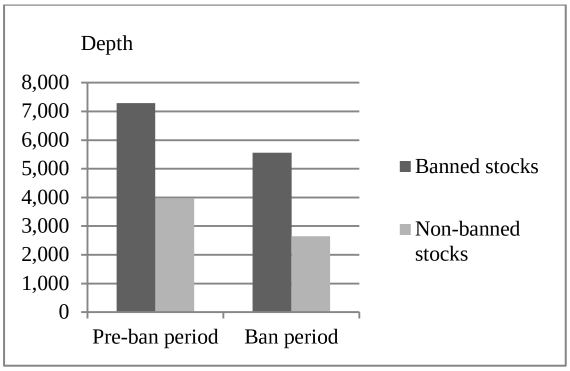

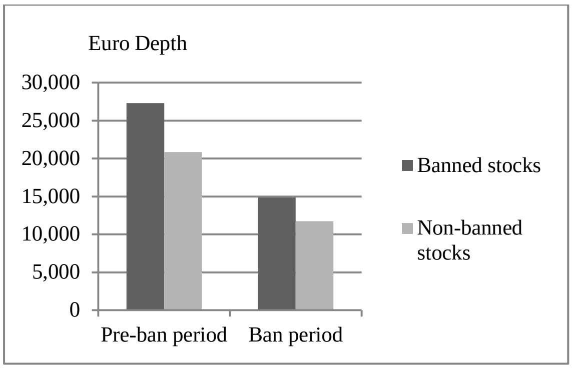

Figures 3B.11–3B.20 Changes in quantity dimension of liquidity: Banned (B) versus non-banned (NB) stocks

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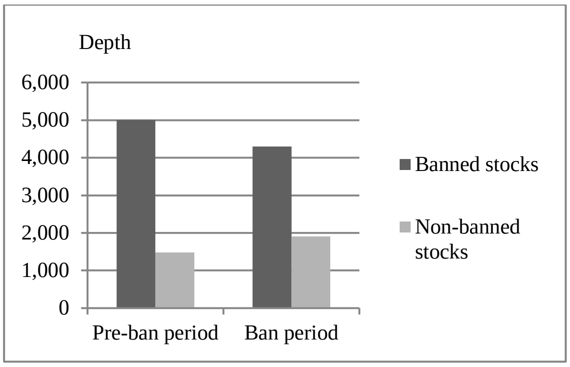
Figure 3B.17: Spain



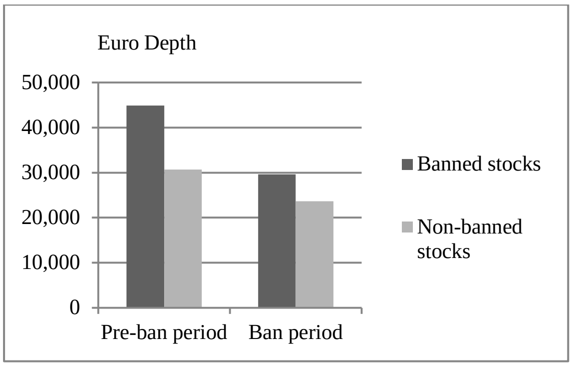
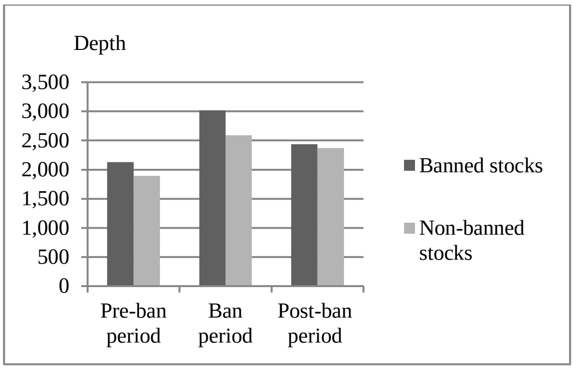


Figure 3B.18: Netherlands



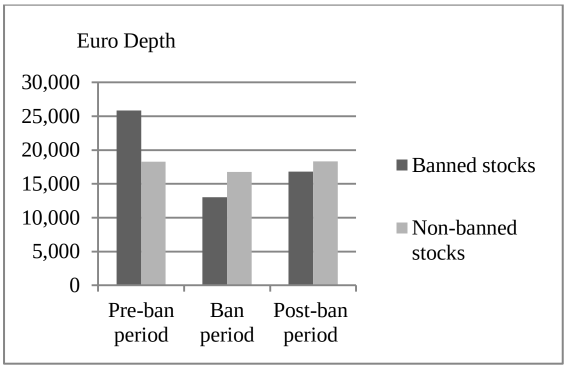
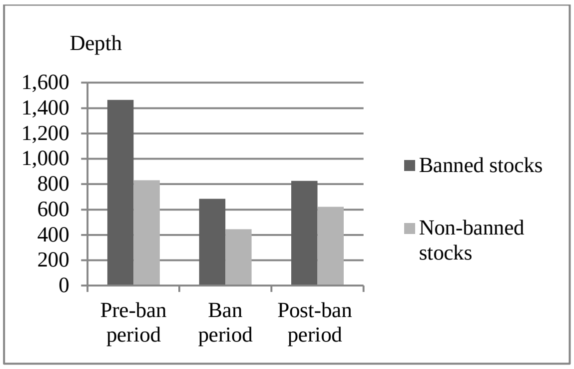
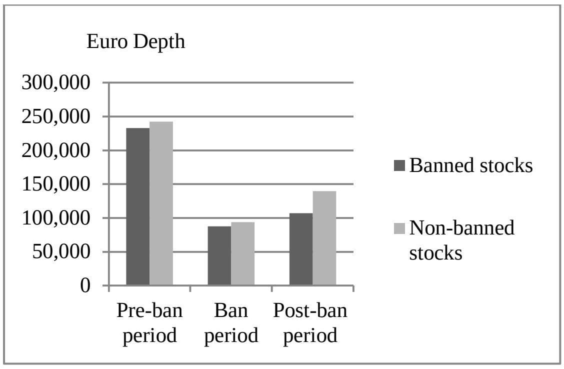


Figure 3B.19: Switzerland





Figures 3B.10–3B.20 Changes in quantity dimension of liquidity: Banned (B) versus non-banned (NB) stocks

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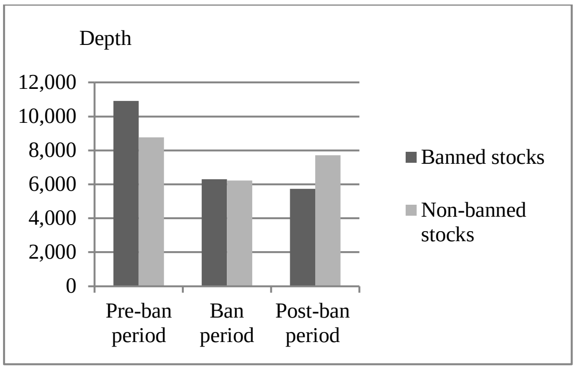
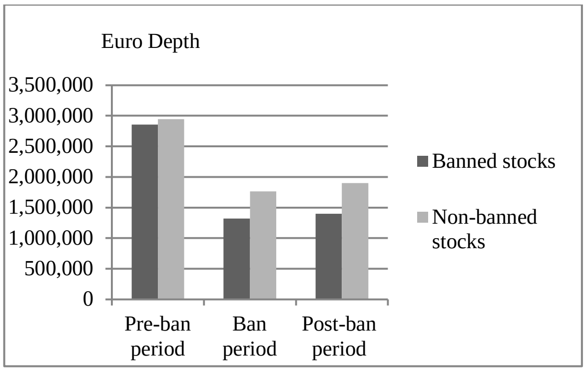
Figure 3B.20: UK





CHAPTER 4

Essay 2: Short-selling regulations and the components of the bid/ask spread

4.1 Introduction

After Lehman Brothers collapsed in September 15, 2008, regulators around the world enforced short-selling regulations to halt the stock price decline of major financial institutions and restore confidence in the safety of the financial system. This essay investigates the impact of the 2008 short-selling regulations on the components of the bid/ask spread in 10 European markets. Prior studies document an increase in bid/ask spread following the introduction of the new short-selling rules (see for example, Bris (2008), Clifton and Snape (2008), Hansson and Fors (2009), Boulton and Braga-Aves (2010), Marsh and Payne (2012), and Beber and Pagano (2013)). This essay extends the earlier studies by decomposing the bid/ask spread into its components to identify the source of the increase noted by prior researchers.

According to theory (see for example, Stoll (1989)) the bid/ask spread can be divided into three components. These are order processing costs, inventory holding costs, and adverse selection costs. Order processing costs arise from matching buying and selling orders and securing transaction completion by liquidity providers (Tinic (1972)). Inventory holding costs refer to the liquidity providers' holding of a portfolio of stocks that is not fully diversified in order to provide immediacy to investors (Stoll (1978a, b)). Lastly, adverse selection costs arise from trading with informed traders and they compensate liquidity suppliers for the risk associated with such trading (Glosten and Milgrom (1985)). Order processing costs and inventory holding costs comprise the non-informational component of the bid/ask spread, which has a temporary impact on stock prices. On the other hand, adverse selection costs are the informational component of the bid/ask spread and their impact on stock prices is permanent (see Hasbrouck (1988)).

Boehmer, Jones, and Zhang (2013) examine the impact of the short-selling ban on the informational (or price impact) and non-informational (or realized spread) components of the bid/ask spread but focus only on the US market. Marsh and Payne (2012) investigate the effects of the short-selling ban on the price impact in the UK. However, the authors use the price impact as a proxy for liquidity rather than the informational component of the bid/ask spread. In contrast, this essay estimates the impact of the short-selling regulations on both components of the bid/ask spread across 10 European countries.

This chapter is organized as follows. Section 4.2 reviews the related literature. Section 4.3 develops the hypotheses. Section 4.4 describes the sample and the data. Section 4.5 describes the methodology. Section 4.6 presents and discusses the empirical results. Section 4.7 provides a summary of the main findings and the concluding remarks.

4.2 Literature Review

4.2.1 Conceptualizing, modeling, and estimating the components of the bid/ask spread: Earlier studies

The bid/ask spread is the cost traders incur in order to trade securities. Market makers or liquidity suppliers need to be compensated for providing immediacy to customers who wish to trade instantly (Demsetz (1968); Tinic (1972)). Moreover, the bid/ask spread reflects the risks associated with managing the liquidity suppliers' inventory (Stoll (1978a, b)) and the lost revenue due to informed trading on the other side of the trade (Copeland and Galai (1983); Easley and O'Hara (1987)). Consequently, there are three main costs involved in the bid/ask spread: order processing costs (OPC), inventory holding costs (IHC), and adverse selection costs (ASC). Changes in all three components may affect the level of the bid/ask spread at the same time. The final, new level of the bid/ask spread depends, on the one hand, on the direction of the effect, and on the other, on the magnitude of the effect (Stoll (1989)).

The two main approaches that model the bid/ask spread and its components are an inventory-based approach and an information-based approach. Stoll (1978a) develops an inventory-based model, in which an individual dealer's trading costs are determined not only by the risk she bears, but also by the way the market is structured. According to the author, a dealer quotes prices to encourage transactions by investors that will move her closer to a desired portfolio. Information costs are peripheral to the analysis, which focuses on inventory management issues and the supply side of the trade. Ho and Stoll (1981) extend the model by adding the demand side. These authors argue that liquidity suppliers bear both uncertainty about the return on their inventory and uncertainty about when future transactions will take place. They conclude that uncertainty of demand to trade with the dealer is not completely eliminated or offset by the dealer's pricing strategy. Ho and Stoll (1983) move one step further by extending a single dealer, stock, and period models to multiple dealers, stocks, and periods, where the dealer's welfare depends on the uncertainty about inventory returns, the arrival of transactions, and the prices set by other dealers.

Copeland and Galai (1983) adopt an information-based approach. The authors argue that there is a trade-off between expected losses to informed traders and expected gains from liquidity traders. In their framework, the presence of informed traders justifies a positive bid/ask spread even when the dealer's IHC are zero. Glosten and Milgrom (1985) argue that bid/ask spreads can be a purely informational phenomenon occurring even when all the specialist's fixed and variable transaction costs are zero and when competition forces the specialist's profit to zero. In contrast to Copeland and Galai (1983) who assume that private information is revealed immediately as soon as the trade occurs, Glosten and Milgrom (1985) allow further trading until the time the information is revealed to resolve the informational differences between insiders and the rest of the market.

A variety of empirical models have been used to estimate the components of the bid/ask spread. One group of models is based on the serial covariance properties of transaction and quote prices. For example, Roll (1984) assumes that market makers face only order processing costs and measures the effective bid/ask spread based on transaction price changes. Stoll (1989) estimates all three components of the bid/ask spread and finds that OPC, ASC, and IHC comprise 47%, 43%, and 10% of the bid/ask spread, respectively. In their model, George, Kaul, and Nimalendran (1991) account for time-varying expected returns and the impact of unexpected returns when estimating the bid/ask spread components. They find that OPC is the predominant component. Another group of spread decomposition models is based on a trade initiation indicator variable rather than on the serial covariance properties of transaction and quote prices. Examples include the models of Glosten and Harris (1988) and Madhavan, Richardson, and Roomans (1997).

Prior empirical research in this area, which applies spread decomposition models, focuses on analysing the behaviour of the bid/ask spread components (i) on an intraday basis (see for example, McInish and Van Ness (2002) and Angelidis and Benos (2009)), (ii) around takeover and earnings announcements (for example, Krinsky and Lee (1996)), (iii) in different stock exchange environments (Porter and Weaver (1996); Wang (1999); Chung, Van Ness, and Van Ness (2004)), (iv) after a change in trading regimes (Giouvris and Philippatos (2008)), and (v) after a change in regulations (McInish, Van Ness, and Van Ness (2001); Frijins, Gilbert, and Tourani-Rad (2008)). However, many of these models generate bid/ask spread component estimates which are theoretically implausible.[134] This essay adopts the approach of Bessembinder and Kaufman (1997), which is based on realized spreads and

[134] See for example, Clarke and Shastri (2000) and Van Ness, Van Ness, and Warr (2001).

price impacts, arguably better measures of the non-information and the informational components of the bid/ask spread.

4.2.2 Short-selling regulations and bid/ask spread components

Boehmer, Jones, and Zhang (2013) examine the impact of the short-selling ban introduced in the US on September 18, 2008 on the components of the bid/ask spread. The authors divide the bid/ask spread into two components – the realized spread and the price impact. The treatment sample comprises 727 NYSE and NASDAQ financial stocks subject to the short-selling ban (banned stocks). The authors match each stock in the banned sample with a stock not subject to the short-selling ban, creating a control sample of non-banned stocks. Since the short-selling ban was applied to most financial institutions, the authors match the banned stocks from the financial sector with non-banned stocks from other sectors by listing exchange, listed options status, market capitalization, and dollar trading volume.

The authors report that in the pre-ban period the 5-minute price impact is 72 basis points for the banned stocks and 77 basis points for the control stocks. During the ban period, the respective figures are 107 and 104 basis points. The realized spread is at 137 basis points for the banned stocks and at 105 basis points for the control stocks during the pre-ban period. In the ban period, the realized spread for the banned stocks increases to 213 basis points while the realized spread for the control stocks increases to 155 basis points. Furthermore, the multivariate analysis shows that during the period of the short-selling ban the price impact and the realized spread of the banned stocks are greater than those of the non-banned stocks by 10 and 15 basis points, respectively. Both results are statistically significant at the 1% level. Overall, the results suggest that both components of the bid/ask spread increase more for the banned stocks relative to the components of the non-banned stocks.

Marsh and Payne (2012) study the impact of the short-selling ban introduced in September 19, 2008 on market quality in the UK. The authors match 10 non-financial non-banned stocks to each of the 23 financial banned stocks by average market values computed over the first half of 2008. Hasbrouck's (1991) VAR approach is used to estimate the change in the price impact of the financial banned stocks relative to the change in the price impact of the matched non-banned stocks. The price impact is estimated as the impulse response of prices to a trade innovation implied by the VAR.

The authors find that during the pre-ban period, there is no significant difference in the price impact between the two samples of stocks. For example, during the pre-ban period, the average price impact for the banned stocks is 0.57 and for the non-banned stocks is 0.55. However, during the ban period, the price impact for the banned stocks increases to 0.95 while the price impact for the non-banned stocks increases to 0.77. The authors argue that the drain in liquidity, as measured by the price impact, is greater for financial banned stocks when compared to that for the non-banned stocks, and this difference is attributed to the short-selling ban.

In summary, there is evidence showing that the wider bid/ask spreads following the introduction of the short-selling ban in the US arose from a significant increase in both the non-informational (i.e., realized spread) and informational (i.e., price impact) components of the bid/ask spread, as documented by Boehmer, Jones, and Zhang (2013). Marsh and Payne (2012) find that during the ban period, the price impact of banned stocks is greater relative to the price impact of non-banned stocks in the UK. However, Marsh and Payne (2012) use a different methodology from Boehmer, Jones, and Zhang (2013). More precisely, Marsh and Payne (2012) do not decompose the bid/ask spread into its two components, and in their framework, price impact proxies for liquidity as a whole rather than the informational component of the bid/ask spread. This essay follows the work of Boehmer, Jones, and Zhang

(2013) by decomposing the bid/ask spread into its non-informational and informational components but in a European context.

4.3 Hypotheses development

One unanticipated – at least from the regulators' point of view – consequence of the introduction of the 2008 short-selling regulations was the increase in bid/ask spreads of stocks subject to these regulations. Beber and Pagano (2013) provide some insights into how the 2008 short-selling bans could potentially affect the non-informational and the informational components of the bid/ask spread. They make the point that in the absence of any exemption, market makers will find it more difficult to hedge their inventory position, and this will tend to widen the bid/ask spreads as the inventory holding costs increase. Even if market makers are exempt from the short-selling bans, the reduced competition they will face from other liquidity suppliers will also tend to widen the non-informational component of the bid/ask spread. Boehmer, Jones, and Zhang (2013) find that the relative non-informational component or the realized spread of the financial banned stocks does increase in the US during the ban period. Based on the insights provided by Beber and Pagano (2013) and the prior findings, this essay hypothesizes that:

Hypothesis 1. The non-informational component or the realized spread of the European financial stocks subject to the 2008 short-selling bans increases (decreases) during the ban (post-ban) period.

Furthermore, Beber and Pagano (2013) argue that prices will tend to be less informative in an environment where short-selling is prohibited, and this will tend to increase the risk to market makers and other liquidity suppliers. Diamond and Verrecchia's (1987) model shows that short-selling prohibitions will hinder the rapid adjustment of prices to new information, particularly bad news. Beber and Pagano (2013) observe that the delayed resolution of uncertainty will put upward pressure on the adverse selection costs. Nevertheless, they also observe that if short sellers are informed traders, then the introduction of short-selling bans will tend to reduce trading activity by informed traders, putting downward pressure on the adverse selection costs. Thus, the effect of the 2008 short-selling regulations on the informational component of the bid/ask spread is an empirical issue. Empirical evidence from the US market shows that the relative informational component of the bid/ask spread (i.e., or the price impact) of the banned stocks increases during the ban period (Boehmer, Jones, and Zhang (2013)). This result implies that the effect of the delayed resolution of uncertainty is stronger than the effect of reduced trading activity by informed traders. Consequently, this essay hypothesizes that:

Hypothesis 2. The informational component or the price impact of the European financial stocks subject to the 2008 short-selling bans increases (decreases) during the ban (post-ban) period.

4.4 Sample and data

4.4.1 Introduction

Section 4.4 is identical to section 3.4 as the same sample and data are used for both essays. Nevertheless, sections 3.4.1 and 3.4.2 are discussed again in this essay for completeness while section 4.4.3 summarizes briefly section 3.4.3.

Thomson Reuters Tick History (TRTH) database of Securities Industry Research Centre of Asia-Pacific (SIRCA) provided intraday quote and trade data on 78 European financial stocks subject to the short-selling regulations. Datastream provided data on the market value of the same stocks. The sample period covers the period from mid-January 2008 to mid-December 2009, which is divided into three sub-periods. The first sub-period (i.e., the pre-ban period) covers the period before the introduction of the new short-selling regulations. The second sub-period (i.e., the ban period) covers the period when the short-selling regulations are in force. Finally, the third sub-period (i.e., the post-ban period) covers the period following the short-selling regulations relaxation (See Table 3.1 on page 64 for a summary of the type of short-selling regulations introduced in each country, the number of financial stocks affected by these short-selling regulations that are included in the sample, and the duration of each sample sub-period. Additionally, see Table 3.2 on page 64 for the trading hours of the stock exchanges in which the sample stocks primarily trade).

4.4.2 Data filtering

The following filters are applied to raw data obtained from the TRTH database of SIRCA. Firstly, the first and last 15 minutes of trading are deleted to control for volatility spikes that occurred during the opening and closing trading sessions (Bessembinder and Kaufman (1997)). Secondly, data are winsorized at 1 and 99 percentiles, respectively, to ameliorate the impact of outliers on the results. Thirdly, data with the following characteristics are excluded from the analysis to minimize data errors (Huang and Stoll (1996)):

(i) Bid/ask quotes if the spread is greater than €3 or negative.[135]

(ii) Trade price p_t when $(p_t-p_{t-1})/p_{t-1} > 0.10$

(iii) Ask quote a_t when $(a_t-a_{t-1})/a_{t-1} > 0.10$

(iv) Bid quote b_t when $(b_t-b_{t-1})/b_{t-1} > 0.10$

After the filtering process, daily data are derived from the intraday data for the analysis.

4.4.3 Control sample

This section briefly presents the process of constructing the control sample (see section 3.4.3 on pages 65–69 for a more detailed description of the construction of the control sample).

[135] For stocks traded in the UK, Switzerland, and Denmark, bid/ask quotes are removed if the spread is greater than 2.5 GBP, 4 CHF, and 23 DKK, respectively.

Each financial stock subject to a short-selling ban (banned stock) is matched to a non-financial stock not subject to a short-selling ban (non-banned or control stock) that (i) is traded on the same stock exchange, and (ii) has similar characteristics such as price, trading volume, and market value, and similar derivatives listings as the financial stock. A matching score is computed as:

$$Score = \left(\frac{P_{banned} - P_{non-banned}}{\frac{P_{banned} + P_{non-banned}}{2}}\right)^2 + \left(\frac{EV_{banned} - EV_{non-banned}}{\frac{EV_{banned} + EV_{non-banned}}{2}}\right)^2 + \left(\frac{MV_{banned} - MV_{non-banned}}{\frac{MV_{banned} + MV_{non-banned}}{2}}\right)^2$$

where P_{banned} is the average price of each banned stock and $P_{non\text{-}banned}$ is the average price of a non-banned stock over the period from January, 2008 to July, 2008. EV_{banned} is the average trading volume in euros (or turnover) for each banned stock and $EV_{non\text{-}banned}$ is the average trading volume in euros of a non-banned stock over the period from January, 2008 to July, 2008.[136] Finally, MV_{banned} is the average market value for each banned stock and $MV_{non\text{-}banned}$ is the average market value of a non-banned stock over the same period. A non-banned stock that best matches the banned stock by generating the lowest score is chosen as a control stock.

4.5 Methodology

4.5.1 Measurement of the dependent variables

This section describes how the non-informational and informational components are measured. Empirical evidence shows that many spread decomposition models generate theoretically implausible component estimates (see for example, Clarke and Shastri (2000) and Van Ness, Van Ness, and Warr (2001)). Consequently, this essay decomposes the effective bid/ask spread into its non-informational component (inventory holding and order

[136] Although the variable is named as EV (i.e., Euro Volume), turnover for British, Danish, and Swiss stocks is expressed in their domestic currencies (i.e., British pound, Danish krone, and Swiss franc, respectively).

processing components) and informational component (adverse selection component) using the method employed by Bessembinder and Kaufman (1997).

The **realized spread**, which measures the non-informational component of the bid/ask spread, is calculated as

$$\%Realized\ spread = D_{it} \times \left[\frac{P_{it} - M_{it+n}}{M_{it}}\right] \times 100 = \%\ Effective\ spread - \%\ Price\ Impact \qquad (4.1)$$

where D_{it} is a trade indicator variable that equals 1 for buyer-initiated orders (i.e., when the transaction price is greater than the quote midpoint) and -1 for seller-initiated orders (i.e., when the transaction price is less than the quote midpoint).[137] P_{it} is the transaction price of stock i at time t while M_{it} is the quote midpoint, which represents the stock's true value. M_{it+n} is the quote midpoint n periods after the transaction. In this essay, n is equal to 5 minutes.

The **adverse selection cost** that the liquidity supplier incurs is estimated as

$$\%\ Price\ impact\ of\ trade = D_{it} \times \left[\frac{M_{it+n} - M_{it}}{M_{it}}\right] \times 100 \qquad (4.2)$$

4.5.2 Regression modeling

Short-selling regulations were introduced as a response to the global financial crisis and mainly applied to financial stocks. This essay uses a difference-in-difference multivariate analysis to reduce the impact of severe market movements during the sample period.[138] To test Hypotheses 1 and 2, the regression analysis is first run at a country level using the model of Boehmer, Jones, and Zhang (2013):

[137] See Ready and Lee (1991).

[138] Belgium, Denmark, and Ireland are excluded from the regression analysis at a country level due to their small sample size. Nevertheless, these countries are included in the regression analysis at an aggregate level where all the countries are pooled together.

$$Diff\_Y_{it} = \alpha_i + \beta D_{it}^{Ban} + \theta X_{it} + e_{it} \qquad (4.3)$$

where $Diff\_Y_{it}$ is the difference between the percentage price impact or PPI (percentage realized spread or PRS) of the banned sample (B) and the PPI (PRS) of the non-banned sample (NB). D_{it}^{Ban} is a dummy variable that takes the value of 1 when short-selling bans are in force and 0 otherwise. This variable is of most interest to the analysis as it attempts to estimate the impact of the short-selling bans on the components of the bid/ask spread of the financial stocks subject to these bans after controlling for other factors that could drive the results. X_{it} is a vector of differences between the control variables of the matched pairs of stocks. The control variables included in the analysis are: market value, trading volume in euros or in their respective domestic currencies for the UK, Switzerland, and Denmark, proportional daily range of transaction prices, and daily volume-weighted average price (VWAP).

OPC are predicted to have a negative relationship with a stock's price and trading activity (e.g., trading euro volume or number of trades). Because OPC are usually fixed costs, they tend to decrease as the price and trading activity of the stock increase (Stoll (1978)). IHC are also predicted to have a negative relationship with trading activity (see for example, Stoll (1978), Tripathy and Peterson (1991), and Menyah and Paudyal (1996)). Consequently, a negative relationship is expected between the non-informational component of the bid/ask spread or realized spread and the euro volume. Stoll (2000) predicts an inverse relationship between IHC and firm's size. He argues that trading stocks of larger firms increase the probability of locating a counterparty, thus reducing the risk of holding inventory. Therefore, a negative relationship is expected between the non-informational component of the bid/ask spread and the market value of the firms in the sample. In contrast, IHC and subsequently the non-informational component of the bid/ask spread are expected to have a positive relationship with the proportional daily range of transaction prices, which proxies for volatility. The more volatile the prices of a stock, the riskier it is for market makers to hold a less than optimal portfolio in order to provide immediacy to their clients (Stoll (1978 a, b)).

The informational component of the spread or price impact is predicted to have a negative relationship with the stock's price and market value. Higher stock price and greater market value of a firm can be associated with greater stability and disclosure and lower probability of informed trading (Stoll (2000)). Also, the informational component of the bid/ask spread is expected to have a negative relationship with trading volume. This is because market makers increase their chances of being compensated for their losses from trading with informed traders through their trading with uninformed traders when the stocks are more actively traded. Finally, a positive relationship is expected between the informational component of the bid/ask spread and volatility. When stock prices are highly volatile informed traders are expected to have more opportunities to trade at the expense of less informed traders, which tends to increase the informational component of the bid/ask spread (Copeland and Galai (1983); Easley and O'Hara (1987)).

A similar analysis is conducted separately for countries that have a post-ban period available.[139] A post-ban period dummy variable – $D_{it}^{Post\text{-}ban}$ – is added to equation (4.3), which now takes the following form:

$$Diff\_Y_{it} = \alpha_i + \beta D_{it}^{Ban} + \gamma D_{it}^{Post\_ban} + \theta X_{it} + e_{it} \tag{4.4}$$

where $D_{it}^{Post\text{-}ban}$ is equal to 1 when short-selling bans are lifted and 0 otherwise.

After this, matched pairs from all countries are pooled together. To test Hypotheses 1 and 2 at an aggregate level, again equation (4.3) is used.

[139] These countries are the Netherlands, Switzerland, and the UK.

Fixed effects are used to account for any differences between the matched stocks in the pre-ban period. The control variables undergo logarithmic transformation to reduce any skewness and heteroskedasticity problems that might occur in the regression analysis (Benston and Hagerman (1974)). Lastly, the Newy-West procedure is used to generate estimates robust to autocorrelation and heteroskedasticity in residuals.

4.6. Empirical Results

4.6.1 Univariate analysis

Table 4.4 reports the average median values of the percentage realized spread (PRS) for the banned (B) and non-banned (NB) stocks, respectively, and the difference in PRS between the two samples across 10 European countries.[140] Panel A reports the results for the seven countries that have two sub-periods, the pre-ban period and the ban period, which are Belgium, Denmark, France, Germany, Ireland, Portugal, and Spain. Panel B reports the results for the three countries that also have a post-ban period available (i.e., Switzerland, the Netherlands, and the UK). PRS (B) – PRS (NB) in column (3) is the average median of the difference between the PRS of the banned stocks and the PRS of the non-banned stocks.[141] Additionally, Table 4.4 shows the statistical significance of the test for equality of medians of the PRS for each sample of stocks separately as well as the relative PRS of the banned stocks or PRS (B) – PRS (NB).

[140] The mean values of the percentage realized spread and the percentage price impact are also provided in Appendix 4A.

[141] Although a robust matching process is used, as described in section 3.4.3 of Chapter 3, stocks in each sample belong to different industries and thus might be affected by different factors, which could lead to discrepancies in their difference variable. Consequently, the average median value of the variable PRS (B) – PRS (NB) is different from the difference simply between the average median of PRS (B) and the average median of PRS (NB).The same holds for the variable PPI (B) – PPI (NB).

Table 4.1 Median percentage realized spread (PRS) for banned (B) and non-banned (NB) stocks

Average median values of the percentage realized spread (PRS) for banned (B) and non-banned (NB) stocks, respectively, as well as their difference (i.e., PRS (B) – PRS (NB)) are reported for 10 European countries. The sample period spans from mid-January 2008 to mid-December 2009 and it is divided into three sub-periods: the pre-ban period, the ban period, and the post-ban period. The results for countries that have two sub-periods, the pre-ban period and the ban period (i.e., Belgium, Denmark, France, Germany, Ireland, Portugal, and Spain) are shown in Panel A. Only Switzerland, the Netherlands, and the UK have a post-ban period available. Their results are reported in Panel B. The sub-periods differ for each country and are specified in Table 3.1 of Chapter 3. Table 4.1 also shows the statistical significance of the test for equality of medians of the PRS for both samples and their difference. Variable Diff is the change in the median PRS between the pre-ban (P1) and ban (P2) periods (P2–P1), and ban and post-ban (P3) periods (P3–P2), respectively. Diff-in-Diff is the change in the difference between the PRS of the banned stocks and the PRS of the non-banned stocks from the pre-ban period to the ban period and from the ban period to the post-ban period, respectively. The Wilcoxon test is used and the p-values of the Z-statistic are given in the parentheses. * indicates statistical significance at the 10% level. ** indicates statistical significance at the 5% level. *** indicates statistical significance at the 1% level.

Panel A: Countries with two sub-periods								
		(1)	(2)	(3)		(4)	(5)	(6)
Country	Sub-periods	PRS(B)	PRS(NB)	PRS(B) – PRS(NB)		Diff PRS(B)	Diff PRS(NB)	Diff-in-Diff PRS(B) – PRS(NB)
Belgium	Pre-ban	0.0047	-0.0052	0.0094	P2-P1	0.0032***	-0.0035***	0.0058***
	Ban	0.0079	-0.0087	0.0152	(p-value)	(0.0011)	(0.0050)	(0.0011)
Denmark	Pre-ban	0.0687	0.0641	0.0037	P2-P1	0.0688***	0.0074***	0.0527***
	Ban	0.1375	0.0715	0.0564	(p-value)	(0.0001)	(0.0001)	(0.0001)
France	Pre-ban	-0.0039	-0.0010	-0.0020	P2-P1	-0.0003	0.0000	-0.0010**
	Ban	-0.0042	-0.0010	-0.0030	(p-value)	(0.7021)	(0.4748)	(0.0494)
Germany	Pre-ban	0.0123	0.0099	0.0001	P2-P1	0.0072***	-0.0014***	0.0103***
	Ban	0.0195	0.0085	0.0104	(p-value)	(0.0001)	(0.0026)	(0.0001)
Ireland	Pre-ban	0.0675	0.0915	-0.0296	P2-P1	0.21***	0.0394***	0.1213***
	Ban	0.2775	0.1309	0.0917	(p-value)	(0.0001)	(0.0001)	(0.0001)
Portugal	Pre-ban	0.0483	0.0337	0.0032	P2-P1	-0.0153	-0.0159***	0.0144***
	Ban	0.0330	0.0178	0.0176	(p-value)	(0.1044)	(0.0001)	(0.0001)
Spain	Pre-ban	0.0126	0.0001	0.0144	P2-P1	0.0089***	0.004***	0.0046***
	Ban	0.0215	0.0041	0.0190	(p-value)	(0.0001)	(0.0009)	(0.0016)
Panel B: Countries with three sub-periods								
Netherlands	Pre-ban	0.0024	0.0013	0.0006	P2-P1	0.0007	-0.0026*	0.0038
					(p-value)	(0.9064)	(0.0531)	(0.6228)
	Ban	0.0031	-0.0013	0.0044	P3-P2	0.0065***	0.0027***	0.0008
	Post-ban	0.0096	0.0014	0.0052	(p-value)	(0.0020)	(0.0012)	(0.7791)
Switzerland	Pre-ban	0.0095	0.0093	0.0012	P2-P1	-0.0087***	-0.0062***	-0.0019
					(p-value)	(0.0001)	(0.0001)	(0.1145)
	Ban	0.0008	0.0031	-0.0007	P3-P2	0.0235***	0.0157***	0.0046***
	Post-ban	0.0243	0.0188	0.0039	(p-value)	(0.0001)	(0.0001)	(0.0012)
UK	Pre-ban	0.0041	0.0066	-0.0011	P2-P1	-0.0023**	-0.0068***	0.0032**
					(p-value)	(0.0464)	(0.0001)	(0.0381)
	Ban	0.0018	-0.0002	0.0021	P3-P2	-0.0051***	-0.0006	-0.0048***
	Post-ban	-0.0033	-0.0008	-0.0027	(p-value)	(0.0001)	(0.9212)	(0.0006)

The data in column (4) of Panel A of Table 4.1 show that PRS (B) increases for five out of seven countries, and this is statistically significant at the 1% level. There is no statistically significant change in the PRS (B) for France and Portugal. For example, PRS (B) for Germany increases significantly by 0.0072 or 72 basis points in the ban period.

Column (5) shows that the PRS (NB) increases for Denmark, Ireland, and Spain and it decreases for Belgium, Germany, and Portugal. All the results are statistically significant at the 1% level. For example, PRS (NB) for Germany decreases significantly by 0.0014 during the ban period.

As shown in column (6) of Panel A of Table 4.1, the relative PRS of the banned stocks increases for six out of seven countries during the ban period. These results are statistically significant at the 1% level. France is the exception because the relative PRS of the French banned stocks decreases in the ban period, and this is significant at the 5% level.

Column (3) of Panel B of Table 4.1 shows that the relative PRS of the British banned stocks increases from -0.0011 in the pre-ban period to 0.0021 in the ban period. This result is statistically significant at the 5% level, as reported in column (6). There is no statistically significant change in the relative PRS of the banned stocks for the Netherlands and Switzerland during the ban period (see column (6)).

In the post-ban period, column (6) of Panel B of Table 4.1 shows that the relative PRS of the banned stocks increases for Switzerland and decreases for the UK. Both results are statistically significant at the 1% level. There is no statistically significant change in the relative PRS of the Dutch banned stocks.

Overall, the univariate analysis shows that the relative PRS of the banned stocks increases significantly for seven out of 10 countries. These results generally support Hypothesis 1 and are consistent with the findings of Boehmer, Jones, and Zhang (2013), who find that the relative PRS of the banned stocks increases significantly during the ban period in the US. However, there is only weak evidence that the relative PRS of the banned stocks declines following the relaxation of the short-selling bans. In particular, this essay finds that the relative PRS of the banned stocks only decreases significantly during the post-ban period in the case of the UK– one of the three countries with a post-ban period available.

Figure 4.1 depicts the time series of the average PRS of the banned and non-banned stocks, respectively, over the period from May 2008 to January 2009. The UK was one of the first countries to both apply short-selling regulations on September 19, 2008, and subsequently lift them on January 16, 2008. These two dates are used as the starting and ending points of the short-selling bans in the figure.[142] Figure 4.1 shows that after the introduction of the short-selling bans, the PRS of the banned stocks was generally higher relative to the PRS of the non-banned stocks.

[142] Other European countries followed the UK and introduced short-selling regulations within the following two weeks.

Figure 4.1 Percentage realized spread: Banned versus non-banned stocks

This figure shows the average percentage realized spread (PRS) aggregated across all banned and non-banned stocks, respectively. The time series for both groups of stocks covers four months before the introduction of the short-selling bans and four months during the short-selling bans. The starting and ending points of the short-selling bans in this figure are September 19, 2008, and January 16, 2009 – the dates of the introduction and end of the short-selling ban in the UK.

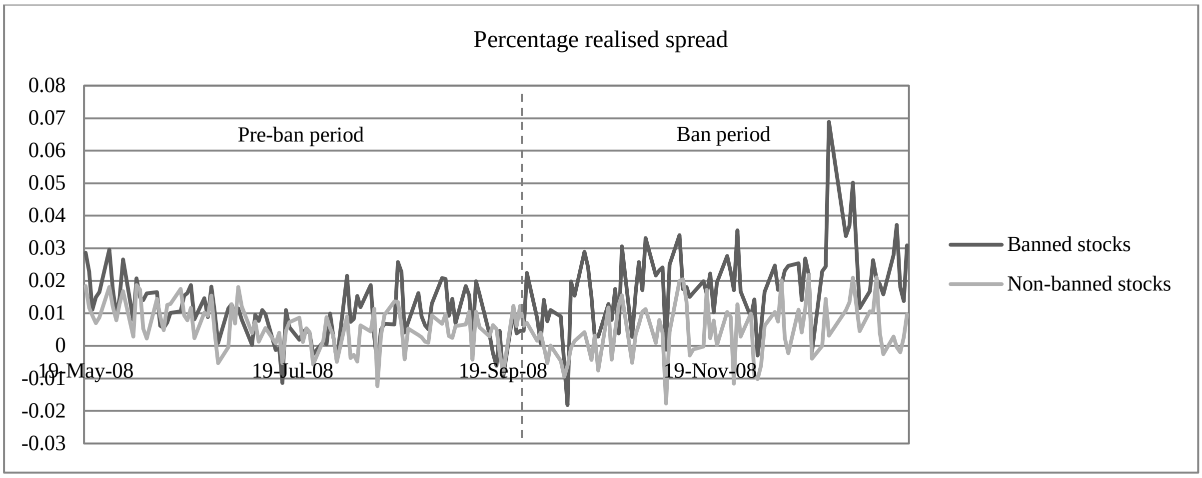


Table 4.2 reports the average median values of the percentage price impact (PPI) for banned (B) and non-banned (NB) stocks, respectively, as well as the median values of the relative PPI of the banned stocks (i.e., PPI (B) – PPI (NB)) across 10 European countries.

As can be seen in column (1) of Panel A of Table 4.2, the PPI (B) increases for six out of seven countries. The increase is statistically significant for all the countries except for Spain (see column (4)). For example, the PPI (B) for France increases from 0.0445 in the pre-ban period to 0.0528 in the ban period. Irish banned stocks experience the most dramatic increase from 0.0546 in the pre-ban period to 0.1285 in the ban period. Both results are statistically significant at the 1% level (see column (4)). In contrast, the PPI (B) for Portugal decreases from 0.0753 in the pre-ban period to 0.0639 in the ban period. This result is also statistically significant at the 1% level as reported in column (4).

Column (5) of Panel A of Table 4.2 shows that the PPI (NB) increases significantly for Belgium, Denmark, Germany, and Spain, decreases for Portugal, and remains nearly unchanged for France and Ireland.

The relative PPI of the banned stocks reported in column (6) of Panel A of Table 4.2 increases for Belgium, France, and Ireland but decreases for Denmark and Spain during the ban period. All the results are statistically significant at the 1% level. There is no significant change in the relative PPI of the banned stocks for Germany and Portugal.

Column (4) of Panel B of Table 4.2 shows that PPI (B) increases significantly for all three countries during the ban period at the 1% significance level. The PPI (NB) also significantly increases during the ban period for Switzerland and the UK. Both results are statistically significant at the 1% level, but there is no significant change in PPI (NB) for the Netherlands during the same period (see column (5)). In the post-ban period, the PPI for both banned and non-banned stocks significantly decreases as reported in columns (4) and (5).

Column (6) of Panel B of Table 4.2 shows that the relative PPI of the banned stocks increases significantly for all countries. All the results are statistically significant at the 1% level. In the post-ban period, the relative PPI of the banned stocks decreases for all three countries; however, the change is statistically significant only for the Dutch stocks. For example, the relative PPI of the Dutch banned stocks increases by 0.0286 in the ban period and subsequently decreases by 0.0233 in the post-ban period.

In summary, the results generally support Hypothesis 2 and the findings of Boehmer, Jones, and Zhang (2013), who find that the relative PPI of the US banned stocks increases during the ban period. Empirical evidence in this essay shows that the relative PPI of the banned stocks increases during the ban period for six out of 10 countries. However, in the post-ban period, only the results for the Netherlands support Hypothesis 2, where the relative PPI of the banned stocks declines significantly following the relaxation of the short-selling ban.

Table 4.2 Median percentage price impact (PPI) for banned (B) and non-banned (NB) stocks

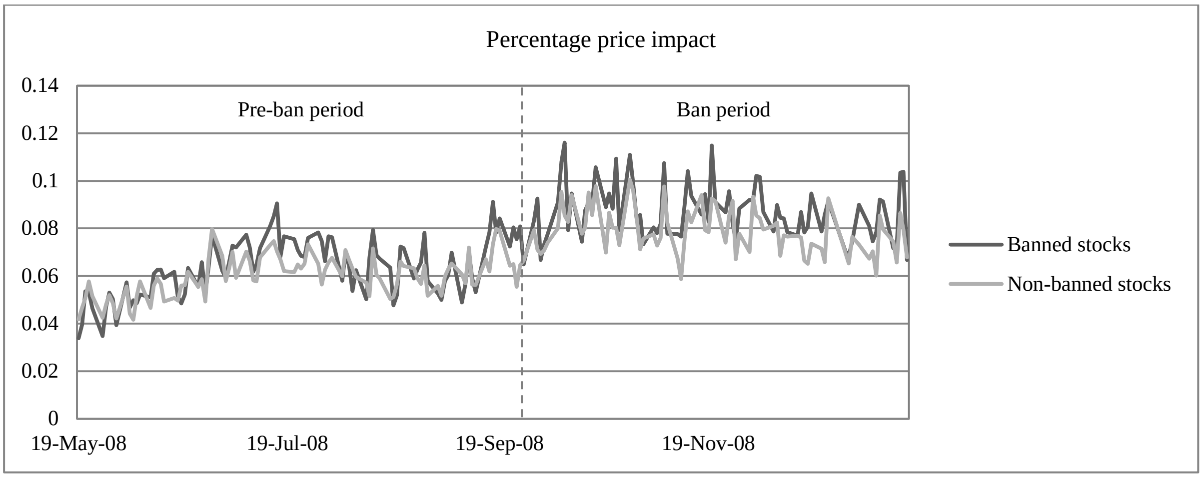
Average median values of the percentage price impact (PPI) for banned (B) and non-banned (NB) stocks, respectively, as well as their difference (i.e., PPI (B) – PPI (NB)) are reported for 10 European countries. The sample period spans from mid-January 2008 to mid-December 2009 and it is divided into three sub-periods: the pre-ban period, the ban period, and the post-ban period. The results for countries that have two sub-periods, the pre-ban period and the ban period (i.e., Belgium, Denmark, France, Germany, Ireland, Portugal, and Spain) are shown in Panel A. Only Switzerland, the Netherlands, and the UK have a post-ban period available. Their results are reported in Panel B. The sub-periods differ for each country and are specified in Table 3.1 of Chapter 3. Table 4.2 also shows the statistical significance of the test for equality of medians of the PPI for both samples and their difference. Variable Diff is the change in the median PPI between the pre-ban (P1) and ban (P2) periods (P2–P1), and ban and post-ban (P3) periods (P3–P2), respectively. Diff-in-Diff is the change in the difference between the PPI of the banned stocks and the PPI of the non-banned stocks from the pre-ban period to the ban period and from the ban period to the post-ban period, respectively. The Wilcoxon test is used and the p-values of the Z-statistic are given in the parentheses. * indicates statistical significance at the 10% level. ** indicates statistical significance at the 5% level. *** indicates statistical significance at the 1% level.

		(1)	(2)	(3)		(4)	(5)	(6)
Country	Sub-periods	PPI(B)	PPI(NB)	PPI(B) – PPI(NB)		Diff PPI(B)	Diff PPI(NB)	Diff-in-Diff PPI(B) – PPI(NB)
Panel A: Countries with two sub-periods								
Belgium	Pre-ban	0.0463	0.0506	-0.0048	P2-P1	0.042***	0.0084***	0.0282***
	Ban	0.0883	0.0590	0.0234	(p-value)	(0.0001)	(0.0001)	(0.0001)
Denmark	Pre-ban	0.0338	0.0359	-0.0017	P2-P1	0.0033**	0.0085***	-0.0064***
	Ban	0.0371	0.0444	-0.0081	(p-value)	(0.0152)	(0.0001)	(0.0001)
France	Pre-ban	0.0445	0.0353	0.0069	P2-P1	0.0083***	-0.0003	0.0092***
	Ban	0.0528	0.0350	0.0161	(p-value)	(0.0001)	(0.0586)	(0.0001)
Germany	Pre-ban	0.0284	0.0270	0.0007	P2-P1	0.0166***	0.0223***	-0.0044
	Ban	0.0450	0.0493	-0.0037	(p-value)	(0.0001)	(0.0001)	(0.0808)
Ireland	Pre-ban	0.0546	0.1164	-0.0621	P2-P1	0.0739***	-0.0083	0.0683***
	Ban	0.1285	0.1081	0.0062	(p-value)	(0.0001)	(0.4369)	(0.0001)
Portugal	Pre-ban	0.0753	0.0803	-0.0007	P2-P1	-0.0114***	-0.0079***	-0.0035
	Ban	0.0639	0.0724	-0.0042	(p-value)	(0.0001)	(0.0006)	(0.8879)
Spain	Pre-ban	0.0580	0.0584	-0.0006	P2-P1	0.0007	0.009***	-0.0026***
	Ban	0.0587	0.0674	-0.0032	(p-value)	(0.0998)	(0.0001)	(0.0076)
Panel B: Countries with three sub-periods								
Netherlands	Pre-ban	0.0490	0.0427	0.0067	P2-P1	0.0304***	-0.0039	0.0286***
	Ban	0.0794	0.0388	0.0353	(p-value)	(0.0001)	(0.5100)	(0.0001)
	Post-ban	0.0434	0.0303	0.0120	P3-P2	-0.036***	-0.0085***	-0.0233***
					(p-value)	(0.0001)	(0.0001)	(0.0001)
Switzerland	Pre-ban	0.0414	0.0384	0.0024	P2-P1	0.0285***	0.0139***	0.0094***
	Ban	0.0699	0.0523	0.0118	(p-value)	(0.0001)	(0.0001)	(0.0001)
	Post-ban	0.0300	0.0238	0.0078	P3-P2	-0.0399***	-0.0285***	-0.004
					(p-value)	(0.0001)	(0.0001)	(0.0606)
UK	Pre-ban	0.0499	0.0495	0.0034	P2-P1	0.0298***	0.0145***	0.0138***
	Ban	0.0797	0.0640	0.0172	(p-value)	(0.0001)	(0.0001)	(0.0001)
	Post-ban	0.0656	0.0523	0.0137	P3-P2	-0.0141***	-0.0117***	-0.0035
					(p-value)	(0.0001)	(0.0001)	(0.4766)

Figure 4.2 plots the time series of the average PPI for banned and non-banned stocks, respectively, covering a period of four months before the introduction of the short-selling bans and four months after the introduction of the short-selling bans. As can be seen in Figure 4.2, the PPI of the banned stocks was slightly higher relative to the PPI of the non-banned stocks, most notably during December 2008.

Figure 4.2 Percentage price impact: Banned versus non-banned stocks

This figure shows the average percentage price impact (PPI) aggregated across all banned and non-banned stocks, respectively. The time series for both groups of stocks covers four months before the introduction of the short-selling bans and four months during the short-selling bans. The starting and ending points of the short-selling bans in this figure are September 19, 2008, and January 16, 2009 – the dates of the introduction and end of the short-selling ban in the UK.



4.6.2 Summary of the univariate results and discussion

The results of the univariate analysis show that the relative non-informational component of the bid/ask spread of financial stocks subject to the 2008 short-selling bans increases during the ban period for seven out of 10 countries in the sample. Only one country, France, shows a decrease in the relative non-informational component of the bid/ask spread of the banned stocks. Two countries in the sample, the Netherlands and Switzerland, show no evidence of any significant change in the relative non-informational component of the bid/ask spread of the banned stocks. Overall, the univariate results support Hypothesis 1 in the ban period. In the post ban period, the univariate results support Hypothesis 1 but only for one out of three countries, in particular, the UK.

Furthermore, the results of the univariate analysis reveal that the relative informational component of the bid/ask spread of the banned stocks increase significantly in the ban period for six out of 10 countries. Two countries, Denmark and Spain, show a significant decrease in the relative informational component of the bid/ask spread of the banned stocks during the ban period. No significant change in the relative informational component of the bid/ask spread of the banned stocks is found for Germany and Portugal. Overall, the univariate results provide some supporting evidence for Hypothesis 2. In the post-ban period, only one out of three countries, the Netherlands, shows a significant decrease in the relative informational component of the bid/ask spread of the banned stocks.

To sum up, the univariate results suggest that the bid/ask spreads of the European financial stocks subject to the short-selling bans increase during the ban period due to an increase in both the non-informational and informational components of the bid/ask spread. However, there is no strong evidence showing that both components of the bid/ask spread decrease following the relaxation of the short-selling bans.

4.6.3 Multivariate analysis

4.6.3.1 Country level results

Table 4.3 reports the coefficient estimates from a multivariate regression analysis where the dependent variable is the difference between the percentage realized spread (PRS) of the banned stocks (B) and the PRS of the non-banned stocks (NB). The coefficient estimates of the ban dummy variable show that the relative PRS of the banned stocks increases significantly for Germany, Portugal, Spain, and the Netherlands, but it does not change significantly for the remaining countries. For example in Portugal, the relative percentage realized spread of the banned stocks rises by 364 basis points in the ban period. In the post-ban period, there is no statistically significant change for any of the three countries. These results provide weak support for Hypothesis 1. The relative PRS of banned stocks increases significantly during the ban period for only four out of seven countries. No support is found for Hypothesis 1 in the post-ban period.

Differences in the euro volume and range of prices are also found to significantly explain the relative PRS of the banned stocks. The coefficient estimate of the Diff_PRANGE variable, which proxies for volatility, is negative and statistically significant at the 1% level for all the countries. This result contradicts the predicted positive relationship between the non-informational component of the bid/ask spread and price volatility (see for example, Stoll (1978a, b)). A negative sign for the coefficient estimate of Diff_PRANGE implies that the more volatile the price of a stock, the smaller the realized spread, which proxies for the non-informational component of the bid/ask spread. However, R^2 values are low for most of the countries with the lowest being the UK (4%) and highest being Spain (37%), suggesting that there are other important factors that determine the realized spread that are not included in the analysis.

Table 4.3 Country-level multivariate results: Percentage realized spread (PRS)

The following regression model has the same form as shown in Boehmer, Jones, and Zhang (2013):

$$Diff\_PRS_{it} = \alpha_i + \beta D_{it}^{Ban} + \theta X_{it} + e_{it} \qquad (1)$$

where $Diff\_PRS_{it}$ is the difference between the PRS of the banned stocks (B) and the PRS of the non-banned stocks (NB). X_{it} is a vector of differences between the control variables of the matched pairs of stocks. These control variables are: market value – MV, trading volume in euros (or in their respective domestic currencies for the UK and Switzerland) – EUROVOL, proportional daily range of transaction prices – PRANGE, and daily value-weighted average price – VWAP. D_{it}^{Ban} is a dummy variable that takes the value of 1 when short-selling bans are in force and 0 otherwise.

For countries with a post-ban period available (i.e., the Netherlands, Switzerland, and the UK), equation (1) takes the following form:

$$Diff\_PRS_{it} = \alpha_i + \beta D_{it}^{Ban} + \gamma D_{it}^{Post\_ban} + \theta X_{it} + e_{it} \qquad (2)$$

where $D_{it}^{Post\text{-}ban}$ is equal to 1 when short-selling bans are lifted and 0 otherwise.

Fixed effects are used to control for any differences between the matched stocks in the pre-ban period. All the control variables are logarithmically transformed. P-values of the coefficient estimates are given in the parentheses below these estimates. * indicates statistical significance at the 10% level. ** indicates statistical significance at the 5% level. *** indicates statistical significance at the 1% level.

Note: Belgium, Denmark, and Ireland are excluded from the regression analysis at a country level due to the small sample size.

		(1)				(2)		
	Predicted sign	France	Germany	Portugal	Spain	Netherlands	Switzerland	UK
Intercept		0.0047 (0.3253)	0.0303*** (0.0064)	0.0151 (0.7554)	0.0100*** (0.0053)	0.0541*** (0.0004)	-0.0021** (0.0431)	0.0021 (0.5120)
Ban dummy variable	+	-0.0016 (0.7195)	0.0133** (0.0318)	0.0364*** (0.0077)	0.0075*** (0.0032)	0.0049*** (0.0060)	-0.0010 (0.7265)	0.0028 (0.4418)
Post-ban dummy variable	-					0.0014 (0.6277)	0.0025 (0.1286)	-0.0022 (0.2875)
Diff_MV	-	-0.0166** (0.0234)	0.0620 (0.1248)	0.0131 (0.7043)	-0.0165 (0.4749)	-0.0075 (0.5808)	0.0113*** (0.0001)	0.0055 (0.1442)
Diff_EUROVOL	-	0.0103*** (0.0082)	-0.0048** (0.0429)	-0.0067 (0.3072)	-0.0018 (0.7146)	0.0101*** (0.0001)	0.0039** (0.0160)	-0.0030 (0.2166)
Diff_PRANGE	+	-0.0046*** (0.0001)	-0.0025*** (0.0001)	-0.0073*** (0.0001)	-0.0046*** (0.0001)	-0.0022*** (0.0001)	-0.0014*** (0.0001)	-0.0017*** (0.0001)
Diff_VWAP	-	-0.0037 (0.5518)	-0.0590 (0.0892)	-0.0262 (0.5664)	-0.0045 (0.5143)	-0.0053 (0.6743)	-0.0171*** (0.0001)	-0.0090 (0.1019)
R^2		0.10	0.05	0.26	0.37	0.06	0.07	0.04

Table 4.4 reports the estimation results for the PPI regression model. The coefficient estimates of the ban dummy variable show that the relative PPI of the banned stocks increases in the ban period in all countries except for Germany and Portugal. All the results are statistically significant at the 5% level with the exception of the Netherlands where the result is statistically significant at the 1% level. For example, in France, the relative PPI of the banned stocks increases by 105 basis points during the ban period. These results generally support Hypothesis 2 and the findings of prior studies, which also find that the relative price

impact of the banned stocks increases following the introduction of the short-selling ban in the US (see Boehmer, Jones, and Zhang (2013)).

Again there is no supporting evidence for Hypothesis 2 in the post-ban period. The coefficient estimate of the post-ban dummy variable is positive and significant at the 1% level for the UK and 10% level for Switzerland. The post-ban dummy variable decreases marginally for the Netherlands. The results for the UK and Switzerland show that the relative PPI of the banned stocks increases rather than decreases as would be expected in the post-ban period.

Differences in market value and range in prices between the two samples are also found to significantly explain the dependent variable. The coefficient estimate of Diff_MV is negative and statistically significant for five out of seven countries while the coefficient estimate of Diff_PRANGE is positive and statistically significant for all seven countries. The sign of the coefficient estimates of both control variables reported in Table 4.4 is in accordance with the predicted sign. R^2 values vary between 7% for Germany and 24% for Switzerland, implying that some other explanatory factors of PPI are omitted from the regression analysis.

Table 4.4 Country-level multivariate results: Percentage price impact (PPI)

The following regression model has the same form as shown in Boehmer, Jones, and Zhang (2013):

$$Diff\_PPI_{it} = \alpha_i + \beta D_{it}^{Ban} + \theta X_{it} + e_{it} \qquad (1)$$

where $Diff\_PPI_{it}$ is the difference between the PPI of the banned stocks (B) and the PPI of the non-banned stocks (NB). X_{it} is a vector of differences between the control variables of the matched pairs of stocks. These control variables are: market value – MV, trading volume in euros (or in their respective domestic currencies for the UK and Switzerland) – EUROVOL, proportional daily range of transaction prices – PRANGE, and daily value-weighted average price – VWAP. D_{it}^{Ban} is a dummy variable that takes the value of 1 when short-selling bans are in force and 0 otherwise.

For countries with a post-ban period available (i.e., the Netherlands, Switzerland, and the UK), equation (1) takes the following form:

$$Diff\_PPI_{it} = \alpha_i + \beta D_{it}^{Ban} + \gamma D_{it}^{Post\_ban} + \theta X_{it} + e_{it} \qquad (2)$$

where $D_{it}^{Post\text{-}ban}$ is equal to 1 when short-selling bans are lifted and 0 otherwise.

Fixed effects are used to control for any differences between the matched stocks in the pre-ban period. All the control variables are logarithmically transformed. P-values of the coefficient estimates are given in the parentheses below these estimates. * indicates statistical significance at the 10% level. ** indicates statistical significance at the 5% level. *** indicates statistical significance at the 1% level.

Note: Belgium, Denmark, and Ireland are excluded from the regression analysis at a country level due to the small sample size.

		(1)				(2)		
	Predicted sign	France	Germany	Portugal	Spain	Netherlands	Switzerland	UK
Intercept		-0.0002 (0.9671)	-0.0452*** (0.0001)	0.0136 (0.5446)	0.0115*** (0.0043)	-0.0720*** (0.0001)	-0.0066*** (0.0001)	-0.0005 (0.8830)
Ban dummy variable	+	0.0105** (0.0335)	-0.0043 (0.5107)	0.0117 (0.3175)	0.0057** (0.0197)	0.0043*** (0.0093)	0.0065** (0.0427)	0.0098** (0.0379)
Post-ban dummy variable	-					-0.0047* (0.0729)	0.0033* (0.0803)	0.0102*** (0.0001)
Diff_MV	-	-0.0130*** (0.0017)	-0.0967** (0.0117)	-0.0427*** (0.0001)	-0.0384 (0.0862)	-0.0151 (0.1016)	-0.0082*** (0.0059)	-0.0126** (0.0349)
Diff_EUROVOL	-	-0.0068 (0.0543)	0.0001 (0.9781)	0.0037 (0.5977)	0.0048 (0.4250)	-0.0108*** (0.0007)	0.0001 (0.9676)	0.0040 (0.0980)
Diff_PRANGE	+	0.0051*** (0.0001)	0.0031*** (0.0001)	0.0100*** (0.0001)	0.0012*** (0.0001)	0.0031*** (0.0001)	0.0021*** (0.0001)	0.0029*** (0.0001)
Diff_VWAP	-	0.0130* (0.0848)	0.0863*** (0.0052)	-0.0005 (0.9464)	0.0039 (0.5131)	0.0032 (0.6703)	-0.0032 (0.2800)	-0.0032 (0.6240)
R^2		0.21	0.07	0.14	0.24	0.20	0.27	0.10

Overall, the country level results show that the relative PRS of the banned stocks increases significantly during the ban period for Germany, Portugal, Spain, and the Netherlands but does not change significantly for France, Switzerland, and the UK. These results provide weak support for Hypothesis 1. Stronger evidence is found for Hypothesis 2. The relative PPI of the banned stocks increases significantly for five out of seven countries in the ban period. The two countries where no significant change in the relative PPI of the banned stocks is found are Germany and Portugal.

4.6.3.2 Aggregate level results

Aggregating the data across all countries generates the results reported in Table 4.5. As only three out of 10 countries have a post-ban period available, the analysis in this section focuses on the impact of short-selling bans on the components of the bid/ask spread from the pre-ban period to the ban period. The coefficient estimate of the ban dummy variable shows that the relative PRS of the banned stocks rises by 94 basis points in the ban period. This estimate is statistically significant at the 1% level. On the other hand, the relative PPI of the banned stocks rises by 34 basis points during the ban period. This result is only statistically significant at the 10% level.

Differences in the proportional range of transaction prices also play an important role in explaining both dependent variables. The sign of the coefficient estimate of Diff_PRANGE for the relative PPI regression model is positive as predicted (Copeland and Galai (1983); Easley and O'Hara (1987)). In contrast, the sign of the coefficient estimate of the Diff_PRANGE variable for the relative PRS regression model is significant but negative rather than positive as would be expected. Although the models' R^2 are not very high (26% for the relative PRS model and 18% for the relative PPI model), a non-negligible proportion of the dependent variables is still explained by these models.

In summary, the aggregate-level results show that the introduction of the short-selling regulations is associated with significantly greater relative PRS of the banned stocks. These results strongly support Hypothesis 1. Furthermore, there is evidence showing that the relative PPI of the banned stocks also increases during the ban period. This result provides weaker support for Hypothesis 2 as it is statistically significant only at the 10% level. Overall, the results reported in this section are qualitatively similar to the results reported by Boehmer, Jones, and Zhang (2013) for the US market.

Table 4.5 Aggregate-level multivariate results

The following regression model has the same form as shown in Boehmer, Jones, and Zhang (2013):

$$Diff\_Y_{it} = \alpha_i + \beta D_{it}^{Ban} + \theta X_{it} + e_{it}$$

where $Diff\_Y_{it}$ is the difference between the percentage realized spread of the banned stocks (B) and non-banned stocks (NB) and the difference between the percentage price impact of the banned stocks and non-banned stocks, respectively. X_{it} is a vector of differences between the control variables of the matched pairs of stocks. These control variables are: market value – MV, trading volume in euros (or in their respective domestic currencies for the UK, Switzerland, and Denmark) – EUROVOL, proportional daily range of transaction prices – PRANGE, and daily volume-weighted average price – VWAP. D_{it}^{Ban} is a dummy variable that takes the value of 1 when short-selling bans are in force and 0 otherwise. Fixed effects are used to control for any differences between the matched stocks in the pre-ban period. All the control variables are logarithmically transformed. P-values of the coefficient estimates are given in the parentheses below these estimates. * indicates statistical significance at the 10% level. ** indicates statistical significance at the 5% level. *** indicates statistical significance at the 1% level.

	Predicted sign	Diff_PRS	Diff_PPI
Intercept		0.0084*	-0.0034
		(0.0751)	(0.2178)
Ban dummy variable	+	0.0094***	0.0034*
		(0.0043)	(0.0551)
Diff_MV	-	-0.0149	-0.0191***
		(0.2196)	(0.0046)
Diff_EUROVOL	-	-0.0008	0.0009
		(0.7173)	(0.6667)
Diff_PRANGE	+	-0.0029***	0.0043***
		(0.0001)	(0.0001)
Diff_VWAP	-	-0.0087	-0.0008
		(0.2398)	(0.8865)
R^2		0.26	0.18

4.6.4 Summary of the multivariate results and discussion

In summary, the country-level multivariate results show that the relative non-informational component of the bid/ask spread of the banned stocks increases significantly for four out of seven countries following the introduction of the short-selling regulations. These countries are Germany, Portugal, Spain, and the Netherlands. The introduction of the short-selling bans is also found to be associated with a greater relative informational component of the bid/ask spread of the banned stocks in five out of seven countries. The two countries where no significant change in the relative PPI of the banned stocks is documented are Germany and Portugal. Consequently, in these two countries, the relative bid/ask spreads of the banned stocks increase during the ban period due to an increase in the relative non-informational component of the bid/ask spread of the banned stocks. In contrast, in France, Switzerland, and the UK, the relative bid/ask spreads of the banned stocks increase during the ban period due to an increase in the relative informational component of the bid/ask spread of the banned stocks. Lastly, in Spain and the Netherlands, the relative bid/ask spreads of the banned stocks increase due to an increase in both relative components of the bid/ask spread of the banned stocks. Overall, these results support Hypotheses 1 and 2.

Aggregating across countries, the multivariate results show that both the relative non-informational and informational components of the bid/ask spread of the banned stocks increase significantly following the introduction of the short-selling regulations supporting Hypothesis 1 and 2. However, the increase in the relative PRS of the banned stocks is greater than the increase in the relative PPI of the banned stocks. These findings suggest that the impact of the short-selling bans is greater on the relative non-informational component of the bid/ask spread of the banned stocks than on the relative informational component of the bid/ask spread of the banned stocks. These results are also in line with the findings of Boehmer, Jones, and Zhang (2013). These authors find that both components of the bid/ask spread increase after the introduction of the short-selling ban in the US. Moreover, their results also show that the relative non-informational component of the bid/ask spread of the US banned stocks increases more than the relative informational component of the bid/ask spread of the US banned stocks.

4.7 Conclusions

This essay investigates the impact of the 2008 short-selling regulations on the non-informational and informational components of the bid/ask spread of the financial stocks subject to these regulations in 10 European countries. The univariate results show that the relative non-informational component of the bid/ask spread of the financial stocks subject to the short-selling regulations increases during the ban period for seven out of 10 countries. The relative informational component of the bid/ask spread of the financial stocks subject to the short-selling bans also increases in the ban period but only for six out of 10 countries.

Furthermore, the aggregate-level multivariate results show that both the relative non-informational and the informational components of the bid/ask spread of the banned stocks increase significantly following the introduction of the short-selling regulations. Although both relative components of the bid/ask spread of the banned stocks increase during the ban period, the results indicate that the short-selling bans have a stronger effect on the non-informational component than on the informational component of the bid/ask spread of the banned stocks.

The country-level multivariate results reveal that the bid/ask spreads of the banned stocks increase through different channels. More precisely, the bid/ask spreads of the banned stocks in Germany and Portugal increase through the non-informational component channel of the bid/ask spread. In contrast, the bid/ask spreads of the banned stocks in France, Switzerland, and the UK increase through the informational component channel of the bid/ask spread. Lastly, in Spain and the Netherlands, the bid/ask spreads of the banned stocks increase through both the non-informational and the informational component channels of the bid/ask spread.

Overall, both univariate and multivariate results provide supporting evidence for Hypotheses 1 and 2. These results also support the findings of Boehmer, Jones, and Zhang (2013), who report a significant increase in both the relative non-informational and the informational components of the bid/ask spread of the banned stocks after the introduction of the short-selling ban in the US.

APPENDIX 4A

Table 4A.1 Mean percentage realized spread (PRS) for banned (B) and non-banned (NB) stocks

Average mean values of the percentage realized spread (PRS) for banned (B) and non-banned (NB) stocks, respectively, as well as their difference (i.e., PRS (B) – PRS (NB)) are reported for 10 European countries. The sample period spans from mid-January 2008 to mid-December 2009 and it is divided into three sub-periods: the pre-ban period, the ban period, and the post-ban period. The results for countries that have two sub-periods, the pre-ban period and the ban period (i.e., Belgium, Denmark, France, Germany, Ireland, Portugal, and Spain) are shown in Panel A. Only Switzerland, the Netherlands, and the UK have a post-ban period available. Their results are illustrated in Panel B. The sub-periods differ for each country and are specified in Table 3.1 of Chapter 3. Table 4A.1 also shows the statistical significance of the test for equality of means of the PRS for both samples and their difference. Variable Diff is the change in the median PRS between the pre-ban (P1) and the ban (P2) periods (P2–P1), and the ban and the post-ban (P3) periods (P3–P2), respectively. Diff-in-Diff is the change in the difference between the PRS of the banned stocks and the PRS of the non-banned stocks from the pre-ban period to the ban period and from the ban period to the post-ban period, respectively. The t-test is used and the p-values of the t-statistic are given in the parentheses. * indicates statistical significance at the 10% level. ** indicates statistical significance at the 5% level. *** indicates statistical significance at the 1% level.

Country	Sub-periods	PRS(B)	PRS(NB)	PRS(B) – PRS(NB)		Diff PRS(B)	Diff PRS(NB)	Diff-in-Diff PRS(B) – PRS(NB)
Panel A: Countries with two sub-periods								
Belgium	Pre-ban	0.0037	-0.0083	0.0120	P2-P1	-0.0014	-0.0049***	0.0035*
	Ban	0.0023	-0.0132	0.0155	(p-value)	(0.4157)	(0.0001)	(0.0720)
Denmark	Pre-ban	0.0731	0.0725	0.0006	P2-P1	0.0704***	0.0486***	0.0218***
	Ban	0.1436	0.1211	0.0224	(p-value)	(0.0001)	(0.0001)	(0.0001)
France	Pre-ban	-0.0114	-0.0109	-0.0004	P2-P1	-0.0113***	0.0043**	-0.0069**
	Ban	-0.0227	-0.0153	-0.0073	(p-value)	(0.0001)	(0.0343)	(0.0431)
Germany	Pre-ban	0.0149	0.0133	0.0016	P2-P1	0.0047*	-0.0088***	0.0136***
	Ban	0.0197	0.0045	0.0152	(p-value)	(0.0986)	(0.0051)	(0.0012)
Ireland	Pre-ban	0.0658	0.1265	-0.0607	P2-P1	0.2591***	0.2455***	0.0136
	Ban	0.3249	0.3720	-0.0471	(p-value)	(0.0001)	(0.0001)	(0.5995)
Portugal	Pre-ban	0.0539	0.0486	0.0053	P2-P1	0.0080	-0.0248***	0.0329***
	Ban	0.0620	0.0237	0.0382	(p-value)	(0.1367)	(0.0001)	(0.0001)
Spain	Pre-ban	0.0135	-0.0080	0.0215	P2-P1	0.0029	0.0013	0.0015
	Ban	0.0164	-0.0066	0.0230	(p-value)	(0.1889)	(0.6311)	(0.6712)
Panel B: Countries with three sub-periods								
Netherlands	Pre-ban	-0.0012	-0.0166	0.0153	P2-P1	-0.0089	-0.0035	-0.0125
	Ban	-0.0101	-0.0130	0.0028	(p-value)	(0.2319)	(0.6813)	(0.2211)
	Post-ban	-0.0001	-0.0084	0.0083	P3-P2	0.0101	0.0046	0.0054
					(p-value)	(0.2544)	(0.6528)	(0.6652)
Switzerland	Pre-ban	0.0097	0.0079	0.0017	P2-P1	-0.0110***	-0.0080***	-0.0030*
	Ban	-0.0013	-0.0000	-0.0012	(p-value)	(0.0001)	(0.0001)	(0.0568)
	Post-ban	0.0232	0.0194	0.0037	P3-P2	0.0245***	0.0196***	0.0049***
					(p-value)	(0.0001)	(0.0001)	(0.0034)
UK	Pre-ban	-0.0000	0.0011	-0.0012	P2-P1	-0.0094	-0.0126***	0.0031
	Ban	-0.0094	-0.0114	0.0019	(p-value)	(0.1275)	(0.0001)	(0.6459)
	Post-ban	-0.0162	-0.0097	-0.0065	P3-P2	-0.0067	0.0017	-0.0084
					(p-value)	(0.2658)	(0.5868)	(0.2107)

Table 4A.2 Mean percentage price impact (PPI) for banned (B) and non-banned (NB) stocks

Average mean values of the percentage price impact (PPI) for banned (B) and non-banned (NB) stocks, respectively, as well as their difference (i.e., PRS (B) – PRS (NB)) are reported for 10 European countries. The sample period spans from mid-January 2008 to mid-December 2009 and it is divided into three sub-periods: the pre-ban period, the ban period, and the post-ban period. The results for countries that have two sub-periods, the pre-ban period and the ban period (i.e., Belgium, Denmark, France, Germany, Ireland, Portugal, and Spain) are shown in Panel A. Only Switzerland, the Netherlands, and the UK have a post-ban period available. Their results are illustrated in Panel B. The sub-periods differ for each country and are specified in Table 3.1 of Chapter 3. Table 4A.2 also shows the statistical significance of the test for equality of means of the PPI for both samples and their difference. Variable Diff is the change in the median PPI between the pre-ban (P1) and the ban (P2) periods (P2–P1), and the ban and the post-ban (P3) periods (P3–P2), respectively. Diff-in-Diff is the change in the difference between the PPI of the banned stocks and the PPI of the non-banned stocks from the pre-ban period to the ban period and from the ban period to the post-ban period, respectively. The t-test is used and the p-values of the t-statistic are given in the parentheses. * indicates statistical significance at the 10% level. ** indicates statistical significance at the 5% level. *** indicates statistical significance at the 1% level.

Country	Sub-periods	PPI(B)	PPI(NB)	PPI(B) – PPI(NB)		Diff PPI(B)	Diff PPI(NB)	Diff-in-Diff PPI(B) – PPI(NB)
Panel A: Countries with two sub-periods								
Belgium	Pre-ban	0.0518	0.0563	-0.0045	P2-P1	0.0536***	0.0169***	0.0367***
	Ban	0.1055	0.0733	0.0321	(p-value)	(0.0001)	(0.0001)	(0.0001)
Denmark	Pre-ban	0.0358	0.0400	-0.0042	P2-P1	0.0072***	0.0280***	-0.0208***
	Ban	0.0430	0.0680	-0.0250	(p-value)	(0.0040)	(0.0001)	(0.0001)
France	Pre-ban	0.0625	0.0538	0.0086	P2-P1	0.0223***	0.0041*	0.0182***
	Ban	0.0847	0.0579	0.0268	(p-value)	(0.0001)	(0.0715)	(0.0001)
Germany	Pre-ban	0.0343	0.0350	-0.0007	P2-P1	0.0199***	0.0251***	-0.0052
	Ban	0.0542	0.0601	-0.0059	(p-value)	(0.0001)	(0.0001)	(0.2274)
Ireland	Pre-ban	0.0630	0.1352	-0.0722	P2-P1	0.0798***	0.0518***	0.0280
	Ban	0.1427	0.1870	-0.0442	(p-value)	(0.0001)	(0.0041)	(0.1290)
Portugal	Pre-ban	0.0973	0.0958	0.0015	P2-P1	0.0042	-0.0071**	0.0114**
	Ban	0.1016	0.0886	0.0129	(p-value)	(0.4301)	(0.0236)	(0.0466)
Spain	Pre-ban	0.0696	0.0736	-0.0039	P2-P1	0.0079***	0.0086***	-0.0006
	Ban	0.0776	0.0822	-0.0046	(p-value)	(0.0003)	(0.0055)	(0.8567)
Panel B: Countries with three sub-periods								
Netherlands	Pre-ban	0.0708	0.0773	-0.0065	P2-P1	0.0338***	-0.0039	0.0378***
					(p-value)	(0.0001)	(0.6641)	(0.0002)
	Ban	0.1046	0.0734	0.0311	P3-P2	-0.0366***	-0.0079	-0.0286**
	Post-ban	0.0680	0.0654	0.0025	(p-value)	(0.0001)	(0.4562)	(0.0257)
Switzerland	Pre-ban	0.0416	0.0412	0.0003	P2-P1	0.0324***	0.0194***	0.0130***
					(p-value)	(0.0001)	(0.0001)	(0.0001)
	Ban	0.0739	0.0606	0.0133	P3-P2	-0.0382***	-0.0344***	-0.0037*
	Post-ban	0.0358	0.0262	0.0095	(p-value)	(0.0001)	(0.0001)	(0.0562)
UK	Pre-ban	0.0674	0.0651	0.0023	P2-P1	0.0391***	0.0294***	0.0097
					(p-value)	(0.0001)	(0.0001)	(0.1627)
	Ban	0.1065	0.0945	0.0120	P3-P2	-0.0118*	-0.0161***	0.0042
	Post-ban	0.0947	0.0784	0.0163	(p-value)	(0.0575)	(0.0001)	(0.5308)

CHAPTER 5

Essay 3: Short-selling regulations and the speed of price adjustment to new information

5.1 Introduction

One consequence of short-selling constraints is the slower adjustment of stock prices to new information and in particular to bad news (Diamond and Verrecchia (1987)). This essay examines the impact of the short-selling regulations that were introduced in the autumn of 2008 on the speed of price adjustment to new information on financial stocks subject to these regulations. Prior empirical evidence shows that short-selling constraints are associated with slower incorporation of new information into stocks prices (see for example, Bris, Goetzmann, and Zhu (2007), Chen and Rhee (2010), and Saffi and Sigurdson (2011) among others). This essay contributes to the literature by using a new dataset associated with the introduction of the 2008 short-selling rules to test the Diamond and Verrecchia (1987) model's prediction. Knowing the exact dates of the introduction and subsequently the relaxation of these short-selling rules provides a unique opportunity to investigate the effects of the changes in short-selling regimes on the information efficiency of the prices of stocks covered under these regimes.

To estimate the speed of price adjustment to new information, this study uses the research method developed by Hasbrouck (1991). Chen and Rhee (2010) also use Hasbrouck's (1991) model, however, their work differs from this study in the following ways. Firstly, the authors examine a different market, Hong Kong. In contrast, this study investigates the impact of the short-selling rules on the speed of stock price adjustment to new information on major financial institutions across 10 European countries.[143] Secondly, Chen and Rhee (2010) focus on the impact of relaxing short-selling regulations while this study focuses mainly on the impact of tightening short-selling regulations on the speed of price adjustment to new information on stocks subject to these regulations. Thirdly, Chen and Rhee (2010) examine

[143] These countries are Belgium, Denmark, France, Germany, Ireland, the Netherlands, Portugal, Spain, Switzerland, and the UK. Appendix 5A also provides empirical results for Norway and Italy–two special cases.

only the stocks that were affected by the changes in the short-selling regulations while this study uses a control sample of stocks that were not affected by the short-selling regulations to control for general market movements. Constructing a control sample is essential in this study as the 2008 short-selling regulations were introduced during a severe financial crisis.

This essay is organized as follows. Section 5.2 reviews the related literature. Section 5.3 develops the hypotheses. Section 5.4 presents the sample and the data. Section 5.5 describes the methodology. Section 5.6 reports and discusses the empirical results. Section 5.7 summarizes the main findings and concludes.

5.2 Literature Review

5.2.1 Short-selling and information processing: Theoretical framework

Diamond and Verrecchia (1987), hereafter DV, provide a theoretical framework for analysing the effects of short-selling constraints on the speed of price adjustment to new information. Their model assumes that traders are rational, which implies that stock prices are not biased upwards when short-selling is constrained. Nonetheless, the arrival of new information is impounded into prices more slowly in the presence of short-selling constraints. Two effects of short-selling constraints are examined: the short-prohibition effect and the short-restriction effect. When short-selling is prohibited, both informed and uninformed traders are not allowed to sell short. On the other hand, when short-selling is restricted, the additional costs of short-selling drive only the uninformed traders out of the market.

The DV model further assumes a competitive and risk-neutral market maker constantly setting bid/ask prices. She faces no inventory costs or constraints and earns zero expected profits.[144] She observes all trades as they occur but does not observe private information. Informed risk-neutral traders observe identical private information while uninformed risk-neutral traders observe only public information. Moreover, trades are generated by either information or liquidity shocks and traders can choose from the following set of actions: buy, sell, sell short, and no trade. The actions traders can observe, however, are: buy, sell or sell short, and no trade. In other words, the traders cannot distinguish whether the sell action they observe is a long sale or a short sale. The decision to sell short depends on the costs associated with the transaction: no cost, restrictions on proceeds from short-selling, or short-selling prohibitions, in which case short-selling is impossible.

In the DV model, informed traders who do not own the share and face no short-selling costs, will sell short the risky asset when they possess negative information. On the other hand, uninformed traders who do not own the share and face no short-selling costs will sell short the risky asset when they desire to consume. However, the situation changes when traders are not allowed to reinvest or consume the proceeds from short-selling immediately. In this case, only the informed traders sell short since they do not need the proceeds for immediate consumption as the uninformed traders do. This implies that proceeds-restrictions on short-selling increase the proportion of informed traders that remain in the pool of "sell-or-sell short" transactions as these restrictions drive the uninformed traders out of the market. Finally, when there are short-selling prohibitions in place, both informed and uninformed traders do not sell short a share unless they own it. Consequently, short-selling prohibitions leave the fraction of informed traders in the pool of "sell-or-sell short" unchanged. Moreover, these prohibitions increase the periods of no-trade for both informed and uninformed traders. This reduces the information that would have been revealed through trades if informed

[144] The losses that a market maker incurs from trading with informed investors are exactly compensated by the gains made from trading with uninformed investors.

traders could sell short. The reduction in the information content of trades is greater when news is bad because that is when informed short-selling is expected to be more frequent.

Overall, an informed trader who possesses negative information but does not own the share sells short only when she faces no costs or proceeds-restrictions. On the other hand, an uninformed trader who needs to consume immediately and does not own the share sells short only when she faces no costs on short-selling. Consequently, the DV model predicts that prices reflect all the available information in the market with a delay when short-selling is prohibited because market participants with negative information cannot reveal their beliefs through trading. This means that, in the presence of short-selling constraints, more time might be required for all the information to be impounded into stock prices.

5.2.2 Short-selling constraints and the speed of price adjustment to new information: Empirical evidence

This sub-section reviews empirical studies that test the DV model's prediction that short-selling constraints will reduce the speed of price adjustment to new information in the spot market. Overall, the results of these studies support the DV model's prediction.

Biais, Bisière, and Décamps (1999) study the effects of short-selling constraints on the frequency of sell orders and the speed of stock price adjustment to bad news during 1996 on the Paris Bourse. In the Paris Bourse the largest and most liquid stocks are traded on the monthly settlement market, which operates as a futures market. Stocks traded on that market are usually not simultaneously traded on the spot market. The remaining stocks are traded on the spot market. The authors hypothesize that (i) market sell orders are less frequent for stocks traded on the spot market than for stocks traded on the monthly settlement market and (ii) market prices adjust faster to positive information than to negative information.[145]

To test their first hypothesis, Biais, Bisière, and Décamps (1999) compare the buy side and sell side of the market for each stock after controlling for market capitalization and volatility. They study the cross-sectional distribution of market sell order frequency which is normalized using the frequency of all market orders for each stock. Consistent with the first hypothesis, the authors find that the frequency of market sell orders is lower in the spot market than in the monthly settlement market. The authors conclude that the more stringent short-selling constraints in the sport market could be the reason for the observed difference in the frequency of sell orders between the two markets.

To test their second hypothesis, the authors use lagged positive and negative returns on the monthly settlement market as proxies for good and bad news, respectively. This is done because stocks traded on this market are the most liquid ones and short-selling constraints are less binding. Spot returns are regressed on lagged spot returns to control for autocorrelation, positive monthly settlement market returns, and negative monthly settlement market returns to test whether there is a significant difference in coefficients based on the sign of monthly settlement market returns. By so doing, the authors attempt to estimate the speed at which the spot prices adjust to the arrival of market information using the delay with which spot market returns follow the monthly settlement market returns. Positive monthly settlement market returns have a significantly positive coefficient at the first lag, while negative monthly settlement returns have a significantly positive coefficient only at the second lag. This result suggests that more time is needed for negative news rather than for positive news to be incorporated in the spot market after it is reflected in the monthly settlement market, which supports the second hypothesis.

[145] Short-selling constraints are less severe in the monthly settlement market.

Overall, both results are consistent with the predictions of the DV model. Short-selling constraints are a possible explanation for the reduced frequency of market sell orders in the spot market as well as for the slower speed with which negative information is impounded into prices in the same market.

Bris, Goetzmann, and Zhu (2007) examine the impact of short-selling restrictions on market efficiency across 46 countries over the period 1990–2001. The authors identify four groups of stocks: (i) domestic stocks that are traded in countries where short-selling is not allowed or practiced, (ii) dual-listed stocks listed both in the domestic market, where short-selling is not allowed or practiced, and in a foreign market, where short-selling is allowed and practiced, (iii) domestic stocks in countries where short-selling is allowed and practised, and (iv) dual-listed stocks listed in the domestic market and in a foreign market where short-selling is allowed and practiced in both markets.

Firstly, the authors estimate two market model regressions of weekly stock returns, one on positive value-weighted market returns and one on negative value-weighted market returns, for each firm in every year.[146] A downside - minus - upside R^2 obtained from estimating the two models measures the relative co-movement of individual stock returns with the market, conditional on the sign of the market return. The authors argue that, when short-selling restrictions are in place, stock returns respond immediately to positive market information but with a lag to negative market information. They hypothesize that the difference between downside R^2 and upside R^2 will drop after the short-selling restrictions are relaxed, which implies a more symmetric incorporation of information under falling and rising markets, respectively. The results show that when short-selling restrictions are lifted the difference

[146] A positive (negative) value-weighted market return is a weighted average of all positive (negative) stock returns. Weights are determined by the market capitalization of the stock at the end of the previous trading period (Bris, Goetzmann, and Zhu (2007)).

between downside - minus - upside R^2 drops, however, the drop is only statistically significant at the 10% level. There is no significant change between downside - minus - upside R^2 for the dual-listed stocks in countries where short-selling was initially restricted and then allowed, which implies that information-based short-selling conveys information through trading in the foreign market. Also, no significant difference in efficiency is found between domestic stocks and dual-listed stocks in countries where short-selling is allowed or practised. These results suggest that efficiency improves for domestic stocks in countries where short-selling was initially restricted and subsequently allowed. However, there is no change in efficiency for domestic stocks in countries where short-selling is allowed and practised during the whole sample period.

Secondly, the authors compute the difference in cross-autocorrelation between individual stock returns and one-week lagged signed (negative versus positive) market returns to measure the speed of price adjustment of individual stocks to market movements. The greater the cross-autocorrelation of the individual stock returns with the market movements, the slower the speed with which new information is incorporated into prices. Bris, Goetzmann, and Zhu (2007) find that the downside - minus - upside cross-autocorrelation for domestic stocks is lower in counties where short-selling is allowed compared to that for domestic stocks where short-selling is not allowed. This result is statistically significant at the 10% level. In countries where short-selling is not allowed, downside cross-autocorrelation is lower for dual-listed stocks compared to the downside cross-autocorrelation for domestic stocks. This result is statistically significant at the 5% level.

Overall, the authors argue that short-selling constraints reduce the speed of price adjustment to bad news. Thus, markets allowing short-selling appear to be more efficient as bad news is incorporated faster into prices.

Saffi and Sigurdson (2011) investigate the impact of short-selling constraints on the speed and accuracy with which prices reflect information for 12,621 firms across 26 countries over the period 2005-2008. Short-selling constraints are proxied by (i) the amount of shares available for lending to short-sellers and (ii) the loan fees charged to short sellers. These two measures are used as explanatory variables in panel regressions to explain cross-sectional differences in information efficiency.

The authors use two measures of efficiency as proposed by Hou and Moskowitz (2005). Firstly, they estimate a regression of weekly stock returns on contemporaneous returns on a world index and a value-weighted domestic index as well as on four weeks lags of the domestic index. Then they re-estimate this equation, imposing a constraint that the coefficients of lagged domestic returns are zero, and compare the differences in R^2 from the two regressions. The authors refer to this difference as delay measure D1. Secondly, the authors construct another delay measure – D2 – which is calculated by dividing the sum of the four lagged market-return betas over the sum of contemporaneous beta and lagged market-return betas. The authors hypothesize that the greater the explanatory power of lagged market returns (i.e., the magnitude of the lagged market-return coefficients relative to the magnitude of all market-return coefficients), the longer it takes for the information to be impounded into prices.

Saffi and Sigurdson (2011) find that both delay measures have an inverse relationship with lending supply, which implies that the fewer the shares available for lending to short-sellers, the more time is needed for the information to be incorporated into prices. For example, the authors report that a one standard-deviation increase in lending supply is accompanied by a decrease in D1 of 0.097 standard deviations and a decrease in D2 of 0.080 standard deviations. Both results are statistically significant at the 1% level. However, there is no evidence showing that larger loan fees are associated with greater delay in information transmission.

Overall, the authors find that prices of stocks with limited lending supply reflect information at a slower pace compared to the prices of stocks that have a higher level of lending supply. This result suggests that relaxing short-selling constraints that are proxied by lending supply improves information efficiency.

5.2.3 Changes in short-selling regulations and the speed of price adjustment to new information: More empirical evidence

This sub-section reviews papers that examine the impact of the changes in short-selling regulations on the speed of price adjustment to new information. Chen and Rhee (2010) and Boehmer and Wu (2013) examine the impact of relaxing short-selling rules on the speed of price adjustment to information. In contrast, the work of Ni and Pan (2010), Kolasinski, Reed, and Thornock (2013), and Beber and Pagano (2013) focuses on the impact of tightening short-selling regulations on information efficiency.

5.2.3.1 Relaxing short-selling regulations

Chen and Rhee (2010) directly measure the information efficiency of prices under various short-selling regimes in the Hong Kong exchange over the period 2000–2004. In this exchange, stocks that satisfy certain requirements may be listed on the D-list and consequently sold short. The speed of price adjustment to new information on non-shortable stocks (before listing on the D-listed) versus shortable stocks (after listing on the D-list) is measured by the autocorrelations in quote revisions and trades using the vector autoregressive model (VAR) as proposed by Hasbrouck (1991). The authors identify the periods when the severity of the short-selling regime changes and compare the speed of price adjustment to new information before and after the regime changes.

Chen and Rhee (2010) argue that the differences between shortable and non-shortable stocks are reflected in differences in the autocorrelation in quote revisions and trades. The more negative the coefficient estimates of the autocorrelation in quote revisions and the less positive the coefficient estimates of the autocorrelation in trades, the faster the speed of price adjustment to new information. The results show that the autocorrelation in quote revisions is more negative for shortable stocks and that the autocorrelation in trades is positive but smaller for shortable stocks. These results suggest that the prices of stocks reflect new information faster when short-selling is allowed.

Furthermore, the authors find that (i) the contemporaneous correlation between trades and quote revisions is positive and greater for shortable stocks, (ii) the quote adjustment subsequent to each trade is positive but decreasing in magnitude at longer lags and less positive for non-shortable stocks, and (iii) the Granger causality coefficient estimates from lagged quote revisions to trades are negative and similar for both shortable and non-shortable stocks.

Overall, Chen and Rhee (2010) show that short-selling enhances the information efficiency of stock prices. When short-selling is allowed, information is impounded into prices at a faster pace under both downward and upward market movements.

Boehmer and Wu (2013) study the impact of daily short-selling flows on the information efficiency of NYSE-listed stocks. The study covers the period 2005–2007 using data published by the NYSE under Regulation SHO.[147]

To measure the relative short-term efficiency of transaction prices, the authors use pricing error, as suggested by Hasbrouck (1993), and the absolute value of intraday quote midpoint

[147] Regulation SHO was introduced by the Securities and Exchange Commission (SEC) in order to study the impact of unrestricted short-selling on market volatility, price efficiency, and liquidity (see Boehmer and Wu (2013: p.291, footnote 3).

return autocorrelations. These measures are then regressed on lagged daily shorting, which is expressed as a percentage of contemporaneous trading volume and control variables, such as relative effective spread, share price, market capitalization, and trading volume. Also, a pilot dummy variable is used for pilot stocks for which short-selling orders are allowed to be executed on downticks. In addition, the authors use a regulation dummy variable that indicates the post Regulation SHO period. This allows the authors to examine how relaxing the short-selling constraints affects information efficiency.[148]

The authors find that the lagged daily short-selling flow has a negative and significant coefficient in all specifications. After controlling for other factors, stocks with more intense non-exempt short-selling are found to have significantly smaller pricing errors and return autocorrelations during regulation SHO. In contrast, exempt short-selling is not found to affect pricing errors or the absolute value of return autocorrelations. The authors argue that this is because it is less likely that exempt short-selling will be information-based.

Boehmer and Wu (2013) also measure low-frequency information efficiency by regressing weekly stock returns on contemporaneous and four weeks of lagged market returns. Then they estimate a second regression, where the coefficients on lagged market returns are restricted to zero, and compare the R^2 from the two models. They find that the prices of pilot stocks incorporate information faster compared to non-pilot stocks, which implies that short-selling restrictions impede the speed with which public information is impounded into prices.

Post-earning-announcement drift is the third measure they use to estimate information efficiency. The authors compute earnings surprises as the difference between actual earnings and monthly I/B/E/S forecasts, standardised by the stock price two days before the announcement date. Abnormal returns are constructed as a stock's raw returns minus value-

[148] Under this regulation, short-selling is allowed on the downtick (i.e., stocks can be sold short at prices that are lower than their previous sale prices) for specified pilot stocks whose short-selling transactions can be observed.

weighted market returns. Lastly, the drift is measured as the cumulative abnormal return following each earnings surprise. Boehmer and Wu (2013) find that an increase in short-selling eliminates the post-earnings announcement drift after negative earnings surprises.

Overall, the results of these studies suggest that short-selling activity increases the speed and accuracy with which information is incorporated into prices.

5.2.3.2 Tightening short-selling regulations

Ni and Pan (2010) investigate the price discovery process between the spot market that was affected by the 2008 short-selling ban and the options market in the US. The sample is comprised of 225 financial stocks subject to the short-selling ban with exchange traded options covering the period from September 19, 2008, the first day of the ban, until October 8, the day the ban was lifted.

The authors argue that as long as trading in options is not extremely expensive, investors with negative information could buy a put option, for example, and their information would be incorporated into the price of the underlying stock in the spot market. To test their hypothesis, Ni and Pan (2010) construct put-call ratios from the put and call option volumes initiated by buyers to open new positions on the banned stocks. Firstly, the banned stocks are sorted into groups by their put-call ratios. Then cumulative excess return for each stock is regressed on option to stock volume conditional on whether the stock is in the top put-call ratio group or the bottom put-call ratio group. The results show that when the short-selling ban is in force, the group of option investors with the most negative information underperform the middle group by 2.84%, suggesting that it takes longer for negative information that is revealed in the options market to be impounded into the underlying stock prices. In contrast, the short-selling ban is not found to affect the incorporation of positive information. When the same analysis is conducted for a control sample of non-banned stocks, again, no significant result is found.

Kolasinski, Reed, and Thornock (2013) examine the impact on the information efficiency in the US market of the Emergency Order (EO) that was introduced in July 2008 and the short-selling ban that was enforced in September 2008.

The EO was introduced by SEC and applied to the stocks of 19 financial institutions. According to the EO, investors had to borrow the selected stocks three days before the short sale took place. The EO led to an increase in demand for the borrowed shares and consequently to an increase in the lending fee that stock lenders charged short-sellers. In order to isolate the impact of the short-selling rules, the authors create two samples of stocks: a banned sample, which comprises the stocks of the 19 financial firms that were subject to the new short-selling rules, and a non-banned sample, which comprises the stocks of financial firms that were not subject to the short-selling rules.

In September 2008, SEC introduced a short-selling ban on nearly 800 financial stocks. The authors again construct two samples of financial stocks, both subject to the short-selling ban; however, one sample includes financial stocks with listed options while the other sample contains financial stocks without listed options.

Both short-selling regulations were associated with a very large increase in the cost of borrowing, which, according to the DV model, should drive the uninformed traders out of the market and increase the ratio of informed to uninformed traders in the market. Also, in the case of the short-selling ban, informed investors could sell short synthetically in the options market. Kolasinski, Reed, and Thornock (2013) use two measures to study the effects of each short-selling regulation on the informativeness of short-selling. They hypothesize that prices react more negatively to unexpected short-interest news for banned stocks under EO and for banned stocks with traded options under the short-selling ban.

The first measure captures the market response to short-interest announcements during the short-selling rule changes. More precisely, the cumulative return around the short interest announcement date is regressed on the unexpected short interest (i.e., current short interest less average short interest) in a difference-in-difference framework controlling for other factors such as size, book to market value, return volatility, and turnover. The results show that during EO the market reaction to unexpected short interest news is significantly more negative for banned stocks relative to non-banned stocks. In contrast, there is no significant difference in market reaction to unexpected short interest announcements between banned firms with options and banned firms without options during the period of the short-selling ban.

The second measure estimates the impact of the short-selling regulations on the relationship between returns and short volume. In particular, the authors conduct a difference-in-difference regression analysis where individual stock returns are regressed on lagged ratio of short volume to total volume, again controlling for other factors such as size, book to market value, turnover, and return volatility. The authors hypothesize that the price impact of short-selling orders should increase if the ratio of informed to uninformed short-sellers increases during the short-selling rules. The results do not provide supporting evidence for the banned stocks under EO; however, the price impact is found to be stronger for banned stocks with listed options relative to the price impact for the banned stocks without listed options under the short-selling ban.

Overall, Kolasinski, Reed, and Thornock (2013) find some evidence supporting the DV model's prediction that the short-selling restrictions increase the ratio of informed to uninformed traders in the market. Furthermore, the authors find that the presence of listed options improves the informativeness of the prices of the financial stocks that were subject to the short-selling ban.

Beber and Pagano (2013) examine the effects of the 2008 short-selling regulations on the speed of price discovery across 30 countries. The speed of price discovery is measured by the degree to which individual stock returns correlate with past market returns rather than contemporaneous market returns. More specifically, the authors estimate a market model in which weekly returns for each individual stock in their sample are regressed on the corresponding broad national stock market index for the period January 2008–June 2009, focusing on residuals. They argue that short-selling bans should impede the incorporation of firm-specific rather than market-wide information into prices. If this is the case, then one would expect higher autocorrelations of residuals during the short-selling ban period, particularly when stock returns are negative. Their results show that (i) the median upside and median downside cross-autocorrelations are positive and greater during the ban, (ii) the median downside cross auto-correlation is larger than the median upside cross-autocorrelation, and (iii) the difference between the median downside cross-autocorrelation and the median upside cross-autocorrelation is larger when short-selling is banned. For example, the median cross-autocorrelation between individual stock returns and lagged market negative return in the pre-ban period is 0.28 and in the ban period 0.35. The respective figures for the median cross-autocorrelation between individual stock returns and market positive or zero returns are 0.23 and 0.26. These results confirm the theoretical prediction by showing that the autocorrelation of residuals is larger for stocks that were affected by the short-selling bans. Furthermore, a variance ratio test is used for stocks subject to a short-selling ban and stocks not subject to a short-selling ban, respectively, as a robustness check. The authors find that for stocks not subject to a short-selling ban, the hypothesis that their returns follow a random walk cannot be rejected in 53% of the cases. In contrast, the comparable figure for stocks subject to a short-selling ban is only 39%. The difference in the result between the two samples is found to be statistically significant at the 1% level. Again, this result supports the view that short-selling constraints reduce the speed of price adjustment to new information.

Overall, the results of these studies suggest that short-selling bans impede the process of information incorporation into prices and that the effect is more pronounced in falling market conditions (Beber and Pagano (2013)). However, the trading of options written on banned stocks might improve the informativeness of the prices of these stocks during short-selling bans (Kolasinski, Reed, and Thornock (2013)).

5.2.4 Summary

The DV model predicts that short-selling constraints reduce the speed of price adjustment to new information and this effect is stronger when markets fall. A number of papers test the DV prediction and provide empirical evidence on this issue. Overall, the studies reviewed in this essay find that when short-selling constraints are in place, the speed and accuracy with which new information is impounded into prices is impaired. However, there is evidence that prices of banned stocks might retain their informativeness under a short-selling ban when traded options are available.

In summary, the speed of price adjustment to new information is found to be lower for stocks with less lending supply, which proxies for short-selling constraints (Saffi and Sigurdson (2011)) and for stocks traded in countries where short-selling is either not practised or prohibited (Bris, Goetzmann, and Zhu (2007); Beber and Pagano (2013)). Alternatively, when the relaxation of short-selling restrictions is examined, it is found that information is incorporated into prices faster for the stocks that are allowed to be shorted (Chen and Rhee

(2010); Boehmer and Wu (2013)). Furthermore, evidence shows that when short-selling is banned, prices of stocks subject to a short-selling ban that have traded options are more informative than the prices of stocks subject to a short-selling ban that do not have traded options (Kolasinski, Reed, and Thornock (2013)). Nevertheless, within the group of stocks that have traded options, information is incorporated into prices at a slower pace in the sub-group in which options investors express the most negative information (Ni and Pan (2010)).

5.3 Hypotheses development

This section develops the hypotheses on the impact of the 2008 short-selling regulations on the speed of price adjustment to new information on the stocks of European financial institutions subject to these short-selling regulations. As discussed in section 5.2, DV argue that the speed of price adjustment to new information is slower when short-selling constraints become more severe.

Chen and Rhee (2010) test DV's model and measure the speed of price adjustment to new information using the methodology proposed by Hasbrouck (1991). This methodology involves estimating the interaction between trades and quotes in a VAR model framework. More precisely, according to Hasbrouck (1991), increasing trade continuity (i.e., more positive autocorrelation in trades) and decreasing quote reversals (i.e., less negative autocorrelation in quotes) are associated with lagged incorporation of information into prices. Stronger autocorrelation in trades leads to less revision in beliefs and consequently reduces the speed of price adjustment to new information (Madhavan, Richardson, and Roomans (1997)). Chen and Rhee (2010) find that new information is incorporated faster into prices when stocks are allowed to be sold short. Thus, the introduction of short-selling regulations in 2008 is expected to reduce the speed of price adjustment to new information measured by trade continuity and quote reversals.

Following Hasbrouck (1991) and Chen and Rhee (2010) this essay formulates the following hypotheses:

Hypothesis 1. Autocorrelation in quotes for European financial stocks subject to a short-selling ban becomes less (more) negative during the short-selling ban (post-ban) period.

Hypothesis 2. Autocorrelation in trades for European financial stocks subject to a short-selling ban becomes more (less) positive during the short-selling ban (post-ban) period.

There were two main forms of short-selling bans introduced in Europe in the autumn of 2008: a naked short-selling ban and a total (naked and covered) short-selling ban. A total short-selling ban is more severe than a naked short-selling ban because the short sellers are not allowed to borrow shares in order to sell them short and they need to actually own them to be able to sell them. Consequently, this essay also hypothesizes that:

Hypothesis 3. The impact of total short-selling bans is stronger that the impact of naked short-selling bans on autocorrelation in quotes and trades.

Empirical evidence shows that the impact of short-selling constraints on price discovery is mitigated by the presence of listed options (Phillips (2011); Kolasinski, Reed, and Thornock (2013)).[149] For example, Kolasinski, Reed, and Thornock (2013) find that the presence of listed options improves the informativeness of prices of stocks that were subject to the 2008 short-selling ban in the US. They argue that informed investors could sell short banned stocks synthetically in the options market during the short-selling ban. In Europe, financial authorities prohibited short-selling in both the spot and the derivatives markets in eight out of

[149] Phillips (2011) investigates whether short-selling was facilitated by the presence of options trading when short-selling was constrained for the US market over the period from 1981 to 1997. The author finds that the introduction of options mitigates short-selling constraints for stocks with low loan supply. However, this result is statistically significant only in the presence of negative information. For example, prior to the introduction of the options, stocks with short-selling constraints are found to adjust to bad news 19% more slowly than stocks without short-selling constraints. In contrast, after the introduction of the options, this difference falls to 4%.

the 10 countries that are examined in this essay.[150] Assuming that the short-selling regulations that were applied to the derivatives markets in Europe were binding, this essay hypothesizes that:

Hypothesis 4. The speed of price adjustment to new information is not affected by the absence or presence of listed derivatives (options and/or futures) for the financial stocks subject to the short-selling bans after the introduction of these bans.

5.4 Sample and Data

5.4.1 Introduction

Intraday quote and trade data on the stocks of 78 financial institutions from 10 European countries were obtained from Thomson Reuters Tick History (TRTH) database of Securities Industry Research Centre of Asia-Pacific (SIRCA).[151,152] The sample period spans from mid-January 2008 to mid-February 2009 and it is divided into three sub-periods. Sub-period 1 (pre-ban period) is defined as the period before the introduction of new short-selling regulations. Sub-period 2 (ban period) is the period when short-selling regulations are introduced and remain in force. Finally, sub-period 3 (post-ban period) is defined as the period when short-selling regulations are lifted (see Table 5.1).

Table 5.1 briefly summarises the short-selling regulations that were introduced in the sample countries and the number of financial stocks affected by their introduction.[153] There are five

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[150] See Chapter 2 for more details.
[151] These countries are Belgium, Denmark, France, Germany, Ireland, the Netherlands, Portugal, Spain, Switzerland, and the UK.
[152] Appendix 5A reports and discusses the results for Norway with two financial stocks in the sample and Italy with 39 financial stocks in the sample–two special cases.
[153] Only 78 financial stocks out of the 109 that were affected by the 2008 short-selling rules are included in the analysis due to data insufficiency for the remaining stocks (Nordea AB-traded on Copenhagen Stock Exchange, AMB General Holdings, Julius Bar, Banco de Andalucia, Banco de Castilla, Banco de Credito Balear, Banco de Galicia, Banco de Vasconia, Caja de Ahorros del Mediterraneo, Inverfiate, Alliance & Leicester, Bradford & Bingley, Arbuthnot Banking Group, European Islamic Investment Bank, Friends Provident, HBOS, Highway Insurance Group, Islamic Bank of Britain, London Scottish Bank, Tawa, HSBC Holdings-traded on Euronext

countries that banned both naked and covered short-selling of financial stocks (i.e., Denmark, Ireland, the Netherlands, Switzerland, and the UK) and five countries that banned only naked short-selling of financial stocks (Belgium, France, Germany, Portugal, and Spain).[154] The three countries with a post-ban period available during the sample period are the Netherlands, Switzerland and the UK.

5.4.2 Control sample

Short-selling regulations were introduced during exceptionally volatile market conditions. Consequently, any results reported for the financial stocks that were subject to the short-selling bans (hereafter banned stocks) could be due to other factors that were not related to the short-selling regulations. This essay uses a control sample of non-financial stocks not subject to the short-selling bans (hereafter non-banned stocks) to control for market movements. The control sample is constructed as described in Chapter 3, section 3.4.3.

Paris, Banco Popular Espanol, Banco Santander-traded on Euronext Lisbon, Just Retirement Holdings, British Insurance Holdings, Finibanco, Banco Guipuzcoano, Banco Pastor, Hypo Real Estate, Irish Life and Permanent, Van der Moolen Holdings, and Swiss Reinsurance).

[154] Refer to Chapter 2 or Appendix 2A for detailed information on the short-selling regulations and stocks that were affected by these regulations in the sample countries.

Table 5.1 Short-selling regulations in Europe

This table describes the type of short-selling regulations that were introduced in 10 European countries, the number of financial stocks that were affected by these short-selling regulations, and the duration of each sample sub-period for each European country that is used in the analysis.

Country	Type of the short-selling ban	The number of banned stocks in the sample	Total number of banned stocks	Pre-ban period	Ban period	Post ban period
Belgium	Naked	4	4	15/01/2008-19/09/2008	22/09/2008-15/12/2009	N/A
France	Naked	10	13	15/01/2008-19/09/2008	22/09/2008-15/12/2009	N/A
Germany	Naked	9	11	15/01/2008-19/09/2008	22/09/2008-15/12/2008	N/A
Portugal	Naked	5	8	15/01/2008-23/09/2008	24/09/2008-15/12/2009	N/A
Spain	Naked	11	20	15/01/2008-23/09/2008	24/09/2008-15/12/2009	N/A
Denmark	Total	3	4	15/01/2008-10/10/2008	13/10/2008-15/12/2009	N/A
Ireland	Total	2	3	15/01/2008-18/09/2008	19/09/2008-15/12/2009	N/A
Netherlands	Total	6	7	15/01/2008-03/10/2008	06/10/2008-29/05/2009	01/06/2009-15/12/2009
Switzerland	Total	5	7	15/01/2008-18/09/2008	19/09/2008-16/01/2009	19/01/2009-15/12/2009
UK	Total	23	32	15/01/2008-18/09/2008	19/09/2008-16/01/2009	19/01/2009-15/12/2009
Total		78	109			

5.5 Methodology

Hasbrouck (1991) argues that the information contained in each trade includes both private and public information; private information is contained in the trade innovation while public information is contained in the quote revision innovation.[155] According to the author, lagged adjustment to information is associated with more positive autocorrelations in trades, and less negative autocorrelations in quote revisions.

This essay uses Hasbrouck's (1991) VAR model to assess the impact of short-selling bans on the speed of price adjustment to new information on the banned stocks.[156] Means and medians of the VAR coefficient estimates are calculated for each lag in each sub-period and then summed to give the aggregate coefficient estimate across all lags for each sub-period.[157] Only the means and medians of the aggregate coefficient estimates are reported in section 5.6.

The following dynamic unrestricted VAR model is used (see also Chen and Rhee (2010)):

$$r_t = \sum_{i=1}^{5} a_i\, r_{t-i} + \sum_{i=0}^{5} b_i\, Q_{t-1} + v_{1,t}$$

$$Q_t = \sum_{i=1}^{5} c_i\, r_{t-i} + \sum_{i=1}^{5} d_i\, Q_{t-1} + v_{2,t}$$

[155] Hasbrouck (1991) develops a model to estimate the interaction between quotes and trades for 80 NYSE/AMEX stocks over the first quarter of 1989 in order to test his theoretical framework. Results show that the quote-midpoint (b_0 coefficient) increases immediately subsequent to a purchase order. Additionally, quote adjustment subsequent to each trade (b_i coefficients) at longer lags is positive but decreasing. Also, the author finds positive autocorrelation in trades (d_i coefficients) and negative autocorrelation in quote revisions (a_i coefficients). Finally, Granger-Sims causality is found to run from quote revisions to trades. Only the last result contradicts the theoretical framework that Hasbrouck (1991) developed. The author attributes it to reporting errors in the quotes that might have led to trade misclassification.

[156] Although Autoregressive Conditional Duration (ACD) models can be used for modeling the intraday behavior of irregularly time-spaced transactions (see for example, Engle and Russell (1998) and Pacurar (2006)), the objective of this essay is to examine whether, on average, there is a change in the speed of price adjustment to new information under various short-selling regimes rather than modeling and comparing the intraday behavior of irregularly time-spaced trades when short-selling regulations change. Thus, Hasbrouck's (1991) approach is chosen to address the research question stated in the present essay.

[157] Specification tests which determine the optimal lag length by minimizing the Akaike Information Criterion indicated five lags for a sample of stocks examined.

where r_t=m_t-m_{t-1} is the log quote midpoint change with m_t being the log quote midpoint for a transaction occurring at time t. Q_t is the buy-sell indicator variable that takes the value of 1 if the trade occurs above m_{t-1}, 0 if the trade occurs at m_{t-1}, and -1 if the trade occurs below m_{t-1}. The a_i coefficients estimate the autocorrelation in the quote revisions while the d_i coefficients estimate the autocorrelation in trades. The b_i coefficients measure the quote adjustment after each trade and the c_i coefficients measure the Granger causality running from quote revisions to trades. Lastly, the b_0 coefficient measures the contemporaneous correlation between trades and quote-midpoint returns (see Chen and Rhee (2010)).

5.6 Empirical results

5.6.1 Country-level results

This section presents the median values of a_i coefficient sums (hereafter aggregate a_i coefficients) and d_i coefficient sums (hereafter aggregate d_i coefficients) aggregated over five lags for both banned and non-banned stocks across 10 European countries.[158]

Table 5.2 reports the median values of aggregate a_i coefficient estimates, which measure the autocorrelation in quote revisions, for banned and non-banned stocks, respectively. Panel A of Table 5.2 reports the results for countries with two sub-periods available (i.e., the pre-ban period and the ban period). These countries are Belgium, Denmark, France, Germany, Ireland, Portugal, and Spain. Accordingly, Panel B of Table 5.2 shows the results for the three countries – the Netherlands, Switzerland, and the UK – with three sub-periods available. These sub-periods are the pre-ban period, the ban period, and the post-ban period.

[158] It was not possible to create a control sample of non-banned stocks for Norway due to data insufficiency and Italy due to frequent changes in short-selling regulations over the sample period. Italian financial authorities alternated between banning short-selling of financial stocks only and banning short-selling of all stocks trading on Italian stock exchanges. Mean and median aggregate a_i and d_i coefficients for the financial banned stocks of these two countries are reported in Appendix 5A. Median values of b_0, aggregate b_i, and aggregate c_i coefficients are reported in Appendix 5B. Appendix 5C reports the mean values of all the coefficients estimated by the Hasbrouck’s (1991) bivariate VAR model.

Prior evidence shows that the aggregate a_i coefficient estimates have a negative sign (Chen and Rhee (2010)). Column (1) of Panel A of Table 5.2 shows that in the pre-ban period, the aggregate a_i coefficient estimates of the banned stocks are negative in four out of seven countries (Belgium, France, Portugal, and Spain) and positive for the remaining countries. However, only two out of seven countries have negative aggregate a_i coefficient estimates for the non-banned stocks during the pre-ban period, as reported in column (2). Columns (1) and (2) of Panel B of Table 5.2 show that the aggregate a_i coefficient estimates are negative for the Dutch and Swiss banned stocks and Swiss non-banned stocks, respectively, during the ban period. One explanation for positive autocorrelation in quote revisions could be that market makers do not immediately revise their quotes to incorporate information. Instead, they choose a smooth price transition to avoid "too discontinuous" prices (Hasbrouck (1991)).

Hypothesis 1 asserts that aggregate a_i coefficient estimates become less negative, or alternatively increase, for the banned stocks during the ban period and more negative, or alternatively decrease, during the post-ban period for the same stocks. As can be seen in column (3) of both Panels of Table 5.2, the aggregate a_i coefficient estimates increase significantly only for the French banned stocks. There is no significant change for Belgium, Germany, Portugal, and Switzerland and the aggregate a_i coefficient estimates decrease significantly for the remaining five countries. For example, the autocorrelation in quote revisions for Spanish banned stocks is -0.0642 in the pre-ban period, which decreases to -0.0865 in the ban period (see column (1)). This result is statistically significant at the 5% level (see column (3)). In contrast, aggregate a_i coefficient estimates increase significantly for the non-banned stocks in four out of 10 countries and decrease significantly only in one country during the ban period (see column (4)). There is no significant change in the aggregate a_i coefficient estimates for the non-banned stocks in Denmark, Ireland, Portugal, Spain, and the UK, as shown in column (4) of Table 5.2.

Overall, there is no evidence supporting Hypothesis 1. However, there is some evidence showing that the aggregate a_i coefficient estimates for the banned stocks change in the opposite direction from what would be expected during the ban period. These results contradict the findings of Chen and Rhee (2010), who report less negative autocorrelation in aggregate a_i coefficient estimates when short-selling is not allowed.

In the post-ban period, the aggregate a_i coefficient estimates increase for both banned and non-banned stocks in all three countries, and these changes are statistically significant at the 1% level as reported in columns (3) and (4), respectively, of Panel B of Table 5.2. For example, the autocorrelation in quote revisions for the British banned stocks increases from -0.0094 in the ban period to 0.0242 in the post-ban period (see column (1)). The respective figures for the autocorrelation in quote revisions for the British non-banned stocks are 0.0124 in the ban period and 0.0901 in the post-ban period (see column (2)). Again, these results contradict Hypothesis 1 in the post-ban period and are not in line with the results reported by Chen and Rhee (2010).

In summary, the results presented in Table 5.2 do not support Hypothesis 1. During the ban period, the aggregate a_i coefficient estimates for the banned stocks decrease significantly rather than increase as would be expected in five out of 10 countries. Furthermore, the aggregate a_i coefficient estimates for the banned stocks increase significantly for all three countries during the post-ban period instead of decreasing as predicted.

Table 5.2 Aggregate a_i coefficient estimates: Country-level results

This table shows the median values of the coefficients $\sum_{i=1}^{5} a_i$ for banned and non-banned stocks, respectively, across 10 European countries. a_i coefficient sums reflect the autocorrelation in quote revisions and are estimated by using the bivariate VAR model as shown in Chen and Rhee (2010):

$$r_t = \sum_{i=1}^{5} a_i\, r_{t-i} + \sum_{i=0}^{5} b_i\, Q_{t-1} + v_{1,t}$$

$$Q_t = \sum_{i=1}^{5} c_i\, r_{t-i} + \sum_{i=1}^{5} d_i\, Q_{t-1} + v_{2,t}$$

The sample period covers the period from mid-January 2008 to mid-December 2009 and it is divided into three sub-periods: the pre-ban period (P1), the ban period (P2), and the post-ban period (P3). Panel A reports the results for countries that have two sub-periods, the pre-ban period and the ban period (i.e., Belgium, Denmark, France, Germany, Ireland, Portugal, and Spain). Panel B shows the results for countries with a post-ban period available (i.e., Switzerland, the Netherlands, and the UK). The Wilcoxon test is used and the p-values of the Z-statistic are given in the parentheses. * indicates statistical significance at the 10% level. ** indicates statistical significance at the 5% level. *** indicates statistical significance at the 1% level.

Panel A: Countries with two sub-periods						
		Aggregate a_i coefficients			Change in aggregate a_i coefficients	
		(1)	(2)		(3)	(4)
Country	Sub-periods	Banned stocks	Non-banned stocks		Banned stocks	Non-banned stocks
Belgium	Pre-ban	-0.0128	0.0270	P2-P1	0.0155	0.0504***
	Ban	0.0026	0.0775	(p-value)	(0.1276)	(0.0001)
Denmark	Pre-ban	0.0858	0.0749	P2-P1	-0.0380***	0.0004
	Ban	0.0478	0.0753	(p-value)	(0.0001)	(0.7478)
France	Pre-ban	-0.0080	0.0030	P2-P1	0.0234***	0.0277***
	Ban	0.0154	0.0307	(p-value)	(0.0001)	(0.0001)
Germany	Pre-ban	0.0183	0.0207	P2-P1	-0.0044	0.0084***
	Ban	0.0139	0.0291	(p-value)	(0.1406)	(0.0060)
Ireland	Pre-ban	0.0397	0.0499	P2-P1	-0.0209***	0.0456
	Ban	0.0188	0.0955	(p-value)	(0.0070)	(0.0632)
Portugal	Pre-ban	-0.0935	-0.0012	P2-P1	0.0329	0.0031
	Ban	-0.0606	0.0019	(p-value)	(0.0896)	(0.5031)
Spain	Pre-ban	-0.0642	-0.0380	P2-P1	-0.0223**	0.0045
	Ban	-0.0865	-0.0335	(p-value)	(0.0271)	(0.2842)
Panel B: Countries with three sub-periods						
Netherlands				P2-P1	-0.0247***	0.0061**
	Pre-ban	-0.0304	0.0018	(p-value)	(0.0003)	(0.0399)
	Ban	-0.0551	0.0079	P3-P2	0.0471***	0.0309***
	Post-ban	-0.0080	0.0388	(p-value)	(0.0001)	(0.0001)
Switzerland				P2-P1	-0.0072	-0.0180**
	Pre-ban	-0.0734	-0.0247	(p-value)	(0.0850)	(0.0285)
	Ban	-0.0806	-0.0427	P3-P2	0.1100***	0.0930***
	Post-ban	0.0295	0.0503	(p-value)	(0.0001)	(0.0001)
UK				P2-P1	-0.0100***	0.0008
	Pre-ban	0.0006	0.0116	(p-value)	(0.0001)	(0.5903)
	Ban	-0.0094	0.0124	P3-P2	0.0337***	0.0777***
	Post-ban	0.0242	0.0901	(p-value)	(0.0001)	(0.0001)

Table 5.3 reports the medians of the aggregate d_i coefficient estimates for banned and non-banned stocks, respectively. Hypothesis 2 states that the autocorrelation in trades, which is measured by the aggregate d_i coefficient estimates, increases during the ban period and subsequently decreases in the post-ban period for the banned stocks. Aggregate d_i coefficient estimates are expected to be positive (Hasbrouck (1991)).

As predicted, the aggregate d_i coefficient estimates are positive in all sub-periods for both samples of stocks (see columns (1) and (2) of Table 5.3). Column (3) of Panels A and B of Table 5.3 shows that during the ban period the aggregate d_i coefficient estimates of the banned stocks increase significantly in only three out of 10 countries (i.e., Germany, Ireland, and the Netherlands). There is no significant change in the aggregate d_i coefficient estimates of the banned stocks for Belgium, Denmark, Portugal, and Switzerland during the ban period. Finally, the aggregate d_i coefficient estimates of banned stocks decrease significantly for France, Portugal, and the UK during the same period. Column (4) of Table 5.3 shows that during the ban period the aggregate d_i coefficient estimates of the non-banned stocks increase significantly only for Portugal. There is a significant decrease in the aggregate d_i coefficient estimates for Belgian, German, Spanish, and British non-banned stocks. Lastly, there is no significant change in the aggregate d_i coefficient estimates of the non-banned stocks for the remaining five countries. For example, the autocorrelation in trades for German banned stocks increases from 0.2476 in the pre-ban period to 0.2618 in the ban period (see column (1)). This result is statistically significant at the 1% level (see column (3)). In contrast, the autocorrelation in trades for German non-banned stocks decreases from 0.2331 in the pre-ban period to 0.2163 in the ban period (see column (2)). This result is also statistically significant at the 1% level (see column (4)). Overall, there is very weak evidence supporting Hypothesis 2 in the ban period.

Column (3) of Panel B of Table 5.3 shows that, in the post-ban period, the aggregate d_i coefficient estimates of the banned stocks decrease significantly for Switzerland as expected but increase significantly for the UK. Both results are statistically significant at the 1% level. There is no significant change in the aggregate d_i coefficient estimates of the Dutch banned stocks during the post-ban period. The aggregate d_i coefficient estimates of the non-banned stocks decrease significantly, at the 1% level, for the Netherlands and Switzerland, while they increase significantly, also at the 1% level, for the UK during the post-ban period. Overall, these results support Hypothesis 2 for banned stocks in only one out of three countries during the post-ban period.

In summary, there is no strong evidence supporting Hypothesis 2 at the country level. The results presented in Table 5.3 are mixed for both samples of stocks with the most predominant result being no significant change in the aggregate d_i coefficient estimates during the ban period.

Table 5.3 Aggregate d_i coefficient estimates: Country-level results

This table shows the median values of the coefficients $\sum_{i=1}^{5} d_i$ for banned and non-banned stocks, respectively, across 10 European countries. d_i coefficient sums reflect the autocorrelation in trades and are estimated by using the bivariate VAR model as shown in Chen and Rhee (2010):

$$r_t = \sum_{i=1}^{5} a_i r_{t-i} + \sum_{i=0}^{5} b_i Q_{t-1} + v_{1,t}$$

$$Q_t = \sum_{i=1}^{5} c_i r_{t-i} + \sum_{i=1}^{5} d_i Q_{t-1} + v_{2,t}$$

The sample period covers the period from mid-January 2008 to mid-December 2009 and it is divided into three sub-periods: the pre-ban period (P1), the ban period (P2), and the post-ban period (P3). Panel A reports the results for countries that have two sub-periods, the pre-ban period and the ban period (i.e., Belgium, Denmark, France, Germany, Ireland, Portugal, and Spain). Panel B shows the results for countries with a post-ban period available (i.e., Switzerland, the Netherlands, and the UK). The Wilcoxon test is used and the p-values of the Z-statistic are given in the parentheses. * indicates statistical significance at the 10% level. ** indicates statistical significance at the 5% level. *** indicates statistical significance at the 1% level.

Panel A: Countries with two sub-periods						
		Aggregate d_i coefficients			Change in aggregate a_i coefficients	
		(1)	(2)		(3)	(4)
Country	Sub-periods	Banned stocks	Non-banned stocks		Banned stocks	Non-banned stocks
Belgium	Pre-ban	0.4278	0.4934	P2-P1	-0.0107	-0.0274***
	Ban	0.4171	0.4659	(p-value)	(0.6761)	(0.0001)
Denmark	Pre-ban	0.2402	0.2546	P2-P1	0.0138	-0.0157
	Ban	0.2541	0.2388	(p-value)	(0.0655)	(0.1499)
France	Pre-ban	0.4462	0.4048	P2-P1	-0.0188***	0.0096
	Ban	0.4274	0.4144	(p-value)	(0.0001)	(0.0728)
Germany	Pre-ban	0.2476	0.2331	P2-P1	0.0141***	-0.0168***
	Ban	0.2618	0.2163	(p-value)	(0.0059)	(0.0002)
Ireland	Pre-ban	0.2345	0.3159	P2-P1	0.1054***	-0.0032
	Ban	0.3400	0.3127	(p-value)	(0.0070)	(0.7042)
Portugal	Pre-ban	0.3526	0.3018	P2-P1	-0.0023	0.0477***
	Ban	0.3504	0.3495	(p-value)	(0.9658)	(0.0001)
Spain	Pre-ban	0.3570	0.3799	P2-P1	-0.0088*	-0.0336***
	Ban	0.3482	0.3462	(p-value)	(0.0154)	(0.0001)
Panel B: Countries with three sub-periods						
Netherlands	Pre-ban	0.3740	0.3494	P2-P1	0.0166**	0.0169
	Ban	0.3906	0.3664	(p-value)	(0.0155)	(0.1992)
	Post-ban	0.3970	0.3221	P3-P2	0.0064	-0.0443***
				(p-value)	(0.0682)	(0.0001)
Switzerland	Pre-ban	0.3728	0.3371	P2-P1	0.0017	0.0151
	Ban	0.3745	0.3522	(p-value)	(0.8746)	(0.2218)
	Post-ban	0.2365	0.2212	P3-P2	-0.1380***	-0.1310***
				(p-value)	(0.0001)	(0.0001)
UK	Pre-ban	0.4528	0.5130	P2-P1	-0.0183***	-0.0729***
	Ban	0.4345	0.4401	(p-value)	(0.0005)	(0.0001)
	Post-ban	0.5639	0.6630	P3-P2	0.1293***	0.2229***
				(p-value)	(0.0001)	(0.0001)

5.6.2 Summary of the country-level results

According to Hasbrouck (1991), less negative autocorrelation in quote revisions and more positive autocorrelation in trades are consistent with a slower speed of price adjustment to new information. Positive autocorrelation in trades implies that sales tend to follow sales and purchases tend to follow purchases. Stronger autocorrelation in trades will be associated with fewer quote revisions, and both suggest that the speed with which new information is incorporated into prices is slower (Madhavan, Richardson, and Roomans (1997)).

Based on prior studies, it was expected that the negative aggregate a_i coefficient estimates, which represent quote revisions, would become less negative during the ban period and more negative during the post-ban period for countries where the post-ban period is available (i.e., the Netherlands, Switzerland, and the UK). It was also expected that the positive aggregate d_i coefficient estimates, which represent autocorrelation in trades, would become more positive during the ban period and less positive in the post-ban period.

Table 5.4 summarizes the findings reported in Tables 5.2 and 5.3 to provide a complete picture of the overall change in the speed of price adjustment to new information during the ban period. Columns (1) and (4) show the changes in the aggregate a_i coefficient estimates for the banned and non-banned stocks, respectively, during the ban period across all countries in the sample. A positive sign indicates an increase in the aggregate a_i coefficient estimates and a negative sign indicates a decrease in the aggregate a_i coefficient estimates. Accordingly, columns (2) and (5) show the changes in the aggregate d_i coefficient estimates for the banned and non-banned stocks during the ban period. This essay hypothesizes that the speed of price adjustment to new information decreases for the banned stocks during the ban period, and consequently expects an increase in both aggregate a_i and d_i coefficient estimates.

As can be seen in column (3) (column (6)) of Table 5.4, at a country level, the speed of price adjustment to new information on banned stocks (non-banned stocks) does not change during the ban period for Belgium, Portugal, and Switzerland (Denmark and Ireland). There is an uncertain impact on the speed of price adjustment to new information on banned stocks (non-banned stocks) during the ban period for France, Ireland, and the Netherlands (Belgium and Germany). The speed of price adjustment to any new information on the banned stocks (non-banned stocks) increases during the ban period for Spain and the UK (Spain and the UK – in terms of lower aggregate d_i coefficient estimates, and Switzerland – in terms of lower aggregate a_i coefficient estimates).

In the post-ban period, figures in Tables 5.2 and 5.3 show that the speed of price adjustment to new information decreases for the banned stocks in the Netherlands and the UK while the impact on the speed of price adjustment to new information for the banned stocks in Switzerland is uncertain. These results contradict Hypotheses 1 and 2.

Overall, the evidence in Table 5.4 shows a difference in behavior in the change of speed of price adjustment to new information between banned and non-banned stocks during the ban period. Spain and the UK are the exception. In these two cases, the speed of price adjustment to new information on banned stocks increases during the ban period in terms of both more negative autocorrelation in quote revisions and less positive autocorrelation in trades. However, for the Spanish and British non-banned stocks, the speed of price adjustment to new information increases only in terms of less positive autocorrelation in trades as there is no significant change in the autocorrelation of quote revisions. These results imply that new information is incorporated faster into the prices of the banned stocks in these two countries.

In summary, country-level results show that the speed of price adjustment to new information is not slower for the banned stocks during the ban period, contradicting Hypothesis 1 and 2.

These results are also not in line with the findings of Chen and Rhee (2010), who show that the speed of price adjustment to new information is slower for stocks that are not allowed to be sold short.

Table 5.4 Change in the speed of price adjustment to new information: A country-level summary of findings

This table summarizes the change in speed of price adjustment to new information from the pre-ban period (P1) to the ban period (P2) across 10 European countries. A positive sign indicates an increase while a negative sign indicates a decrease of the aggregate a_i and d_i coefficient estimates during the ban period. {} indicates statistical significance at the 5% level. [] indicates statistical significance at the 1% level. A significant increase in either or both aggregate a_i and d_i coefficients indicates a slower speed of price adjustment to new information whereas a significant decrease in the same coefficients indicates a faster speed of price adjustment to new information. If the aggregate a_i and d_i coefficients show opposite signs and are significant, the change in speed of price adjustment to new information becomes uncertain.

	Banned stocks (P2 - P1)			Non-banned stocks (P2 - P1)		
	(1)	(2)	(3)	(4)	(5)	(6)
Country	Change in aggregate a_i coefficients	Change in aggregate d_i coefficients	Overall change in speed of price adjustment	Change in aggregate a_i coefficients	Change in aggregate d_i coefficients	Overall change in speed of price adjustment
Belgium	+	-	No change	[+]	[-]	Uncertain
Denmark	[-]	+	Faster	+	-	No change
France	[+]	[-]	Uncertain	[+]	+	Slower
Germany	-	[+]	Slower	[+]	[-]	Uncertain
Ireland	[-]	[+]	Uncertain	+	-	No change
Portugal	+	-	No change	+	[+]	Slower
Spain	{-}	{-}	Faster	+	[-]	Faster
Netherlands	[-]	{+}	Uncertain	{+}	+	Slower
Switzerland	-	+	No change	{-}	+	Faster
UK	[-]	[-]	Faster	+	[-]	Faster

5.6.2 Aggregate-level results

5.6.2.1 Unconditional aggregate-level results

This sub-section presents the results of aggregate a_i and d_i coefficient estimates aggregated across all countries and focuses on the changes of these estimates between the pre-ban and ban periods as reported in Table 5.5. The next two sub-sections present the results of aggregate a_i and d_i coefficient estimates at an aggregate level conditioning on the stringency of the short-selling bans and the presence of listed derivatives, respectively.

Columns (3) and (4) of Table 5.5 show that both aggregate a_i and d_i coefficient estimates for the banned stocks decrease significantly during the ban period. The change in the aggregate a_i coefficient estimates is statistically significant at the 1% level, while the change in the aggregate d_i coefficient estimates is statistically significant at the 10% level. These results imply a faster speed of price adjustment to new information for the banned stocks during the ban period and contradict both Hypothesis 1 and 2 (see column (5)). In contrast, there is uncertain change in the speed of price adjustment to new information for the non-banned stocks, as there is a significant increase in the aggregate a_i coefficient estimates but a significant decrease in the aggregate d_i coefficient estimates in the ban period (see columns (3) and (4)).

Overall, these results do not support Hypothesis 1 and 2 at an aggregate level. Also, these results are not consistent with the findings reported by Chen and Rhee (2010). These authors apply the same methodology as this essay but focus on the Hong Kong market during a different sample period. They find that the aggregate a_i coefficient estimates are less negative and the d_i coefficient estimates are more positive for stocks that are not allowed to be sold short. Other studies that examine the impact of changes in short-selling regulations on the speed of price adjustment to new information, but use different research methods and frequency of data, also find that short-selling restrictions are associated with slower speed of price adjustment to new information (see for example, Bris, Goetzman, and Zhu (2007) and Beber and Pagano (2013)).

Table 5.5 Aggregate a_i and d_i coefficient estimates: Aggregate-level results

This table reports the median values of the coefficients $\sum_{i=1}^{5} a_i$ and $\sum_{i=1}^{5} d_i$ and their changes from the pre-ban period (P1) to the ban period (P2) for banned and non-banned stocks, respectively, at an aggregate level. a_i coefficient sums reflect the autocorrelation in quote revisions and d_i coefficient sums reflect the autocorrelation in trades. The aggregate coefficients are estimated by using the bivariate VAR model as shown in Chen and Rhee (2010):

$$r_t = \sum_{i=1}^{5} a_i \, r_{t-i} + \sum_{i=0}^{5} b_i \, Q_{t-1} + v_{1,t}$$

$$Q_t = \sum_{i=1}^{5} c_i \, r_{t-i} + \sum_{i=1}^{5} d_i \, Q_{t-1} + v_{2,t}$$

The sample period covers the period from mid-January 2008 to mid-December 2009. The Wilcoxon test is used and the p-values of the Z-statistic are given in the parentheses. * indicates statistical significance at the 5% level. ** indicates statistical significance at the 1% level.

		(1)	(2)		(3)	(4)	(5)
Sample	Sub-periods	Aggregate a_i coefficients	Aggregate d_i coefficients		Change in aggregate a_i coefficients	Change in aggregate d_i coefficients	Overall change in speed of price adjustment
Banned	Pre-ban	-0.0065	0.3868	P2-P1	-0.0095***	-0.0021*	Faster
stocks	Ban	-0.0160	0.3847	(p-value)	(0.0001)	(0.0673)	
Non-banned	Pre-ban	0.0050	0.4049	P2-P1	0.0114***	-0.0242***	Uncertain
stocks	Ban	0.0164	0.3807	(p-value)	(0.0001)	(0.0001)	

5.6.2.2 Stringency of short-selling bans

This sub-section presents results conditional on the stringency of the short-selling bans. Hypothesis 3 asserts that under total short-selling bans, the speed of price adjustment to new information on the banned stocks is slower than that of the banned stocks under naked short-selling bans during the ban period.

Table 5.6 reports the aggregate a_i and d_i coefficient estimates for banned and non-banned stocks at an aggregate level. Panel A of Table 5.6 shows the results for countries that introduced naked short-selling bans while Panel B shows the results for countries which applied total (naked and covered) short-selling bans. Because only three out of 10 countries have a post-ban period available, the analysis in this sub-section focuses on the change in the speed of price adjustment from the pre-ban period to the ban period.

Column (1) of Panels A and B of Table 5.6 shows that the aggregate a_i coefficient estimates are negative in both sub-periods for the banned stocks while they are positive in both sub-periods for the non-banned stocks. As can be seen in column (3) of Panel A, there is no significant change in the aggregate a_i coefficient estimates for the banned stocks whereas there is a significant increase, at the 1% level, in the aggregate a_i coefficient estimates for the non-banned stocks.

Aggregate d_i coefficient estimates are positive in both sub-periods for both samples (see column (2) of Panels A and B of Table 5.6). Column (4) of Panel A of Table 5.6 shows that there is a significant increase in the aggregate d_i coefficient estimates for both samples. These results imply that the speed of price adjustment to new information is slower during the ban period for both samples under naked short-selling bans. However, the speed of price adjustment to new information decreases during the ban period only in terms of stronger positive autocorrelation in trades for the banned stocks and in terms of stronger positive autocorrelation in both quote revisions and trades for the non-banned stocks.

Under total short-selling bans, the speed of price adjustment to new information increases significantly during the ban period for the banned stocks, while the change in the speed of price adjustment to new information is uncertain for the non-banned stocks (see column (5) of Panel B of Table 5.6). For example, the aggregate a_i coefficient estimate for the banned stocks decreases from -0.0022 in the pre-ban period to -0.0130 in the ban period, implying stronger quote revisions in the ban period (see column (1) of Panel B of Table 5.6). The aggregate d_i coefficient estimate for the banned stocks also decreases from 0.4090 in the pre-ban period to 0.3868 in the ban period (see column (2) of Panel B of Table 5.6). Both changes are statistically significant at the 1% level as reported in columns (3) and (4), respectively, of Table 5.6.

These results contradict Hypothesis 3, which predicts a greater increase in both aggregate coefficient estimates under total short-selling bans than under naked short-selling bans. There is no empirical evidence in prior studies showing the impact of the short-selling bans on the speed of price adjustment to new information based on the severity of these bans. However, theory suggests that short-selling prohibitions (total short-selling bans) rather than simply short-selling restrictions (naked short-selling bans) should be associated with slower incorporation of information into prices (see DV (1987)).

Table 5.6 Aggregate a_i and d_i coefficient estimates: Naked versus total short-selling bans

This table reports the median values of the coefficients $\sum_{i=1}^{5} a_i$ and $\sum_{i=1}^{5} d_i$ and their changes from the pre-ban period (P1) to the ban period (P2) for banned and non-banned stocks, respectively, at an aggregate level. a_i coefficient sums reflect the autocorrelation in quote revisions and d_i coefficient sums reflect the autocorrelation in trades. The aggregate coefficients are estimated by using the bivariate VAR model as shown in Chen and Rhee (2010):

$$r_t = \sum_{i=1}^{5} a_i\, r_{t-i} + \sum_{i=0}^{5} b_i\, Q_{t-1} + v_{1,t}$$

$$Q_t = \sum_{i=1}^{5} c_i\, r_{t-i} + \sum_{i=1}^{5} d_i\, Q_{t-1} + v_{2,t}$$

The sample period covers the period from mid-January 2008 to mid-December 2009. Panel A shows the results for countries where only naked short-selling bans were introduced, while Panel B shows the results for countries which introduced total (naked and covered) short-selling bans. The Wilcoxon test is used and the p-values of the Z-statistic are given in the parentheses. * indicates statistical significance at the 10% level. ** indicates statistical significance at the 5% level. *** indicates statistical significance at the 1% level.

Panel A: Naked short-selling bans							
		(1)	(2)		(3)	(4)	(5)
Group	Sub-periods	Aggregate a_i coefficients	Aggregate d_i coefficients		Change in aggregate a_i coefficients	Change in aggregate d_i coefficients	Overall change in speed of price adjustment
Banned	Pre-ban	-0.0142	0.3561	P2-P1	-0.0042	0.0273***	Slower
stocks	Ban	-0.0184	0.3834	(p-value)	(0.4994)	(0.0001)	
Non-banned	Pre-ban	0.0023	0.3547	P2-P1	0.0134***	0.0261***	Slower
stocks	Ban	0.0157	0.3808	(p-value)	(0.0001)	(0.0001)	
Panel B: Total short-selling bans							
Banned	Pre-ban	-0.0022	0.4090	P2-P1	-0.0108***	-0.0221***	Faster
stocks	Ban	-0.0130	0.3868	(p-value)	(0.0001)	(0.0001)	
Non-banned	Pre-ban	0.0081	0.4579	P2-P1	0.0088***	-0.0774***	Uncertain
stocks	Ban	0.0169	0.3806	(p-value)	(0.0001)	(0.0001)	

5.6.2.3 Derivatives on underlying stocks

Table 5.7 shows the results for banned and non-banned stocks, respectively, conditional on the presence and absence of listed derivatives. Panel A of Table 5.7 reports the results for stocks with listed derivatives while Panel B of Table 5.7 presents the results for stocks without listed derivatives.

Hypothesis 4 asserts that presence or absence of derivatives does not affect the impact of short-selling bans on the speed of price adjustment to new information. As can be seen in columns (3) and (4) of Panel A of Table 5.7, the aggregate a_i coefficient estimate of the banned stocks decreases significantly at the 1% level and the aggregate d_i coefficient estimate of the banned stocks decreases marginally at the 5% level. These results imply a faster speed of price adjustment to new information for the banned stocks with listed derivatives during the ban period. The change in the speed of price adjustment to new information for the non-banned stocks with listed derivatives is uncertain during the ban period, as the aggregate a_i coefficient estimate for these stocks significantly increases while the aggregate d_i coefficient estimate significantly decreases. In contrast, columns (3) and (4) of Panel B of Table 5.7 show that there is a slower speed of price adjustment to new information for the banned stocks without listed derivatives in terms of a greater aggregate a_i coefficient estimate during the ban period. However, this result is statistically significant only at the 10% level. Again there is an uncertain change in the speed of price adjustment to information for the non-banned stocks without listed derivatives.

Overall, the results in this sub-section do not support Hypothesis 4. On the contrary, the results show that listing of derivatives affects the impact of short-selling regulations on the speed of price adjustment to new information of the banned stocks. There is an indication that short-selling bans lead to a slower speed of price adjustment to new information for the banned stocks without listed derivatives. However, there is stronger evidence showing that the presence of listed derivatives improves the informativeness of the prices of banned stocks during the ban period. This result supports the findings of Kolasinski, Reed, and Thornock (2013) who find that the informativeness of the prices of banned stocks is better if these stocks have listed options.

Table 5.7 Aggregate a_i and d_i coefficient estimates: Presence versus absence of listed derivatives

This table reports the median values of the coefficients $\sum_{i=1}^{5} a_i$ and $\sum_{i=1}^{5} d_i$ and their changes from the pre-ban period (P1) to the ban period (P2) for banned and non-banned stocks, respectively, at an aggregate level. a_i coefficient sums reflect the autocorrelation in quote revisions and d_i coefficient sums reflect the autocorrelation in trades and are estimated by using the bivariate VAR model as shown in Chen and Rhee (2010):

$$r_t = \sum_{i=1}^{5} a_i\, r_{t-i} + \sum_{i=0}^{5} b_i\, Q_{t-1} + v_{1,t}$$

$$Q_t = \sum_{i=1}^{5} c_i\, r_{t-i} + \sum_{i=1}^{5} d_i\, Q_{t-1} + v_{2,t}$$

The sample period covers the period from mid-January 2008 to mid-December 2009. Panel A shows the results for stocks with listed derivatives, while Panel B shows the results for stocks without listed derivatives. The Wilcoxon test is used and the p-values of the Z-statistic are given in the parentheses. * indicates statistical significance at the 10% level. ** indicates statistical significance at the 5% level. *** indicates statistical significance at the 1% level.

Group	Sub-periods	(1) Aggregate a_i coefficients	(2) Aggregate d_i coefficients		(3) Change in aggregate a_i coefficients	(4) Change in aggregate d_i coefficients	(5) Overall change in speed of price adjustment
Panel A: Stocks with listed derivatives							
Banned stocks	Pre-ban	-0.0121	0.3817	P2-P1	-0.0114***	-0.0013*	Faster
	Ban	-0.0235	0.3803	(p-value)	(0.0001)	(0.0554)	
Non-banned stocks	Pre-ban	0.0022	0.3940	P2-P1	0.0096***	-0.0149***	Uncertain
	Ban	0.0118	0.3792	(p-value)	(0.0001)	(0.0001)	
Panel B: Stocks without listed derivatives							
Banned stocks	Pre-ban	0.0112	0.4103	P2-P1	0.0167*	0.0105	Slower
	Ban	0.0280	0.4208	(p-value)	(0.0565)	(0.1915)	
Non-banned stocks	Pre-ban	0.0208	0.4453	P2-P1	0.0239***	-0.0540***	Uncertain
	Ban	0.0448	0.3914	(p-value)	(0.0001)	(0.0001)	

5.6.4 Summary of the aggregate-level results and discussion

The aggregate-level results show that the speed of price adjustment to new information on the European banned stocks increases significantly while there is uncertain change in the speed of price adjustment to new information for the European non-banned stocks during the ban period. These results contradict Hypotheses 1 and 2 and the findings of prior studies (see for example, Chen and Rhee (2010)). Moreover, there is no evidence showing that the speed of price adjustment to new information is slower under total short-selling bans. This result does not support Hypothesis 3 and contradicts the theoretical prediction of the DV model, which states that the incorporation of information is slower under more severe short-selling constraints. Lastly, empirical evidence suggests that listed derivatives improve the information efficiency of the prices of banned stocks, contradicting Hypothesis 4. However, this result is in line with the findings of Kolasinski, Reed, and Thornock (2013), who also find that information is incorporated faster into the prices of banned stocks with listed options relative to banned stocks without listed options during the ban period. Again, there is an uncertain change in the speed of price adjustment to information for non-banned stocks with and without listed derivatives during the ban period.

Overall, empirical evidence shows that the trading of derivatives written on banned stocks is associated with faster speed of price adjustment to new information on these stocks during the ban period. In contrast, the speed of price adjustment to new information decreases during the ban period for the banned stocks without listed derivatives, as expected. However, this result is only statistically significant at the 10% level.

5.7 Conclusions

This essay examines the impact of the 2008 short-selling bans that were introduced in 10 European countries on the speed of price adjustment to new information on financial stocks subject to these bans. The country-level results show that the speed of price adjustment to new information decreases in only one out of 10 countries, Germany. In contrast, the speed of price adjustment to new information increases during the ban period for three out of 10 countries, Denmark, Spain, and the UK. For the remaining six countries, either the speed of price adjustment to new information does not change (Belgium, Portugal, and Switzerland) or the impact of the short-selling bans on the speed of price adjustment to new information is uncertain (France, Ireland, and the Netherlands). These results contradict Hypotheses 1 and 2.

Furthermore, the unconditional aggregate-level results show that the speed of price adjustment to new information increases during the ban period for the banned stocks. These results contradict the findings of prior studies which report slower speed of price adjustment to new information when short-selling is not allowed (see for example, Bris, Goetzman, and Zhu (2007) and Chen and Rhee (2010)). Conditional on the stringency of the short-selling bans, empirical evidence shows that the speed of price adjustment to new information is faster under total short-selling bans and slower under naked short-selling bans. These results do not support Hypothesis 3 and the theoretical predictions of the DV model, which predicts slower speed of price adjustment to new information under more severe short-selling constraints (i.e., total short-selling bans).

Lastly, the aggregate-level results, conditional on the presence and the absence of listed derivatives, reveal that the speed of price adjustment to new information is faster for banned stocks with listed derivatives and slower for banned stocks without listed derivatives. Although these results do not support Hypothesis 4, they are in line with the findings reported by Kolasinski, Reed, and Thornock (2013). The authors also find that the presence of listed options improves the informativeness of the prices of banned stocks in the US. To sum up, in most of the cases, the change in the speed of price adjustment to new information on the non-banned stocks is uncertain during the ban period. In contrast, there are clearer patterns in the change in the speed of price adjustment to new information for the banned stocks during the same period.

APPENDIX 5A

5A.1 Norway

Table 5A.1 shows the results of the aggregate a_i and d_i coefficient estimates for Norwegian financial banned stocks subject to a total (naked and covered) short-selling ban.[159] Data in column (4) of Panel A of Table 5A.1 show that the median value of the aggregate a_i coefficient estimate increases during the ban period at the 5% significance level (see column (5)). The median value of the aggregate d_i coefficient estimate also increases, as can be seen in column (4) of Panel B of Table 5A.1. This result is significant at the 1% level, as reported in column (5) of Panel B of Table 5A.1. Overall, the speed of price adjustment to new information is slower during the ban period for the Norwegian banned stocks, which supports Hypotheses 1 and 2. In the post-ban period, although both the median value of the aggregate a_i coefficient estimate and the median value of the d_i coefficient estimate decrease, the changes are not statistically significant, as reported in columns (6) and (7), respectively.

[159] Stocks of five Norwegian financial institutions were subject to the short-selling ban (see Appendix 2A). However, due to data insufficiency only two out of five stocks of the above financial institutions were used in the analysis. These financial institutions are DnB NOR ASA and Storebrand ASA.

Table 5A.1 Speed of price adjustment to new information: Norway

This table reports the mean and median values of the coefficients $\sum_{i=1}^{5} a_i$ and $\sum_{i=1}^{5} d_i$ and their changes from the pre-ban period to the ban period and from the ban period to the post-ban period, respectively, for Norwegian banned stocks. a_i coefficient sums reflect the autocorrelation in quote revisions and d_i coefficient sums reflect the autocorrelation in trades. The aggregate coefficients are estimated by using the bivariate VAR model as shown in Chen and Rhee (2010):

$$r_t = \sum_{i=1}^{5} a_i\, r_{t-i} + \sum_{i=0}^{5} b_i\, Q_{t-1} + v_{1,t}$$

$$Q_t = \sum_{i=1}^{5} c_i\, r_{t-i} + \sum_{i=1}^{5} d_i\, Q_{t-1} + v_{2,t}$$

The sample period covers the period from mid-January 2008 to mid-December 2009. Panel A shows the results for the aggregate a_i coefficient estimates, while Panel B shows the results for the aggregate d_i coefficient estimates. The t-test and the Wilcoxon test are used to test for the equality of means and medians. P-values of the t-statistic and the Z-statistic are also provided. * indicates statistical significance at the 10% level. ** indicates statistical significance at the 5% level. *** indicates statistical significance at the 1% level.

Panel A: Aggregate a_i coefficient estimates							
	(1)	(2)	(3)	(4)	(5)	(6)	(7)
	Pre-ban	Ban	Post-ban	Pre-ban vs. ban	P-value	Ban vs. post-ban	P-value
Means	0.0586	0.0780	0.0821	0.0194	0.0681	0.0041	0.8208
Medians	0.0556	0.0813	0.0766	0.0257**	0.0240	-0.0048	0.9043
Panel B: Aggregate d_i coefficient estimates							
Means	0.3740	0.4382	0.4363	0.0643***	0.0001	-0.0020	0.8761
Medians	0.3824	0.4486	0.4420	0.0661***	0.0001	-0.0065	0.5970

5A.2 Italy–A special case

Table 5A.2 reports the results for Italy. Panels A and C of Table 5A.2 show the changes in the mean and median values of the aggregate a_i and d_i coefficient estimates, respectively. The statistical significance of the changes in the mean and median values of the aggregate a_i and d_i coefficient estimates are reported in Panels B and D of the same table respectively. Italy comprises a special case in the analysis as financial regulators in this country changed short-selling rules several times during the ban period. Consequently, the ban period is further divided into three ban sub-periods. In ban period 1, a naked short-selling ban is introduced on all financial stocks traded on Italian stock exchanges. In ban period 2, the short-selling ban becomes more stringent and is applied to both naked and covered short-selling. Finally, in ban period 3, the ban on covered short-selling is relaxed, and only the naked short-selling ban is prohibited. These changes in the stringency of short-selling bans provide an opportunity to examine the changes in the speed of price adjustment to new information within the ban period. The aggregate a_i coefficient estimate is expected to be less negative in ban period 2 and subsequently more negative during the ban period 3. On the other hand, the aggregate d_i

coefficient estimate is expected to be more positive in ban period 2 and then decrease in ban period 3.

As can be seen in Panel A of Table 5A.2, the aggregate a_i coefficient estimate behaves as predicted. For example, the mean aggregate a_i coefficient estimate increases from -0.1117 in the pre-ban period to -0.0953 in ban period 1, then increases to -0.0777 in ban period 2, and finally, decreases to -0.0800 in ban period 3. Nevertheless, the results reported in Panel B show that only the change between ban period1 and ban period 2 is statistically significant. In contrast, the median value of the aggregate a_i coefficient estimate is more negative after the introduction of the ban (decreasing from -0.0577 to -0.0754). This result is statistically significant at the 1% level. The median value of the aggregate a_i coefficient estimate increases in ban period 2 as expected, from -0.0754 to -0.0700. This result is only statistically significant at the 10% level. There is no statistically significant change in the median values of the aggregate a_i coefficient estimates between other sub-periods.

Panel C of Table 5A.2 shows that the mean and median values of the aggregate d_i coefficient estimates increase during ban period 1 and then increase even more in ban period 2. These differences are statistically significant at the 1% level as reported in Panel D of the same table. For example, the mean (median) value of the aggregate d_i coefficient estimate increases significantly from 0.2572 (0.2648) in the pre-ban period to 0.3289 (0.3344) in ban period 1 and then to 0.3702 (0.3619) in ban period 2. The result is not statistically significant in ban period 3. Overall, the results for Italy imply uncertain change in the speed of price adjustment to new information during ban period 1 and slower speed of price adjustment to new information during ban period 2.

Table 5A.2 Speed of price adjustment to new information: Italy

This table reports the mean and median values of the coefficients $\sum_{i=1}^{5} a_i$ and $\sum_{i=1}^{5} d_i$ and their changes from the pre-ban period to the ban period and from the ban period to the post-ban period, respectively, for Italian banned stocks. The ban period is further divided into three ban sub-periods: ban period 1 (introduction of a naked short-selling ban on stocks of all financial firms listed on Italian stock exchanges); ban period 2 (introduction of a total (naked and covered) short-selling ban on stocks of 43 specified financial institutions; and ban period 3 (introduction of a naked short-selling ban on stocks of 43 specified financial institutions). a_i coefficient sums reflect the autocorrelation in quote revisions and d_i coefficient sums reflect the autocorrelation in trades. The aggregate coefficients are estimated by using the bivariate VAR model as shown in Chen and Rhee (2010):

$$r_t = \sum_{i=1}^{5} a_i\, r_{t-i} + \sum_{i=0}^{5} b_i\, Q_{t-1} + v_{1,t}$$

$$Q_t = \sum_{i=1}^{5} c_i\, r_{t-i} + \sum_{i=1}^{5} d_i\, Q_{t-1} + v_{2,t}$$

The sample period covers the period from mid-January 2008 to mid-December 2009. Panel A shows the results for the aggregate a_i coefficient estimates, while Panel C shows the results for the aggregate d_i coefficient estimates. Panels B and D report the statistical significance of the equality of means and medians of aggregate a_i and d_i coefficient estimates, respectively. The t-test and the Wilcoxon test are used and the p-values of their statistics are given in the parentheses. * indicates statistical significance at the 10% level. ** indicates statistical significance at the 5% level. *** indicates statistical significance at the 5% level.

Note: Due to data insufficiency only 39 out of 43 stocks were used in the analysis. The four financial institutions excluded from the analysis are Banca Carige Risp., Banco Sardegna Risp., Banca Popolare Di Spaleto, and Banco Desio Brianza Risp.

Panel A: Aggregate a_i coefficient estimates					
	Pre-ban	Ban 1	Ban 2	Ban 3	Post-ban
Means	-0.1117	-0.0953	-0.0777	-0.0800	-0.0914
Medians	-0.0577	-0.0754	-0.0700	-0.0657	-0.0708
Panel B: Test for equality of means and medians of aggregate a_i coefficient estimates					
	Pre-ban vs. Ban 1 (p-value)	Ban 1vs. Ban 2 (p-value)	Ban 2 vs. Ban 3 (p-value)	Ban 3 vs. Post-ban (p-value)	
Means	0.0164	0.0176*	-0.0023	-0.0113	
	(0.5554)	(0.0576)	(0.8352)	(0.1763)	
Medians	-0.0177**	0.0054*	0.0042	-0.0051	
	(0.0001)	(0.0845)	(0.5988)	(0.1126)	
Panel C: Aggregate d_i coefficient estimates					
	Pre-ban	Ban 1	Ban 2	Ban 3	Post-ban
Means	0.2572	0.3289	0.3702	0.3713	0.3541
Medians	0.2648	0.3344	0.3619	0.3789	0.3641
Panel D: Test for equality of means and medians of aggregate d_i coefficient estimates					
	Pre-ban vs. Ban 1 (p-value)	Ban 1vs. Ban 2 (p-value)	Ban 2 vs. Ban 3 (p-value)	Ban 3 vs. Post-ban (p-value)	
Means	0.0718**	0.0412**	0.0012	-0.0173*	
	(0.0001)	(0.0001)	(0.8680)	(0.0135)	
Medians	0.0696**	0.0275**	0.0170	-0.0148*	
	(0.0001)	(0.0001)	(0.3381)	(0.0222)	

APPENDIX 5B

Tables 5B.1–5B.10 Median values of the coefficient estimates from Hasbrouck's (1991) bivariate VAR model: Country-level results

These tables show the median values of the coefficient estimates aggregated over five lags for banned stocks (Panel A)and non-banned stocks (Panel B), at a country level, using the following bivariate VAR model:

$$r_t = \sum_{i=1}^{5} a_i\, r_{t-i} + \sum_{i=0}^{5} b_i\, Q_{t-1} + v_{1,t}$$

$$Q_t = \sum_{i=1}^{5} c_i\, r_{t-i} + \sum_{i=1}^{5} d_i\, Q_{t-1} + v_{2,t}$$

where a_j coefficient sums estimate autocorrelation in the quote revisions; b_{oj} coefficient estimate measures the contemporaneous correlation between trades and quote-midpoint return; b_j coefficient sums estimate quote adjustments subsequent to each trade; c_j coefficient sums estimate the Granger causality from lagged quote revisions to trades; d_j coefficient sums estimate autocorrelation in trades. The sample period covers the period from mid-January 2008 to mid-December 2009 and it is divided into three sub-periods: the pre-ban period, the ban period, and the post-ban period. The subscript *j* takes the value of 1, 2, and 3 in the pre-ban, ban, and post period, respectively. Countries that have two sub-periods – the pre-ban period and the ban period – are Belgium, Denmark, France, Germany, Ireland, Portugal, and Spain. Countries with a post-ban period available are Switzerland, the Netherlands, and the UK. The Wilcoxon test is used and the p-values of the Z-statistic are given in the parentheses. * indicates statistical significance at the 10% level. ** indicates statistical significance at the 5% level. *** indicates statistical significance at the 1% level.

Table 5B.1 Belgium

Aggregate coefficients	Pre-ban period	Ban period	Pre-ban vs. ban period	Difference	P-value
Panel A: Banned stocks					
a_j	-0.0128	0.0026	$a_2 - a_1$	0.0155	0.1276
b_{0j}	0.0393	0.0583	$b_{02} - b_{01}$	0.0190***	0.0001
b_j	-0.0022	-0.0047	$b_2 - b_1$	-0.0024***	0.0001
c_j	-4.0590	-1.7638	$c_2 - c_1$	2.2952***	0.0001
d_j	0.4278	0.4171	$d_2 - d_1$	-0.0107	0.6761
Panel B: Non-banned stocks					
a_j	0.0270	0.0775	$a_2 - a_1$	0.0504***	0.0001
b_{0j}	0.0328	0.0342	$b_{02} - b_{01}$	0.0014***	0.0008
b_j	-0.0017	-0.0033	$b_2 - b_1$	-0.0016***	0.0001
c_j	-3.9910	-3.0092	$c_2 - c_1$	0.9817***	0.0001
d_j	0.4934	0.4659	$d_2 - d_1$	-0.0274***	0.0001

Table 5B.2 Denmark

Aggregate coefficients	Pre-ban period	Ban period	Pre-ban vs. ban period	Difference	P-value
Panel A: Banned stocks					
a_j	0.0858	0.0478	$a_2 - a_1$	-0.0380***	0.0001
b_{0j}	0.0287	0.0472	$b_{02} - b_{01}$	0.0185***	0.0001
b_j	0.0077	0.0097	$b_2 - b_1$	0.0020**	0.0370
c_j	-0.5964	-0.3213	$c_2 - c_1$	0.2751***	0.0001
d_j	0.2402	0.2541	$d_2 - d_1$	0.0138	0.0655
Panel B: Non-banned stocks					
a_j	0.0749	0.0753	$a_2 - a_1$	0.0004	0.7478
b_{0j}	0.0361	0.0385	$b_{02} - b_{01}$	0.0023	0.1689
b_j	0.0049	0.0036	$b_2 - b_1$	-0.0013	0.6527
c_j	-0.8507	-0.6202	$c_2 - c_1$	0.2305	0.4602
d_j	0.2546	0.2388	$d_2 - d_1$	-0.0157	0.1499

Tables 5B.1–5B.10 Median values of the coefficient estimates from Hasbrouck's (1991) bivariate VAR model: Country-level results

Tables continued from previous page

Table 5B.3 France

Aggregate coefficients	Pre-ban period	Ban period	Pre-ban vs. ban period	Difference	P-value
Panel A: Banned stocks					
a_j	-0.0080	0.0154	$a_2 - a_1$	0.0234***	0.0001
b_{0j}	0.0398	0.0341	$b_{02} - b_{01}$	-0.0056***	0.0025
b_j	-0.0004	-0.0017	$b_2 - b_1$	-0.0014***	0.0001
c_j	-3.9660	-3.1386	$c_2 - c_1$	0.8274***	0.0001
d_j	0.4462	0.4274	$d_2 - d_1$	-0.0188***	0.0001
Panel B: Non-banned stocks					
a_j	0.0030	0.0307	$a_2 - a_1$	0.0277***	0.0001
b_{0j}	0.0369	0.0283	$b_{02} - b_{01}$	-0.0086***	0.0001
b_j	-0.0013	-0.0012	$b_2 - b_1$	0.0001	0.2440
c_j	-3.6577	-3.7966	$c_2 - c_1$	-0.1389***	0.0053
d_j	0.4048	0.4144	$d_2 - d_1$	0.0096*	0.0728

Table 5B.4 Germany

Aggregate coefficients	Pre-ban period	Ban period	Pre-ban vs. ban period	Difference	P-value
Panel A: Banned stocks					
a_j	0.0183	0.0139	$a_2 - a_1$	-0.0044	0.1406
b_{0j}	0.0100	0.0171	$b_{02} - b_{01}$	0.0071***	0.0001
b_j	0.0005	0.0007	$b_2 - b_1$	0.0002	0.8598
c_j	-4.4173	-2.2910	$c_2 - c_1$	2.1263***	0.0001
d_j	0.2476	0.2618	$d_2 - d_1$	0.0141***	0.0059
Panel B: Non-banned stocks					
a_j	0.0207	0.0291	$a_2 - a_1$	0.0084***	0.0060
b_{0j}	0.0097	0.0145	$b_{02} - b_{01}$	0.0048***	0.0001
b_j	0.0003	0.0010	$b_2 - b_1$	0.0007**	0.0001
c_j	-4.4118	-2.2175	$c_2 - c_1$	2.1942***	0.0001
d_j	0.2331	0.2163	$d_2 - d_1$	-0.0168***	0.0002

Table 5B.5 Ireland

Aggregate coefficients	Pre-ban period	Ban period	Pre-ban vs. ban period	Difference	P-value
Panel A: Banned stocks					
a_j	0.0397	0.0188	$a_2 - a_1$	-0.0209***	0.0070
b_{0j}	0.0239	0.0815	$b_{02} - b_{01}$	0.0576***	0.0001
b_j	-0.0003	-0.0126	$b_2 - b_1$	-0.0123***	0.0001
c_j	-1.4575	-0.4080	$c_2 - c_1$	1.0495***	0.0001
d_j	0.2345	0.3400	$d_2 - d_1$	0.1054***	0.0001
Panel B: Non-banned stocks					
a_j	0.0499	0.0955	$a_2 - a_1$	0.0456*	0.0632
b_{0j}	0.0561	0.0435	$b_{02} - b_{01}$	-0.0126***	0.0001
b_j	-0.0047	-0.0048	$b_2 - b_1$	-0.0002	0.6495
c_j	-0.5856	-0.5677	$c_2 - c_1$	0.0179	0.6276
d_j	0.3159	0.3127	$d_2 - d_1$	-0.0032	0.7042

Tables 5B.1–5B.10 Median values of the coefficient estimates from Hasbrouck's (1991) bivariate VAR model: Country-level results

Tables continued from previous page

Table 5B.6 Portugal

Aggregate coefficients	Pre-ban period	Ban period	Pre-ban vs. ban period	Difference	P-value
Panel A: Banned stocks					
a_j	-0.0935	-0.0606	$a_2 - a_1$	0.0329*	0.0896
b_{0j}	0.0866	0.0635	$b_{02} - b_{01}$	-0.0231***	0.0001
b_j	-0.0034	-0.0054	$b_2 - b_1$	-0.0020	0.2203
c_j	-1.5508	-1.1080	$c_2 - c_1$	0.4428***	0.0001
d_j	0.3526	0.3504	$d_2 - d_1$	-0.0023	0.9658
Panel B: Non-banned stocks					
a_j	-0.0012	0.0019	$a_2 - a_1$	0.0031	0.5031
b_{0j}	0.0881	0.0612	$b_{02} - b_{01}$	-0.0269***	0.0001
b_j	-0.0072	-0.0045	$b_2 - b_1$	0.0028**	0.0282
c_j	-0.6845	-0.9305	$c_2 - c_1$	-0.2459**	0.0265
d_j	0.3018	0.3495	$d_2 - d_1$	0.0477***	0.0001

Table 5B.7 Spain

Aggregate coefficients	Pre-ban period	Ban period	Pre-ban vs. ban period	Difference	P-value
Panel A: Banned stocks					
a_j	-0.0642	-0.0865	$a_2 - a_1$	-0.0223**	0.0271
b_{0j}	0.0706	0.0771	$b_{02} - b_{01}$	0.0065	0.8643
b_j	-0.0050	-0.0045	$b_2 - b_1$	0.0005	0.4553
c_j	-2.4344	-1.6629	$c_2 - c_1$	0.7716***	0.0001
d_j	0.3570	0.3482	$d_2 - d_1$	-0.0088**	0.0154
Panel B: Non-banned stocks					
a_j	-0.0380	-0.0335	$a_2 - a_1$	0.0045	0.2842
b_{0j}	0.0537	0.0601	$b_{02} - b_{01}$	0.0064***	0.0001
b_j	-0.0032	-0.0040	$b_2 - b_1$	-0.0008***	0.0061
c_j	-2.3601	-1.7813	$c_2 - c_1$	0.5788***	0.0001
d_j	0.3799	0.3462	$d_2 - d_1$	-0.0336***	0.0001

Table 5B.8 Netherlands

Aggregate coefficients	Pre-ban period	Ban period	Post-ban period	Pre-ban vs. ban period	Difference	P-value	Ban vs. post-ban period	Difference	P-value
Panel A: Banned stocks									
a_j	-0.0304	-0.0551	-0.0080	$a_2 - a_1$	-0.0247***	0.0003	$a_3 - a_2$	0.0471***	0.0001
b_{0j}	0.0548	0.0663	0.0260	$b_{02} - b_{01}$	0.0116***	0.0001	$b_{03} - b_{02}$	-0.0404***	0.0001
b_j	-0.0006	-0.0034	-0.0014	$b_2 - b_1$	-0.0028**	0.0001	$b_3 - b_2$	0.0021**	0.0152
c_j	-4.1007	-2.1074	-4.0828	$c_2 - c_1$	1.9933***	0.0001	$c_3 - c_2$	-1.9754***	0.0001
d_j	0.3740	0.3906	0.3970	$d_2 - d_1$	0.0166**	0.0155	$d_3 - d_2$	0.0064*	0.0682
Panel B: Non-banned stocks									
a_j	0.0018	0.0079	0.0388	$a_2 - a_1$	0.0061**	0.0399	$a_3 - a_2$	0.0309***	0.0001
b_{0j}	0.0461	0.0372	0.0221	$b_{02} - b_{01}$	-0.0089	0.6641	$b_{03} - b_{02}$	-0.0151***	0.0001
b_j	-0.0015	-0.0018	-0.0016	$b_2 - b_1$	-0.0003	0.3065	$b_3 - b_2$	0.0002	0.9135
c_j	-3.1859	-3.1899	-2.8535	$c_2 - c_1$	-0.0040	0.7918	$c_3 - c_2$	0.3364**	0.0352
d_j	0.3494	0.3664	0.3221	$d_2 - d_1$	0.0169	0.1992	$d_3 - d_2$	-0.0443***	0.0001

Tables 5B.1–5B.10 Median values of the coefficient estimates from Hasbrouck's (1991) bivariate VAR model: Country-level results

Tables continued from previous page

Table 5B.9 Switzerland

Aggregate coefficients	Pre-ban period	Ban period	Post-ban period	Pre-ban vs. ban period	Difference	P-value	Ban vs. post-ban period	Difference	P-value
Panel A: Banned stocks									
a_j	-0.0734	-0.0806	0.0295	$a_2 - a_1$	-0.0072*	0.0850	$a_3 - a_2$	0.1100***	0.0001
b_{0j}	0.0497	0.0585	0.0204	$b_{02} - b_{01}$	0.0088***	0.0001	$b_{03} - b_{02}$	-0.0381***	0.0001
b_j	-0.0001	0.0015	0.0012	$b_2 - b_1$	0.0016***	0.0002	$b_3 - b_2$	-0.0003	0.3723
c_j	-5.7913	-2.6268	-1.1810	$c_2 - c_1$	3.1645***	0.0001	$c_3 - c_2$	1.4458***	0.0001
d_j	0.3728	0.3745	0.2365	$d_2 - d_1$	0.0017	0.8746	$d_3 - d_2$	-0.1380***	0.0001
Panel B: Non-banned stocks									
a_j	-0.0247	-0.0427	0.0503	$a_2 - a_1$	-0.0180**	0.0285	$a_3 - a_2$	0.0930***	0.0001
b_{0j}	0.0476	0.0545	0.0202	$b_{02} - b_{01}$	0.0069***	0.0001	$b_{03} - b_{02}$	-0.0343***	0.0001
b_j	-0.0031	-0.0013	-0.0001	$b_2 - b_1$	0.0018***	0.0001	$b_3 - b_2$	0.0012***	0.0061
c_j	-4.4972	-3.2000	-1.8458	$c_2 - c_1$	1.2972***	0.0001	$c_3 - c_2$	1.3542***	0.0001
d_j	0.3371	0.3522	0.2212	$d_2 - d_1$	0.0151	0.2218	$d_3 - d_2$	-0.1310***	0.0001

Table 5B.10 UK

Aggregate coefficients	Pre-ban period	Ban period	Post-ban period	Pre-ban vs. ban period	Difference	P-value	Ban vs. post-ban period	Difference	P-value
Panel A: Banned stocks									
a_j	0.0006	-0.0094	0.0242	$a_2 - a_1$	-0.0100***	0.0001	$a_3 - a_2$	0.0337***	0.0001
b_{0j}	0.0193	0.0418	0.0281	$b_{02} - b_{01}$	0.0225***	0.0001	$b_{03} - b_{02}$	-0.0137***	0.0001
b_j	-0.0004	-0.0038	-0.0029	$b_2 - b_1$	-0.0034***	0.0001	$b_3 - b_2$	0.0009***	0.0034
c_j	-3.2393	-1.1219	-2.9739	$c_2 - c_1$	2.1174***	0.0001	$c_3 - c_2$	-1.8520***	0.0001
d_j	0.4528	0.4345	0.5639	$d_2 - d_1$	-0.0183***	0.0005	$d_3 - d_2$	0.1293***	0.0001
Panel B: Non-banned stocks									
a_j	0.0116	0.0124	0.0901	$a_2 - a_1$	0.0008	0.5903	$a_3 - a_2$	0.0777***	0.0001
b_{0j}	0.0220	0.0403	0.0263	$b_{02} - b_{01}$	0.0183***	0.0001	$b_{03} - b_{02}$	-0.0139***	0.0001
b_j	-0.0004	-0.0007	-0.0026	$b_2 - b_1$	-0.0003***	0.0048	$b_3 - b_2$	-0.0019***	0.0001
c_j	-5.0200	-2.7310	-6.4373	$c_2 - c_1$	2.2890***	0.0001	$c_3 - c_2$	-3.7063***	0.0001
d_j	0.5130	0.4401	0.6630	$d_2 - d_1$	-0.0729***	0.0001	$d_3 - d_2$	0.2229***	0.0001

Tables 5B.11 Median values of the coefficient estimates from Hasbrouck's (1991) bivariate VAR model: Aggregate-level results

These tables show the median values of the coefficient estimates aggregated over five lags for banned stocks (Panel A)and non-banned stocks (Panel B), at an aggregate level, using the following bivariate VAR model:

$$r_t = \sum_{i=1}^{5} a_i\, r_{t-i} + \sum_{i=0}^{5} b_i\, Q_{t-1} + v_{1,t}$$

$$Q_t = \sum_{i=1}^{5} c_i\, r_{t-i} + \sum_{i=1}^{5} d_i\, Q_{t-1} + v_{2,t}$$

where a_j coefficient sums estimate autocorrelation in the quote revisions; b_{oj} coefficient estimate measures the contemporaneous correlation between trades and quote-midpoint return; b_j coefficient sums estimate quote adjustments subsequent to each trade; c_j coefficient sums estimate the Granger causality from lagged quote revisions to trades; d_j coefficient sums estimate autocorrelation in trades. The sample period covers the period from mid-January 2008 to mid-December 2009. The analysis focuses on the pre-ban and the ban periods. The subscript *j* takes the value of 1 and 2in the pre-ban and the ban period, respectively. The Wilcoxon test is used and the p-values of the Z-statistic are given in the parentheses. * indicates statistical significance at the 10% level. ** indicates statistical significance at the 5% level. *** indicates statistical significance at the 1% level.

Aggregate coefficients	Pre-ban period	Ban period	Pre-ban vs. ban period	Difference	P-value
Panel A: Banned stocks					
a_j	-0.0065	-0.0160	$a_2 - a_1$	-0.0095***	0.0001
b_{0j}	0.0340	0.0533	$b_{02} - b_{01}$	0.0193***	0.0001
b_j	-0.0003	-0.0025	$b_2 - b_1$	-0.0022***	0.0001
c_j	-3.3135	-1.5583	$c_2 - c_1$	1.7552***	0.0001
d_j	0.3868	0.3847	$d_2 - d_1$	-0.0021*	0.0673
Panel B: Non-banned stocks					
a_j	0.0050	0.0164	$a_2 - a_1$	0.0114***	0.0001
b_{0j}	0.0348	0.0437	$b_{02} - b_{01}$	0.0089***	0.0001
b_j	-0.0008	-0.0017	$b_2 - b_1$	-0.0009***	0.0001
c_j	-3.4541	-2.1816	$c_2 - c_1$	1.2725***	0.0001
d_j	0.4049	0.3807	$d_2 - d_1$	-0.0242***	0.0001

Tables 5B.12–5B.13 Stringency of short-selling bans

These tables show the median values of the coefficient estimates aggregated over five lags for banned stocks (Panel A) and non-banned stocks (Panel B), conditional on the stringency of the short-selling bans, using the following bivariate VAR model:

$$r_t = \sum_{i=1}^{5} a_i \, r_{t-i} + \sum_{i=0}^{5} b_i \, Q_{t-1} + v_{1,t}$$

$$Q_t = \sum_{i=1}^{5} c_i \, r_{t-i} + \sum_{i=1}^{5} d_i \, Q_{t-1} + v_{2,t}$$

where a_j coefficient sums estimate autocorrelation in the quote revisions; b_{oj} coefficient estimate measures the contemporaneous correlation between trades and quote-midpoint return; b_j coefficient sums estimate quote adjustments subsequent to each trade; c_j coefficient sums estimate the Granger causality from lagged quote revisions to trades; d_j coefficient sums estimate autocorrelation in trades. Table 5B.12 reports the results for countries that introduced naked short-selling bans while Table 5B.13 shows the results for countries that introduced total (naked and covered) short-selling bans. The sample period covers the period from mid-January 2008 to mid-December 2009. The analysis focuses on the pre-ban and the ban periods. The subscript j takes the value of 1and 2 in the pre-ban and the ban period, respectively. The Wilcoxon test is used and the p-values of the Z-statistic are given in the parentheses. * indicates statistical significance at the 10% level. ** indicates statistical significance at the 5% level. *** indicates statistical significance at the 1% level.

Table 5B.12 Naked short-selling bans

Aggregate coefficients	Pre-ban period	Ban period	Pre-ban vs. ban period	Difference	P-value
Panel A: Banned stocks					
a_j	-0.0142	-0.0184	$a_2 - a_1$	-0.0042	0.4994
b_{0j}	0.0442	0.0526	$b_{02} - b_{01}$	0.0084***	0.0001
b_j	-0.0005	-0.0026	$b_2 - b_1$	-0.0021***	0.0001
c_j	-3.3244	-1.9577	$c_2 - c_1$	1.3667***	0.0001
d_j	0.3561	0.3834	$d_2 - d_1$	0.0273***	0.0001
Panel B: Non-banned stocks					
a_j	0.0023	0.0157	$a_2 - a_1$	0.0134***	0.0001
b_{0j}	0.0390	0.0437	$b_{02} - b_{01}$	0.0047***	0.0001
b_j	-0.0008	-0.0021	$b_2 - b_1$	-0.0012***	0.0001
c_j	-3.0620	-2.2706	$c_2 - c_1$	0.7914***	0.0001
d_j	0.3547	0.3808	$d_2 - d_1$	0.0261***	0.0001

Table 5B.13 Total short-selling bans

Aggregate coefficients	Pre-ban period	Ban period	Pre-ban vs. ban period	Difference	P-value
Panel A: Banned stocks					
a_j	-0.0022	-0.0130	$a_2 - a_1$	-0.0108***	0.0001
b_{0j}	0.0267	0.0542	$b_{02} - b_{01}$	0.0275***	0.0001
b_j	-0.0002	-0.0022	$b_2 - b_1$	-0.0020***	0.0001
c_j	-3.2998	-0.9034	$c_2 - c_1$	2.3963***	0.0001
d_j	0.4090	0.3868	$d_2 - d_1$	-0.0221***	0.0001
Panel B: Non-banned stocks					
a_j	0.0081	0.0169	$a_2 - a_1$	0.0088***	0.0001
b_{0j}	0.0306	0.0438	$b_{02} - b_{01}$	0.0131***	0.0001
b_j	-0.0008	-0.0007	$b_2 - b_1$	0.0001	0.7987
c_j	-3.9542	-1.9943	$c_2 - c_1$	1.9599***	0.0001
d_j	0.4579	0.3806	$d_2 - d_1$	-0.0774***	0.0001

Tables 5B.14–5B.15 Derivatives on underlying stocks

These tables show the median values of the coefficient estimates aggregated over five lags for banned stocks (Panel A) and non-banned stocks (Panel B), conditional on the presence or absence of listed derivatives, using the following bivariate VAR model:

$$r_t = \sum_{i=1}^{5} a_i\, r_{t-i} + \sum_{i=0}^{5} b_i\, Q_{t-1} + v_{1,t}$$

$$Q_t = \sum_{i=1}^{5} c_i\, r_{t-i} + \sum_{i=1}^{5} d_i\, Q_{t-1} + v_{2,t}$$

where a_j coefficient sums estimate autocorrelation in the quote revisions; b_{oj} coefficient estimate measures the contemporaneous correlation between trades and quote-midpoint return; b_j coefficient sums estimate quote adjustments subsequent to each trade; c_j coefficient sums estimate the Granger causality from lagged quote revisions to trades; d_j coefficient sums estimate autocorrelation in trades. Table 5B.14 reports the results for stocks with listed derivatives while Table 5B.15 reports the results for stocks without listed derivatives. The sample period covers the period from mid-January 2008 to mid-December 2009. The analysis focuses on the pre-ban and the ban periods. The subscript j takes the value of 1and 2 in the pre-ban and the ban period, respectively. The Wilcoxon test is used and the p-values of the Z-statistic are given in the parentheses. * indicates statistical significance at the 10% level. ** indicates statistical significance at the 5% level. *** indicates statistical significance at the 1% level.

Table 5B.14 Stocks with listed derivatives

Aggregate coefficients	Pre-ban period	Ban period	Pre-ban vs. Ban periods	Difference	P-value
Panel A: Banned stocks					
a_j	-0.0121	-0.0235	$a_2 - a_1$	-0.0114***	0.0001
b_{0j}	0.0312	0.0490	$b_{02} - b_{01}$	0.0178***	0.0001
b_j	-0.0003	-0.0021	$b_2 - b_1$	-0.0018***	0.0001
c_j	-4.0661	-1.9567	$c_2 - c_1$	2.1095***	0.0001
d_j	0.3817	0.3803	$d_2 - d_1$	-0.0013*	0.0554
Panel B: Non-banned stocks					
a_j	0.0022	0.0118	$a_2 - a_1$	0.0096***	0.0001
b_{0j}	0.0304	0.0386	$b_{02} - b_{01}$	0.0082***	0.0001
b_j	-0.0006	-0.0014	$b_2 - b_1$	-0.0007***	0.0001
c_j	-4.1995	-2.7259	$c_2 - c_1$	1.4736***	0.0001
d_j	0.3940	0.3792	$d_2 - d_1$	-0.0149***	0.0001

Table 5B.15 Stocks without listed derivatives

Aggregate coefficients	Pre-ban period	Ban period	Pre-ban vs. Ban periods	Difference	P-value
Panel A: Banned stocks					
a_j	0.0112	0.0280	$a_2 - a_1$	0.0167*	0.0565
b_{0j}	0.0424	0.0854	$b_{02} - b_{01}$	0.0429***	0.0001
b_j	-0.0016	-0.0142	$b_2 - b_1$	-0.0126***	0.0001
c_j	-1.2077	-0.4602	$c_2 - c_1$	0.7475***	0.0001
d_j	0.4103	0.4208	$d_2 - d_1$	0.0105	0.1915
Panel B: Non-banned stocks					
a_j	0.0208	0.0448	$a_2 - a_1$	0.0239***	0.0001
b_{0j}	0.0488	0.0822	$b_{02} - b_{01}$	0.0334***	0.0001
b_j	-0.0028	-0.0067	$b_2 - b_1$	-0.0039***	0.0001
c_j	-1.7531	-0.8998	$c_2 - c_1$	0.8533***	0.0001
d_j	0.4453	0.3914	$d_2 - d_1$	-0.0540***	0.0001

APPENDIX 5C

Tables 5C.1–5C.10 Mean values of the coefficient estimates from Hasbrouck's (1991) bivariate VAR model: Country-level results

These tables show the mean values of the coefficient estimates aggregated over five lags for banned stocks (Panel A)and non-banned stocks (Panel B), at a country level, using the following bivariate VAR model:

$$r_t = \sum_{i=1}^{5} a_i\, r_{t-i} + \sum_{i=0}^{5} b_i\, Q_{t-1} + v_{1,t}$$

$$Q_t = \sum_{i=1}^{5} c_i\, r_{t-i} + \sum_{i=1}^{5} d_i\, Q_{t-1} + v_{2,t}$$

where a_j coefficient sums estimate autocorrelation in the quote revisions; b_{oj} coefficient estimate measures the contemporaneous correlation between trades and quote-midpoint return; b_j coefficient sums estimate quote adjustments subsequent to each trade; c_j coefficient sums estimate the Granger causality from lagged quote revisions to trades; d_j coefficient sums estimate autocorrelation in trades. The sample period covers the period from mid-January 2008 to mid-December 2009 and it is divided into three sub-periods: the pre-ban period, the ban period, and the post-ban period. The subscript *j* takes the value of 1, 2, and 3 in the pre-ban, ban, and post period, respectively. Countries that have two sub-periods – the pre-ban period and the ban period – are Belgium, Denmark, France, Germany, Ireland, Portugal, and Spain. Countries with a post-ban period available are Switzerland, the Netherlands, and the UK. The t-test is used and the p-values of the t-statistic are given in the parentheses. * indicates statistical significance at the 10% level. ** indicates statistical significance at the 5% level. *** indicates statistical significance at the 1% level

Table 5C.1 Belgium

Aggregate coefficients	Pre-ban period	Ban period	Pre-ban vs. Ban periods	Difference	P-value
Panel A: Banned stocks					
a_j	-0.0080	-0.0041	$a_2 - a_1$	0.0039	0.5970
b_{0j}	0.0410	0.0773	$b_{02} - b_{01}$	0.0364***	0.0001
b_j	-0.0035	-0.0092	$b_2 - b_1$	-0.0058***	0.0001
c_j	-4.2658	-2.1633	$c_2 - c_1$	2.1025***	0.0001
d_j	0.4126	0.4179	$d_2 - d_1$	0.0053	0.3666
Panel B: Non-banned stocks					
a_j	0.0258	0.0688	$a_2 - a_1$	0.0429***	0.0001
b_{0j}	0.0374	0.0536	$b_{02} - b_{01}$	0.0162***	0.0001
b_j	-0.0014	-0.0071	$b_2 - b_1$	-0.0057***	0.0001
c_j	-4.3747	-3.4978	$c_2 - c_1$	0.8769***	0.0001
d_j	0.4762	0.4514	$d_2 - d_1$	-0.0248***	0.0002

Table 5C.2 Denmark

Aggregate coefficients	Pre-ban period	Ban period	Pre-ban vs. Ban periods	Difference	P-value
Panel A: Banned stocks					
a_j	0.0771	0.0269	$a_2 - a_1$	-0.0502***	0.0001
b_{0j}	0.0288	0.0524	$b_{02} - b_{01}$	0.0236***	0.0001
b_j	0.0075	0.0117	$b_2 - b_1$	0.0042**	0.0225
c_j	-0.7025	-0.3475	$c_2 - c_1$	0.3550***	0.0001
d_j	0.2302	0.2486	$d_2 - d_1$	0.0184	0.1037
Panel B: Non-banned stocks					
a_j	0.0560	0.0614	$a_2 - a_1$	0.0054	0.6502
b_{0j}	0.0417	0.0555	$b_{02} - b_{01}$	0.0138***	0.0001
b_j	0.0056	0.0059	$b_2 - b_1$	0.0004	0.8951
c_j	-0.7792	-0.7645	$c_2 - c_1$	0.0147	0.8805
d_j	0.2465	0.2282	$d_2 - d_1$	-0.0182	0.1434

Tables 5C.1–5C.10 Mean values of the coefficient estimates from Hasbrouck's (1991) bivariate VAR model: Country-level results

Tables continued from previous page

Table 5C.3 France

Aggregate coefficients	Pre-ban period	Ban period	Pre-ban vs. Ban periods	Difference	P-value
Panel A: Banned stocks					
a_j	-0.0211	0.0146	$a_2 - a_1$	0.0357***	0.0001
b_{0j}	0.0497	0.0605	$b_{02} - b_{01}$	0.0107***	0.0001
b_j	-0.0012	-0.0062	$b_2 - b_1$	-0.0049***	0.0001
c_j	-4.9557	-3.6698	$c_2 - c_1$	1.2859***	0.0001
d_j	0.4404	0.4233	$d_2 - d_1$	-0.0171***	0.0001
Panel B: Non-banned stocks					
a_j	0.0053	0.0385	$a_2 - a_1$	0.0332***	0.0001
b_{0j}	0.0464	0.0440	$b_{02} - b_{01}$	-0.0025*	0.0512
b_j	-0.0031	-0.0045	$b_2 - b_1$	-0.0014*	0.0597
c_j	-4.4043	-5.1896	$c_2 - c_1$	-0.7853***	0.0001
d_j	0.3963	0.4063	$d_2 - d_1$	0.0100**	0.0229

Table 5C.4 Germany

Aggregate coefficients	Pre-ban period	Ban period	Pre-ban vs. Ban periods	Difference	P-value
Panel A: Banned stocks					
a_j	0.0160	0.0109	$a_2 - a_1$	-0.0051	0.2887
b_{0j}	0.0139	0.0264	$b_{02} - b_{01}$	0.0125***	0.0001
b_j	0.0005	-0.0007	$b_2 - b_1$	-0.0012**	0.0144
c_j	-5.5623	-2.6107	$c_2 - c_1$	2.9516***	0.0001
d_j	0.2561	0.2688	$d_2 - d_1$	0.0127***	0.0089
Panel B: Non-banned stocks					
a_j	0.0200	0.0426	$a_2 - a_1$	0.0226***	0.0006
b_{0j}	0.0159	0.0254	$b_{02} - b_{01}$	0.0095***	0.0001
b_j	0.0005	0.0023	$b_2 - b_1$	0.0018**	0.0386
c_j	-4.8384	-2.7469	$c_2 - c_1$	2.0915***	0.0001
d_j	0.2373	0.2209	$d_2 - d_1$	-0.0164***	0.0064

Table 5C.5 Ireland

Aggregate coefficients	Pre-ban period	Ban period	Pre-ban vs. Ban periods	Difference	P-value
Panel A: Banned stocks					
a_j	0.0546	-0.0172	$a_2 - a_1$	-0.0719**	0.0165
b_{0j}	0.0269	0.1102	$b_{02} - b_{01}$	0.0833***	0.0001
b_j	-0.0003	-0.0149	$b_2 - b_1$	-0.0146	0.1104
c_j	-1.7157	-0.5047	$c_2 - c_1$	1.2110***	0.0001
d_j	0.2365	0.3359	$d_2 - d_1$	0.0994***	0.0001
Panel B: Non-banned stocks					
a_j	0.0564	0.0712	$a_2 - a_1$	0.0147	0.6588
b_{0j}	0.0708	0.1123	$b_{02} - b_{01}$	0.0416*	0.0859
b_j	-0.0075	-0.0059	$b_2 - b_1$	0.0016	0.9196
c_j	-0.7771	-0.7732	$c_2 - c_1$	0.0039	0.9746
d_j	0.2901	0.3073	$d_2 - d_1$	0.0173	0.3871

Tables 5C.1–5C.10 Mean values of the coefficient estimates from Hasbrouck's (1991) bivariate VAR model: Country-level results

Tables continued from previous page

Table 5C.6 Portugal

Aggregate coefficients	Pre-ban period	Ban period	Pre-ban vs. Ban periods	Difference	P-value
Panel A: Banned stocks					
a_j	-0.1155	-0.0909	$a_2 - a_1$	0.0246	0.2264
b_{0j}	0.1107	0.0688	$b_{02} - b_{01}$	-0.0419***	0.0001
b_j	-0.0049	-0.0074	$b_2 - b_1$	-0.0025	0.2883
c_j	-1.7524	-1.3064	$c_2 - c_1$	0.4460***	0.0003
d_j	0.3199	0.3320	$d_2 - d_1$	0.0121	0.3469
Panel B: Non-banned stocks					
a_j	-0.0354	-0.0374	$a_2 - a_1$	-0.0020	0.9198
b_{0j}	0.1179	0.0738	$b_{02} - b_{01}$	-0.0441***	0.0001
b_j	-0.0090	-0.0026	$b_2 - b_1$	0.0065**	0.0269
c_j	-0.9571	-1.2063	$c_2 - c_1$	-0.2491**	0.0309
d_j	0.2557	0.3164	$d_2 - d_1$	0.0608***	0.0001

Table 5C.7 Spain

Aggregate coefficients	Pre-ban period	Ban period	Pre-ban vs. Ban periods	Difference	P-value
Panel A: Banned stocks					
a_j	-0.0722	-0.0865	$a_2 - a_1$	-0.0143**	0.0334
b_{0j}	0.0874	0.0966	$b_{02} - b_{01}$	0.0092***	0.0001
b_j	-0.0058	-0.0089	$b_2 - b_1$	-0.0031**	0.0169
c_j	-2.8185	-2.1208	$c_2 - c_1$	0.6977***	0.0001
d_j	0.3546	0.3381	$d_2 - d_1$	-0.0165***	0.0042
Panel B: Non-banned stocks					
a_j	-0.0461	-0.0330	$a_2 - a_1$	0.0131**	0.0358
b_{0j}	0.0628	0.0714	$b_{02} - b_{01}$	0.0086***	0.0001
b_j	-0.0032	-0.0048	$b_2 - b_1$	-0.0016	0.1575
c_j	-2.6473	-2.2553	$c_2 - c_1$	0.3920***	0.0001
d_j	0.3714	0.3448	$d_2 - d_1$	-0.0266***	0.0001

Table 5C.8 Netherlands

Aggregate coefficients	Pre-ban period	Ban period	Post-ban period	Pre-ban vs. Ban periods	Difference	P-value	Ban vs. Post-ban periods	Difference	P-value
Panel A: Banned stocks									
a_j	-0.0469	-0.0803	-0.0136	$a_2 - a_1$	-0.0334***	0.0009	$a_3 - a_2$	0.0667***	0.0001
b_{0j}	0.0620	0.1007	0.0396	$b_{02} - b_{01}$	0.0387***	0.0001	$b_{03} - b_{02}$	-0.0612***	0.0001
b_j	-0.0016	-0.0095	-0.0068	$b_2 - b_1$	-0.0080***	0.0001	$b_3 - b_2$	0.0027	0.2141
c_j	-5.1584	-2.0724	-4.3941	$c_2 - c_1$	3.0861***	0.0001	$c_3 - c_2$	-2.3218***	0.0001
d_j	0.3614	0.3734	0.3933	$d_2 - d_1$	0.0119	0.1670	$d_3 - d_2$	0.0199**	0.0194
Panel B: Non-banned stocks									
a_j	-0.0145	0.0076	0.0363	$a_2 - a_1$	0.0221**	0.0191	$a_3 - a_2$	0.0287***	0.0025
b_{0j}	0.0680	0.0713	0.0386	$b_{02} - b_{01}$	0.0033	0.3890	$b_{03} - b_{02}$	-0.0327***	0.0001
b_j	-0.0038	-0.0062	-0.0057	$b_2 - b_1$	-0.0024	0.2727	$b_3 - b_2$	0.0005	0.7872
c_j	-4.1733	-4.1487	-3.8397	$c_2 - c_1$	0.0246	0.9005	$c_3 - c_2$	0.3090	0.1749
d_j	0.3440	0.3562	0.3247	$d_2 - d_1$	0.0122	0.1812	$d_3 - d_2$	-0.0315***	0.0007

Tables 5C.1–5C.10 Mean values of the coefficient estimates from Hasbrouck's (1991) bivariate VAR model: Country-level results

Tables continued from previous page

Table 5C.9 Switzerland

Aggregate coefficients	Pre-ban period	Ban period	Post-ban period	Pre-ban vs. Ban periods	Difference	P-value	Ban vs. Post-ban periods	Difference	P-value
Panel A: Banned stocks									
a_j	-0.0652	-0.1184	0.0231	$a_2 - a_1$	-0.0532	0.1779	$a_3 - a_2$	0.1415***	0.0001
b_{0j}	0.0492	0.0792	0.0246	$b_{02} - b_{01}$	0.0300***	0.0001	$b_{03} - b_{02}$	-0.0546***	0.0001
b_j	0.0002	0.0192	0.0021	$b_2 - b_1$	0.0189***	0.0004	$b_3 - b_2$	-0.0171***	0.0005
c_j	-5.7735	-2.5774	-1.2082	$c_2 - c_1$	3.1961***	0.0001	$c_3 - c_2$	1.3692***	0.0001
d_j	0.3646	0.3654	0.2455	$d_2 - d_1$	0.0008	0.9136	$d_3 - d_2$	-0.1199***	0.0001
Panel B: Non-banned stocks									
a_j	-0.0043	-0.0301	0.0520	$a_2 - a_1$	-0.0257	0.4646	$a_3 - a_2$	0.0821***	0.0001
b_{0j}	0.0495	0.0603	0.0232	$b_{02} - b_{01}$	0.0107**	0.0001	$b_{03} - b_{02}$	-0.0371***	0.0001
b_j	-0.0040	-0.0032	-0.0007	$b_2 - b_1$	0.0008	0.6953	$b_3 - b_2$	0.0026*	0.0959
c_j	-4.5994	-3.3683	-2.0722	$c_2 - c_1$	1.2311**	0.0001	$c_3 - c_2$	1.2960***	0.0001
d_j	0.3382	0.3499	0.2210	$d_2 - d_1$	0.0117	0.1673	$d_3 - d_2$	-0.1289***	0.0007

Table 5C.10 UK

Aggregate coefficients	Pre-ban period	Ban period	Post-ban period	Pre-ban vs. Ban periods	Difference	P-value	Ban vs. Post-ban periods	Difference	P-value
Panel A: Banned stocks									
a_j	-0.0104	-0.0529	0.0810	$a_2 - a_1$	-0.0425	0.3193	$a_3 - a_2$	0.1339	0.2096
b_{0j}	0.0269	0.0665	0.0432	$b_{02} - b_{01}$	0.0396**	0.0001	$b_{03} - b_{02}$	-0.0233***	0.0001
b_j	-0.0023	-0.0148	-0.0110	$b_2 - b_1$	-0.0125*	0.0220	$b_3 - b_2$	0.0038	0.5426
c_j	-4.3106	-1.7105	-4.5039	$c_2 - c_1$	2.6002**	0.0001	$c_3 - c_2$	-2.7935***	0.0001
d_j	0.4493	0.4417	0.5581	$d_2 - d_1$	-0.0076*	0.0475	$d_3 - d_2$	0.1164***	0.0001
Panel B: Non-banned stocks									
a_j	0.0072	0.0092	0.1015	$a_2 - a_1$	0.0020	0.6297	$a_3 - a_2$	0.0923***	0.0001
b_{0j}	0.0277	0.0541	0.0426	$b_{02} - b_{01}$	0.0264**	0.0001	$b_{03} - b_{02}$	-0.0115***	0.0001
b_j	-0.0022	-0.0073	-0.0090	$b_2 - b_1$	-0.0051**	0.0001	$b_3 - b_2$	-0.0017*	0.0936
c_j	-6.3691	-4.3571	-9.3438	$c_2 - c_1$	2.0120**	0.0001	$c_3 - c_2$	-4.9867***	0.0001
d_j	0.5095	0.4443	0.6339	$d_2 - d_1$	-0.0652**	0.0001	$d_3 - d_2$	0.1896***	0.0001

Tables 5C.11 Mean values of the coefficient estimates from Hasbrouck's (1991) bivariate VAR model: Aggregate-level results

These tables show the mean values of the coefficient estimates aggregated over five lags for banned stocks (Panel A)and non-banned stocks (Panel B), at an aggregate level, using the following bivariate VAR model:

$$r_t = \sum_{i=1}^{5} a_i \, r_{t-i} + \sum_{i=0}^{5} b_i \, Q_{t-1} + v_{1,t}$$

$$Q_t = \sum_{i=1}^{5} c_i \, r_{t-i} + \sum_{i=1}^{5} d_i \, Q_{t-1} + v_{2,t}$$

where a_j coefficient sums estimate autocorrelation in the quote revisions; b_{oj} coefficient estimate measures the contemporaneous correlation between trades and quote-midpoint return; b_j coefficient sums estimate quote adjustments subsequent to each trade; c_j coefficient sums estimate the Granger causality from lagged quote revisions to trades; d_j coefficient sums estimate autocorrelation in trades. The sample period covers the period from mid-January 2008 to mid-December 2009. The analysis focuses on the pre-ban and the ban periods. The subscript *j* takes the value of 1 and 2in the pre-ban and the ban period, respectively. The t-test is used and the p-values of the t-statistic are given in the parentheses. * indicates statistical significance at the 10% level. ** indicates statistical significance at the 5% level. *** indicates statistical significance at the 1% level.

Aggregate coefficients	Pre-ban period	Ban period	Pre-ban vs. ban period	Difference	P-value
Panel A: Banned stocks					
a_j	-0.0220	-0.0384	$a_2 - a_1$	-0.0164*	0.0652
b_{0j}	0.0452	0.0753	$b_{02} - b_{01}$	0.0301***	0.0001
b_j	-0.0018	-0.0067	$b_2 - b_1$	-0.0050***	0.0001
c_j	-4.1563	-2.1691	$c_2 - c_1$	1.9872***	0.0001
d_j	0.3757	0.3705	$d_2 - d_1$	-0.0052**	0.0166
Panel B: Non-banned stocks					
a_j	0.0004	0.0121	$a_2 - a_1$	0.0118***	0.0001
b_{0j}	0.0449	0.0593	$b_{02} - b_{01}$	0.0144***	0.0001
b_j	-0.0024	-0.0042	$b_2 - b_1$	-0.0018***	0.0011
c_j	-4.3188	-3.2609	$c_2 - c_1$	1.0578***	0.0001
d_j	0.3840	0.3678	$d_2 - d_1$	-0.0163***	0.0001

Tables 5C.12–5C.13 Stringency of short-selling bans

These tables show the mean values of the coefficient estimates aggregated over five lags for banned stocks (Panel A) and non-banned stocks (Panel B), conditional on the stringency of the short-selling bans, using the following bivariate VAR model:

$$r_t = \sum_{i=1}^{5} a_i\, r_{t-i} + \sum_{i=0}^{5} b_i\, Q_{t-1} + v_{1,t}$$

$$Q_t = \sum_{i=1}^{5} c_i\, r_{t-i} + \sum_{i=1}^{5} d_i\, Q_{t-1} + v_{2,t}$$

where a_j coefficient sums estimate autocorrelation in the quote revisions; b_{oj} coefficient estimate measures the contemporaneous correlation between trades and quote-midpoint return; b_j coefficient sums estimate quote adjustments subsequent to each trade; c_j coefficient sums estimate the Granger causality from lagged quote revisions to trades; d_j coefficient sums estimate autocorrelation in trades. Table 5C.12 reports the results for countries that introduced naked short-selling bans while Table 5C.13 shows the results for countries that introduced total (naked and covered) short-selling bans. The sample period covers the period from mid-January 2008 to mid-December 2009. The analysis focuses on the pre-ban and the ban periods. The subscript *j* takes the value of 1and 2 in the pre-ban and the ban period, respectively. The t-test is used and the p-values of the t-statistic are given in the parentheses. * indicates statistical significance at the 10% level. ** indicates statistical significance at the 5% level. *** indicates statistical significance at the 1% level.

Table 5C.12 Naked short-selling bans

Aggregate coefficients	Pre-ban period	Ban period	Pre-ban vs. Ban periods	Difference	P-value
Panel A: Banned stocks					
a_j	-0.0326	-0.0367	$a_2 - a_1$	-0.0041	0.2478
b_{0j}	0.0555	0.0747	$b_{02} - b_{01}$	0.0193***	0.0001
b_j	-0.0026	-0.0074	$b_2 - b_1$	-0.0048***	0.0001
c_j	-4.1648	-2.5558	$c_2 - c_1$	1.6090***	0.0001
d_j	0.3574	0.3715	$d_2 - d_1$	0.0141***	0.0001
Panel B: Non-banned stocks					
a_j	-0.0072	0.0085	$a_2 - a_1$	0.0157***	0.0001
b_{0j}	0.0493	0.0575	$b_{02} - b_{01}$	0.0082***	0.0001
b_j	-0.0026	-0.0043	$b_2 - b_1$	-0.0017***	0.0024
c_j	-3.6895	-3.2863	$c_2 - c_1$	0.4032***	0.0001
d_j	0.3470	0.3688	$d_2 - d_1$	0.0218***	0.0001

Table 5C.13 Total short-selling bans

Aggregate coefficients	Pre-ban period	Ban period	Pre-ban vs. Ban periods	Difference	P-value
Panel A: Banned stocks					
a_j	-0.0111	-0.0417	$a_2 - a_1$	-0.0306	0.1516
b_{0j}	0.0346	0.0766	$b_{02} - b_{01}$	0.0420***	0.0001
b_j	-0.0008	-0.0053	$b_2 - b_1$	-0.0045	0.1186
c_j	-4.1476	-1.3904	$c_2 - c_1$	2.7572***	0.0001
d_j	0.3945	0.3685	$d_2 - d_1$	-0.0260***	0.0001
Panel B: Non-banned stocks					
a_j	0.0089	0.0206	$a_2 - a_1$	0.0117**	0.0402
b_{0j}	0.0399	0.0634	$b_{02} - b_{01}$	0.0235***	0.0001
b_j	-0.0021	-0.0040	$b_2 - b_1$	-0.0019	0.1121
c_j	-5.0252	-3.2017	$c_2 - c_1$	1.8236***	0.0001
d_j	0.4256	0.3654	$d_2 - d_1$	-0.0602***	0.0001

Tables 5C.14–5C.15 Derivatives on underlying stocks

These tables show the mean values of the coefficient estimates aggregated over five lags for banned stocks (Panel A) and non-banned stocks (Panel B), conditional on the presence or absence of listed derivatives, using the following bivariate VAR model:

$$r_t = \sum_{i=1}^{5} a_i \, r_{t-i} + \sum_{i=0}^{5} b_i \, Q_{t-1} + v_{1,t}$$

$$Q_t = \sum_{i=1}^{5} c_i \, r_{t-i} + \sum_{i=1}^{5} d_i \, Q_{t-1} + v_{2,t}$$

where a_j coefficient sums estimate autocorrelation in the quote revisions; b_{oj} coefficient estimate measures the contemporaneous correlation between trades and quote-midpoint return; b_j coefficient sums estimate quote adjustments subsequent to each trade; c_j coefficient sums estimate the Granger causality from lagged quote revisions to trades; d_j coefficient sums estimate autocorrelation in trades. Table 5C.14 reports the results for stocks with listed derivatives while Table 5C.15 reports the results for stocks without listed derivatives. The sample period covers the period from mid-January 2008 to mid-December 2009. The analysis focuses on the pre-ban and the ban periods. The subscript j takes the value of 1and 2 in the pre-ban and the ban period, respectively. The t-test is used and the p-values of the t-statistic are given in the parentheses. * indicates statistical significance at the 10% level. ** indicates statistical significance at the 5% level. *** indicates statistical significance at the 1% level.

Table 5C.14 Stocks with listed derivatives

Aggregate coefficients	Pre-ban period	Ban period	Pre-ban vs. Ban periods	Difference	P-value
Panel A: Banned stocks					
a_j	-0.0220	-0.0479	$a_2 - a_1$	-0.0258***	0.0002
b_{0j}	0.0429	0.0655	$b_{02} - b_{01}$	0.0226***	0.0001
b_j	-0.0018	-0.0039	$b_2 - b_1$	-0.0021***	0.0019
c_j	-4.8082	-2.5040	$c_2 - c_1$	2.3041***	0.0001
d_j	0.3696	0.3638	$d_2 - d_1$	-0.0057**	0.0106
Panel B: Non-banned stocks					
a_j	0.0000	0.0094	$a_2 - a_1$	0.0095***	0.0054
b_{0j}	0.0388	0.0496	$b_{02} - b_{01}$	0.0108***	0.0001
b_j	-0.0018	-0.0027	$b_2 - b_1$	-0.0010*	0.0871
c_j	-5.0387	-3.7958	$c_2 - c_1$	1.2429***	0.0001
d_j	0.3781	0.3679	$d_2 - d_1$	-0.0102***	0.0001

Table 5C.15 Stocks without listed derivatives

Aggregate coefficients	Pre-ban period	Ban period	Pre-ban vs. Ban periods	Difference	P-value
Panel A: Banned stocks					
a_j	-0.0218	0.0057	$a_2 - a_1$	0.0274	0.4451
b_{0j}	0.0541	0.1208	$b_{02} - b_{01}$	0.0667***	0.0001
b_j	-0.0015	-0.0195	$b_2 - b_1$	-0.0180***	0.0014
c_j	-1.5836	-0.6195	$c_2 - c_1$	0.9641***	0.0001
d_j	0.3997	0.4012	$d_2 - d_1$	0.0015	0.8122
Panel B: Non-banned stocks					
a_j	0.0018	0.0229	$a_2 - a_1$	0.0211***	0.0029
b_{0j}	0.0651	0.0982	$b_{02} - b_{01}$	0.0331***	0.0001
b_j	-0.0044	-0.0101	$b_2 - b_1$	-0.0057***	0.0004
c_j	-1.9238	-1.1125	$c_2 - c_1$	0.8112***	0.0001
d_j	0.4037	0.3672	$d_2 - d_1$	-0.0365***	0.0001

CHAPTER 6

Essay 4: Short-selling regulations and spot-futures informational dynamics

6.1 Introduction

After the Lehman Brothers collapse in the autumn of 2008, financial authorities banned naked short-selling or both naked and covered short-selling in many countries around the world. This essay examines the effects of the short-selling bans on the spot-futures informational dynamics in Europe. According to Diamond and Verrecchia's (1987) theoretical framework, short-selling constraints in the spot market will reduce the contribution from that market to the price discovery process. Empirical evidence shows that short-selling constraints in the spot market have an impact on the informational linkage between the spot and futures markets (see for example, Fung and Jiang (1999), Jiang, Fung, and Cheng (2001), and Tse and Chan (2010)). This study is closely related to the work of Fung and Jiang (1999) and Jiang, Fung, and Cheng (2001). However, there are two main differences. Firstly, Fung and Jiang (1999) and Jiang, Fung, and Cheng (2001) investigate the impact of relaxing short-selling regulations on the spot-futures informational dynamics in Hong Kong while this essay examines the impact of both enforcing and subsequently relaxing the short-selling regulations on the spot-futures informational dynamics in Europe. Secondly, although this essay also uses a vector error correction model (VECM) as in Fung and Jiang (1999), it focuses on measuring the information share of the spot market in price discovery, as proposed by Hasbrouck (1995), under various short-selling regimes. This information share is measured by the estimated VECM's long-run coefficients. McMillan and Philip (2012) also look at the impact of the short-selling bans on the spot-futures dynamics in Europe at a country level. In contrast, this essay investigates the spot-futures dynamics under short-selling bans at a sector level.

Overall, this essay contributes to the literature by testing Diamond and Verrecchia's (1987) prediction in European spot and futures markets in the following ways. A new data set that was generated by the introduction of the 2008 short-selling bans is used focusing on sector level rather than country level. Furthermore, a different research method is applied to estimate the contribution from the spot market to the price discovery process.

This chapter is organized as follows. Section 6.2 reviews the related literature. Section 6.3 develops the hypotheses. Section 6.4 presents the sample and the data. Section 6.5 describes the methodology. Section 6.6 reports and discusses the results. Section 6.7 summarizes the main findings and concludes.

6.2 Literature Review

6.2.1 Short-selling constraints and the lead/lag relationship between spot and futures markets under good versus bad news

Short-selling constraints have been identified as a potential cause of the lead/lag relationship between spot and futures markets (see for example, Abhyankar (1995)). The rationale behind it is that since there are no short-selling constraints in the futures market, futures prices should reflect good news and bad news symmetrically. If binding short-selling constraints are present in the spot market, futures prices should lead the spot market price to a greater degree when the news is bad.

In one of the first studies to examine this issue, Chan (1992) investigates the lead/lag relationship between spot and futures markets in the US using tests for Granger causality. Firstly, the author sorts the observations by the sign and size of five-minute returns on the Major Market Index (MMI) over the period 1984–1987. Secondly, he regresses these returns on five-minute MMI futures returns and S&P 500 index futures returns conditional on bad or good news. Thirdly, the author compares the regression coefficients of the extreme groups of returns. Periods of bad news (or when MMI returns are negative) and periods of good news (or when MMI returns are positive) proxy for the level of short-selling constraints. The rationale behind this is that during periods of bad news, short-selling constraints are expected to be more binding as selling sentiment prevails. Short sellers profit from price declines, thus, demand for borrowed to-be-sold-short stocks is expected to increase in falling markets, pushing up the lending fee. Furthermore, the author observes that the inability to use short-selling proceeds and the up-tick rule make it more costly for investors with negative information to sell short.[160]

Chan (1992) finds that, overall, the feedback from the futures market to the spot market is greater than the feedback in the opposite direction, and that the feedback strength does not change under bad news compared to good news. In particular, index prices (i) do not lag futures prices longer under bad news and (ii) continue lagging futures prices under good news. The author argues that short-selling restrictions are not a constraint to marginal arbitrageurs, who are able to exploit their information by selling stocks under bad news.

Abhyankar (1995) also examines whether the lead/lag relationship between the FTSE 100 index futures and the underlying spot index differs under good news versus bad news over the period 1986–1990. Abhyankar (1995) sorts hourly returns on the spot index and futures into quartiles based on the size and sign of the spot index returns. Negative returns on the index proxy for bad news while positive returns on the index proxy for good news. Next, the author regresses these spot returns on the lead and lag hourly futures returns using Hansen's (1982) variance-co-variance estimator to generate *t*-ratios consistent in the presence of serial correlation and heteroskedasticity. The results show that neither the futures returns nor the spot returns lead one another in the presence of good or bad news. In contrast, during "moderate" news the futures market is found to lead the spot market. Assuming that short-selling constraints are more binding during periods of bad news, the futures market would be

[160] Under the up-tick rule, the seller is not allowed to sell short the stock if the current price is lower than the previous sale price. The purpose of this rule is to prevent short sellers from inducing prices declines.

expected to lead the spot market under bearish market conditions. However, the results do not support that hypothesis.

Tse and Chan (2010) examine the lead-lag relationship between S&P 500 spot index and the futures contracts written on that index from March 2004 until July 2004. The authors propose a threshold regression model (TRM) where threshold variables proxy for different market conditions and segment the sample data into different linear regimes. More specifically, the basis is used as a threshold variable to examine the effect of short-selling constraints on the lead/lag relationship between the spot and futures markets.[161]

According to the cost-of-carry model, a high basis implies that the futures price is high relative to the spot price, which triggers "cash and carry" arbitrage. In this situation, arbitrageurs would hold short positions in futures and long positions in the spot market to exploit the price discrepancy. A low basis, on the other hand, implies that the futures contract is undervalued compared to the spot price, which triggers a "reverse cash and carry" arbitrage. More precisely, when the basis is low, arbitrageurs would go long on the futures market and go short on the spot market. If short-selling constraints are binding and hamper the arbitrage process, the spot market should adjust more slowly during periods when the basis is low. The results show bidirectional causality between the spot market and the futures market when the basis is high. In contrast, there is only a unidirectional causality running from the futures market to the spot market when the basis is low. This latter result is consistent with the presence of short-selling constraints on the spot market. However, conditional on good or bad news, Tse and Chen (2010) find that the futures market tends to lead the spot market under good, bad, and neutral news. The spot market only appears to significantly lead the futures market when there is no trading direction in the market.

[161] The basis is the difference between the logarithmic futures price and the logarithmic spot price.

Overall, there is mixed evidence on the impact of the short-selling constraints on the lead/lag relationship between the spot and futures markets. All papers discussed in this sub-section examine the lead/lag relationship between the two markets under bad and good news, which are proxied by the negative and positive spot index returns, respectively. The results do not show that the lead of the futures market over the spot market is stronger under bad news. However, when Tse and Chan (2010) use the basis as a proxy for short-selling constraints, empirical evidence shows that there is asymmetry in the lead/lag relationship, which is consistent with the short-selling constraints argument.

6.2.2 The impact of changes in short-selling regulations on the price discovery process between the spot and futures markets

6.2.2.1 Relaxing short-selling regulations

Fung and Jiang (1999) investigate the impact of relaxing short-selling regulations on the dynamic relationship between the Hang Seng index and its related futures contracts. The authors examine the lead/lag relationship between the spot and futures markets employing a VECM and minute-by-minute quote and trade data. The model estimates the speed of adjustment towards the long-run equilibrium and identifies which market plays a more significant role in correcting pricing errors. This is done by examining two sets of coefficients. The coefficients on the lagged terms of changes in spot and futures prices capture the short-run effects and thus identify the direction of the short-run causal relationship. The coefficients on the error correction term represent the long-run dynamics between the two prices and thus describe the price discovery process.

In order to study how the spot-futures lead/lag relationship is affected by the changes in the short-selling regulations, the authors divide the sample period into three sub-periods: (i) a period of no short-selling where short-selling of all stocks is prohibited (period 1), (ii) a period of limited short-selling where only 17 out of 33 constituent stocks of the Hang Seng index can be sold short but must be sold short at prices higher than their prices of previous sale trades (period 2), and (iii) a period of unlimited short-selling where all stocks can be sold short without any restrictions (period 3). For the whole sample period, feedback is found between both markets, although the lead of the futures market is stronger. After short-selling regulations are lifted in period 3, the two markets become more integrated and each market adjusts its prices faster if a change in prices occurs in the other market. The speed of price adjustment to new information in the spot market is higher in period 3 than in period 1. Both results imply that relaxing short-selling regulations improves the information efficiency of spot and futures markets.

The authors also sort the market data by "bearishness" or bullishness" within each sub-period by constructing a relative strength index (RSI). This allows them to investigate any differences in the dynamic relationship between the Hang Seng index and its associated futures contracts in falling and rising markets, respectively. The lowest RSI decile contains observations on the largest downward price movements so it is more likely to be affected by the changes in short-selling regulations. The highest RSI decile corresponds to observations on the largest upward price movements so it is less likely to be affected by the changes in short-selling regulations.

The results indicate that in period 1, when short-selling is banned and the market is bearish, the error correction coefficients in the spot market equation are insignificant, implying that during this period the spot index does not seem to adjust to the long-run equilibrium. Furthermore, the spot market is not found to Granger-cause the futures market. In period 3, however, the spot market is found to adjust to the long-run equilibrium and Granger-cause the futures market under a bearish market. On the other hand, there is no evidence of any significant change in the causality from futures to spot, and vice versa, for all three periods in a rising market.

Overall, the results suggest that when short-selling regulations are lifted, the lead of the spot market over the futures market strengthens, as does the informational linkage between the two markets. These results are found to be significant during falling markets when short-selling regulations are expected to be more binding.

Jiang, Fung, and Cheng (2001) use a different research method to examine the lead/lag relationship between the Hang Seng index futures and its underlying index under different market conditions over the period 1994–1996. The lead/lag relationship between the returns of the Hang Seng Index (HSI) and the HSI futures is estimated conditional on the falling and rising markets over the same three periods employed by Fung and Jiang (1999). However, the authors note that Fung and Jiang (1999) use a VECM that does not incorporate the contemporaneous relationship between the spot and futures markets. To correct for this, Jiang, Fung, and Cheng (2001) use a Sims-style Granger causality test to evaluate the contemporaneous informational linkage between the two markets. Each price of the index futures is matched to its related index quote within a 3-minute interval. Then the authors regress the pre-whitened spot index returns on the contemporaneous, lagged, and lead pre-whitened futures returns.[162]

The authors find that in period 3, when short-selling regulations are lifted the contemporaneous inter-market linkage is strengthened. Also, the lead time of the futures market over the spot market is significantly reduced when (i) falling market conditions prevail and (ii) the futures price is underpriced relative to the spot price. The first result implies that when short-selling regulations are relaxed, information is reflected more rapidly

[162] To pre-whiten index returns, the authors start with an autoregressive model of high order and reduce the autoregressive terms gradually until the residuals are white noise.

in both spot and futures markets and the two markets are more likely to react simultaneously to the same information. The second result is consistent with the view that short-selling constraints are more binding in a falling market and when "reverse cash and carry" arbitrage is impeded.

Overall, the findings of the papers discussed in this sub-section show that relaxing short-selling regulations strengthens the contemporaneous relationship between the spot and futures markets. Furthermore, the results also show that when short-selling is unconstrained, the price discovery process is enhanced in the spot market (i) under falling market conditions when one would expect more selling pressures, and (ii) when arbitrage between the two markets requires going long the futures and going short the spot.

6.2.2.2 The 2008 short-selling bans and the spot-futures dynamics

McMillan and Philip (2012) investigate the impact of the 2008 short-selling bans on the spot-futures dynamics across nine European countries. The authors use data on national indexes and their respective futures over the period January 2006–August 2010.[163] Firstly, the change in basis (i.e., logarithm of the futures price minus the logarithm of the spot price) is investigated from the pre-ban period to the ban period using the cost-of-carry model. The results show that the equilibrium position of the basis decreases during the ban period, which implies that the relative spot price compared to the futures price increases after the introduction of the short-selling bans in all countries with the exception of Germany. The authors argue that the spot market is "overpriced" due to the short-selling constraints.

To estimate the impact of the short-selling bans on the speed of adjustment back to equilibrium or alternatively, the dynamics of equilibrium reversion, the authors use an Augmented Dickey-Fuller equation by adding a ban dummy variable. Overall, the results

[163] These countries are Austria, Belgium, France, Germany, the Netherlands, Portugal, Spain, Switzerland, and the UK.

show that for the majority of the markets the speed of adjustment to equilibrium decreases during the ban period. Also, the authors examine whether the short-selling bans had an impact on the no-arbitrage bands using the quadratic-logistic smooth-transition model. The results reveal that, during the ban period, the lower no-arbitrage bound widens. This suggests that after the introduction of the short-selling bans arbitrageurs found it more costly to exploit futures underpricing via "reverse cash and carry" arbitrage, which involves buying the futures and selling short the spot. Finally, the authors conduct the same analysis by (i) creating value-weighted portfolios where portfolio 1 is comprised of banned stocks and portfolio 2 is comprised of non-banned stocks to control for market movements and (ii) examining national indexes in countries which did not introduce short-selling regulations. The results show that the decrease in the equilibrium position of the basis is greater for the banned stocks relative to the non-banned stocks during the ban period. In contrast, the speed of adjustment to equilibrium is faster for the non-banned stocks relative to the banned stocks. There is no significant change in the equilibrium position of the basis or the speed of adjustment to equilibrium in three the European countries that did not enforce short-selling regulations.[164] These results imply that short-selling bans had an impact on the spot-futures dynamics and in particular on the "reverse cash and carry" arbitrage.

6.2.3 Summary of the literature review

Prior studies that were discussed in section 6.2.1 do not support Diamond and Verrecchia's (1987) prediction that short-selling constraints reduce the information content of trades when there is bad news (see for example, Chan (1992) and Abhyankar (1995)). One interpretation is that the good or bad news criterion may not be the best way of testing the impact of short-selling constraints on information incorporation into prices. In falling markets, for example, investors could establish short positions through derivatives markets if they had a strong

[164] These countries are Finland, Hungary, and Sweden.

negative view about the prospects of a firm. Consequently, sorting the data by good or bad market conditions does not seem to be a good way to capture the impact of short-selling constraints on the lead/lag relationship between the spot and futures markets when there is no explicit change in short-selling regulations. In contrast, partitioning the data by the magnitude of the basis appears to provide greater insight on how short-selling constraints might affect the dynamic relationship between the two markets (Tse and Chan (2010)).

Section 6.2.2 presented papers which directly tested how changes in the severity of short-selling regulations affected the price discovery process between the spot and futures markets. The results show that when short-selling regulations are relaxed, the lead of the spot market over the futures market as well as the contemporaneous relationship between the spot and futures markets strengthen. These results are more pronounced in falling markets and when futures are underpriced (Fung and Jiang (1999); Jiang, Fung, and Cheng (2001)). Finally, McMillan and Philip (2012) find that the 2008 short-selling bans that were introduced in Europe hampered the arbitrage process between the spot and futures markets and reduced the speed of equilibrium reversion of the basis.

Overall, under explicit changes in short-selling regulations, the empirical evidence is consistent with Diamond and Verrecchia's (1987) prediction that short-selling constraints in the spot market reduce the contribution from that market to the price discovery process.

6.3 Hypothesis development

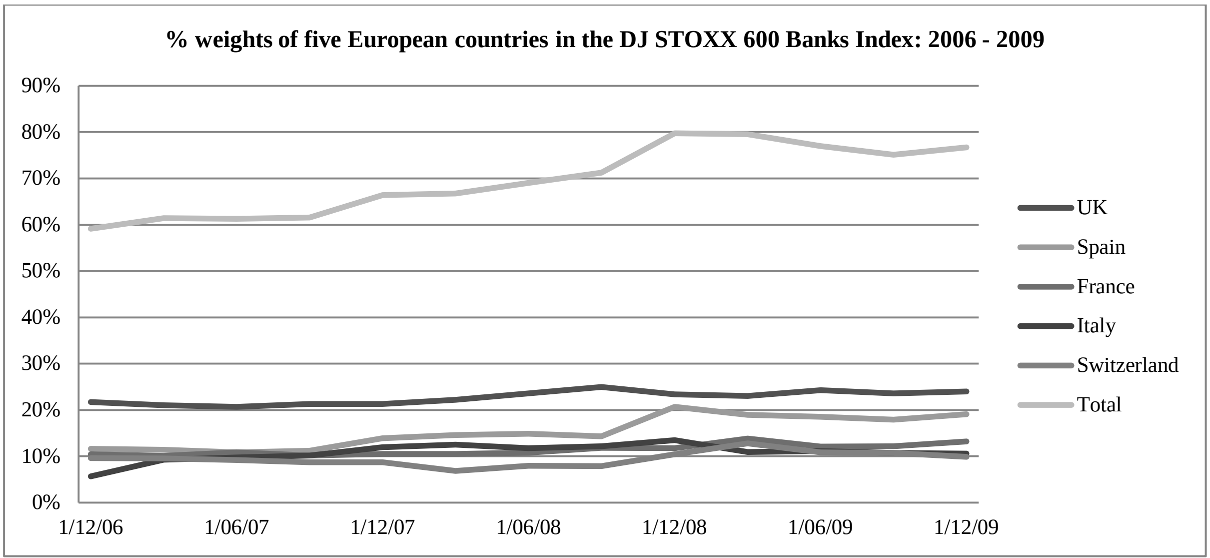
This section develops the hypotheses on the impact of the short-selling regulations on the informational interaction between the futures and spot markets in Europe. More specifically, the informational linkage between the Dow Jones STOXX 600 Banks Index futures (hereafter banks index futures) and its underlying index (hereafter banks index) is examined under various short-selling regimes. Theory predicts that short-selling constraints in the spot market will reduce its contribution to the price discovery process (Diamond and Verrechhia (1987)). Fung and Jiang (1999) and Jiang, Fung, and Cheng (2001) find that relaxing short-selling regulations increases the contribution from the spot market to the price discovery process. In contrast, McMillan and Philip (2012) find that the introduction of the short-selling bans is associated with an impaired "reverse cash and carry" arbitrage process between national spot indexes and their related futures contracts in nine European countries.

As of December 2009, the banks index comprised 55 stocks from 15 European countries.[165] During the sample period of this study, countries that on average have the greatest percentage weight in the index are the UK (24%), Spain (16%), France (11%), Italy (11%), and Switzerland (10%). The average percentage weight in the banks index of the top five countries is greater than 70% (see Figure 6.1). Consequently, the short-selling regulations that were introduced in these countries are expected to have a strong impact on the price discovery process between the banks index futures and the underlying index. More precisely, the UK and Switzerland banned both naked and covered short-selling of specified financial stocks from September 19, 2008 to January 16, 2009. France prohibited the naked short-selling of specified financial stocks on September 22, 2008. Spain introduced a disclosure regime for net short positions of specified financial institutions with economic interest greater than 0.25% of issued share capital on September 24, 2008.[166] Between September 23, 2008 and February 1, 2009, Italy changed its short-selling rules several times, with the most stringent short-selling regime from October 10, 2008 to December 31, 2008, when both

[165] These countries are Austria, Belgium, Denmark, Finland, France, Germany, Greece, Ireland, Italy, Norway, Portugal, Spain, Sweden, Switzerland, and the UK.
[166] Naked short-selling has been banned in this country since 1967.

naked and covered short-selling bans were applied to all the stocks traded on the Italian stock exchanges.[167]

Figure 6.1 Percentage weights of five European countries in the Dow Jones STOXX 600 Banks index through the period 2006–2009



The sample period is divided into two sub-periods: the pre-ban period and the ban period. In the pre-ban period, the percentage of index components affected by some form of short-selling regulations is around 16%, which represents the weight of Spanish stocks in the banks index as Spain has banned naked short-selling since 1967. The ban period is further divided into three sub-periods (see Figure 6.2). In the first ban sub-period, almost all the constituent stocks are subject to some form of short-selling regulations. The only countries that do not ban short-selling are Sweden and Finland, and their average percentage weight in the index is less than 3%. Based on theoretical predictions and prior empirical evidence, this essay hypothesizes that:

Hypothesis 1. The contribution from the banks index to price discovery decreases in the first ban sub-period (Ban period I versus pre-ban period).

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[167] See Appendix 2A for more details on short-selling rules in European countries as provided by Gruenewald, Wagner, and Weber (2009b) and updated by the author of this thesis.

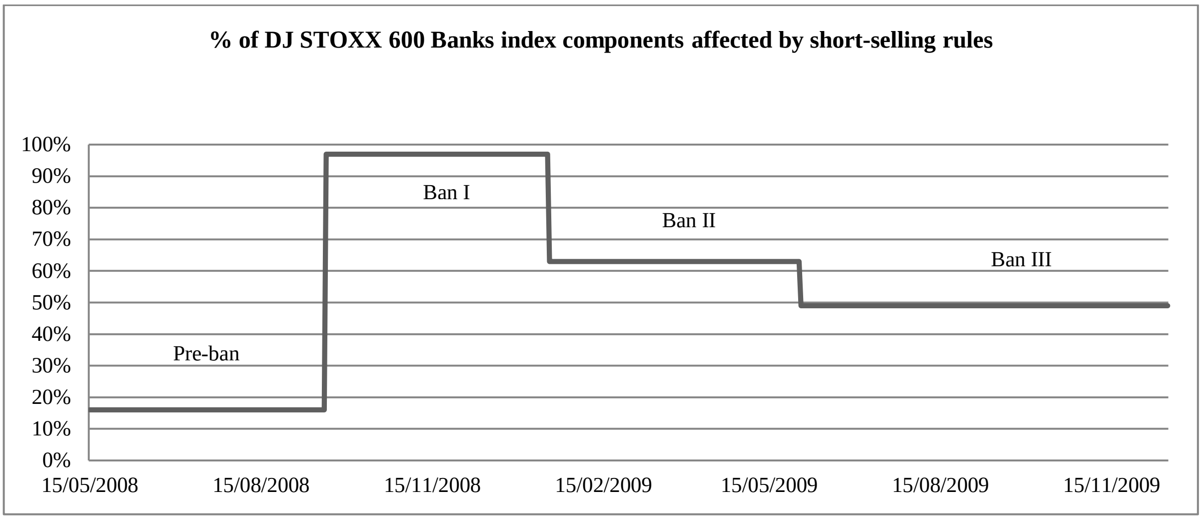
In the second ban sub-period, the UK and Switzerland, with an average percentage weight of 34% in the index, lift their short-selling bans. This means that during the second ban sub-period, 37% of the index components are allowed to be sold short as opposed to 3% in the first ban sub-period. Consequently, this essay hypothesizes that:

Hypothesis 2. The contribution from the banks index to price discovery increases in the second ban sub-period (Ban period II versus ban period I).

In the third ban sub-period two more countries, Greece and Italy, lift their short-selling bans. Their average percentage weight in the index is nearly 14%. Thus, this essay expects that:

Hypothesis 3. The contribution from the banks index to price discovery increases in the third ban sub-period (Ban period III versus ban period II).

Figure 6.2 Percentage of the Dow Jones STOXX 600 Banks index components covered by short-selling rules in the three ban sub-periods



6.4 Sample and Data

The sample period covers the period from January 15, 2008 to December 15, 2009 and it is divided into two sub-periods i.e., (i) the pre-ban period of January 15, 2008–September 18, 2008, and (ii) the ban period of September 19, 2008–December 15, 2009. To account for the changes in the short-selling regulations, the ban period is further divided into three sub-periods i.e., (i) the first ban sub-period of September 19, 2008–January 16, 2009 (Ban I), (ii) the second ban sub-period of January 19, 2009–May 29, 2009 (Ban II), and (iii) the third ban sub-period of June 1, 2009–December 15, 2009 (Ban III).

The Dow Jones STOXX 600 indexes family provides 19 narrow-based super-sector indexes including the banks super-sector index. High-frequency quote and trade data are available for 16 super-sector indexes and are obtained from from Thomson Reuters Tick History (TRTH) database of Securities Industry Research Centre of Asia-Pacific (SIRCA). This essay also examines the price discovery process between other Dow Jones STOXX 600 super-sector indexes and their respective futures contracts. These super-sector indexes are used as a control sample because the majority of their components were not affected by the short-selling rules. Thus, it would be interesting to examine the differences (if any) in their price discovery processes.

As already noted in section 6.3, the Dow Jones STOXX 600 banks index comprises the stock prices of 55 European banks. This index is a sub-index of the STOXX Europe 600 Index, which covers the largest 600 European stocks.[168] The Dow Jones STOXX 600 banks index was introduced on June 15, 1998 to gauge the performance of European bank stocks. The base value of the banks index is 100 as of December 1991. Each bank stock's weight in the index depends on its free float market capitalization. The composition of the banks index is reviewed quarterly in March, June, September, and December and the index is updated every 15 seconds during local trading hours.[169]

[168] Information regarding STOXX indexes was obtained from www.stoxx.com.
[169] The number of constituent stocks in the banks index varies between 54 and 62 during the sample period. This information is obtained from TRTH of SIRCA database (www.tickhistory.thomsonreuters.com)

A futures contract written on the Dow Jones STOXX 600 banks index is traded on EUREX. This contract, which is cash settled with payment on the first exchange day following the final settlement day, is known as the Dow Jones STOXX 600 banks futures contract. Moreover, the contract unit per index point is €50 and the minimum price change is 0.1 index points. The last trading day and the final settlement day is the third Friday of each maturity month if this is an exchange day; otherwise it is the exchange day immediately preceding that day. The futures contracts follow the March/June/September/December quarterly expiration cycle. The daily settlement price for the current maturity month is calculated from the volume-weighted average of the prices of all transactions during the last minute before 17:30 CET (Central European Time), provided that more than five trades are transacted within this period. The final settlement price is established by EUREX on the final settlement day and is derived from the average of the banks index values calculated between 11:50am and 12:00pm CET.[170] Additionally, the index futures contracts are traded electronically, and its underlying spot index is also reported electronically. The trading hours for the banks index are 9:00am–6:00pm (CET), while the related futures contracts trade from 7:50am to 10:00pm (CET).

To test the hypotheses 1 to 3, time series of spot index and futures returns at 15-second intervals are constructed, resulting in approximately 1.5 million returns per series. The 15-second frequency has been chosen because it matches the highest reporting frequency for the spot index data. Midpoint quotes of the nearby Dow Jones STOXX 600 banks index futures contract and the value of the underlying index are used to calculate 15-second returns. The value of the index is constructed from the transaction prices of its constituent stocks; thus, any new information that arrives to the market and manifests itself through trading in index components is recorded and made known to the public every 15 seconds.

[170] See www.eurexchange.com.

A VECM is estimated using the 15-second returns data from the common trading hours from 9:30am to 5:30pm CET. The first 30 minutes as well as the last 30 minutes of the trading day are excluded from the analysis as data errors in the time series can be observed during the opening and closing hours of the market.[171]

6.5 Methodology

Studies of price discovery assume that prices of a financial asset that is traded in multiple markets are driven by the same underlying information. These prices are linked by one or more linear arbitrage relationships. Any divergence from these price relationships will trigger arbitrage trading, which will bring the prices back to their long-run equilibrium (see for example, Hasbrouck (1995)).

There are two common factor models widely used in the literature, the permanent-transitory model presented by Gonzalo and Granger (1995) and the information share model proposed by Hasbrouck (1995). Both models estimate the contribution from each market to the price discovery process.[172] Before proceeding to the price discovery process analysis, these studies examine whether there is a long-run stable relationship between the spot and futures prices.[173] Firstly, a non-stationarity test on individual series is employed. Then, co-integration tests are carried out to investigate whether the two I(1) price series are co-integrated. The next step is to estimate a VECM and use the long-run estimated coefficients to measure the contribution from each market to the price discovery process.

Hasbrouck (1995) and Gonzalo and Granger (1995) define price discovery in different ways. Hasbrouck (1995) assumes that price volatility reflects the flow of information and attributes

[171] SIRCA data appeared to have technical errors. All STOXX index values and futures values that generated returns in excess of 2.5% were deleted. STOXX index values that were lower (greater) than their reported daily low (high) prices were also removed.
[172] Baillie, Booth, Tse, and Zabotina (2002), De Jong (2002), and Lehmann (2002) examine the relation between the two models.
[173] See for example Quan (1992).

a greater share of the efficient price discovery to the market that contributes the greatest share to this volatility.[174] On the other hand, Gonzalo and Granger (1995) decompose the common factor itself. The model attributes the leading role solely to the market that adjusts least to the price movement in the other markets and ignores the correlation among these markets.

To test hypotheses 1 to 3, the following VECM is estimated[175]:

$$\Delta F_t = \varphi_F + \gamma_F u_{t-1} + \sum_{i=1}^{n} \delta_{1i}\, \Delta S_{t-1} + \sum_{i=1}^{n} \lambda_{1i}\, \Delta F_{t-i} + z_{F,t} \tag{6.1}$$

$$\Delta S_t = \varphi_S + \gamma_S u_{t-1} + \sum_{i=1}^{n} \delta_{2i}\, \Delta F_{t-1} + \sum_{i=1}^{n} \lambda_{2i}\, \Delta S_{t-i} + z_{S,t} \tag{6.2}$$

where F_t is the logarithm of the futures price and S_t is the logarithm of the index value; φ_F and φ_S are the constants; $\Delta F_t = F_t - F_{t-1}$ is the current return on the futures contracts; $\Delta S_t = S_t - S_{t-1}$ is the current return on the underlying index; $\sum_{i=1}^{n} \delta_{1i}\, \Delta S_{t-1}$ are the past returns on the stock index that explain the current return on the futures market; $\sum_{i=1}^{n} \lambda_{1i}\, \Delta F_{t-i}$ are the past returns on the futures market that explain the current return on the futures market (i.e., account for persistence in the futures market); $\sum_{i=1}^{n} \delta_{2i}\, \Delta F_{t-1}$are the past returns on the futures market that explain the current return on the stock market; $\sum_{i=1}^{n} \lambda_{2i}\, \Delta S_{t-i}$are the past returns on the stock market that explain the current return on the stock market (i.e., account for persistence in the spot market); $z_{F,t}\ and\ \ z_{S,t}$ are zero-mean and serially uncorrelated innovations in the futures and the stock market, respectively; $u_{t-1} = [(F_{t-1} - S_{t-1}) - E(F_{t-1} - S_{t-1})]$ is the error correction term, which represents the long-run dynamics between the two prices and describe the price discovery process; $E(F_{t-1} - S_{t-1})$is the cost of carry; $\gamma_F\ and\ \gamma_S$ are the error correction coefficients in the futures and the stock market, respectively, which estimate the speed of adjustment back to equilibrium.

[174] See Baillie, Booth, Tse, and Zabotina (2002: p.320).
[175] See Fung and Jiang (1999).

When the innovations in the two markets are uncorrelated the information share (IS) is given by[176]:

$$IS_j = \frac{\gamma_j^2 \sigma_j^2}{\gamma_F^2 \sigma_F^2 + \gamma_S^2 \sigma_S^2} \tag{6.3}$$

for *j=S, F*

If price innovations are correlated, Hasbrouck (1995) uses the Cholesky factorization of the covariance matrix to establish upper and lower bounds.[177] The lower the correlation, the closer the upper and lower bounds. Hasbrouck (1995) argues that most of the contemporaneous correlation in practical applications is due to time aggregation. Using 15-second intervals to measure returns on the stock index and index futures contracts in this paper is an attempt to ameliorate the correlation issue.

The upper and lower bounds of the spot index contribution to price discovery are defined as[178]:

$$IS_U = \frac{\left(\gamma_F\, \sigma_S - \gamma_S \frac{\sigma_{FS}}{\sigma_S}\right)^2}{\gamma_F^2 \sigma_S^2 - 2\gamma_S \gamma_F \sigma_{FS} + \gamma_S^2 \sigma_F^2} \tag{6.4}$$

$$IS_L = \frac{\gamma_F^2 \left(\sigma_S^2 - \frac{\sigma_{FS}^2}{\sigma_F^2}\right)}{\gamma_F^2 \sigma_S^2 - 2\gamma_S \gamma_F \sigma_{FS} + \gamma_S^2 \sigma_S^2} \tag{6.5}$$

The sum of the futures and spot market information shares equals one.

[176] See Amatatsu and Baba (2008).
[177] Baillie, Booth, Tse, and Zabotina (2002) show that the mean of the upper and lower bound estimates provides a good estimate of a market's share to price discovery.
[178] See Amatatsu and Baba (2008).

6.6 Empirical results

6.6.1 Non-stationarity and co-integration tests

Before proceeding to the analysis of the price discovery process, it is essential to examine whether there is a long-run stable relationship between the spot and futures prices.[179] Firstly, the Augmented Dickey-Fuller (ADF) non-stationarity test is employed. Table 6.1 shows that the logarithmically transformed levels of both the spot index and index futures time series are non-stationary for all the Dow Jones STOXX 600 super-sector indexes. The null hypothesis of a unit root in either the spot or futures time series cannot be rejected for any sector. The statistics in Table 6.1 also show that their first differences, or return time series, are stationary. The null hypothesis of a unit root in the difference in either the spot or futures time series is strongly rejected for all sectors. Next, the Johansen test is used to investigate whether the time series, which are both non-stationary of order one, I(1), are co-integrated. The null hypothesis that there are no co-integrating vectors is rejected in favour of the alternative hypothesis that there is at least one co-integrating vector (see Table 6.2).

[179] See for example Quan (1992).

Table 6.1 Augmented Dickey-Fuller non-stationarity test statistics

This table reports the Augmented Dickey-Fuller non-stationarity test statistics for all the Dow Jones STOXX 600 super-sector indexes. S is the logarithm of the index value and F is the logarithm of the index futures price. ΔS is the first difference of the logarithmically transformed index value and ΔF is the first difference of the logarithmically transformed index futures price. The p-values are given in the parentheses. The null hypothesis is that the two time series are non-stationary. * indicates statistical significance at the 10% level. ** indicates statistical significance at the 5% level. *** indicates statistical significance at the 1% level.

	Levels		First differences (returns)		Basis
Time series	S	F	ΔS	ΔF	F-S
Dow Jones STOXX 600 super-sectors			ADF test statistic (p-value)		
Banks	-1.368 (0.5991)	-1.113 (0.7122)	-276.634*** (0.0001)	-96.598*** (0.0001)	-3.193** (0.0209)
Automobiles & Parts	-1.776 (0.3928)	-0.7922 (0.8208)	-104.556*** (0.0001)	-99.879*** (0.0001)	-5.261*** (0.0001)
Basic Resources	-1.801 (0.3801)	-1.487 (0.5399)	-410.321*** (0.0001)	-446.291*** (0.0001)	-8.671*** (0.0001)
Chemicals	-1.299 (0.6320)	-1.073 (0.7285)	-436.363*** (0.0001)	-105.838*** (0.0001)	-4.009*** (0.0014)
Constructions & Materials	-0.6131 (0.8654)	-0.9090 (0.7861)	-87.426*** (0.0001)	-102.194*** (0.0001)	-3.996*** (0.0014)
Financial Services	-0.5567 (0.8776)	-0.8320 (0.8095)	-93.186*** (0.0001)	-100.843*** (0.0001)	-5.718*** (0.0001)
Food & Beverage	-1.309 (0.6276)	-1.181 (0.6850)	-594.400*** (0.0001)	-106.557*** (0.0001)	-5.759*** (0.0001)
Health Care	-1.745 (0.4083)	-0.946 (0.7738)	-545.016*** (0.0001)	-103.664*** (0.0001)	-5.336*** (0.0001)
Industrial Goods & Services	-1.606 (0.4794)	-1.0389 (0.7414)	-313.633*** (0.0001)	-101.811*** (0.0001)	-6.224*** (0.0001)
Insurance	-1.509 (0.5293)	-1.012 (0.7511)	-355.510*** (0.0001)	-101.946*** (0.0001)	-6.045*** (0.0001)
Media	-1.454 (0.5568)	-1.593 (0.4857)	-352.157*** (0.0001)	-110.0947*** (0.0001)	-4.172*** (0.0007)
Oil & Gas	-1.912 (0.3269)	-1.143 (0.7005)	-383.848*** (0.0001)	-103.315*** (0.0001)	-4.545*** (0.0001)
Personal & Household Goods	-0.859 (0.8013)	-1.277 (0.6421)	-378.535*** (0.0001)	-103.535*** (0.0001)	-6.123*** (0.0001)
Technology	-1.737 (0.4121)	-0.5608 (0.8767)	-585.722*** (0.0001)	-102.340*** (0.0001)	-6.513*** (0.0001)
Travel & Leisure	-0.879 (0.7953)	-1.462 (0.5527)	-268.935*** (0.0001)	-99.159*** (0.0001)	-3.435*** (0.0098)
Utilities	-0.8834 (0.7941)	-0.5421 (0.8806)	-583.278*** (0.0001)	-106.201*** (0.0001)	-4.630*** (0.0001)

Table 6.2 Johansen co-integration test statistics

This table reports the Trace statistic from the Johansen co-integration test for all the Dow Jones STOXX 600 super-sector indexes. The null hypothesis is that the two time series are not co-integrated. * indicates statistical significance at the 10% level. ** indicates statistical significance at the 5% level. *** indicates statistical significance at the 1% level.

Dow Jones STOXX 600 super-sectors	Trace statistic	p-value
Banks	51.082***	0.0001
Automobiles & Parts	36.352***	0.0001
Basic Resources	95.364***	0.0001
Chemicals	30.602***	0.0001
Constructions & Materials	31.673***	0.0001
Financial Services	69.573***	0.0001
Food & Beverage	54.530***	0.0001
Health Care	34.885***	0.0001
Industrial Goods & Services	74.477***	0.0001
Insurance	45.310***	0.0001
Media	24.959***	0.0001
Oil & Gas	33.238***	0.0001
Personal & Household Goods	72.899***	0.0001
Technology	57.605***	0.0001
Travel & Leisure	31.271***	0.0001
Utilities	35.510***	0.0001

6.6.2 The information share of the Dow Jones STOXX 600 Banks index

Table 6.3 reports the descriptive statistics of the Hasbrouck's (1995) information shares of the Dow Jones STOXX 600 Banks index in the price discovery process and provides the statistics of the tests for equality of means and medians between different sub-periods.

As shown in Panel A of Table 6.3, the mean (median) banks index's (spot market) contribution to price discovery is 68.15% (72.84%) during the pre-ban period. The information share of the banks index decreases slightly to 65.78% (72.26%) in ban period I; however, the difference is not statistically significant (Panel B of Table 6.3). This result does not support Hypothesis 1. Also, the data in Panel A of Table 6.3 show that the mean spot market's information share increases to 77.23% (86.17%) in ban period II, as some countries with significant weights in the index relax their short-selling regulations. This result is statistically significant at the 1% level as reported in Panel B of Table 6.3 and it supports Hypothesis 2. Lastly, data in Panel A of Table 6.3 show that in ban period III, where other countries abandon stringent short-selling rules, the mean spot information share decreases dramatically to 59.38% (64.08%). Panel B of Table 6.3 shows that this decrease is statistically significant at the 1% level. This result is contrary to Hypothesis 3, according to which the contribution from the banks index to price discovery is expected to increase in the third ban sub-period.

Table 6.3 Information Shares (IS) of the Dow Jones STOXX 600 Banks index

Panel A of Table 6.3 shows the descriptive statistics of the Hasbrouck's (1995) Information Shares (IS) for the Dow Jones STOXX 600 Banks index. The IS are measured using 15 second returns and 10 lags. The whole sample period spans from 15/01/2008 to 15/12/2009 and it is divided into four sub-periods: the pre-ban period (September 19, 2008–December 15, 2009), ban I period (September 19, 2008–January 16, 2009), ban II period (January 19, 2009–May 29, 2009), and ban III period (June 1, 2009–December 15, 2009). Observations refer to the number of days included in each sub-period. Panel B shows the statistical significance of the test for equality of means and medians of the IS between different sub-periods. T-test is used to test the significance of equality of means while Wilcoxon test is used to test the significance of equality of medians. P-values are given in the parentheses below the differences. *indicates statistical significance at the 10% level. ** indicates statistical significance at the 5% level. *** indicates statistical significance at the 1% level.

Panel A: Descriptive Statistics							Panel B: Tests for equality of means and medians		
Sector: Banks									
Period	Mean	Median	Maximum	Minimum	Standard Deviation	Observations	Difference (Period)	Difference (Means)	Difference (Medians)
Pre-ban	0.6815	0.7284	0.9999	0.0204	0.2409	171	Ban I vs. pre-ban	-0.0237 (0.476)	-0.0058 (0.5445)
Ban I	0.6578	0.7226	0.9918	0.0031	0.2571	81	Ban II vs. ban I	0.1145*** (0.0020)	0.1391*** (0.0011)
Ban II	0.7723	0.8617	0.9987	0.0573	0.2225	91	Ban III vs. ban II	-0.1785*** (0.0001)	-0.2209*** (0.0001)
Ban III	0.5938	0.6408	0.9917	0.0982	0.2357	142	Ban III vs. pre-ban	-0.0877*** (0.0013)	-0.0876*** (0.0008)

In summary, the results show that there was no significant migration of informed traders to the futures market for the European banks super-sector immediately after the introduction of the short-selling bans on the stocks of major European banks. These results do not support Diamond and Verrecchia's (1987) theoretical prediction of a reduced contribution from the spot market to the price discovery process due to the short-selling constraints. Nor are these results in line with the findings of prior studies, which find that short-selling regulations in the spot market hamper the contribution from that market to price discovery (see for example, Fung and Jiang (1999) and Jiang, Fung, and Cheng (2001)).

Li (2009) shows that under less volatile market conditions the futures price leads the spot price in the price discovery process. On the other hand, under highly volatile market conditions, the spot market appears to be more informationally efficient than the futures market. Short-selling regulations were introduced during extremely volatile market conditions, which could potentially explain the lead of the spot market over the futures market in the price discovery process during the first ban sub-period. Furthermore, prior studies document increased transaction costs in trading derivatives written on stocks subject to the short-selling bans in the US (Grundy, Lim, and Verwijmeren (2012); Battalio and Schultz (2011)). In Europe, additional transaction costs could be associated with the introduction of short-selling bans in both spot and derivatives markets. For example, some countries such as the UK, France, and Spain banned short-selling in spot and derivatives markets, while other countries such as Germany, Italy, and Switzerland banned short-selling only in the spot market.[180] Consequently, the imposition of similar short-selling regulations and higher transaction costs in the futures market could also explain why there was no migration of informed traders from the spot market to the futures market in the first ban sub-period.

[180] See Chapter 2 for a more comprehensive description of the short-selling regulations that were introduced in European countries.

In the third ban sub-period, the results again contradict the findings of prior studies. Fung and Jiang (1999) and Jiang, Fung, and Cheng (2001)) find that the contribution from the spot market to the price discovery process increases when short-selling regulations are lifted in Hong Kong. In contrast, this essay finds that the contribution from the spot market to price discovery decreases significantly after the relaxation of the short-selling bans. These results suggest that there could be other factors that affected the informational dynamics between the spot and futures markets in Europe, which had a stronger impact on these dynamics than the short-selling bans. To examine that possibility, this essay looks at other super-sector indexes of the Dow Jones STOXX 600 family during the same sub-periods. Most of the components of these indexes were not subject to short-selling regulations.

6.6.3 The information shares of the Dow Jones STOXX 600 super-sector indexes

Figures 6.3 and 6.4 depict the mean and median percentage information shares, respectively, for 16 super-sector indexes of the Dow Jones STOXX 600 family in different sub-periods. More detailed results are reported on a sector-by-sector basis in Table 6.4. A common trend is apparent in all indexes except for Basic Resources. The results show a significant increase in the spot market's contribution to the price discovery process between the months of January and May, 2009 and a significant decrease in the spot market's contribution to the price discovery process between the months of June and December, 2009. These changes mirror those of the bank sector.

Figure 6.3 Mean values of Hasbrouck's (1995) spot market information shares in percentage terms

This figure illustrates the mean values of the Hasbrouck's (1995) spot market information shares across 16 Dow Jones STOXX 600 super-sector indexes. The sample period spans from January, 15 of 2008 to December 15, of 2009 and it is divided into four sub-periods: the pre-ban period (September 19, 2008–December 15, 2009), ban I period (September 19, 2008–January 16, 2009), ban II period (January 19, 2009–May 29, 2009), and ban III period (June 1, 2009–December 15, 2009).

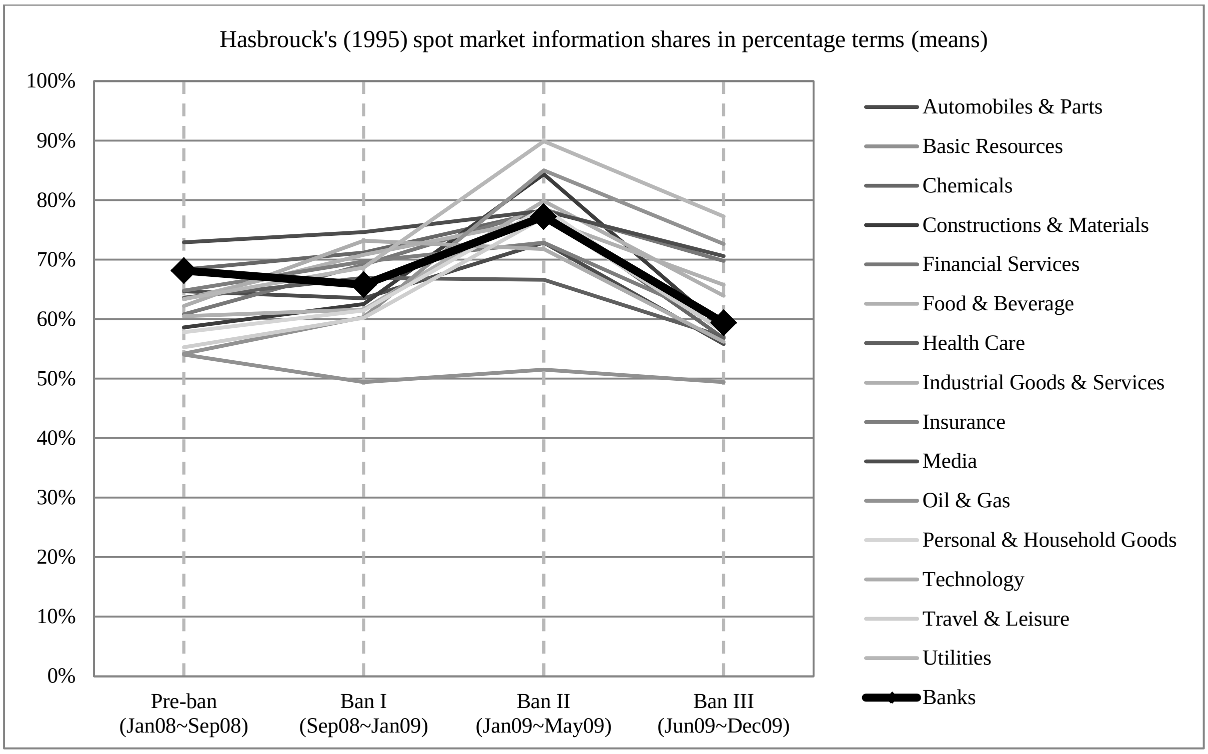
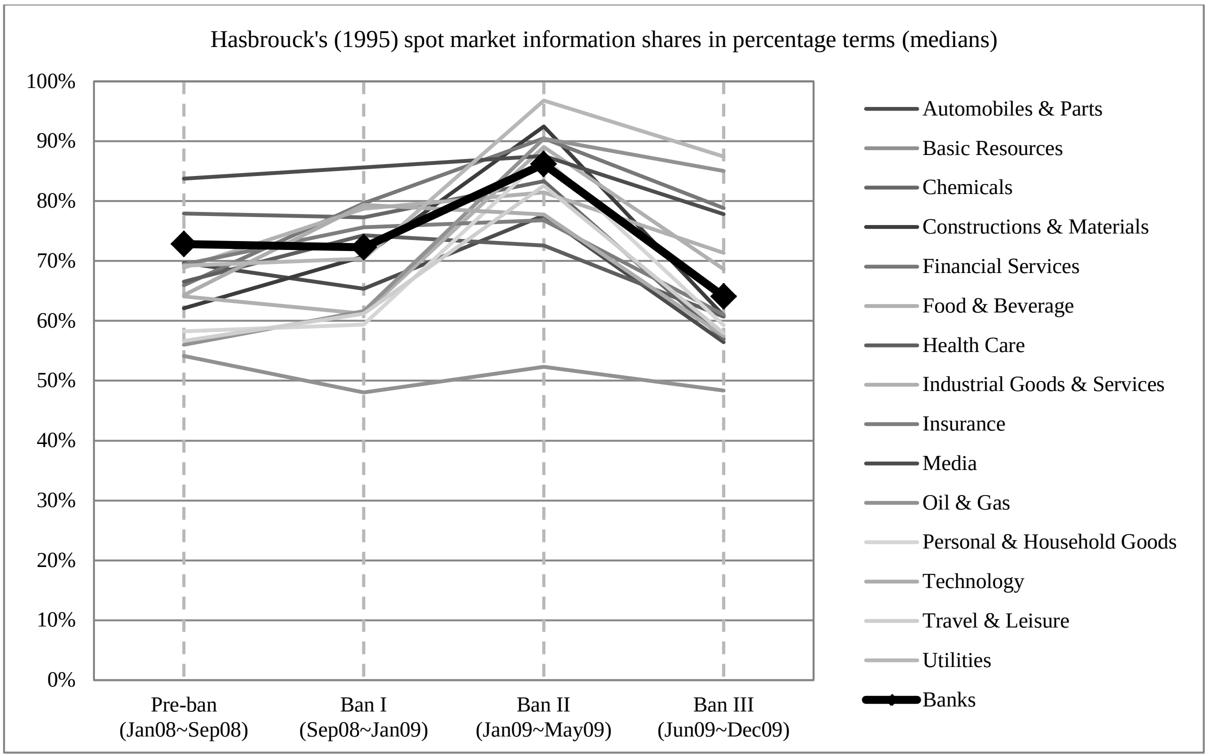


Figure 6.4 Median values of Hasbrouck's (1995) spot market information shares in percentage terms

This figure illustrates the median values of the Hasbrouck's (1995) spot market information shares across 16 Dow Jones STOXX 600 super-sector indexes. The sample period spans from January, 15 of 2008 to December 15, of 2009 and it is divided into four sub-periods: the pre-ban period (September 19, 2008–December 15, 2009), ban I period (September 19, 2008–January 16, 2009), ban II period (January 19, 2009–May 29, 2009), and ban III period (June 1, 2009–December 15, 2009).



Panel A of Table 6.4 shows the descriptive statistics of the Hasbrouck's (1995) information shares of other Dow Jones STOXX 600 super-sectors, and Panel B of Table 6.4 reports the statistics of tests for equality of means and medians between different sub-periods.

The data in Panel A of Table 6.4 show that between the pre-ban period and ban period I there is no statistically significant change in both mean and median values of the information share of the spot market for 11 out of 15 super-sector indexes. The information share of the spot market increases significantly for the remaining four super-sector indexes.[181] Furthermore, there is a statistically significant increase in the spot market information share from ban period I to ban period II and a subsequent statistically significant decrease in the spot market

[181] These four super-sectors are Financial Services, Food & Beverage, Oil & Gas, and Technology.

information share from ban period II to ban period III for 10 out of 15 super-sector indexes.[182] For example, as shown in Panel A of Table 6.4, the mean value of the information share of Oil and Gas index increases by 24.66%, from 60.33% to 84.99% between the periods ban I and ban II. Construction & Materials, Utilities, and Industrial Goods & Services follow with an increase of 21.82%, 21.23%, and 18.22%, respectively. The greatest decrease of 26.05% in the mean value of the spot market information share between period ban II and ban III is found for Construction & Materials. The Personal & Household Goods index contribution to price discovery decreases by 20.9%, from 78.63% to 57.73%, between the same sub-periods. The data in Panel B of Table 6.4 show that all the results are statistically significant at the 1% level.

Overall, the results show that the information shares of the Dow Jones STOXX 600 super-sector indexes trend upwards until the end of the second ban sub-period and subsequently drop in the third ban sub-period. The changes in the information shares of the Dow Jones STOXX 600 super-sector indexes are similar to the movements of the information share of the Dow Jones STOXX 600 Banks index. In summary, these sector-level results suggest that there were other market forces that drove the spot-futures informational dynamics during the sample period.

[182] These super-sectors are Automobiles & Parts, Construction & Materials, Financial Services, Food & Beverage, Industrial Goods & Services, Oil & Gas, Personal & Household Goods, Technology, Travel & Leisure, and Utilities.

Table 6.4 Information Shares (IS) of the Dow Jones STOXX 600 super-sector indexes

Panel A of Table 6.4 shows the descriptive statistics of the Hasbrouck's (1995) Information Shares (IS) for 15 Dow Jones STOXX 600 super-sector indexes. The IS are measured using 15 second returns and 10 lags. The whole sample period spans from January, 15 of 2008 to December 15, of 2009 and it is divided into four sub-periods: the pre-ban period (September 19, 2008–December 15, 2009), ban I period (September 19, 2008–January 16, 2009), ban II period (January 19, 2009–May 29, 2009), and ban III period (June 1, 2009–December 15, 2009). Observations refer to the number of days included in each sub-period. Panel B shows the statistical significance of the test for equality of means and medians of the IS between different sub-periods. T-test is used to test the significance of equality of means while Wilcoxon test is used to test the significance of equality of the medians. P-values are given in the parentheses below the differences. *indicates statistical significance at the 10% level. ** indicates statistical significance at the 5% level. *** indicates statistical significance at the 1% level.

Panel A: Descriptive Statistics							Panel B: Tests for equality of means and medians		
Sector: Automobiles & Parts									
Period	Mean	Median	Maximum	Minimum	Standard Deviation	Observations	Difference (Period)	Difference (Means)	Difference (Medians)
Pre-ban	0.6467	0.6972	0.9991	0.0096	0.2852	171	Ban I vs. pre-ban	-0.012 (0.7361)	-0.0433 (0.2812)
Ban I	0.6347	0.6539	0.9905	0.1723	0.2115	81	Ban II vs. ban I	0.0939*** (0.0057)	0.1222*** (0.0020)
Ban II	0.7286	0.7761	0.9952	0.1062	0.2265	91	Ban III vs. ban II	-0.1706*** (0.0001)	-0.2119*** (0.0001)
Ban III	0.5580	0.5642	0.9850	0.0841	0.2102	142	Ban III vs. pre-ban	-0.0887*** (0.0023)	-0.133*** (0.0003)
Sector: Basic Resources									
Pre-ban	0.5406	0.5413	0.9806	0.1052	0.1924	171	Ban I vs. pre-ban	-0.0464 (0.0599)	-0.0606 (0.0789)
Ban I	0.4942	0.4807	0.8587	0.0031	0.1568	81	Ban II vs. ban I	0.0211 (0.3732)	0.0426 (0.3614)
Ban II	0.5153	0.5233	0.8200	0.1681	0.1533	91	Ban III vs. ban II	-0.0212 (0.2958)	-0.0399 (0.2157)
Ban III	0.4941	0.4834	0.9753	0.1604	0.1490	142	Ban III vs. pre-ban	-0.0465** (0.0195)	-0.0579** (0.0256)
Sector: Chemicals									
Pre-ban	0.6834	0.779	0.9994	0.0167	0.2844	171	Ban I vs. pre-ban	0.0286 (0.4374)	-0.006 (0.7420)
Ban I	0.712	0.773	0.9985	0.1467	0.2443	81	Ban II vs. ban I	0.0684 (0.0520)	0.0602** (0.0390)
Ban II	0.7804	0.8332	0.9999	0.2233	0.2143	91	Ban III vs. ban II	-0.2127*** (0.0001)	-0.2633*** (0.0001)
Ban III	0.5677	0.5699	0.9475	0.1015	0.1811	142	Ban III vs. pre-ban	-0.1157*** (0.0001)	-0.2091*** (0.0001)

Table 6.4 Information Shares (IS) of the Dow Jones STOXX 600 super-sector indexes

Table continued from previous page

Panel A: Descriptive Statistics							Panel B: Tests for equality of means and medians		
Sector: Constructions & Materials									
Period	Mean	Median	Maximum	Minimum	Standard Deviation	Observations	Difference (Period)	Difference (Means)	Difference (Medians)
Pre-ban	0.5860	0.6209	0.9994	0.0113	0.3230	171	Ban I vs. pre-ban	0.0392 (0.3491)	0.0865 (0.6400)
Ban I	0.6252	0.7074	0.9988	0.0356	0.2793	81	Ban II vs. ban I	0.2182*** (0.0001)	0.2172*** (0.0001)
Ban II	0.8434	0.9246	0.9992	0.0223	0.2136	91	Ban III vs. ban II	-0.2605*** (0.0001)	-0.3157*** (0.0001)
Ban III	0.5829	0.6089	0.9996	0.0097	0.3151	142	Ban III vs. pre-ban	-0.0031 (0.9312)	-0.012 (0.8992)
Sector: Financial Services									
Pre-ban	0.6079	0.6595	0.9978	0.0054	0.2769	171	Ban I vs. pre-ban	0.0835** (0.0276)	0.137** (0.0177)
Ban I	0.6914	0.7965	0.9923	0.0023	0.2844	81	Ban II vs. ban I	0.0931** (0.0249)	0.1084*** (0.0013)
Ban II	0.7845	0.9049	0.9994	0.0668	0.2552	91	Ban III vs. ban II	-0.0868** (0.0170)	-0.1166*** (0.0029)
Ban III	0.6977	0.7883	0.9989	0.0328	0.2776	142	Ban III vs. pre-ban	0.0898*** (0.0047)	0.1288*** (0.0015)
Sector: Food & Beverage									
Pre-ban	0.6333	0.6889	0.9997	0.024	0.2826	171	Ban I vs. pre-ban	0.0747** (0.0396)	0.0984 (0.1144)
Ban I	0.7080	0.7873	0.9937	0.0131	0.2328	81	Ban II vs. ban I	0.0605 (0.823)	0.027** (0.0370)
Ban II	0.7685	0.8143	0.9994	0.0823	0.2211	91	Ban III vs. ban II	-0.1107*** (0.0008)	-0.1007*** (0.0004)
Ban III	0.6578	0.7136	0.9957	0.0217	0.2556	142	Ban III vs. pre-ban	0.0245 (0.4255)	0.0247 (0.6946)

Table 6.4 Information Shares (IS) of the Dow Jones STOXX 600 super-sector indexes

Table continued from previous page

Panel A: Descriptive Statistics							Panel B: Tests for equality of means and medians		
Sector: Health Care									
Period	Mean	Median	Maximum	Minimum	Standard Deviation	Observations	Difference (Period)	Difference (Means)	Difference (Medians)
Pre-ban	0.6362	0.6656	0.9982	0.0268	0.2685	171	Ban I vs. pre-ban	0.033 (0.3539)	0.0773 (0.4615)
Ban I	0.6692	0.7429	0.9970	0.1066	0.2528	81	Ban II vs. ban I	-0.003 (0.9402)	-0.0171 (0.8192)
Ban II	0.6662	0.7258	0.9983	0.0582	0.2765	91	Ban III vs. ban II	-0.0964** (0.0103)	-0.1184*** (0.0077)
Ban III	0.5698	0.6074	0.9975	0.0402	0.2777	142	Ban III vs. pre-ban	-0.0664** (0.0332)	-0.0582** (0.0303)
Sector: Industrial Goods & Services									
Pre-ban	0.6049	0.6407	0.9995	0.0110	0.2874	171	Ban I vs. pre-ban	0.0112 (0.7592)	-0.029 (0.9963)
Ban I	0.6161	0.6117	0.9816	0.0225	0.2282	81	Ban II vs. ban I	0.1822*** (0.0001)	0.279*** (0.0001)
Ban II	0.7983	0.8907	0.9985	0.0401	0.2335	91	Ban III vs. ban II	-0.1593*** (0.0001)	-0.2042*** (0.0001)
Ban III	0.6390	0.6865	0.9914	0.0671	0.2389	142	Ban III vs. pre-ban	0.0341 (0.2609)	0.0458 (0.5086)
Sector: Insurance									
Pre-ban	0.6477	0.6945	0.9993	0.0223	0.2642	171	Ban I vs. pre-ban	0.049 (0.1662)	0.0616 (0.1846)
Ban I	0.6967	0.7561	0.9957	0.0390	0.2551	81	Ban II vs. ban I	0.0316 (0.4189)	0.0121 (0.2304)
Ban II	0.7283	0.7682	0.9999	0.1229	0.2543	91	Ban III vs. ban II	-0.1344*** (0.0001)	-0.1576*** (0.0001)
Ban III	0.5939	0.6106	0.9657	0.0034	0.2159	142	Ban III vs. pre-ban	-0.0538 (0.2243)	-0.0839** (0.0225)

Table 6.4 Information Shares (IS) of the Dow Jones STOXX 600 super-sector indexes

Table continued from previous page

Panel A: Descriptive Statistics							Panel B: Tests for equality of means and medians		
Sector: Media									
Period	Mean	Median	Maximum	Minimum	Standard Deviation	Observations	Difference (Period)	Difference (Means)	Difference (Medians)
Pre-ban	0.7289	0.8375	0.9994	0.0505	0.2797	171	Ban I vs. pre-ban	0.0175 (0.6322)	0.0188 (0.7266)
Ban I	0.7464	0.8563	0.9965	0.0772	0.2483	81	Ban II vs. ban I	0.0359 (0.3401)	0.0194 (0.1004)
Ban II	0.7823	0.8757	0.9996	0.0487	0.2437	91	Ban III vs. ban II	-0.0764** (0.0196)	-0.0973** (0.0008)
Ban III	0.7059	0.7784	0.9949	0.0139	0.2412	142	Ban III vs. pre-ban	-0.023 (0.4419)	-0.0591** (0.0290)
Sector: Oil & Gas									
Pre-ban	0.5419	0.5603	0.9945	0.0301	0.2715	171	Ban I vs. pre-ban	0.0614* (0.0722)	0.0559* (0.0901)
Ban I	0.6033	0.6162	0.9690	0.1673	0.2047	81	Ban II vs. ban I	0.2466*** (0.0001)	0.2879*** (0.0001)
Ban II	0.8499	0.9041	0.9986	0.2010	0.1482	91	Ban III vs. ban II	-0.1235*** (0.0001)	-0.0535* (0.0895)
Ban III	0.7264	0.8506	0.9984	0.1173	0.2769	142	Ban III vs. pre-ban	0.1845** (0.0001)	0.2903*** (0.0001)
Sector: Personal & Household goods									
Pre-ban	0.5779	0.5823	0.9989	0.0040	0.2858	171	Ban I vs. pre-ban	0.0368 (0.3077)	0.0115 (0.4949)
Ban I	0.6147	0.5938	0.9919	0.1100	0.2191	81	Ban II vs. ban I	0.1716*** (0.0001)	0.2797*** (0.0001)
Ban II	0.7863	0.8735	0.9994	0.0766	0.2368	91	Ban III vs. ban II	-0.209*** (0.0001)	-0.2808** (0.0001)
Ban III	0.5773	0.5927	0.9938	0.0552	0.2724	142	Ban III vs. pre-ban	-0.0006 (0.9838)	0.0104 (0.9894)

Table 6.4 Information Shares (IS) of the Dow Jones STOXX 600 super-sector indexes

Table continued from previous page

Panel A: Descriptive Statistics							Panel B: Tests for equality of means and medians		
Sector: Technology									
Period	Mean	Median	Maximum	Minimum	Standard Deviation	Observations	Difference (Period)	Difference (Means)	Difference (Medians)
Pre-ban	0.6219	0.6432	0.9983	0.0407	0.2561	171	Ban I vs. pre-ban	0.11*** (0.0010)	0.1507*** (0.0017)
Ban I	0.7319	0.7939	0.9966	0.1241	0.2164	81	Ban II vs. ban I	-0.0143 (0.6971)	-0.0164 (0.6743)
Ban II	0.7176	0.7775	0.9988	0.0961	0.2629	91	Ban III vs. ban II	-0.1556*** (0.0001)	-0.203** (0.0001)
Ban III	0.562	0.5745	0.9584	0.0273	0.1964	142	Ban III vs. pre-ban	-0.0599** (0.0232)	-0.0687** (0.0230)
Sector: Travel & Leisure									
Pre-ban	0.5529	0.5663	0.9974	0.0098	0.2841	171	Ban I vs. pre-ban	0.0494 (0.1832)	0.0459 (0.2329)
Ban I	0.6023	0.6122	0.9922	0.0893	0.2498	81	Ban II vs. ban I	0.1684*** (0.0001)	0.2138*** (0.0001)
Ban II	0.7707	0.826	0.9996	0.0325	0.2222	91	Ban III vs. ban II	-0.1909*** (0.0001)	-0.2464*** (0.0001)
Ban III	0.5798	0.5796	0.9991	0.0067	0.3181	142	Ban III vs. pre-ban	0.0269 (0.4318)	0.0133 (0.3442)
Sector: Utilities									
Pre-ban	0.6339	0.6922	0.9961	0.0163	0.2717	171	Ban I vs. pre-ban	0.0525 (0.1378)	0.0119 (0.2107)
Ban I	0.6864	0.7041	0.9932	0.034	0.2364	81	Ban II vs. ban I	0.2123*** (0.0001)	0.2637*** (0.0001)
Ban II	0.8987	0.9678	0.9992	0.0137	0.1825	91	Ban III vs. ban II	-0.1259*** (0.0001)	-0.0935*** (0.0001)
Ban III	0.7728	0.8743	0.9992	0.1553	0.2429	142	Ban III vs. pre-ban	0.1389*** (0.0001)	0.1821*** (0.0001)

6.6.4 Price discovery in rising versus falling markets

Chan (1992) and Tse and Chan (2010) examine the informational interaction between the spot and futures markets conditional on the sign of market returns. The rationale behind it is that if short-selling constraints are binding then there should be a slower incorporation of new information into prices under falling markets because there is a greater selling pressure in the markets. However, the authors do not find a significant difference in the lead/lag relationship between the spot and the futures markets under falling versus rising markets.

Table 6.5 reports the average Hasbrouck's (1995) spot information shares for 16 Dow Jones STOXX 600 super-sector indexes in each sub-period conditional on whether the spot returns are negative or positive. More specifically, the sample is partitioned into two sub-samples: a sample of days when index returns are negative and a sample of days when index returns are positive. These results are shown in Panel A of Table 6.5. Panel B of Table 6.5 shows the results of the statistical significance of the difference between the information shares under falling markets and the information shares under rising markets across all the super-sector indexes in the sample in each sub-period. In the presence of short-selling constraints, assuming that they are binding, it would be expected to find a lower contribution from the spot market to the price discovery under falling markets. Furthermore, after the introduction of the short-selling bans, the difference between the information share of the spot market under falling and rising markets would be expected to increase for the banks index compared to other indexes. This is because the banking sector was the most affected by the new short-selling regulations.

The results reported in Table 6.5 do not find a significant difference in the spot information shares under different market conditions for all the super-sector indexes and all the sub-periods. These results are in line with the findings of Chan (1992) and Tse and Chan (2010).

However, the main difference is that in these studies there are no explicit changes in the short-selling regulations. Rather, the authors use market conditions to proxy for short-selling constraints.

Table 6.5 Information shares in falling markets versus information shares in rising markets

Panel A of Table 6.5 reports the average Hasbrouck's (1995) information shares for 16 Dow Jones STOXX 600 super-sector indexes over each sub-period. The sample is partitioned into two sub-samples conditional on negative and positive index returns, respectively. The whole sample period spans from January, 15 of 2008 to December 15, of 2009 and it is divided into four sub-periods: the pre-ban period (September 19, 2008–December 15, 2009), ban I period (September 19, 2008–January 16, 2009), ban II period (January 19, 2009–May 29, 2009), and ban III period (June 1, 2009–December 15, 2009). Panel B shows the statistical significance of the difference between the information shares in falling markets and the information shares in rising markets for each Dow Jones STOXX 600 super-sector index in each sub-period. T-test is used and p-values are given in the parentheses below the differences. *indicates statistical significance at the 10% level. ** indicates statistical significance at the 5% level. *** indicates statistical significance at the 1% level.

	Panel A: Average information shares										Panel B: Statistical significance of the difference				
	Falling market					Rising market					Falling vs. rising market				
Dow Jones STOXX 600 super-sectors	Period					Period					Period				
	Whole	Pre-ban	Ban I	Ban II	Ban III	Whole	Pre-ban	Ban I	Ban II	Ban III	Whole	Pre-ban	Ban I	Ban II	Ban III
Banks	0.6596	0.6843	0.6454	0.7848	0.5710	0.6811	0.6994	0.6758	0.7635	0.6137	-0.0215 (0.3402)	-0.0151 (0.4998)	-0.0304 (0.6048)	0.0213 (0.6551)	-0.0427 (0.2832)
Automobiles & Parts	0.6321	0.6398	0.6513	0.7325	0.5898	0.6381	0.6564	0.5952	0.7369	0.5277	-0.006 (0.7927)	-0.0166 (0.7086)	0.0561 (0.2474)	-0.0044 (0.9258)	0.0621 (0.0787)
Basic Resources	0.5072	0.5381	0.5129	0.5006	0.4753	0.5210	0.5425	0.4742	0.5269	0.5107	-0.0138 (0.3680)	-0.0044 (0.8820)	0.0387 (0.2845)	-0.0263 (0.4202)	-0.0354 (0.1637)
Chemicals	0.6950	0.7121	0.7276	0.8057	0.5826	0.6518	0.6560	0.6960	0.7587	0.5548	0.0432 (0.0569)	0.0561 (0.1996)	0.0316 (0.5632)	0.047 (0.3001)	0.0278 (0.3640)
Constructions & Materials	0.6516	0.6217	0.6391	0.8358	0.5812	0.6283	0.5450	0.6276	0.8503	0.5843	0.0233 (0.4112)	0.0767 (0.1227)	0.0115 (0.8596)	-0.0145 (0.7483)	-0.0031 (0.9532)
Financial Services	0.6675	0.5821	0.7045	0.7816	0.7006	0.6952	0.6398	0.6759	0.7865	0.6951	-0.0277 (0.2792)	-0.0577 (0.1777)	0.0286 (0.6556)	-0.0049 (0.9284)	0.0055 (0.9054)
Food & Beverage	0.6728	0.6109	0.7145	0.7785	0.6354	0.6824	0.6503	0.7013	0.7480	0.6734	-0.0096 (0.6854)	-0.0394 (0.3757)	0.0132 (0.8003)	0.0305 (0.5223)	-0.038 (0.3860)
Health Care	0.6360	0.6315	0.6944	0.6760	0.5654	0.6201	0.6357	0.6564	0.6587	0.5732	0.0159 (0.5221)	-0.0042 (0.9200)	0.038 (0.4988)	0.0173 (0.7686)	-0.0078 (0.8699)
Industrial Goods & Services	0.6485	0.5914	0.6209	0.8162	0.6308	0.6575	0.6178	0.6114	0.7830	0.6466	-0.009 (0.7075)	-0.0264 (0.5514)	0.0095 (0.8536)	0.0332 (0.5031)	-0.0158 (0.6965)
Insurance	0.6689	0.6365	0.7139	0.7610	0.6201	0.6422	0.6613	0.6742	0.6976	0.5747	0.0267 (0.2439)	-0.0248 (0.5449)	0.0397 (0.4908)	0.0634 (0.2362)	0.0454 (0.2151)
Media	0.7278	0.7301	0.7230	0.7642	0.7047	0.7417	0.7277	0.7704	0.7964	0.7070	-0.0139 (0.5541)	0.0024 (0.9568)	-0.0474 (0.3947)	-0.0322 (0.5375)	-0.0023 (0.9553)
Oil & Gas	0.6645	0.5428	0.6189	0.8577	0.7288	0.6640	0.5410	0.5847	0.8439	0.7241	0.0005 (0.9833)	0.0018 (0.9663)	0.0342 (0.4578)	0.0138 (0.6621)	0.0047 (0.9204)
Personal & Household Goods	0.6363	0.5810	0.6173	0.8304	0.5876	0.6101	0.5749	0.6115	0.7432	0.5689	0.0262 (0.2930)	0.0061 (0.8901)	0.0058 (0.9072)	0.0872 (0.0791)	0.0187 (0.6848)
Technology	0.6463	0.6175	0.7519	0.7447	0.5559	0.6348	0.6273	0.7058	0.6943	0.5680	0.0115 (0.6070)	-0.0098 (0.8051)	0.0461 (0.3447)	0.0504 (0.3645)	-0.0121 (0.7156)
Utilities	0.7262	0.6391	0.6733	0.8851	0.7675	0.7401	0.6285	0.6981	0.9109	0.7778	-0.0139 (0.5607)	0.0106 (0.7990)	-0.0248 (0.6402)	-0.0258 (0.5036)	-0.0103 (0.8027)

6.6.5 Summary of the results and discussion

6.6.5.1 Overview

Overall, the results are not in line with theory, which predicts that short-selling constraints will reduce the contribution from the spot market to the price discovery process. Nor do these results support the theoretical prediction of an increased contribution from the spot market to the price discovery process when short-selling constraints are relaxed (Diamond and Verrecchia (1987)). In contrast, this essay finds that the information share of the banks index in the price discovery process does not change in the first ban sub-period when short-selling bans cover nearly 97% of the components of the banks index. Alternatively, in the third ban sub-period, when short-selling bans continue to be relaxed, the information share of the banks index in the price discovery process decreases rather than increases as would be expected. These results also contradict the findings of prior studies. For example, Fung and Jiang (1999) and Jiang, Fung, and Cheng (2001) find that the contribution from the spot market to price discovery increases after the relaxation of the short-selling regulations in Hong Kong. To sum up, these results imply that there was no migration of informed traders from the spot market to the futures markets.

Many European countries introduced short-selling regulations in both spot and derivatives markets in the autumn of 2008. Increased compliance costs in the futures market associated with these short-selling regulations could potentially explain why there was no migration of informed traders from the spot to the futures market.

A closer look at the spot-futures informational dynamics of other DJ STOXX 600 super-sector indexes reveals that the spot market information share of the Banks super-sector index follows the same trend as the spot market information share of the other super-sector indexes. This suggests that the factors driving the spot market information share of price discovery in these super-sectors also drive the spot market information share of price discovery in the Banks super-sector.

Li (2009) combines a VECM with a Markov-switching model to examine the lead/lag relationship between spot and futures markets. He finds that under a low variance regime, the futures price leads the spot price in the price discovery process. On the contrary, under a high variance regime, the spot market appears to lead the futures market. The recent financial crisis was characterized by elevated volatility, which could explain the lead of the spot market over the futures market in the price discovery process.

Furthermore, Grundy, Lim, and Verwijmeren (2012) and Battalio and Schultz (2011) find that during the recent financial crisis the transaction costs in trading derivatives increase significantly in the US. Moreover, the increased transaction costs in derivatives markets is more pronounced for derivatives written on stocks subject to the short-selling ban. Consequently, increased transaction costs in the European super-sector futures markets could also be one of the factors that determined the observed spot-futures informational dynamics.

6.6.5.2 Spot market information shares and price volatility

The more volatile the market conditions, the more informationally efficient the spot market over the futures market (Li (2009)). The data in Tables 6.3 and 6.4 show that, overall, there is no significant change in the information shares of the spot market for the majority of the Dow Jones STOXX super-sector indexes in the first ban sub-period. In the second ban sub-period, the contribution from the spot market to price discovery increases significantly and subsequently decrease significantly in the third ban sub-period. Tables 6.6 and 6.7 report the mean and median values of price volatility, respectively, as measured by the difference between the logarithmic daily highest price and logarithmic daily lowest price of the Dow Jones STOXX super-sector indexes across all sub-periods. Furthermore, Figures 6.5 and 6.6

illustrate the changes in mean and median values of price volatility, respectively, across all super-sectors in the sample over all sub-periods.[183] Overall, the results show that price volatility increases significantly in the first ban sub-period and then decreases significantly in the next two sub-periods. These results are consistent with the volatility argument only in the third ban sub-period, where the decreased information shares in the spot market are associated with lower price volatility of all Dow Jones STOXX 600 super-sector indexes. In contrast, the information shares of the spot market increase dramatically while the price volatility exhibits a substantial drop during the second ban sub-period, which implies that there were other factors contributing to the observed spot-futures informational dynamics during the sample period.

[183] The Travel & Leisure super-sector is excluded from the sample due to reporting errors in the data.

Table 6.6 Mean values of price volatility

Panel A of Table 6.6 shows the mean values of the price volatility for 15 Dow Jones STOXX 600 super-sector indexes in each sub-period. Price volatility is measured as the difference between the logarithmic daily highest value and logarithmic daily lowest value of each index averaged over each sub-period. The whole sample period spans from January, 15 of 2008 to December 15, of 2009 and it is divided into four sub-periods: the pre-ban period (September 19, 2008–December 15, 2009), ban I period (September 19, 2008–January 16, 2009), ban II period (January 19, 2009–May 29, 2009), and ban III period (June 1, 2009–December 15, 2009). Panel B shows the statistical significance of the test for equality of means of the price volatility between different sub-periods. T-test is used and P-values are given in the parentheses below the differences. *indicates statistical significance at the 10% level. ** indicates statistical significance at the 5% level. *** indicates statistical significance at the 1% level.

Panel A: Univariate results						Panel B: Tests for equality of means		
Sector	Pre-ban	Ban I	Ban II	Ban III		Ban I vs. pre-ban	Ban II vs. ban I	Ban III vs. ban II
Banks	0.0305	0.0515	0.0491	0.0224	Diff	0.021***	-0.0024	-0.0267***
					(p-value)	(0.0001)	(0.5224)	(0.0001)
Automobiles & Parts	0.0279	0.0771	0.0434	0.0286	Diff	0.0492***	-0.0337***	-0.0148***
					(p-value)	(0.0001)	(0.0016)	(0.0001)
Basic Resources	0.0341	0.0838	0.0518	0.0329	Diff	0.0497***	-0.032***	-0.0189***
					(p-value)	(0.0001)	(0.0001)	(0.0001)
Chemicals	0.0208	0.0465	0.0286	0.0211	Diff	0.0257***	-0.0179***	-0.0075***
					(p-value)	(0.0001)	(0.0001)	(0.0001)
Constructions & Materials	0.0246	0.0467	0.0346	0.0216	Diff	0.0221***	-0.0121***	-0.013***
					(p-value)	(0.0001)	(0.0001)	(0.0001)
Financial Services	0.0239	0.0464	0.0331	0.0185	Diff	0.0225***	-0.0133***	-0.0146***
					(p-value)	(0.0001)	(0.0001)	(0.0001)
Food & Beverage	0.0161	0.0323	0.0206	0.013	Diff	0.0162***	-0.0117***	-0.0076***
					(p-value)	(0.0001)	(0.0001)	(0.0001)
Health Care	0.0171	0.0333	0.0189	0.0119	Diff	0.0162***	-0.0144***	-0.007***
					(p-value)	(0.0001)	(0.0001)	(0.0001)
Industrial Goods & Services	0.0226	0.0464	0.0295	0.0178	Diff	0.0238***	-0.0169***	-0.0117***
					(p-value)	(0.0001)	(0.0001)	(0.0001)
Insurance	0.0281	0.0536	0.0463	0.0219	Diff	0.0255***	-0.0073*	-0.0244***
					(p-value)	(0.0001)	(0.0664)	(0.0001)
Media	0.0208	0.0378	0.0221	0.0148	Diff	0.017***	-0.0157***	-0.0073***
					(p-value)	(0.0001)	(0.0001)	(0.0001)
Oil & Gas	0.0221	0.0515	0.028	0.0184	Diff	0.0294***	-0.0235***	-0.0096***
					(p-value)	(0.0001)	(0.0001)	(0.0001)
Personal & Household Goods	0.0194	0.0354	0.0226	0.0149	Diff	0.016***	-0.0128***	-0.0077***
					(p-value)	(0.0001)	(0.0001)	(0.0001)
Technology	0.0257	0.0446	0.0291	0.0177	Diff	0.0189***	-0.0155***	-0.0114***
					(p-value)	(0.0001)	(0.0001)	(0.0001)
Utilities	0.0184	0.0407	0.0247	0.015	Diff	0.0223***	-0.016***	-0.0097***
					(p-value)	(0.0001)	(0.0001)	(0.0001)

Figure 6.5 Mean values of price volatility

This figure illustrates the mean values of the price volatility as measured by the difference between the logarithmic daily highest price and logarithmic daily lowest price averaged over each sub-period for 15 Dow Jones STOXX 600 super-sector indexes. The sample period spans from January, 15 of 2008 to December 15, of 2009 and it is divided into four sub-periods: the pre-ban period (September 19, 2008–December 15, 2009), ban I period (September 19, 2008–January 16, 2009), ban II period (January 19, 2009–May 29, 2009), the ban III period (June 1, 2009–December 15, 2009).

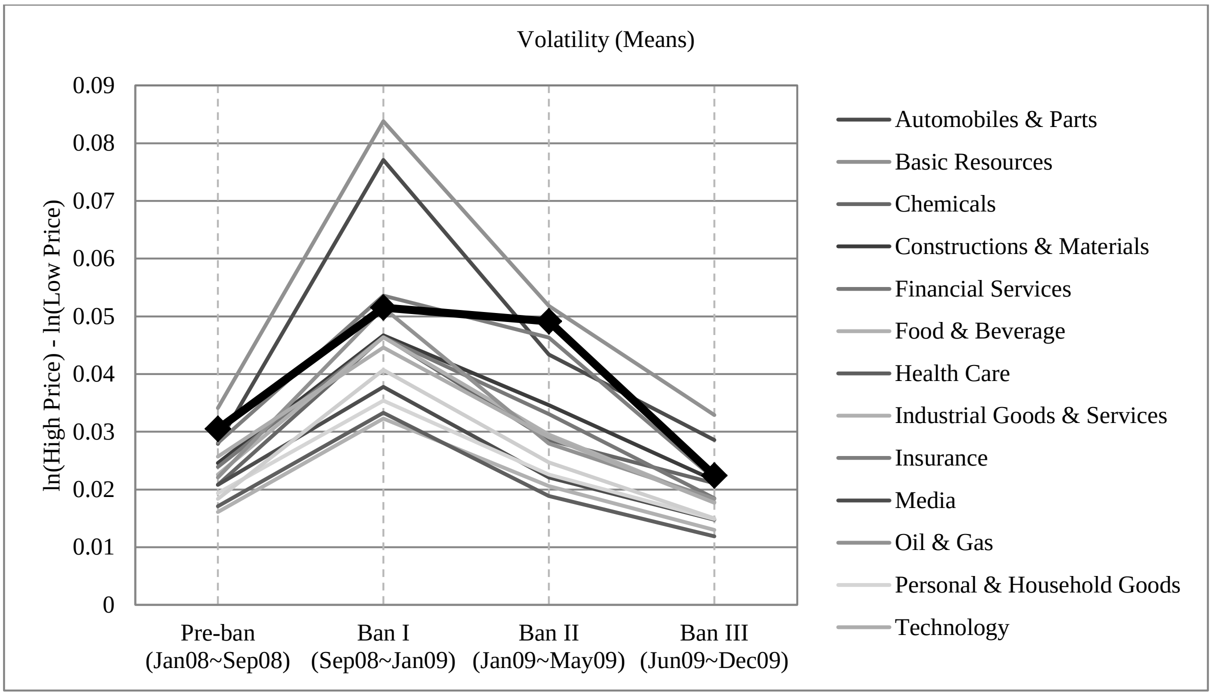


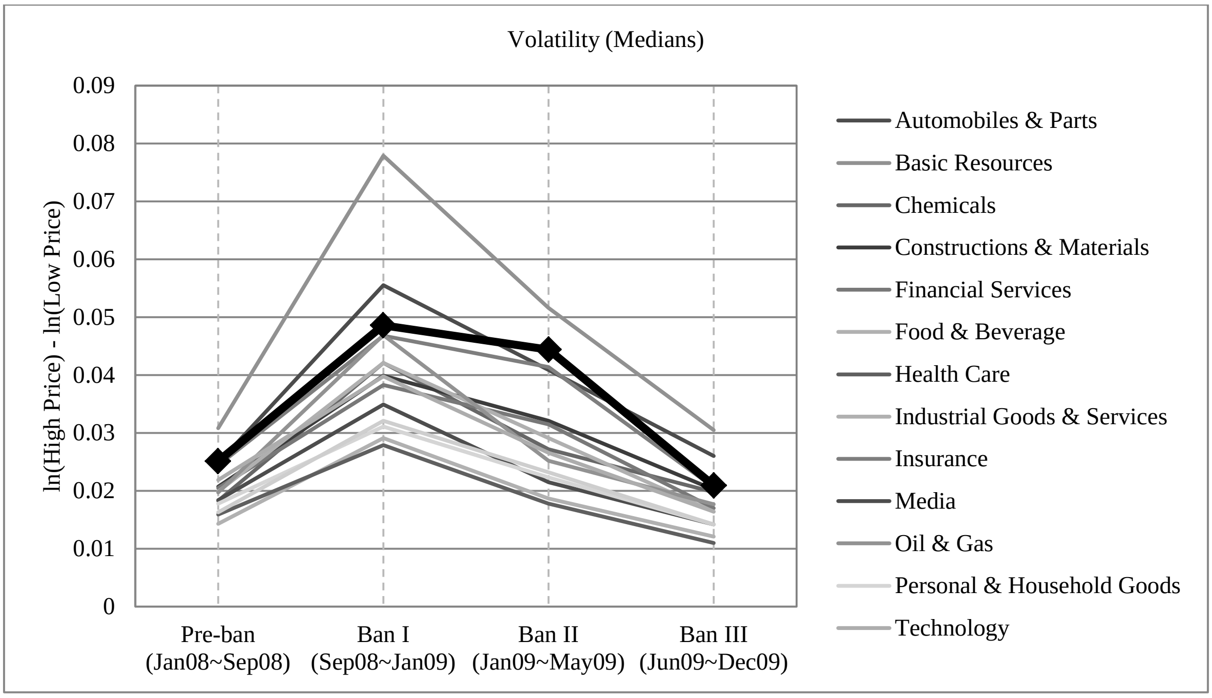
Table 6.7 Median values of price volatility

Panel A of Table 6.7 shows the median values of the price volatility for 15 Dow Jones STOXX 600 super-sector indexes in each sub-period. Price volatility is measured as the difference between the logarithmic daily highest value and logarithmic daily lowest value of each index averaged over each sub-period. The whole sample period spans from January, 15 of 2008 to December 15, of 2009 and it is divided into four sub-periods: the pre-ban period (September 19, 2008–December 15, 2009), ban I period (September 19, 2008–January 16, 2009), ban II period (January 19, 2009–May 29, 2009), and ban III period (June 1, 2009–December 15, 2009). Panel B shows the statistical significance of the test for equality of medians of the price volatility between different sub-periods. Wilcoxon test is used and P-values are given in the parentheses below the differences. *indicates statistical significance at the 10% level. ** indicates statistical significance at the 5% level. *** indicates statistical significance at the 1% level.

Panel A: Univariate results						Panel B: Tests for equality of medians		
Sector	Pre-ban	Ban I	Ban II	Ban III		Ban I vs. pre-ban	Ban II vs. ban I	Ban III vs. ban II
Banks	0.0251	0.0486	0.0444	0.0209	Diff	0.0235***	-0.0042	-0.0235***
					(p-value)	(0.0001)	(0.6273)	(0.0001)
Automobiles & Parts	0.0248	0.0555	0.0409	0.026	Diff	0.0307***	-0.0146***	-0.0149***
					(p-value)	(0.0001)	(0.0001)	(0.0001)
Basic Resources	0.0308	0.0779	0.0516	0.0305	Diff	0.0471***	-0.0263***	-0.0211***
					(p-value)	(0.0001)	(0.0001)	(0.0001)
Chemicals	0.0183	0.0421	0.0272	0.0198	Diff	0.0238***	-0.0149***	-0.0074***
					(p-value)	(0.0001)	(0.0001)	(0.0001)
Constructions & Materials	0.0207	0.0399	0.0321	0.0202	Diff	0.0192***	-0.0078***	-0.0119***
					(p-value)	(0.0001)	(0.0012)	(0.0001)
Financial Services	0.0205	0.0383	0.0315	0.017	Diff	0.0178***	-0.0068***	-0.0145***
					(p-value)	(0.0001)	(0.0013)	(0.0001)
Food & Beverage	0.0143	0.0291	0.0187	0.0121	Diff	0.0148***	-0.0104***	-0.0066***
					(p-value)	(0.0001)	(0.0001)	(0.0001)
Health Care	0.0159	0.0279	0.0178	0.011	Diff	0.012***	-0.0101***	-0.0068***
					(p-value)	(0.0001)	(0.0001)	(0.0001)
Industrial Goods & Services	0.0199	0.0421	0.0291	0.0164	Diff	0.0222***	-0.013***	-0.0127***
					(p-value)	(0.0001)	(0.0001)	(0.0001)
Insurance	0.0244	0.0468	0.0414	0.0206	Diff	0.0224***	-0.0054	-0.0208***
					(p-value)	(0.0001)	(0.1444)	(0.0001)
Media	0.0184	0.0349	0.0215	0.0142	Diff	0.0165***	-0.0134***	-0.0073***
					(p-value)	(0.0001)	(0.0001)	(0.0001)
Oil & Gas	0.0197	0.0469	0.0252	0.0177	Diff	0.0272***	-0.0217***	-0.0075***
					(p-value)	(0.0001)	(0.0001)	(0.0001)
Personal & Household Goods	0.0178	0.0311	0.0223	0.0142	Diff	0.0133***	-0.0088***	-0.0081***
					(p-value)	(0.0001)	(0.0001)	(0.0001)
Technology	0.0219	0.0398	0.0266	0.0164	Diff	0.0179***	-0.0132***	-0.0102***
					(p-value)	(0.0001)	(0.0001)	(0.0001)
Utilities	0.0163	0.0321	0.0232	0.0142	Diff	0.0158***	-0.0089***	-0.009***
					(p-value)	(0.0001)	(0.0001)	(0.0001)

Figure 6.6 Median values of price volatility

This figure illustrates the median values of price volatility as measured by the difference between the logarithmic daily highest price and logarithmic daily lowest price averaged over each sub-period for 15 Dow Jones STOXX 600 super-sector indexes. The sample period spans from January, 15 of 2008 to December 15, of 2009 and it is divided into four sub-periods: the pre-ban period (September 19, 2008–December 15, 2009), ban I period (September 19, 2008–January 16, 2009), ban II period (January 19, 2009–May 29, 2009), and ban III period (June 1, 2009–December 15, 2009).



6.7 Conclusions

This essay examines the impact of the short-selling regulations that were introduced in Europe in the autumn of 2008 on the informational dynamics between the spot (Dow Jones STOXX 600 super-sector indexes) and futures markets. The introduction of the short-selling regulations on stocks of major European financial institutions – mostly banks – is expected to reduce the importance of the spot market in the price discovery process. Alternatively, the relaxation of the short-selling bans is expected to increase the contribution from the spot market to the price discovery process. The results show that there is no significant decrease in the information share of the spot market immediately after the introduction of the short-selling bans for the banks index. Nor is there a significant increase in the information share of the banks index when short-selling regulations are further relaxed in the last ban sub-period. Furthermore, empirical evidence shows that the information share of the banks super-sector index follows a similar trend to the one that is documented for the majority of the other Dow Jones STOXX 600 super-sector indexes during the sample period. These results suggest that there are other factors that drove the spot-futures informational dynamics in Europe during that period. Changes in volatility in the spot market could explain the behaviour of the information share of that market. According to Li (2009), greater volatility in the spot market would be associated with a greater contribution from the spot market to the price discovery process. However, the results show that this prediction holds only in the last ban sub-period where the lower volatility in the spot market is accompanied by a lower information share of that market in price discovery.

Overall, the results do not support Diamond and Verrecchia's (1987) prediction that short-selling constraints in the spot market reduce the contribution from that market to the price discovery process and vice versa. Furthermore, these results are not in line with the findings of prior studies, which find that relaxing short-selling regulations increases the contribution from the spot market to the price discovery process (Fung and Jiang (1999) and Jiang, Fung, and Cheng (2001)). Conditional on the sign of the returns on the indexes, this essay does not find that the spot market is less informationally efficient relative to the futures market under falling markets. These results are consistent with the findings of Chan (1992) and Tse and Chan (2010).

APPENDIX 6A

Table 6A.1 Dow Jones STOXX 600 Banks Index components (as of December, 2009)

No.	Country	Bank	Weight (%)
1	IE	Allied Irish Banks	0.13
2	GR	Alpha Bank	0.49
3	GR	Bank of Greece	0.10
4	IE	Bank of Ireland	0.15
5	ES	Bankinter	0.25
6	GB	Barclays	4.34
7	IT	BCA Carige	0.18
8	IT	BCA Monte dei Paschi di Siena	0.46
9	IT	BCA Popolare di Milano	0.26
10	IT	BCA Popolare di Sondrio	0.27
11	IT	BCA Popolare Emilia Romagna	0.33
12	ES	BBVA	5.91
13	PT	BCO Comercial Portugues	0.39
14	ES	BCO de Valencia	0.16
15	PT	BCO Espirito Santo	0.29
16	IT	BCO Popolare	0.42
17	ES	BCO Popular Espanol	0.56
18	ES	BCO Sabadell	0.47
19	ES	BCO Santander	11.7
20	FR	BNP Paribas	6.86
21	DE	Commerzbank	0.54
22	FR	Credit Agricole	1.62
23	CH	Credit Suisse GRP	4.64
24	IT	Credito Valtellines	0.14
25	DK	Danske Bank	0.90
26	DE	Deutsche Bank	3.96
27	DE	Deutsche Postbank	0.24
28	BE	Dexia	0.41
29	NO	DNB NOR	0.86
30	GR	EFG Eurobank Ergasias	0.30
31	AT	Erste Group Bank	0.67
32	FR	GRP Societe Generale	4.38
33	GB	HSBC	17.1
34	IT	Intesa Sanpaolo	3.70
35	CH	Julius Baer Grp	0.60
36	DK	Jyske Bank	0.19
37	BE	KBC Grp	0.56
38	GB	Lloyds Banking Grp	2.53

Table 6A.1 Dow Jones STOXX 600 Banks Index components (as of December, 2009)

Table continued from previous page

No.	Country	Bank	Weight (%)
39	IT	Mediobanca	0.55
40	GR	National Bank of Greece	1.35
41	FR	Natixis	0.35
42	SE	Nordea Bank	2.24
43	GR	Piraeus Bank	0.34
44	FI	Pohjola Bank	0.20
45	AT	Raiffeisen International Bank	0.24
46	GB	Royal Bank of Scotland Grp	0.72
47	SE	Skandinaviska Enskilda Bka	0.93
48	GB	Standard Chartered	3.50
49	SE	Svenska Handelsbankena	1.18
50	SE	Swedbank	0.75
51	DK	Sydbank	0.16
52	IT	UBI BCA	0.79
53	CH	UBS	4.79
54	IT	UNICREDIT	4.62
55	CH	Valiant	0.27

Table 6A.2 Countries' weight in Dow Jones STOXX 600 Banks Index (as of December, 2009)

Country	Number of banks	Total weight in index
UK	5	28.15%
Spain	6	19.1%
France	4	13.21%
Italy	11	11.72%
Switzerland	4	10.28%
Sweden	4	5.10%
Germany	3	4.74%
Greece	5	2.58%
Denmark	3	1.24%
Belgium	2	0.96%
Austria	2	0.92%
Norway	1	0.86%
Portugal	2	0.68%
Ireland	2	0.28%
Finland	1	0.19%

CHAPTER 7

Conclusions

7.1 Introduction

At the outset of the Global Financial Crisis, financial authorities in many countries introduced various forms of short-selling regulations on the stocks of financial institutions in an attempt to halt further stock price declines and restore confidence in the financial system. This study examined the impact of these short-selling regulations on (i) market liquidity, (ii) the components of the bid/ask spread, (iii) the speed of price adjustment to new information, and (iv) the price discovery process between spot and futures markets in Europe over the period from January 2008 to December 2009.

This chapter is organised as follows. Section 7.2 summarises and discusses the main findings for each empirical chapter (i.e., chapters 3–6). Section 7.3 reviews the contributions of this study. Section 7.4 discusses the limitations of this study and provides suggestions for future research.

7.2 Summary and discussion of key findings

7.2.1 Liquidity

Chapter 3 (Essay 1) examines the impact of the 2008 short-selling bans on the liquidity of stocks of 78 major financial institutions across 10 European countries. Essay 1 hypothesizes that the bid/ask spreads increase (decrease) and the two measures of depth – depth and euro depth – decrease (increase) for financial stocks subject to the short-selling bans following the introduction (relaxation) of these bans (Hypotheses 1, 2A, and 2B). Essay 1 also hypothesizes that the liquidity of the financial stocks subject to the short-selling bans deteriorates more if these bans prohibit both naked and covered short-selling (Hypothesis 3). Lastly, this essay hypothesizes that the liquidity of the financial stocks subject to the 2008 short-selling bans is not affected by the absence or presence of listed derivatives following the introduction these bans (Hypothesis 4).

The sample is comprised of two sub-samples of stocks: the banned stocks (i.e., financial stocks that were subject to the short-selling bans) and the non-banned stocks (i.e., non-financial stocks that were not subject to the short-selling bans). The non-banned stocks are used to control for general market movements during the sample period. Bid/ask spreads, which proxy for the price dimension of liquidity, are measured as proposed by Stoll (2000). Additionally, the two measures of depth that represent the quantity dimension of liquidity, as defined in Heflin, Shaw, and Wild (2005), are used. Finally, the multivariate analysis follows that of Boehmer, Jones, and Zhang (2013).

The country-level univariate results show that both the quoted and effective spreads of the banned stocks increase relative to those of the non-banned stocks during the ban period for all countries with the exception of Spain. Moreover, the relative depth in currency terms (or euro depth) of the banned stocks decreases significantly in the ban period for every country apart from Germany and Portugal. In summary, the univariate results show that the introduction of the short-selling bans is associated with an unambiguous deterioration in liquidity for eight out of 10 countries where the increase in relative bid/ask spreads is accompanied by a decrease in relative euro depth. The impact of the short-selling bans on the liquidity of German and Spanish banned stocks is ambiguous. More precisely, in the case of Germany, an increase in the relative bid/ask spreads of the banned stocks is accompanied by an increase in the relative euro depth of the banned stocks during the ban period. The opposite is observed for Spain where both the relative bid/ask spreads and the euro depth of banned stocks decrease during the ban period. In the post-ban period, the univariate results show that liquidity improves unambiguously for two out of the three countries – the Netherlands and the UK – with a post-ban period available. The impact of relaxing the short-selling regulations on liquidity for Switzerland during the post-ban period is ambiguous. These univariate results provide supporting evidence for Hypotheses 1 and 2B and are in line with the findings of prior studies (see for example, Hansson and Fors (2009) and Beber and Pagano (2013)).

The multivariate analysis is conducted for seven out of 10 countries. Belgium, Denmark, and Ireland were excluded from the regression analysis due to their very small sample sizes. The country-level multivariate results show that the relative liquidity of the banned stocks deteriorates for Portugal, Spain, the Netherlands, and the UK during the ban period in terms of higher effective bid/ask spreads and lower euro depth. Liquidity also deteriorates for Germany, France, and Switzerland during the ban period, but only in terms of higher effective bid/ask spreads. These results strongly support Hypothesis 1, as the relative effective bid/ask spreads of banned stocks increase significantly for all countries. However, there is weaker support for Hypothesis 2B, as the relative euro depth of the banned stocks decreases significantly for only four out of seven countries during the ban period. In contrast, there is no support for Hypothesis 2A. In the post-ban period, the country-level multivariate results contradict all the hypotheses. In particular, the relative liquidity of the banned stocks is found to deteriorate rather than improve after the relaxation of the short-selling bans.

The aggregate-level multivariate results show that the relative liquidity of the banned stocks deteriorates after the introduction of the short-selling bans, strongly supporting Hypotheses 1, 2A, and 2B. Furthermore, conditional on the severity of the short-selling bans, the aggregate-level multivariate results show that the relative liquidity of the banned stocks deteriorates in countries with more stringent short-selling bans. These results support Hypothesis 3 and are consistent with the findings of Beber and Pagano (2013), who find a stronger impact of the short-selling regulations on liquidity under more stringent short-selling regimes. However, conditional on trading derivatives, the aggregate-level multivariate results reveal that the relative liquidity of the banned stocks deteriorates significantly for stocks with listed derivatives. In contrast, there is no significant deterioration in the relative liquidity of the banned stocks without listed derivatives. This result contradicts Hypothesis 4 as well as the findings of Beber and Pagano (2013), who find that the bid/ask spreads of the banned stocks without listed options increase more relative to the bid/ask spreads of the banned stocks with listed options.

Overall, empirical evidence shows that short-selling regulations are associated with deterioration in the relative liquidity of the banned stocks in Europe. The impact of the short-selling regulations on liquidity is stronger for banned stocks under more stringent short-selling regimes and for banned stocks with listed derivatives. However, there is no evidence showing that the liquidity of the banned stocks improves following the relaxation of the short-selling bans in countries with a post-ban period available. On the contrary, the multivariate analysis reveals that the relative liquidity of the banned stocks continues to deteriorate in the post-ban period.

7.2.2 Components of the bid/ask spread

Chapter 4 (Essay 2) investigates the effects of the 2008 short-selling bans on the components of the bid/ask spread of stocks of 78 financial institutions across10 European countries. Essay 2 hypothesizes that the non-informational component of the European financial stocks subject to the 2008 short-selling bans increases (decreases) following the introduction (relaxation) of these bans (Hypothesis 1). Essay 2 also hypothesizes that the informational component of the European financial stocks subject to short-selling bans increases (decreases) following the introduction (relaxation) of the short-selling bans (Hypothesis 2). Again, the analysis is conducted on two samples of stocks – the 78 banned stocks and the 78 non-banned stocks – to control for market movements during the sample period. The bid/ask spread is decomposed into its non-informational and informational components as proposed by Bessembinder and Kaufman (1997). Furthermore, a multivariate regression analysis is conducted following the approach of Boehmer, Jones, and Zhang (2013).[184] The country-level univariate results show that the relative non-informational component of the bid/ask spread of the banned stocks increases significantly after the introduction of the short-selling bans for seven out of 10 countries. The relative non-informational component of the bid/ask spread of the banned stocks decreases significantly for France but there is no significant change in the relative non-informational component of the bid/ask spread of the banned stocks for the Netherlands and Switzerland during the ban period. The relative informational component of the bid/ask spread of the banned stocks also increases significantly in the ban period, but only for six out of 10 countries. Denmark and Spain show a significant decrease in the relative informational component of the bid/ask spread of the banned stocks during the ban period but there is no significant change in the relative informational component of the bid/ask spread of the banned stocks for Germany and Portugal. Overall, these results support Hypotheses 1 and 2 and they are in line with the findings of prior studies (see for example, Boehmer, Jones, and Zhang (2013)). However, there is no supporting evidence for Hypotheses 1 and 2 in the post-ban period.

The country-level multivariate results show that the relative non-informational component of the bid/ask spread of the banned stocks increases significantly for four out of seven countries during the ban period (Germany, Portugal, Spain, and the Netherlands). Furthermore, following the introduction of the short-selling bans, the relative informational component of

[184] Belgium, Denmark, and Ireland were excluded from the regression analysis due to their very small sample sizes.

the bid/ask spread of the banned stocks increases significantly for five out of seven countries (France, Spain, the Netherlands, Switzerland and the UK). Overall, these results provide some supporting evidence for Hypotheses 1 and 2 at the country level. However, the country-level multivariate results do not support Hypotheses 1 and 2 in the post-ban period.

The aggregate-level multivariate results show that both the relative non-informational and the informational components of the bid/ask spread of the banned stocks increase significantly in the ban period. Although, both relative components of the bid/ask spread of the banned stocks increase in the ban period, the increase in the relative non-informational component of the bid/ask spread of the banned stocks is greater than the increase in the relative informational component of the bid/ask spread of the banned stocks. These results support Hypotheses 1 and 2 as well as the results that were reported by Boehmer, Jones, and Zhang (2013) for the US market.

Overall, the results suggest that bid/ask spreads on European banned stocks increase during the ban period due to an increase in both the non-informational and the informational components of the bid/ask spread, with the former increasing more than the latter. However, there is no evidence showing that both components of the bid/ask spread decrease following the relaxation of the short-selling bans in the post-ban period.

7.2.3 Speed of price adjustment to new information

Chapter 5 (Essay 3) examines the impact of the 2008 short-selling bans on the speed of price adjustment to new information for the stocks of 78 major European financial institutions across 10 countries. Essay 3 hypothesizes that the autocorrelation in quotes and trades for the European financial stocks subject to the short-selling bans increases (decreases) following the introduction (relaxation) of these bans (Hypotheses 1 and 2). Essay 3 also hypothesizes that the impact of the short-selling bans on the speed of price adjustment to new information is stronger if these bans prohibit both naked and covered short-selling (Hypothesis 3). Finally, essay 3 hypothesizes that the speed of price adjustment to new information for the European financial stocks subject to the short-selling bans is not affected by the absence or presence of listed derivatives after the introduction of these bans (Hypothesis 4).

The same analysis is also conducted for 78 non-financial stocks not subject to the short-selling bans, which are used as a control sample. To estimate the speed of price adjustment to new information, the bivariate VAR model is used as shown in the papers of Hasbrouck (1991) and Chen and Rhee (2010).

The country-level results show that the speed of price adjustment to new information does not change during the ban period for the banned stocks in Belgium, Portugal, and Switzerland. In contrast, the speed of price adjustment to new information increases for the banned stocks in Denmark, Spain and the UK, while there is an uncertain impact on the speed of price adjustment to new information for the banned stocks in France, Ireland, and the Netherlands during the same period. Lastly, the speed of price adjustment to new information is slower during the ban period, as hypothesized, but only for the banned stocks in Germany. Overall, these results do not support Hypotheses 1 and 2 at the country level.

The aggregate-level results show that the speed of price adjustment to new information increases significantly for the European banned stocks during the ban period. These results contradict Hypotheses 1 and 2 and the findings of prior studies which find that the speed of price adjustment to new information is slower for stocks that are not allowed to be sold short (see for example, Chen and Rhee (2010)). Furthermore, conditional on the severity of the short-selling bans, the evidence shows that the speed of price adjustment to new information is not slower under more stringent short-selling regimes, as predicted by Diamond and Verrecchia's (1987) model. On the contrary, the results show that the speed of price adjustment to new information is faster for the banned stocks that were subject to more severe short-selling regulations. These results are not in line with Hypothesis 3. Also, empirical evidence suggests that the speed of price adjustment to new information is faster for the banned stocks with listed derivatives relative to the speed of price adjustment to new information of the banned stocks without listed derivatives during the ban period. These results do not support Hypothesis 4; however, they are in line with the findings of Kolasinski, Reed, and Thornock (2013), who find that the information is incorporated faster into the prices of the banned stocks with listed options relative to the banned stocks without listed options during the ban period in the US.

Overall, the results suggest that the trading of derivatives improves the informativeness of the prices of the European banned stocks during the ban period. In contrast, the speed of price adjustment to new information is found to decrease for the European banned stocks without listed derivatives, as hypothesized.

7.2.4 Price discovery

Chapter 6 (Essay 4) examines the effects of the 2008 short-selling bans on the spot-futures informational dynamics at a European super-sector level. More specifically, this essay investigates the price discovery process between the Dow Jones STOXX 600 super-sector indexes (spot market) and the futures contracts (futures market) written on these indexes under various short-selling regimes. Diamond and Verracchia's (1987) model predicts that short-selling constraints in the spot market will reduce the contribution from that market to the price discovery process. Consequently, this essay hypothesizes that the contribution from the spot market for the European Banks super-sector decreases under tight short-selling regulations (Hypothesis 1) and then gradually increases as these short-selling regulations are relaxed (Hypotheses 2 and 3). This is because short-selling regulations had the potential to affect nearly all the components of that index.

Hasbrouck's (1995) model is used to estimate the information share of the spot market in the price discovery process under tight versus relaxed short-selling regulations. The results show that the information share of the Banks super-sector index does not change significantly when short-selling regulations tighten. This result does not support Hypothesis 1. Following the relaxation of the short-selling bans in some countries, the information share of the Banks super-sector index increases, as predicted, supporting Hypothesis 2. However, when short-selling regulations are relaxed in additional countries, the results show that the information share of the Banks super-sector spot market decreases significantly rather than increases as predicted. This result contradicts Hypothesis 3. In summary, empirical evidence is not in line with the findings of prior studies which find that the contribution from the spot market to the price discovery process increases after the relaxation of the short-selling regulations (see for example, Fung and Jiang (1999) and Jiang, Fung, and Cheng (2001)).

Examining the spot-futures informational dynamics of other European super-sector indexes, most components of which were not subject to the short-selling regulations, reveals that the information share of the Banks super-sector index behaves in a similar way to the information shares of the other super-sector indexes. This implies that other factors were driving the spot market information share of price discovery in most European super-sectors and these were also driving the spot market information share of price discovery in the Banks super-sector.

Overall, the results suggest that the short-selling regulations had little impact on the spot-futures informational dynamics in the European Banks super-sector.

7.3 Contributions of this study

Short-selling and subsequently short-selling regulations have been controversial for many years. This study attempts to shed light on the impact of the 2008 short-selling regulations on (i) the liquidity, (ii) the components of the bid/ask spread, and (iii) the speed of price adjustment to new information for the European financial stocks subject to these regulations. This study also investigates the effects of the short-selling bans on the spot-futures informational dynamics at a super-sector level in Europe. The recent financial crisis was triggered by problems in the banking sector. Thus, financial authorities in many European countries banned short-selling mainly on stocks of banks and financial institutions. This study focuses on the financial stocks that were subject to the short-selling regulations using a control sample of non-financial stocks not subject to the short-selling regulations to control for market movements.

This thesis makes a number of contributions to the literature. The main contribution of *Chapter 3* is to incorporate two measures of depth into the analysis of the liquidity impact of the short-selling regulations. Prior studies examine the impact of the short-selling regulations mostly on the bid/ask spreads (see for example, Bris (2008), Boulton and Braga-Aves (2010), Boehmer, Jones, and Zhang (2013), and Beber and Pagano, 2013)). However, to provide a more complete picture of the impact of the short-selling regulations on liquidity, which is the main purpose of *Chapter 3*, the simultaneous examination of bid/ask spreads and depth is considered to be more appropriate (Lee, Mucklow, and Ready (1993)).

Chapter 4 contributes to the literature by exploring the channel through which the bid/ask spreads increased following the introduction of the short-selling bans. The bid/ask spread can be decomposed into two components: a non-informational component, which is comprised of the order processing costs and inventory holding costs, and an informational component, which captures the information asymmetry costs (Bessembinder and Kaufman (1997)). *Chapter 4* adds to the literature by showing that both the non-informational and the informational components increase during the ban period and that that the former increase more than the latter.

Chapters 5 and *6* contribute to the literature by testing the existing theory on the effects of the short-selling constraints on information efficiency in a new setting. Both chapters use a new dataset that was generated by the introduction of the 2008 short-selling regulations to test Diamond and Verrecchia's (1987) theoretical prediction in Europe. *Chapter 5* focuses on the information efficiency at an individual stock level while *Chapter 6* focuses on the information efficiency between spot and futures markets at a super-sector index level.

7.4 Limitations and future research

An important challenge and consequently limitation of this study is that short-selling regulations were introduced during a severe financial crisis. Although a control sample of stocks was constructed to control for market movements, this sample is comprised of non-financial stocks that belong to other than financial sectors. Therefore, these control stocks might not ideally match the financial stocks that are the focus of this study. A better control sample would be comprised of financial stocks that were not subject to the short-selling bans. However, such data was not available.

The weak results on the contribution from the spot market of the Banks super-sector index to price discovery under various short-selling regimes in *Chapter 6* could also be due to excessive turbulence in the markets during the sample period. In this situation, simple 'before versus after' univariate analysis might not be capable of isolating the impact of the changes in the short-selling regulations.

A possible extension to the empirical work of *Chapters 5* and *6* would be to conduct a multivariate regression analysis to control for other factors that affect the speed of price adjustment to new information and the information share of the spot market.

Some European countries banned short-selling in both spot and derivatives markets while other countries chose to prohibit short-selling only in the spot market. Consequently, another extension to this study would be to examine the potential differences in the impact of the short-selling regulations conditional on whether the countries banned short-selling only in the spot markets or extended the bans to the derivatives markets as well.

Furthermore, this study does not examine the case of the cross-listed stocks that were subject to different short-selling regimes. Providing insights into the regulatory arbitrage that might have taken place would be an interesting topic for further research.

References

Abhyankar, A. (1995). Return and volatility dynamics in the FTSE 100 stock index and stock index futures markets. *Journal of Futures Markets*, *15*(4), 547–488.

Ackert, L., Charupat, N., Church, B., & Deaves, R. (2002). Bubbles in experimental asset markets: Irrational Exuberance no more. *Working Paper.*

Aitken, M., & Frino, A. (1996). The determinants of market bid ask spreads on the Australian stock exchange: Cross-sectional analysis. *AAANZ Conference.*

Aitken, M., Frino, A., McCorry, M., & Swan, P. (1998). Short sales are almost instantaneously bad news: Evidence from Australian stock exchange. *Journal of Finance, 53*(6), 2205–2223.

Amatatsu, Y., & Baba, N. (2008). Price discovery from cross-currency and FX swaps: A structural analysis. *Working Paper*. Bank of International Settlements.

Angelidis, T., & Benos, A. (2009). The components of the bid-ask spread: the case of the Athens stock exchange. *European Financial Management, 15*(1), 112–144.

Bai, Y., Chang, E. C., & Wang, J. (2006). Asset prices under short-sale constraints. *Working Paper.* Hong Kong Institute of Economics and Business Strategy.

Baillie, R., Booth, G., Tse, Y., & Zabotina, T. (2002). Price discovery and common factor models. *Journal of Financial Markets, 5*(3), 309–321.

Barnea, A. (1974). Performance evaluation of New York Stock Exchange Specialists. *Journal of Financial & Quantitative Analysis, 9*(4), 511–535.

Battalio, R., & Schultz, P. (2011). Regulatory Uncertainty and Market Liquidity: The 2008 Short Sale Ban's Impact on Equity Option Markets. *Journal of Finance, 66*(6), 2013–2053.

Beber, A., & Pagano, M. (2013). Short-Selling Bans Around the World: Evidence from the 2007-09 Crisis. *Journal of Finance, 68*(1), 343–381.

Benston, G.J., & Hagerman, R. L. (1974). Determinants of Bid-Asked Spreads in the Over-the-Counter Market. *Journal of Financial Economics, 1*(4), 353–364.

Bessembinder, H., & Kaufman , H. M. (1997). A Comparison of Trade Execution Costs for NYSE and NASDAQ-Listed Stocks. *Journal of Financial & Quantitative Analysis, 32*(3), 287–310.

Biais, B., Bisière, C., & Décamps, J-P. (1999). Short sales constraints, liquidity and price discovery: an empirical analysis on the Paris Bourse. *European Financial Management, 5*(3), 395–409.

Blau, B., Van Ness, B., Van Ness, R., & Wood, R. (2010). Short selling during extreme market movements. *The Journal of Trading, 5*(4), 14–27.

Boehmer, E., Jones, C., & Zhang, X. (2013). Shackling short sellers: The 2008 shorting ban. *The Review of Financial Studies, 26*(6), 1363–1400.

Boehmer, E. and J. Wu (2013). Short Selling and the Price Discovery Process. *The Review of Financial Studies, 26*(2), 287–322.

Boulton, T. J., & Braga-Alves, M. V. (2010). The skinny on the 2008 naked short-sale restrictions. *Journal of Financial Markets*, *13*(4), 397–421.

Bris, A. (2008). Short selling activity in financial stocks and the SEC July 15th emergency order. *Working Paper*.

Bris, A., Goetzmann, W., & Zhu, N. (2007). Efficiency and the bear: Short sales and markets around the world. *Journal of Finance, 62*(3), 1029–1079.

Brunnermeier,M. (2009). Deciphering the liquidity and credit crunch 2007–2008. *Journal of Economic Perspectives, 23*(1), 77–100.

Brunnermeier, M., & Pedersen, L. H. (2005). Predatory Trading. *Journal of Finance, 60*(4), 1825–1863.

CESR–The Committee of European Securities Regulators. (2009). Measures adopted by CESR members on short selling-updated (www.cesr.eu).

CESR–The Committee of European Securities Regulators. (2010). Model for a Pan-European Short Selling Disclosure Regime. *Report* (www.esma.europa.eu).

Chan, K. (1992). A further analysis of the lead-lag relationship between the cash market and the index futures market. *Review of Financial Studies, 5*(1), 123–152.

Chen, C. X. & Rhee, S. G. (2010). Short sales and speed of price adjustment: Evidence from the Hong Kong stock market. *Journal of Banking & Finance 34*(2), 471–483.

Chiu, J., Chung, H., Ho, K-Y., & Wang, G.H.K. (2012). Funding liquidity and equity liquidity in the subprime crisis period: Evidence from the ETF market. *Journal of Banking & Finance, 36*(9), 2660–2671.

Chung, K., Van Ness, B., & Van Ness, R. (2004). Specialists, Limit-Order Traders, and the Components of the Bid-Ask Spread. *The Financial Review, 39*(2), 255–270.

Clarke, J., & Shastri, K. (2000). On information asymmetry metrics. *Working Paper.*

Clifford Chance. (2008a). New short selling rules: global review. *Client briefing* (www.cliffordchance.com).

Clifford Chance. (2008b). New short selling rules: global review-update. *Client briefing* (www.cliffordchance.com).

Clifford Chance. (2008c). Short selling global review: Iceland and Norway update. *Client briefing* (www.cliffordchance.com).

Clifford Chance. (2008d). Short selling global review: update (3). *Client briefing* (www.cliffordchance.com).

Clifford Chance. (2008e). Short selling global review: update (4). *Client briefing* (www.cliffordchance.com).

Clifford Chance. (2009a). Short selling global review: update (6). *Client briefing* (www.cliffordchance.com).

Clifford Chance. (2009b). Short selling global review: update (7). *Client briefing* (www.cliffordchance.com).

Clifford Chance. (2009c). Shot selling rules: the global picture. *Client briefing* (www.cliffordchance.com).

Clifton, M., & Snape, M. (2008). The effect of short-selling restrictions on liquidity: Evidence from the London Stock Exchange. *Working Paper*.

Copeland, T. E., & Galai, M. (1983). Information effects on the bid-ask spread. *Journal of Finance, 38*(5), 1457–1469.

De Jong, F. (2002). Measures of contributions to price discovery: a comparison. *Journal of Financial Markets, 5*(3), 323–327.

Dechow, P. M., Hutton, A. P., Meulbroek, L., & Sloan, R.G. (2001). Short-sellers, fundamental analysis, and stock returns. *Journal of Financial Economics, 61*(1), 77–106.

Demsetz, H. (1968). The cost of transacting. *Quarterly Journal of Economics, 82*(1), 33–53.

Desai, H., Krishnamurthy, S., & Venkataraman, K. (2006). Do short sellers target firms with poor earnings quality? Evidence from earnings restatements. *Review of Accounting Studies, 11*(1), 71–90.

Diamond, D., & Verrecchia, R. (1987). Constraints on short-selling and asset price adjustment to private information. *Journal of Financial Economics*, *18*(2), 277–311.

Dupont, D. (2000). Market Making, Prices, and Quantity Limits. *The Review of Financial Studies*, *13*(4), 1129–1151.

Easley, D., & O'Hara, M. (1987). Price, trade size, and information in securities markets. *Journal of Financial Economics, 19*(1), 69–90.

Engle, R.F., & Russell, J. R. (1998). Autoregressive conditional duration model: a new model for irregularly spaced transaction data. *Econometrica*, *66*(5), 1127–1162.

Fabozzi, F. (2004). Short selling strategies, risks, and rewards. *Hoboken, New Jersey: John Wiley & Sons, Inc.*

Frijins, B., Gilbert, A., & Tourani-Rad, A. (2008). Insider trading, regulation, and the components of the bid-ask spread. *The Journal of Financial Research, 31*(3), 225–246.

FSA–Financial Services Authority. (2009). Short selling. *Discussion Paper*.

Fung, J.K.W., & Jiang, L. (1999). Restrictions on Short-Selling and Spot-Futures Dynamics. *The Journal of Futures Markets, 19*(6), 695–715.

George, T., Kaul, G., & Nimalendran, M. (1991). Estimation of the bid/ask spread components: A new approach. *The Review of Financial Studies, 4*(4), 623–656.

Giouvris, E., & Philippatos, G. (2008). Determinants of the components of the bid-ask spreads on the London Stock Exchange. *Journal of Money, Investment & Banking*, (1), 49–61.

Glosten, L., & Milgrom, P. (1985). Bid, ask, and transaction prices in a specialist market with heterogeneously informed traders. *Journal of Financial Economics, 14*(1), 71–100.

Gonzalo, J., & Granger, C. W. J. (1995). Estimation of common long-memory components in cointegrated systems. *Journal of Business & Economic Statistics, 13*(1), 27–36.

Gruenewald, S., Wagner, A., & Weber, R. (2009a). Emergency short selling restrictions in the course of the financial crisis. *Working Paper*.

Gruenewald, S., Wagner, A., & Weber, R. (2009b). Short selling regulation after financial crisis-First principles revisited. *Working Paper.*

Grundy, B. D., Lim, B., & Verwijmeren, P. (2012). Do option markets undo restrictions on short sales? Evidence from the 2008 short-sale ban. *Journal of Financial Economics, 106*(2), 331–348.

Hansson, F., & Fors, E. R. (2009). Get shorty? Market impact of the 2008-09 U.K. short selling ban. *Working Paper*.

Haruvy, E., & Noussair, C. (2006). The effect of short selling on bubbles and crashes in experimental spot asset markets. *Working Paper.*

Hasbrouck, J. (1988). Trades, quotes, inventories, and information. *Journal of Financial Economics, 22*(2), 229–252.

Hasbrouck, J. (1991). Measuring the information content of stock trades. *Journal of Finance, 46*(1), 179–207.

Hasbrouck, J. (1993). Assessing the quality of a security market: A new approach to transaction-cost measurement. *Review of Financial Studies, 6*(1), 191–212.

Hasbrouck, J. (1995). One security, many markets: Determining the contributions to price discovery. *Journal of Finance, 50*(4), 1175–1199.

Heflin, F. L., Shaw, K. W., & Wild, J.J. (2005). Disclosure Policy and Market Liquidity: Impact of Depth Quotes and Order Sizes. *Contemporary Accounting Research, 22*(4), 829–865.

Ho, T., & Stoll, H. (1981). Optimal dealer pricing under transactions and return uncertainty. *Journal of Financial Economics, 9*(1), 47–73.

Ho, T., & Stoll, H. (1983). The dynamics of dealer markets under competition. *Journal of Finance, 38*(4), 1053–1074.

Hong, H., & Stein, J. (2003). Differences of opinion, short sales constraints, and market crashes. *The Review of Financial Studies, 16*(2), 487–525.

Hou, K. & Moskowitz, T. J. (2005). Market Frictions, Price Delay, and the Cross-Section of Expected Returns. *Review of Financial Studies, 18*(3), 981–1020.

Huang, R. D., & Stoll, H. R. (1996). Dealer versus auction markets: A paired comparison of execution costs on NASDAQ and the NYSE. *Journal of Financial Economics, 41*(3), 313–357.

IOSCO–Technical Committee of the International Organization of Securities Commissions. (2009). Regulations of short selling. *Consultation Report* (www.iosco.org).

ISLA–International Securities Lending Association. (2009). Securities lending and short selling (www.asla.com.au).

Jiang, L., Fung, J. K. W., & Cheng, L.T.W. (2001). The Lead-Lag Relation Between Spot and Futures Markets Under Different Short-Selling Regimes. *The Financial Review, 36*(3), 63–88.

Kolasinski, A., Reed, A., & Thornock, J. (2013). Can Short Restrictions Actually Increase Informed Short Selling? *Financial Management*, *42*(1), 155–181.

Krinsky, I., & Lee, J. (1996). Earnings Announcements and the Components of the Bid-Ask Spread. *Journal of Finance, 51*(4), 1523–1535.

Kyle, A. (1985). Continuous auctions and insider trading. *Econometrica*, *53*(6), 1315–1335.

Lee, C. M. C., Mucklow, B., & Ready, M. J. (1993). Spreads, Depths, and the Impact of Earnings Information: An Intraday Analysis. *The Review of Financial Studies, 6*(2), 345–374.

Lee, C. M. C., & Ready, M. J. (1991). Inferring Trade Direction from Intraday Data. *Journal of Finance, 46*(2), 733–746.

Lehmann, B. (2002). Some desiderata for the measurement of price discovery across markets. *Journal of Financial Markets, 5*(3), 259–276.

Li, M-Y. L. (2009). The dynamics of the relationship between spot and futures markets under high and low variance regimes. *Applied Stochastic Models in Business & Industry, 25*(6), 696–718.

Madhavan, A., Richardson, M., & Roomans, A. (1997). Why do security prices change? A transaction-level analysis of NYSE stocks. *The Review of Financial Studies, 10*(4), 1035–1064.

Mann, S. V., & Ramanlal, P. (1996). The dealers' price/size quote and market liquidity. *Journal of Financial Research, 19*(2), 243–271.

Marsh, I. W., & Payne, R. (2012). Banning short sales and market quality: The UK's experience. *Journal of Banking & Finance, 36*(7), 1975–1986.

McInish, T., & Van Ness, B. (2002). An intraday Examination of the Components of the Bid-Ask Spread. *The Financial Review, 37*(4), 507–524.

McInish, T., Van Ness, B., & Van Ness, R. (2001). Market changes and spread components, implications for international markets. *Journal of International Financial Markets, Institutions & Money, 11*(1), 65–73.

McInish, T., & Wood, R. (1992). An analysis of intraday patterns in bid/ask spreads for NYSE stocks. *Journal of Finance, 47*(2), 753–764.

McMillan, D.G., & Philip, D. (2012). Short-sale constraints and efficiency of the spot-futures dynamics. *International Review of Financial Analysis, 24*, 129–136.

Menyah, K., & Paudyal, K. (1996). The determinants and dynamics of bid-ask spreads on the London Stock Exchange. *Journal of Financial Research, 19*(3), 377–394.

Ni, S. X., & Pan, J. (2010). Trading puts and CDS on stocks with short sale ban. *Working Paper*.

Pacurar, M. (2006). Autoregressive Conditional Duration (ACD) Models in Finance: A Survey of the Theoretical and Empirical Literature. *Working Paper*.

Phillips, B. (2011). Options, short-sale constraints and market efficiency: A new perspective. *Journal of Banking & Finance, 35*(2), 430–442.

Porter, D., & Weaver, D. (1996). Estimating bid-ask spread components: specialist versus multiple market maker systems. *Review of Quantitative Finance & Accounting, 6*(2), 167–180.

Quan, J. (1992). Two-step testing procedure for price discovery role of futures prices. *The Journal of Futures Markets, 12*(2), 139–149.

Saffi, P. A. C., & Sigurdsson, K. (2011). Price Efficiency and Short Selling. *Review of Financial Studies, 24*(3), 821–852.

Scheinkman, J., & Xiong, W. (2003). Overconfidence and speculative bubbles. *Journal of Political Economy, 111*(6), 1183–1219.

Shkilko, A., Van Ness, B., & Vann Ness, R. (2012). Short-selling and intraday price pressures. *Financial Management, 41*(1), 345–370.

Stoll, H. (1978a). The supply of dealer services in securities markets. *Journal of Finance, 33*(4), 1133–1151.

Stoll, H. (1978b). The pricing of security dealer services: An empirical study of NASDAQ stocks. *Journal of Finance*, *33*(4), 1153–1172.

Stoll, H. (1989). Inferring the components of the bid/ask spread: Theory and empirical tests. *Journal of Finance*, *44*(1), 115–134.

Stoll, H. (2000). Friction. *Journal of Finance, 55*(4), 1478–1514.

Tauli, T. (2004). What is short-selling? *New York: McGraw-Hill, Inc*.

Tinic, S. (1972). The economics of liquidity services. *Quarterly Journal of Economics*, *86*(1), 79–93.

Tinic, S., & West, R. (1972). Competition and the pricing of dealer services in the over-the-counter market. *Journal of Financial & Quantitative Analysis, 7*(3), 1707–1727.

Tripathy, N., & Peterson, R.L. (1991). The relationship between OTC bid ask spreads and dealer size: The impact of order processing and diversification costs. *Journal of Financial Research, 14*(2), 117–127.

Tse, Y., & Chan, W. (2010). The lead-lag relation between the S&P 500 spot and futures markets: An intraday-data analysis using a threshold regression model. *The Japanese Economic Review, 61*(1), 133–143.

Van Ness, B., Van Ness, R., & Warr, R. (2001). How Well Do Adverse Selection Components Measure Adverse Selection? *Financial Management*, *30*(3), 77–98.

Wang, J. (1999). Asymmetric information and the bid-ask spread: an empirical comparison between automated order execution and open outcry auction. *Journal of International Financial Markets, Institutions and Money*, *9*(2), 115–128.