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ABSTRACT

New Zealand panel data, which provide extensive information on the benefit histories of children and their parents, is used to estimate an intergenerational correlation coefficient in welfare participation. Recent estimation techniques for addressing issues of measurement error are applied in this analysis (Zimmerman 1992, Solon 1992, Bjorklund and Jantti 1997, Couch, D. and T. Dunn 1997, Auginbaugh, 2000). The long-term benefit histories of parents and instrumental variable techniques provide lower and upper-bound estimates of the true intergenerational correlation. A remarkably narrow band is estimated for this parameter, placing this correlation coefficient at slightly less than 0.4. Approximately one-third of this effect appears to operate through the lower educational attainment of children reared in families receiving social welfare benefits.

I. Introduction

Is benefit dependency transmitted from parents to children? Do cycles of disadvantage persist from one generation to the next? Questions over the transfer of economic status, especially income, between generations have recently received considerable attention in the economics literature (Zimmerman 1992, Solon 1992, Bjorklund and Jantti 1997, Couch, D. and T. Dunn 1997 and Auginbaugh, 2000). For example, once various corrections have been made for measurement error, these studies find that the correlation between the earnings of fathers and their sons is approximately 0.4. This suggests that there is less mobility across generations than indicated in previous studies (e.g., see the earlier survey by Becker and Tomes, 1986). Our goal is to apply estimation techniques developed in this recent literature to the study of intergenerational transmission of welfare participation in New Zealand.

To our knowledge, there has been no empirical analysis to date of the relationship between the participation of parents and their offspring in social welfare programmes in New Zealand. This is unfortunate for two reasons. Firstly, the social welfare system in this country is particularly conducive to this type of research. It is comprised of several negative-income-tax-type programmes that provide, in essence, universal coverage of the adult population. All individuals are 'categorically eligible' for some income transfers. These benefits simply abate away with other income. The basic structure of this system has remained essentially intact over our sample period.[1]

Secondly, recently available panel data in this country provide the necessary information on the benefit histories of both parents and their children. The Christchurch Health and Development Study (CHDS) follows the progress of around 1,000 children from birth through age 21. Annual data on the benefit status of parents was collected while these children were between the ages of one and fourteen. This detailed information is critical for eliminating measurement error in this key independent variable. We show that a more accurate measure of the long-term benefit propensity of parents substantially raises the estimated intergenerational correlation of welfare participation. Instrumental variable estimates also place this correlation coefficient at just under 0.4. Approximately one-third of this effect appears to operate through the lower educational attainment of children reared in families receiving social welfare benefits.

The plan of this paper is as follows: Section II provides a review of earlier studies in this literature. Section III models the approach taken in this study, and surveys the data used in this estimation. Section IV presents our econometric results. Section V summarises our empirical findings and provides some concluding remarks.

II. Earlier Studies

The overall intergenerational link between the welfare participation of parents and their offspring is expected to reflect both the direct and indirect pathways. There are a number of direct mechanisms through which welfare participation may pass from one generation to the next (e.g., see Antel 1992). Firstly, parental welfare participation may lower the distaste or stigma experienced by offspring in receiving social welfare benefits. Secondly, welfare dependency by parents may reduce labour market opportunities of their children. For example, children of welfare recipients may have less exposure to on-the-job experience, and fewer job search skills and informal job contacts through their parents' lack of participation in the labour market. Finally, transaction costs may be reduced for families receiving welfare benefits, because these children learn how the 'system works'. These effects are expected to operate through resource-deficient home environments, poor educational outcomes and greater propensities of single parenting among the children from welfare families.[2] Indirect effects may operate through myriad of observable and unobservable personal and economic circumstances that leave both parents and their children more at risk of welfare dependency

As indicated earlier, previous empirical research on intergenerational welfare dependency is practically non-existent in New Zealand.[3] One exception is a recent study by Seth-Purdie (2000), which uses data from the CHDS to examine various risk factors behind the welfare participation of youth at age 21. The author finds that youth raised in families receiving some welfare benefits are more than twice as likely, compared to youth raised in families that never receive welfare benefits, to receive income from these same transfer programs at the time of the survey at age 21. This study was quite broad, examining a wide array of childhood adversity measures that could potentially influence welfare participation in early adulthood. No attempt was made to provide an in-depth analysis of the intergenerational link in welfare dependency. It was even more disconcerting that the author did not take full advantage of the extensive welfare histories of both parents and their offspring available in the CHDS for this analysis.

Consequently, most of the literature relevant to this present study comes from overseas. There are two sets of earlier studies that are important here. Solon (1992) and Zimmerman (1992) examine the effects of measurement error, due to short periods of observation on parental income, in estimating intergenerational income correlations. Their studies show that short observation 'windows' on parents' income may result in estimates of intergenerational correlations that are biased downward due to errors-in-variables. Cross-sectional data on parental income may be a poor proxy for their permanent income, containing a high ‘noise’ to ‘signal’ ratio.

These studies contend that averaging parental income over several cross sections and the use of instrumental variables provide both ‘lower’ and ‘upper’ bound estimates of the true intergenerational correlation coefficient. Three subsequent studies (Bjorklund and Jantti, 1997; Couch, D. and T. Dunn 1997; and Auginbaugh, 2000) provide further support for the ‘Solon method’.

A second set of studies estimates the link between the welfare participation of parents and their offspring. While the literature on intergenerational income correlations has undergone significant change, studies on intergenerational welfare outcomes have received attention only in the past decade, and mainly based on US data. This is partly because longitudinal data sets have only recently spanned a sufficient number of years to allow the estimation of the relationship between the experiences of parents and their offspring.[4] Due to data restrictions, these studies have generally relied on data from two relatively short periods of observation (e.g., see Gottschalk 1992, and Rainwater 1987). The first window of observation is often used to determine the mother’s welfare exposure, and the second window of observation generally examines the daughter’s welfare experience.

The literature on intergenerational welfare transmission in the US mainly focuses on mother and daughter participation under the now defunct Aid to Families with Dependent Children (AFDC) program (e.g. Duncan, Hill and Hoffman 1988, Gottschalk 1990, 1992, 1996, Antel 1992, Levine and Zimmerman 1996, 2000 and Ratcliffe 1996).[5] These studies generally find that daughters of welfare recipients are more likely to receive welfare than daughters of non-recipients. For example, Rainwater (1987) estimates that daughters of mothers who received AFDC are more than twice as likely to be welfare recipients. Gottschalk (1990) shows that daughters raised in households that received AFDC are disproportionately non-white and disadvantaged (as measured either by the family’s income or mother’s education). Levine and Zimmerman (1996) find that 23% of the daughters of welfare recipients, and at most 7% of the daughters of non-recipients, are themselves AFDC recipients. Because of the demise of the AFDC program over the last few years, this line of research will become increasingly difficult to pursue in the U.S.

One problem in interpreting the results from this US literature is that the intergenerational relationship in the welfare participation of mothers and their daughters is inextricably tied to their ‘categorical eligibility’ for these programs. Although there are some exceptions, income support benefits in the US are received almost entirely by single parents. The structure of the old AFDC program is similar to the structure of the on-going Domestic Purposes Benefit (DPB) in New Zealand. Both are basically negative income tax (NIT) programs for single-parented families, where cash benefits abate away as other income exceeds some maximum threshold. The difference between the countries is that New Zealand also has an Unemployment Benefit program that provides income support, funded out of general tax revenue, to the remainder of the non-disabled, working-age population. This program also has a NIT structure, with both a similar level of benefits and an absence of term limits like DPB, but a relatively stronger work requirement. Thus, it would be fair to characterise the social welfare system as providing something close to a universal NIT structure, with little role for categorical eligibility issues. This should make it easier to measure the extent of intergenerational welfare participation in this country.

In addition, empirical results from US studies are based on relatively limited information on the welfare histories of parents and their offspring. Our study applies the Solon method in estimating the intergenerational correlation of welfare participation to a panel with more extensive information on the welfare histories of parents and their offspring. Since the welfare eligibility system in New Zealand is essentially gender neutral, we are able to examine the intergenerational correlations for both males and females, and other subsamples by ethnicity and other demographic characteristics.

III. The Model

We start with a simple reduced-form model. Let Y_i represent the 'permanent propensity' of welfare participation for a son or daughter at any point in time. Let X_i represent a similar permanent propensity of welfare participation for the youth's parents. These are the long-term benefit states for adjacent generations. If both variables are standardised to have zero means and unit variances (now expressed as lower-case variables), then a single parameter ρ in a two-variable regression is the intergenerational correlation coefficient we want to estimate.

$$y_i = \rho x_i + \varepsilon_i \tag{1}$$

A key methodological issue involves potential measurement error in the two permanent propensities. The problem is that y_i and x_i are not directly observable. Suppose both benefit propensities are measured only at a point in time, indexed by t for the parents and s for the offspring. What we observe is in each case is a 'noisey' indicator of the true permanent propensities:

$$y_{is} = y_i + \nu_{is} \tag{2}$$

$$x_{it} = x_i + \omega_{it} \tag{3}$$

where ν_{is} and ω_{it} are the transitory components. Denoting the population variances of these disturbances as $\sigma^2{}_\nu$ and $\sigma^2{}_\omega$, and assuming that they are uncorrelated with each other and with y_i and x_i, we can estimate the following regression model:

$$y_{is} = \beta x_{it} + u_{is} \tag{4}$$

where the probability limit on the estimated coefficient is:

$$plim\ \hat{\beta} = \frac{\rho}{1+\sigma_\nu^2} < \rho \tag{5}$$

Any measurement error associated with the long-run benefit status of the parents will result in an estimated correlation coefficient that is biased downward. Yet, measurement error in the long-run benefit status of offspring will *not* necessarily result in an underestimate of ρ.

Yet, multiple observations on the benefit status of the parents should mitigate the bias associated with this measurement error. Substituting in the mean frequency of welfare participation over T periods for the parents raises the 'signal' to 'noise' ratio by 'averaging out' these transitory figures:

$$y_{is} = \beta' \bar{x}_{iT} + u_{is} \tag{6}$$

where

$$\bar{x}_{iT} = \frac{1}{T}\sum_{t=1}^{T} x_{it} \tag{7}$$

This results in an ever-shrinking bias as T increases.

$$plim\ \hat{\beta}' = \frac{\rho}{1+\frac{\sigma_v^2}{T}} < \rho \tag{8}$$

Of course, even the extensive social welfare histories of parents available in the CHDS won't entirely eliminate this problem. These are just 'snapshots' of the continuous benefit histories of these parents (up to 14 discrete observations), and these transitory outcomes may not be uncorrelated with each other. However, this procedure should provide a better 'lower bound' estimate for the true intergenerational correlation coefficient.

An 'upper bound' estimate can be formed by the use of an instrumental variable (IV) technique. This can be easily motivated. Suppose the long-run benefit status of the son or daughter is a function of the permanent benefit propensity of the parents and some other background variable z_i (the education of the parents, family structure, etc.), which is also written for convenience as a single, standardised variable with zero mean and unit variance.

$$y_i = \rho_1 x_i + \rho_2 z_i + \psi_i \tag{9}$$

The goal of this analysis is to obtain an estimate of ρ, the overall correlation between the benefit propensities of adjacent generations. Yet, the well-known formula for omitted-variable bias gives the relationship between these parameters:

$$\rho = \rho_1 + \rho_2 \rho_{12} \tag{10}$$

where ρ_{12} is the correlation coefficient between the parents' long-term welfare propensity and this other background variable. The 'partial' correlation coefficient ρ_1 holds constant the separate influence of the other background variable. The two parameters ρ and ρ_1 are identical if either $\rho_2 = 0$ (i.e., z_i has no *direct* impact on the offspring's benefit propensity) or $\rho_{12} = 0$ (i.e., the two regressors x_i and z_i are uncorrelated with one another).

We already know that measurement error in x_i will result in an underestimate ρ under OLS estimation of both equations (4) or (6). Suppose we use z_i as an instrumental variable. If this instrument is uncorrelated with the disturbances in equations (2) and (3), then the IV estimator of the coefficient in equation (4) is:[6]

$$plim\ \hat{\beta}_{IV} = \frac{\rho + \rho_2(1-\rho_{12}^2)}{\rho_{12}} > \rho \tag{11}$$

This IV estimator provides a consistent estimate of ρ only if either $\rho_2 = 0$ (i.e., the background variable doesn't directly influence the welfare propensity of the offspring), or $|\rho_{12}| = 1$ (i.e., the background variable is perfectly correlated with the parents' welfare propensity). Of course, if $0 < \rho_{12} < 1$ (i.e., x_i and z_i are positively, but not perfectly correlated), then this IV estimator will tend to overestimate the true correlation coefficient (if $\rho_2 > 0$, as expected). This bias may also be mitigated by multiple observations on the benefit outcomes for parents, if this reduces the direct effect of z_i in the determination of the welfare participation for their offspring.

Our hope is that these two estimation procedures will produce lower and upper-bound estimates of the true intergenerational correlation coefficient. Better information on the welfare histories of the parents should substantially reduce the distance between these limits, and produce a more accurate picture of the true link between the welfare dependency of parents and their offspring.

IV. Empirical Analysis

Data from the Christchurch Health and Development Study (CHDS) is used to produce the OLS and IV estimators for the correlation coefficient in welfare participation between parents and their offspring. The CHDS is an excellent data source for the task at hand. It is a longitudinal study of approximately 1,200 children born in Christchurch area hospitals between April and August 1977. Christchurch is the largest metropolitan area on New Zealand's South Island. These families were interviewed annually from birth to age 14 of the children in this cohort. These youth were also surveyed at ages 16, 18 and 21.

The main dependent variable for this study is the computed frequency of welfare participation for youth in the five-year period between their 16th and 21st birthdays (Y_i). This variable is constructed primarily from retrospective data taken from the CHDS surveys at ages 18 and 21. We know the number of months in which these youth received social welfare benefits between their 16th and 18th birthdays from either the Unemployment Benefit (UB) or Domestic Purposes Benefit (DPB) programme.[7] We also know the number of months in which they received benefits from the UB programme between their 18th and 21st birthdays.[8]

Table 1 provides descriptive statistics on our welfare participation measure of youth over the five years preceding their 21st birthdays. Our sample consists of 847 youth.

It consists of all parents and their offspring in the CHDS who provided valid information for the purposes of this study. These youth completed the interviews at ages 18 and 21, and their parents completed all earlier annual interviews when these young people were aged 1 through 14. The mean benefit propensity for youth in this sample is 0.101. There is approximately a 10% probability that youth in the CHDS will be receiving a social welfare benefit at any point in time between their 16th and 21st birthdays.

[Table 1]

Yet, this mean benefit propensity varies substantially with respect to several demographic characteristics. Females (11.7%) are more likely than males (8.5%) to be welfare participants. These frequencies are significantly different from one another at better than a 5% level. Maori (19.1%) are more than twice as likely as non-Maori (8.8%) to be on a benefit. Youth with no school or post-school qualification (29.3%) are more than four-times as likely as those with a qualification (6.4%) to be a welfare participant. These means are significantly different from one another at better than a 1% level.

Table 2 shows the extensive heterogeneity in benefit participation among CHDS youth. More than one-half of these individuals did *not* receive social welfare benefits in any month over this five-year period (52.2%). Nearly one-quarter (22.8%) received benefits no more than 10% of the time. Only about one in twenty youth (5.3%) were on a benefit in more than one-half of the months over the sample period. Slightly more than one-half of the time spent on welfare among these young people was concentrated among 77 individuals (9.1% of the sample).

[Table 2]

The main independent variable for this study is the computed frequency of welfare participation for the parents of youth in this sample (X_i).[9] The CHDS contains annual snapshots on the benefit status of up to two parents from ages 1 through 14 of the child. We measure this variable as the proportion of years in which the child lived in a household that was receiving social welfare benefits at the time of each survey.[10] The mean benefit propensity for parents in this sample is 0.137. Like the benefit propensity for youth, there is a great deal of heterogeneity in welfare participation among parents. Nearly two-thirds of parents (62.3%) never received a benefit over this fourteen-year period. More than one-sixth of families (17.6%) received a benefit between one and three years. More than one-ninth of families received a benefit between four and eight years (12.0%). Only about one in twelve families were on a benefit for nine or more years (8.0%). Exactly one-half of the time spent on welfare by these parents was concentrated among 75 families (8.9% of the sample).

OLS Estimation: The Advantage of Long-Run Welfare Histories for Parents

With the construction of our main dependent and independent variables, we can begin our regression analysis. Table 3 shows the extent of the potential downward bias associated with measurement error in the 'permanent' benefit propensities of the families in which these children were raised. These results come from two-variable OLS regressions, where both the dependent and independent variables (y_i and x_i) have been standardised to have zero means and unit variances within the sample.

The first column of numbers in this table displays the results from 14 separate regressions where only a single 'snapshot' of the parents' benefit status is used from one interview at ages 1 through 14 of the child. The estimated correlation coefficients range from 0.214 at age 2 to 0.320 at age 13. All of these estimated parameters are significantly different than zero at better than a 1% level. The mean of the estimated correlation coefficients across the 14 years is 0.262. There appears to be a slight increase in these intergenerational correlations as the time interval narrows between the observation of the benefit histories of parents and their children.

[Table 3]

As we broaden our observation window to capture additional years of possible benefit participation of the parents, the estimated correlation coefficients increase substantially. The mean estimated correlation is 0.300 if either 3 or 4-year averages are used in the construction of the parents' benefit propensity. This mean intergenerational correlation increases to 0.334 if 7-year averages are used, and finally to 0.373 when the entire 14-year window is exploited.

These results are consistent with a gradual weakening in the bias associated with measurement error in the permanent benefit propensity of parents. The message is that long panels are necessary to fully capture the extent of the intergenerational relationship in welfare histories. In fact, some measurement error undoubtedly remains even with the 14 years of information on the benefit histories of the parents. The reason is that these are just 14 separate snapshots of benefit incidence at the time of the annual surveys. More continuous information on welfare histories between the annual surveys might result in even higher estimated correlation coefficients.

IV Estimation: A Possible Upper-Bound Estimate on the Intergenerational Correlation Coefficient

Now consider the use of family background characteristics as possible instrumental variables for this analysis. The problem is that valid instruments are difficult to find a priori. They should influence the benefit propensity of the parents, but have no direct impact on the benefit propensity of their offspring. The CHDS provides a wide array of personal and family background characteristics that might serve in this capacity. We choose to focus on the structure of the child's household and the qualifications of the parents. The first variable is the proportion of years between the ages of 1 and 14 for the child in which there was only a single parent in the household. Six dummy variables are used to indicate the school and post-school qualifications obtained by both parents.[11] We expect that the benefit propensity of the parents will be closely related to household structure and educational attainment. The question is whether or not these same factors affect the benefit propensity of the offspring, once we control for the welfare participation of the parents.

Table 4 displays the results from both OLS regressions. The key explanatory variable in the regression on the benefit propensity for parents, reported in the first column of this table, is the presence of a single parent in the household. The estimated coefficient on this variable is 3.538, with an estimated standard error of only 0.101. As expected, the qualifications of parents have consistently negative effects on the benefit propensity of the family. A school or post-school qualification by the mother, and a post-school qualification by the father, significantly reduce the probability that the family will be on the benefit at any point in time. In total, these few background factors account for nearly two-thirds of the variation (R^2 = 0.637) in the benefit propensity of the parents.

Yet, when these same measures of family structure and parental education are included in a multiple regression model on benefit propensity of youth, there is virtually no change on the coefficient attached to the benefit propensity of parents. This estimated correlation coefficient declines slightly from 0.373 in the two-variable regression model (last column of Table 3) to 0.372 in the multiple regression model. In both cases, it's significant at better than a 1% level. Only post-school qualifications for the mother and father have a direct, negative impact on the benefit propensity of youth. Overall, these family background measures capture some of the residual variation in the benefit propensities of youth. The R^2 statistic increases slightly from 0.139 in the earlier two-variable regression (0.373 squared) to 0.151 in this multiple regression.

Two important conclusions can be deduced from these regressions. Firstly, these family background measures appear on the surface to be fairly appropriate instruments. They influence the benefit history of the parents, but have little direct effect on the subsequent benefit histories of the children, once we control for the benefit propensity of the parents. Secondly, the positive association between the benefit propensities of parents and youth was not merely a proxy for omitted measures of family structure and parents education. The actual benefit history of the parents has a much more important and direct impact on what happens to their children.

[Table 4]

The first row of Table 5 compares the results from the OLS and IV estimation of the intergenerational correlation coefficient on the welfare participation of parents and their offspring in the simple two-variable model. The instruments used here are those included in the regressions reported in Table 4 (Proportion of Years with a Single Parent, and the six dummy variables on the school and post-school qualifications of parents). Note that the estimated coefficients under OLS and IV are both 0.373. Although the estimated standard error is slightly larger with the IV technique, both estimated coefficients are significantly different from zero at better than a 1% level. It was expected that these two estimation techniques would form lower and upper-bound estimates around the true value of ρ. This approach suggests that the true correlation coefficient is indeed around 0.373.[12]

[Table 5]

Separate regressions were estimated for the subamples demarcated by the same demographic characteristics of youth included earlier in Table 1. We had noted that mean benefit propensities vary significantly by gender, ethnicity and educational attainment. We now ask whether or not the intergenerational correlation coefficients also vary by these characteristics. Is there any evidence that the statistical relationships between the welfare participation of parents and their offspring vary by the personal characteristics of youth?

The remaining results reported in Table 5 for the separate regressions on these three sets of subsamples show a slightly stronger intergenerational link in benefit propensities for females, Maori and unqualified youth. The estimated correlation coefficients are greater than 0.4 for both females and Maori under both OLS and IV. This suggests that it may be important to look separately at the intergenerational transmission mechanism for welfare dependency across demographic groups. Finally, the relatively smaller correlations within the two educational groups for youth suggest that the benefit propensities of parents may influence that of their offspring indirectly through educational attainment.

We have very little information from this analysis thus far on exactly how this ‘transmission mechanism’ between the benefit histories of parents and their children might work. What intermediate steps might exist in this intergenerational relationship? No attempts are made here to estimate a formal structural model of this behavioural outcome. Instead, we experiment with a slightly less reduced-form model that allows the parents’ benefit propensity to influence the child’s benefit propensity indirectly through the educational attainment of the child.

Consider the following simple, recursive model.

$$q_i = \alpha_1 x_i + u_i \tag{12}$$
$$y_i = \alpha_2 x_i + \alpha_3 q_i + v_i \tag{13}$$

The dependent variable in equation (12) measures the quantity of education completed by youth (q_i). This is the number of years of school and tertiary education completed by age 21. If this variable is standardised to have a zero mean and unit variance, then the parameter (α_1) on the parents' benefit propensity is a correlation coefficient.

Equation (13) allows the parents' benefit propensity to have both a direct effect and an indirect effect, through the educational attainment of the offspring, on the youth's own benefit propensity. The parameters α_2 and α_3 are partial correlation coefficients.

If equation (12) is substituted into (13), we end up with the original reduced-form equation (1).

$$\begin{aligned} y_i &= \alpha_2 x_i + \alpha_3(\alpha_1 x_i + u_i) + v_i \\ &= \underbrace{(\alpha_2 + \alpha_1\alpha_3)}_{\rho} x_i + \underbrace{(v_i + \alpha_3 u_i)}_{\varepsilon_i} \end{aligned} \tag{14}$$

With estimates of α_1, α_2 and α_3, we can calculate the proportion of the overall intergenerational correlation between benefit propensities of parents and offspring that operates through the lower educational attainment of youth.

Table 6 presents the results from the separate estimation of equations (12) and (13). The benefit propensity of parents is negatively related to the educational attainment of their children. The estimated correlation is –0.306, and is significantly different from zero at better than a 1% level. The educational attainment of youth also negatively influences their benefit propensity. The estimated correlation is –0.315, and is significantly different from zero at a 1% level. Holding the years of education of youth constant, the *partial* correlation coefficient between the benefit propensities of parents and their offspring (0.276) is positive and significant.

[Table 6]

The product of these estimated correlation coefficients of α_1 and α_3 indicates the strength of this 'indirect' transmission mechanism. The result is 0.096 (-0.306 • –0.315), which accounts for approximately one-quarter (0.096/0.373 or 25.7%) of the overall correlation between the benefit propensities of parents and their offspring. The estimated correlation coefficient of α_2 measures the strength of the 'direct' transmission mechanism. Slightly less than three-quarters of this overall intergenerational relationship (0.276/0.373 or 74.1%) appears to operate outside of the educational attainment of youth.

V. Conclusions

This study uses recently available panel data in New Zealand, which provides extensive information on the benefit histories of both children and their parents, to estimate the intergenerational correlation coefficient in welfare participation. The Christchurch Health and Development Study (CHDS) follows the progress of around 1,000 children from birth to age 21.

Two estimation techniques are used to produce lower and upper-bound estimates for this parameter. The first procedure uses the average of the annual benefit participation of parents over fourteen consecutive years. This independent variable substantially increases the estimated correlation in benefit histories between parents and their children, and suggests that relatively long panels are necessary to accurately measure the extent of the transmission in welfare dependency from one generation to the next. Single snapshots on the welfare participation of parents and their children may substantially underestimate the true intergenerational correlation coefficient.

The second estimation procedure uses other family background characteristics as instrumental variables for the benefit propensity of the parents. We recognise that these variables may also directly influence the welfare participation of youth. As a result, this instrumental variable technique may tend to overestimate this intergenerational correlation coefficient. The idea is that this places an upper-bound estimate on the true parameter. The proportion of years in a single-parent household and the educational attainment of both parents explain nearly two-thirds of the variation in the benefit propensity of the family in which the child was raised.

When these same background factors are directly included in a regression on the benefit propensity of the youth, the estimated correlation coefficient on the benefit propensity of the parents is 0.372 and significant at better than a 1% level. This is virtually identical to the estimated correlation coefficient of 0.373 in the two-variable regression involving benefit propensities of parents and their offspring. The IV estimator is also estimated to be 0.373, and significant at better than a 1% level.

Thus, all of our estimation techniques point to an intergenerational correlation coefficient of just over 0.37 in New Zealand. Other regression results suggest that this transmission of welfare dependency from one generation to the next may be somewhat stronger among females and Maori youth. Furthermore, approximately one-third of this effect appears to operate through the lower educational attainment of children reared in families receiving social welfare benefits.

Table 1

Mean Benefit Propensities Among CHDS Youth

	N	$\overline{Y}$
Entire Sample	847	0.101
By Gender:		
Males	411	0.085*
Females	436	0.117*
By Ethnicity:		
Non-Maori	735	0.088**
Maori	112	0.191**
By Education:		
Some Qualification	708	0.064**
No Qualification	139	0.293**

** Means significantly different at 1% level.
* Means significantly different at 5% level.

Notes: The 'benefit propensity' Y_i is the proportion of months over the five-year interval between the 16th and 21st birthdays of these youth in which they received social welfare benefits.

Table 2

Dispersion in Benefit Propensities Among CHDS Youth

	N	Proportion of Sample
$Y_i = 0$	442	0.522
$0 < Y_i \leq 0.05$	115	0.136
$0.05 < Y_i \leq 0.10$	78	0.092
$0.10 < Y_i \leq 0.20$	73	0.086
$0.20 < Y_i \leq 0.30$	46	0.054
$0.30 < Y_i \leq 0.50$	48	0.057
$0.50 < Y_i \leq 0.75$	33	0.039
$0.75 < Y_i \leq 0.99$	8	0.009
$Y_i = 1$	4	0.005
Total	847	1.000

Notes: The 'benefit propensity' Y_i is the proportion of months over the five-year interval between the 16^{th} and 21^{st} birthdays of these youth in which they received social welfare benefits.

Table 3

Estimated Correlation Coefficients on
Benefit Propensities Between Parents and Youth
The Importance of Measurement Error in the Independent Variable

Child's Age When Parents' Benefit Propensity X_i is Measured:	Using a Single Year	Using 3 or 4-Year Means	Using 7-Year Means	Using a 14-Year Mean
1	0.234** (0.033)			
2	0.214** (0.034)	0.265** (0.033)		
3	0.240** (0.033)			
4	0.277** (0.033)		0.321** (0.033)	
5	0.270** (0.033)	0.302** (0.033)		
6	0.231** (0.033)			
7	0.253** (0.033)			0.373** (0.032)
8	0.278** (0.033)			
9	0.229** (0.033)	0.282** (0.033)		
10	0.254** (0.033)			
11	0.263** (0.033)		0.347** (0.032)	
12	0.313** (0.033)	0.349** (0.032)		
13	0.320** (0.033)			
14	0.292** (0.033)			
Column Means	0.262	0.300	0.334	0.373

** Significantly different from zero at 1% level.
* Significantly different from zero at 10% level.

Notes: Both the dependent variable (y_i) and independent variables (x_i) have been adjusted to have zero means and unit standard deviations in this sample. As a result, the intercept term in each regression has been suppressed, and the single parameter estimate can be interpreted as a correlation coefficient.

Table 4

OLS Regression Results on Benefit Propensities of Parents and Youth

Independent Variables	Dependent Variables: Benefit Propensity of Parents (x_i)	Dependent Variables: Benefit Propensity of Youth (y_i)
Constant	-0.231** (0.041)	0.128* (0.064)
Benefit Propensity of Parents (x_i)	---	0.372** (0.053)
Proportion of Years with Single Parent	3.538** (0.101)	-0.181 (0.242)
Mother has School Qualification	-0.160** (0.050)	-0.061 (0.077)
Mother has Post-School Qualification	-0.253** (0.061)	-0.183* (0.095)
Mother has University Degree	-0.102 (0.107)	-0.188 (0.164)
Father has School Qualification	-0.052 (0.050)	-0.032 (0.076)
Father has Post-School Qualification	-0.130* (0.069)	-0.197* (0.107)
Father has University Degree	-0.106 (0.077)	-0.069 (0.118)
N	847	
R^2	0.637	0.151
Adjusted R^2	0.634	0.143

** Significantly different from zero at 1% level.
* Significantly different from zero at 10% level.

Table 5

Estimated Correlation Coefficients on Benefit Propensities Between Parents and Youth: OLS and IV Estimation Techniques

	N	Parameter Estimates OLS	IV
Entire Sample	847	0.373** (0.032)	0.373** (0.040)
By Gender:			
Males	411	0.321** (0.047)	0.325** (0.059)
Females	436	0.432** (0.043)	0.420** (0.054)
By Ethnicity:			
Non-Maori	735	0.315** (0.035)	0.307** (0.045)
Maori	112	0.454** (0.085)	0.420** (0.103)
By Education:			
Some Qualification	708	0.198** (0.037)	0.231** (0.047)
No Qualification	139	0.329** (0.080)	0.317** (0.099)

** Significantly different from zero at 1% level.
* Significantly different from zero at 10% level.

Notes: The six instrumental variables used in the regression results reported in the second column are the Proportion of Years with a Single Parent, and the six dummy variables on the school and post-school qualifications of mothers and fathers used in the regressions reported in Table 4.

Table 6

OLS Regression Results on the Educational Attainment and Benefit Propensity of Youth

Independent Variables	Dependent Variables: Years of Education of Youth (q_i)	Dependent Variables: Benefit Propensity of Youth (y_i)
Benefit Propensity of Parents (x_i)	-0.306** (0.033)	0.276** (0.032)
Years of Education of Youth (q_i)	---	-0.315** (0.032)
N	847	
R^2	0.094	0.229
Adjusted R^2	0.094	0.228

** Significantly different from zero at 1% level.
* Significantly different from zero at 10% level.

Notes: All three variables used in this estimation (y_i, x_i and q_i) are adjusted to have zero means and unit standard deviations. The estimated parameters are thus partial correlation coefficients. The variable q_i is the number of years of formal education completed by the youth at age 21.

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Footnotes

1. Maloney (2000) for a description of the relatively minor reforms made to these welfare programmes in the 1980's and 1990's. Substantial changes to a welfare system over time might reduce the measured intergenerational correlation coefficient.

2 Yet, Duncan et. al (1988) note that welfare receipt may be seen as an 'investment' in children that provides parents with more resources at their disposal to help improve various aspects of the lives of their children. This would tend to weaken, and even reverse, the hypothesised positive link between the welfare participation of parents and their children. Both the sign and magnitude of this intergenerational correlation coefficient are ultimately empirical issues.

3 This is the case in Australia as well. However, the reason for the lack of research of this type in Australia probably can be attributed to the lack of data from suitably long panels.

4 Two examples are the National Longitudinal Survey of Youth (NLSY) and the Panel Study of Income Dynamics (PSID).

5 This is because there were essentially no universal income transfer programs for men in the US. Restrictive eligibility criteria and term limits exist under the Unemployment Insurance program. The Food Stamp program provides only in-kind benefits, and historically few men have received cash benefits under the AFDC program for 'unemployed parents'.

6 See the more general derivation of this IV estimator in the Appendix to Solon (1992).

7 Nearly all youth receiving a social welfare benefit were participating in the UB or DPB programmes. Thus, we can largely ignore other minor income transfer programmes in New Zealand. Both UB and DPB are essentially negative income tax programmes. Maximum weekly benefit levels vary by age, marital status and number of dependent children in the family. Weekly benefits abate away at benefit reduction rates that rise from 30% to 70% once earned income exceeds certain thresholds. DPB is available to single parents, and does not generally carry a work requirement. Everyone else is at least potentially eligible for UB, which does carry a work requirement.

8 The problem is that no direct information is available on the receipt of benefits under the DPB programme between the ages of 18 and 21. However, we use other information in the CHDS to infer DPB recipiency over these three years. Firstly, we know whether or not the youth received income from DPB at the time of the interview at age 21. Secondly, we know the 'events' that would trigger the eligibility for this programme over the previous three years. We assume that DPB recipients at age 21 were receiving benefits under this

programme since age 18 during the months in which a dependent child was living in their household, they were unmarried and they were not working full-time. This decision rule altered the measured welfare participation of 33 individuals (all female) in our sample (3.9% of our total sample).

9 The term ‘parents’ refers more generally to the parental figures living in the child’s household at the time of each annual interview. This could include either stepparents or other custodial adults.

10 All of the social welfare benefits received by parents in this sample came from either the Unemployment Benefit (UB) or Domestic Purposes Benefit (DPB). See footnote 7 for a brief description of the structure of these income transfer programs.

11 Information is available on the educational qualifications of both parents in this sample, even if the child subsequently lived in a single-parent household.

12 This does not eliminate the possibility that other forms of bias may exist in this situation. For example, misspecification of the overall structural model determining the welfare participation of parents and their offspring could overestimate this correlation coefficient.