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Intergenerational Income Mobility in a Less-Developed, High-Inequality Context: The Case of Chile*

Javier I. Nunez and Leslie Miranda

Abstract

This paper studies the magnitude of intergenerational income mobility in less developed, high inequality Chile. Following a known methodology where fathers' incomes are predicted from standard income determinants such as education and occupation, we get comparable estimates of the intergenerational income elasticity in the range of 0.57 to 0.74 and 0.63 to 0.76 for ages 25-40 and 31-40, respectively. These values place Chile at the high end of the available international evidence. Considering Chile's high income inequality, this finding supports the hypothesis proposed in the literature of an inverse relationship between cross-sectional income inequality and intergenerational income mobility.

KEYWORDS: intergenerational income mobility, social mobility, equality of opportunity

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1 Introduction

In the last two decades there has been a significant increase in the amount of research devoted to the study of intergenerational income mobility in a variety of countries. This research has expanded the understanding of the dynamic and intergenerational aspects of social inequality, and has enabled assessing the degree of social mobility and equality of opportunity in these countries in comparative perspective.[1] Nonetheless, a large proportion of these research efforts have focused on a restricted number of mostly developed and relatively egalitarian countries, and therefore little is yet known about the features of intergenerational income mobility in the less-developed world. This paper attempts to contribute to this literature by studying the magnitude of intergenerational income mobility in Chile, a less-developed country that is among the nations with the highest income inequality in the world.[2]

One reason for the relative neglect of the less developed countries in the research of intergenerational mobility has been the limited availability of adequate country-level intergenerational datasets. To overcome this problem, in this paper we follow the methodology first developed by Björklund and Jäntti (1997) for studying intergenerational income mobility when parents' incomes are not directly available. This methodology requires datasets containing determinants of the parents' income (such as schooling and occupation), whereby the fathers' income predictions can be obtained. This approach has since been employed in many countries, providing valuable empirical evidence for studying intergenerational income mobility in comparative perspective. Following this approach, this paper explores the magnitude of intergenerational income mobility in Chile, and discusses the results in the context of the available international evidence and Chile's high income inequality.

The rest of the paper is organized as follows. Section two presents the theoretical framework, the empirical strategy and the dataset employed in this study. Section three presents and discusses the main results in comparative perspective, and section four concludes.

[1] The notions of intergenerational social mobility and equality of opportunity are related concepts, as more equality of opportunity reduces the influence of an individual's socioeconomic background on his economic achievements in adulthood, which would be reflected in a higher level of intergenerational economic mobility. See for example Bourguignon, Ferreira and Menendez (2007) for a discussion on empirical approaches to "equality of opportunity", and Núñez and Tartakowsky (2007, 2010) and Contreras et al. (2009) for empirical assessments in Chile.

[2] Chile's income inequality is on the high end of the international spectrum, with a Gini coefficient of 0.55. For an excellent account of economic and social inequality in Chile and Latin America, see De Ferranti et al. (2003).

2 Empirical strategy and data

The purpose of this paper is to establish the level of intergenerational economic mobility in Chile employing a methodology that allows international comparisons, which we describe below.

Assume that the permanent incomes of a sample of fathers and their offspring were observed. Then, the following log-linear relationship between the permanent income of father and son could be estimated by OLS:

$$Y_{si} = \beta_0 + \beta_1 Y_{fi} + \varepsilon_i \tag{1}$$

where Y_{si} denotes the log of the son's permanent income in family i and Y_{fi} the log of his father's permanent income, and ε_i is an error term independent of Y_{fi}. The parameter of interest β_1 represents the intergenerational income elasticity, that is, the elasticity of a son's permanent income with respect to his father's permanent income. Equation (1) can illustrate two extreme cases of interest. First, $\beta_1=0$ would depict a situation involving high intergenerational mobility, as the permanent income of sons in adulthood would show no statistical association with the permanent income of their fathers. At the other extreme, if $\beta_1=1$ there would be a situation of low intergenerational mobility, since a son born of a parent with an income of, say, x per cent above the mean will have, in expected value, an income exactly x per cent above the mean of his own generation. Hence, $1-\beta_1$ can be interpreted as a summary measure of the degree of intergenerational income mobility, or alternatively, as a measure of the "regression-to-the-mean" effect in the transmission of the socioeconomic status from parents to their offspring.

However, long-run incomes are not directly observed. Instead, data sets usually provide measures of *current* incomes or earnings. Solon (1992) and Zimmerman (1992) have shown that the use of income in a single year can underestimate the true intergenerational income elasticities due to the presence of transitory components in current income, especially in combination with the use of a homogeneous sample. A solution for reducing this bias relies on panel data on the income of fathers in order to obtain an average of their current income over several periods as a proxy of their permanent income. Solon (1992) shows that the inconsistency of the intergenerational elasticity coefficient diminishes with the number of years over which incomes are averaged.

Another methodological issue emerges when, as in this paper, the incomes of fathers are not directly available. In this context, a methodology proposed by Björklund and Jäntti (1997) for studying intergenerational mobility, and followed thereafter by several other studies, addresses this issue by using two separate samples in order to obtain predictions of fathers' incomes.[3] The first step of this approach consists in estimating earnings equations using an older sample of men in order to obtain estimated coefficients of key earnings determinants, such as schooling and occupation, for example. Then, the estimated coefficients can be employed to predict the income of the fathers of a sample of sons who have reported the relevant information about their fathers. Using these predicted incomes, the intergenerational income elasticity can then be estimated. This methodology is often referred to in the related literature as two-sample instrumental variables estimation (TSIV), or two-sample, two-stage least squares (TSTSLS).[4] To describe these procedures more formally, assume that the log of the current income of father of family *i* and his son at date *t* and *p* respectively, can be written as:

$$Y_{fit} = Y_{fi} + \alpha_1 Age_{fit} + \alpha_2 Age_{fit}^2 + \mu_{fit} \tag{2}$$

$$Y_{sip} = Y_{si} + \beta_2 Age_{sip} + \beta_3 Age_{sip}^2 + \mu_{sip} \tag{3}$$

where μ_{fit} and μ_{sip} incorporate transitory fluctuations in the current income of fathers and sons as well as measurement errors, and where *Age* of father and son is included to control for life-cycle effects in earnings. Let Z_{fi} denote a set of socio-demographic characteristics associated with permanent income (like education and occupation) of fathers from a sample of families $i \in I$ such that the father's permanent income can be described as $Y_{fi} = Z_{fi}\gamma + v_{fi}$, where v_{fi} is an unobserved term affecting permanent income independent of Z_{fi}. Then, from equation (2) the father's current income at time *t*, Y_{fit} can be written as:

$$Y_{fit} = Z_{fi}\gamma + v_{fi} + \alpha_1 Age_{fit} + \alpha_2 Age_{fit}^2 + \mu_{fit} \tag{4}$$

The term Y_{fit} is not observed in sample *I*. However, if there is a separate sample of adult men *J* from the same population as *I*, sample *J* can be used to provide an estimate of γ, namely $\hat{\gamma}$, which would be derived from estimation of equation (5) using the sample of adult men *J*, that is,

$$Y_{jt} = Z_{jt}\gamma + v_{jt} + \alpha_1 Age_{jt} + \alpha_2 Age_{jt}^2 + \mu_{jt} \tag{5}$$

[3] Björklund and Jäntti (1997) follow the previous contributions by Angrist and Krueger (1992) and Arellano and Meghir (1992).
[4] See for example Dunn (2007).

for $j \in J$. From an OLS estimation of (5) one can obtain predictions of the fathers' earnings in sample I from $\hat{Y}_{fit} = Z_{fi}\hat{\gamma} + \hat{\alpha}_1 Age_t + \hat{\alpha}_2 Age_t^2$, where Age_t denotes a standardized age across fathers.[5] This prediction can then be used in a second stage to estimate the intergenerational income elasticity coefficient β_I from:

$$Y_{sip} = \beta_0 + \beta_1(\hat{Y}_{fit}) + \beta_2 Age_{sip} + \beta_3 Age_{sip}^2 + \eta_i \tag{6}$$

where equation (6) controls for life-cycle effects in the son's current income at time p.

In this paper, the estimates of β_I are based on the estimation of equations (5) and (6) on separate samples as described in the following section. In particular, in the first stage we estimate an earnings equation as in (5), which allows for different schooling returns for different educational levels:

$$Y_{jt} = \gamma_0 + \gamma_1 dP_{jt} + \gamma_2 dS_{jt} + \gamma_3 dT_{jt} + \gamma_4 dU_{jt} + \alpha_1 Age_{jt} + \alpha_2 Age_{jt}^2 + \varepsilon_{jt} \tag{7}$$

where dP_{jt} is a dummy variable for complete primary education, dS_{jt} for secondary education, dT_{jt} for technical education, dU_{jt} for university education and ε_{jt} is a random error term. The reference educational category in (7) is no education or incomplete primary education. In another specification we also include four types of fathers' occupations under the assumption that occupation is a good instrument, in addition to schooling, for estimating the father's permanent income.

In a second stage, we use the estimated parameters in (7) and the information on fathers reported by their sons to predict their income at year t, as follows:

$$\hat{Y}_{fit} = \hat{\gamma}_0 + \hat{\gamma}_1 dP_{fit} + \hat{\gamma}_2 dS_{fit} + \hat{\gamma}_3 dT_{fit} + \hat{\gamma}_4 dU_{fit} + \hat{\alpha}_1 Age_t + \hat{\alpha}_2 Age_t^2 \tag{8}$$

Finally, we obtain the intergenerational income elasticity β_1 as in equation (6) above.

The methodology described above is subject to some well-known biases that have been identified in the related literature. As shown in Solon (1992, 2002), a first bias may arise if the father's schooling and occupation, apart from being correlated with the father's earnings, are also positive predictors of the son's earnings in their own right. Thus, in the second-stage regression, where schooling and occupation are used to predict the father's earnings but are not included as

[5] Accordingly, the prediction of the father's income is based only on the father's "permanent" socioeconomic characteristics in Z_{fi}, and not on his age at moment t.

separate explanatory variables of the son's earnings, the resulting omitted-variable problem would yield an upward bias in the intergenerational income elasticity.

Another source of bias is related to the ages of sons being considered for estimating equation (6). In particular, various studies have found that the estimated intergenerational elasticities increase substantially as sons' earnings are observed further on in their careers. Accordingly, studies that use earnings data of sons in the early stages of their life-cycle -- as in this study -- tend to underestimate the intergenerational income elasticity. This arises if the measurement error in the son's early earnings is negatively correlated with the long-run income, as can be expected.[6] These biases are thus expected to influence in opposite directions. However, due to the existence of these potential biases, below we compare the results obtained in this paper for Chile with the results of international studies that follow a similar methodology and are accordingly subject to the same kind of biases (see Table 2).

The data employed to estimate (5) and (6) comes from the *Encuesta de Caracterización Socioeconómica* (Socioeconomic Survey) (CASEN), a survey representative at national and regional levels, conducted regularly in Chile since its earliest version in 1987. The 2006 version of this survey covered approximately 70,000 households nationwide. The CASEN survey provides standard socioeconomic information on the heads of households and other adult members thereof, including gender, age, educational attainment, employment status, occupations, economic sectors and monthly incomes from wages, salaries and self-employment across the different economic sectors of the economy, in both formal and informal sectors, in urban and rural areas. We use the two earliest versions of CASEN (1987 and 1990) to estimate the earnings equations as in (7) in order to obtain the regression coefficients used for predicting the fathers' incomes. These surveys contain about 20,000 observations respectively, as shown in the Appendix Tables A and B.

In order to avoid selectivity issues associated with female participation in the labor market, we restrict our attention to intergenerational income mobility between fathers and their sons. The analysis of intergenerational income mobility between parents and their daughters is a subject for future research.

Earnings equations as in (7) were estimated for male individuals in the labor force in the 15-55 age range, with positive income and working at least 30 hours per week. Our sample of sons comes from the 2006 version of the CASEN survey. In this version, in addition to the regular demographic and socioeconomic questions, respondents were asked to provide information about the educational attainment (grouped into five categories: i) without education or incomplete primary education, ii) complete primary education, iii) complete secondary

[6] See for example Solon (2002), Haider and Solon (2006), Grawe (2006) and Dunn (2007).

education, iv) complete technical education and v) complete university education), types of occupation (grouped in four categories: i) employers; ii) employees and blue-collar workers; iii) self-employed; and iv) domestic workers) and other individual characteristics of their parents.

Our sample considered sons with positive incomes, working at least 30 hours per week, in the 25-40 age range, in order to avoid potential selectivity problems with individuals outside this age range.[7] Our final sample consisted of 11,186 pairs of fathers and sons.

The predicted incomes of fathers were estimated dividing the sample of sons into three sub-samples by age groups: 25-30, 31-35, and 36-40. We selected the samples to predict the fathers' incomes by assuming that the most important father-offspring socioeconomic transmissions mechanisms, in particular the influence of the father's socioeconomic status on his son's educational attainment,[8] occur when the son is about 6 to 21 years old. Hence, in order to estimate the father's income, we employ the 1990 version of the CASEN survey for the 25-30 and 31-35 age groups, and the 1987 version for the 36-40 age group.[9] Finally, the second-stage regression was estimated considering fathers up to 55 years of age in order to avoid potential selectivity issues in fathers above that age.

3 Results

Table 1 shows estimates of the intergenerational income elasticity coefficient β_1. Estimates in column 2 are obtained from the predicted income of fathers derived from schooling, and estimates in column 4 employ schooling and occupation. The auxiliary first-step earnings regressions employed are provided in the Appendix Tables A and B. Table 1 indicates that the predicted log incomes of fathers have a significant positive effect on their sons' (log) incomes.[10] For the whole sample

[7] In Chile the male participation rate is limited prior to age 25, and increases rapidly thereafter. On the other hand, we restrict our attention to sons who were 21 or younger in the first CASEN survey of 1987, most of whom would have still been affected by their parents' socioeconomic condition, in particular in relation to their involvement, continuation in (or exclusion from) secondary and tertiary education. In addition, most sons in that age range were still living with their parents.

[8] As suggested, for example, in Becker and Tomes (1979) and Solon (2004).

[9] The fathers' log incomes predicted from both CASEN surveys were transformed into deviations from their respective means, in order to make them comparable in the second stage of the methodology. The sons' observed incomes (in logs) in the 2006 survey were also expressed as deviations from the mean value.

[10] Following Murphy and Topel (1985) all the standard errors are corrected to address the fact that imputed regressors are measured with sampling error (see Murphy and Topel (1985) for more details about inference in two-step econometric models).

comprising ages 25 to 40, the estimated intergenerational income elasticity is around 0.57-0.74, depending on whether occupation is employed to predict the fathers' income in addition to schooling.

Table 1 also shows that the intergenerational income elasticity is lower for the 25-30 age group than for the 31-40 group. Although it might be tempting to interpret this as an indication of increasing social mobility in Chile in the last decades, this may also be the result of life-cycle effects in the sons' earnings that may yield a lower intergenerational income elasticity for younger individuals, as argued earlier. The intergenerational income elasticity for the 31-40 age group is indeed higher, in the range of 0.63 to 0.76.[11]

Table 1. Estimates of the intergenerational income elasticity, Chile.

Sons' age group	Father's income estimated from schooling	Obs.	Father's income estimated from schooling and occupation	Obs.
25-30	0.72	3,028	0.45	3,028
	[0.059]		[0.066]	
31-35	0.73	3,557	0.59	3,557
	[0.076]		[0.063]	
36-40	0.79	4,601	0.66	4,601
	[0.054]		[0.051]	
25-40	0.74	11,186	0.57	11,186
	[0.065]		[0.054]	
31-40	0.76	8,158	0.63	8,158
	[0.055]		[0.050]	

Note: Robust standard errors with Murphy and Topel (1985) correction in brackets.

Table 2 presents some international evidence on intergenerational income mobility. Column 5 shows the elasticities obtained from methodologies similar to the one employed in this work. This evidence indicates that the estimates of the intergenerational income elasticity for Chile reported in Table 1 are on the high

[11] These elasticities are higher than those previously obtained by Núñez and Risco (2004) and Núñez and Miranda (2010) for Greater Santiago (Chile's capital city), which are in the range of 0.52-0.58, and those estimated in Contreras, Fuenzalida and Núñez (2006) based on urban areas in Chile (0.67). However, these differences are to be expected, since large urban areas are likely to provide higher educational and labor opportunities for intergenerational mobility than the rural and small urban areas included in the nationwide representative data. Table C in the appendix reports results of intergenerational income mobility studies in urban Chile.

end of the spectrum of the available international evidence.[12] The estimated values for Chile are substantially higher than those of the United States and the United Kingdom, often regarded as being among the countries with the lowest intergenerational mobility among the developed nations.[13]

Table 2. Reported estimates of international intergenerational income elasticity.

Country	Study	Son's ages	Method	
			OLS	IV-TSTSLS
Australia	Leigh (2007)	25-54		0.2-0.3
Brazil	Dunn (2004)	25-34	0.53	0.69
Brazil	Ferreira and Veloso (2006)	25-64		0.58
Canada	Corak and Heisz (1999)	29-32	0.23	
Canada	Fortin and Lefebvre (1998)	17-59		0.19-0.22
Malaysia	Grawe (2001)	-		0.54
Finland	Osterbacka (2001)	25-45	0.13	
France	Lefranc and Trannoy (2005)	30-40		0.36-0.43
Germany	Wiegand (1997)	27-33	0.34	
Italy	Piraino (2007)	30-45		0.48
United Kingdom	Dearden, Machin, Reed (1997)	33		0.39-0.59
United States	Solon (1992)	25-33	0.29-0.39	
United States	Solon (1992)	25-33		0.45-0.53
United States	Björklund and Jänti (1997)	28-36		0.52
Sweden	Björklund and Jänti (1997)	29-38		0.28
Nepal	Grawe (2001)	-		0.44
Pakistan	Grawe (2001)	-		0.46

Source: Individual papers.

The figures for Chile also exceed the available estimates for other less-developed countries, namely Nepal, Pakistan and Malaysia. The one country that shows intergenerational income elasticities of a similar order of magnitude to Chile is Brazil. This is suggestive, considering the inverse relationship between cross-sectional income inequality and intergenerational social mobility that has

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[12] It is interesting to contrast this finding with evidence on *occupational* intergenerational mobility in Chile. Torche (2005) finds a significant degree of intergenerational mobility among non-elite occupations. Yet, this greater degree of intergenerational mobility in these occupational classes is claimed to be "largely inconsequential, because it takes place among classes that share similar positions in the social hierarchy of resources and rewards" (p. 422).

[13] See for example Piraino (2007) and Björklund and Jäntti (1997).

been proposed in the literature, along with some supporting evidence.[14] The comparatively low level of intergenerational income mobility for Chile and Brazil would be consistent with this hypothesis, considering the high income inequality of both countries compared with the international evidence.[15] It is also consistent with this hypothesis that less-developed Nepal, Pakistan and Malaysia have lower intergenerational income elasticities than Brazil and Chile, as well as lower Gini coefficients (47.2 , 30.6 and 49.2, respectively, UNDP, 2007).

4 Conclusions

This paper has studied the degree of intergenerational income mobility in Chile, thus providing an assessment of intergenerational mobility in the context of a less developed country with high income inequality. Following the methodology developed by Björklund and Jäntti (1997) that has been employed by several international studies, we find comparable estimates of the intergenerational income elasticity for Chile in the range of 0.57 to 0.74 for the 25-40 age group and 0.63 to 0.76 for the 31-40 age group. These values place Chile at the high end of the available international evidence, indicating a modest degree of intergenerational income mobility in comparative perspective. These values are of a similar order of magnitude to the ones reported for Brazil, which shares with Chile the feature of having a high level of income inequality in comparison with the international evidence. This finding is coherent with the hypothesis of an inverse relationship between cross-sectional income inequality and intergenerational income mobility that has been proposed in the theoretical and empirical literature.

The methodology used in this paper overcomes many of the limitations arising from the unavailability of parent-offspring income data, a common problem in less developed countries. The data required for this methodology is, however, more widely available. Hence, a more widespread use of this methodology across different countries and regions would significantly contribute to a better understanding of the dynamic and intergenerational aspects of social inequality worldwide in comparative perspective.

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[14] Expanding the theoretical framework in Becker and Tomes (1979), Solon (2002, 2004) finds that cross sectional income inequality and intergenerational mobility can be related because both dimensions depend positively on the earnings return of human capital investment and the mechanical heritability of income-relevant traits between generations. Yet, the connection between cross sectional inequality and intergenerational mobility is less than exact, for example if societies differ in the heterogeneity of ability or other endowments, which would increase inequality but not necessarily intergenerational mobility. For empirical evidence, see Björklund and Jäntti (1997), Corak (2006), Solon (2002), Dunn (2007) and Andrews and Leigh (2008).

[15] See for example De Ferranti et al. (2003) and Inter-American Development Bank (1999).

5 Appendix

Table A. Estimates of earnings equations using education and age as regressors.
Dependent variable: Log earnings

	CASEN 1987	CASEN 1990
Complete primary education	0.3256	0.3034
	[0.0158]	[0.0171]
Complete secondary education	0.9068	0.7651
	[0.0201]	[0.0201]
Complete technical education	1.1570	1.1343
	[0.0729]	[0.0716]
Complete university education	1.9301	1.6910
	[0.0335]	[0.0366]
Age	0.0875	0.0657
	[0.0045]	[0.0045]
Age^2	-0.0009	-0.0006
	[0.0001]	[0.0000]
Constant	7.8026	8.9858
	[0.0742]	[0.0750]
Observations	19,192	20,378
Adj. R-squared	0.34	0.29

Note: Robust standard errors are in brackets.

Table B. Estimates of earnings equations using education, age and occupation as regressors. Dependent variable: Log earnings

	CASEN 1987	CASEN 1990
Complete primary education	0.3203	0.2937
	[0.0156]	[0.0166]
Complete secondary education	0.8805	0.7437
	[0.0197]	[0.0191]
Complete technical education	1.1234	1.0932
	[0.0716]	[0.0661]
Complete university education	1.8753	1.6100
	[0.0326]	[0.0350]
Age	0.0886	0.0687
	[0.0043]	[0.0043]
Age^2	-0.0009	-0.0007
	[0.0001]	[0.0000]
Employer	1.2442	1.2646
	[0.0621]	[0.0531]
Self-employed	0.1046	0.2529
	[0.0154]	[0.0168]
Domestic worker	-0.4185	-0.3042
	[0.1113]	[0.1187]
Constant	7.7819	8.9307
	[0.0722]	[0.0721]
Observations	19,192	20,378
Adj. R-squared	0.37	0.35

Notes: Robust standard errors are in brackets.
“Employees and blue-collar workers” are the default occupations.

Table C. Estimates of the intergenerational income elasticity in urban Chile

Study	Database	Father's income predictors	Population	Son's cohort	Elasticity
Núñez and Risco (2004)	Employment and Unemployment Survey	Schooling	Greater Santiago	23-55	0.55
Contreras, Fuenzalida and Núñez (2006)	IALS	Schooling	National urban	23-55	0.67
Núñez and Miranda (2010)	Employment and Unemployment Survey	Schooling	Greater Santiago	23-65	0.54
Núñez and Miranda (2010)	Employment and Unemployment Survey	Schooling and occupation	Greater Santiago	23-65	0.52

Source: Individual papers.

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