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RECENT FINDINGS ON INTERGENERATIONAL INCOME AND EDUCATIONAL MOBILITY IN CHILE

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Recent Findings on Intergenerational Income and Educational Mobility in Chile

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Abstract

This paper provides new evidence on intergenerational mobility in Chile. Income mobility elasticities for Chile are the range of 0.52 to 0.67, which stand as fairly high in comparison with the international evidence. We also find that educational mobility is lower for the younger cohorts, suggesting an increase of intergenerational educational mobility in the last decades. Finally, we find evidence of a higher degree of intergenerational persistence at the two extremes of the income distribution, particularly at the top of the distribution. We suggest this mirrors the unusually high income concentration at the top of the Chilean income distribution.

Keywords:

Intergenerational mobility, schooling, mobility patterns.

JEL: D3, I2, J6.

Recent Findings on Intergenerational Income and Educational Mobility in Chile[1]

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This paper provides new evidence on intergenerational mobility in Chile. Income mobility elasticities for Chile are the range of 0.52 to 0.67, which stand as fairly high in comparison with the international evidence. We also find that educational mobility is lower for the younger cohorts, suggesting an increase of intergenerational educational mobility in the last decades. Finally, we find evidence of a higher degree of intergenerational persistence at the two extremes of the income distribution, particularly at the top of the distribution. We suggest this mirrors the unusually high income concentration at the top of the Chilean income distribution.

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Introduction

Much of the literature on inequality in Chile has focused on the inequality of outcomes, typically the distribution of income, but little is known about the country's levels

[1] We thank Dante Contreras, Florencia Torche, Claudio Sapelli, Marcela Perticara, Osvaldo Larranaga, David Bravo, Claudia Sanhueza, Cristina Risco and Marcelo Fuenzalida for previous discussions, comments and insights. This research has been funded by the Interamerican Development Bank.

of inequality of opportunities and its evolution in recent times. A common approach to assess a country's degree of equality of opportunities is the notion of intergenerational mobility: a higher level of equality of opportunities is expected to decrease the effect of an individual's early socioeconomic background on his economic achievement in adulthood, implying therefore a higher level of intergenerational economic mobility, or alternatively, a lower level of intergenerational transmission of the relative socioeconomic status from parents to their offspring.

This paper attempts to present and examine part of the recent research on intergenerational income mobility in Chile. This paper is mostly descriptive in nature. The main questions that we address are the following. What is the level of intergenerational income mobility in Chile? How does it compare with the existing international evidence? Can anything be suggested about the trend in intergenerational mobility in Chile in the past decades? Finally, what are some of the salient patterns of the process of intergenerational mobility in Chile? and how do these patters compare with the international evidence? Since the research in intergenerational mobility is rather recent in Chile, there is a significant amount of issues ahead yet to be addressed. This work proposes and discusses some of them.

The next section motivates the paper by providing evidence on the conceptual and empirical distinctions between the notions of inequality of income vs. inequality of opportunities. Section 2 presents the theoretical framework, section 3 describes the empirical strategy, section 4 describes the dataset we use in this study, and section 5 presents the mail results. Section 6 concludes.

1. Inequality of Outcomes vs. inequality of Opportunities in Chile

There has been a long debate about whether inequality, and the policies designed to deal with it, must focus on the inequality of outcomes or on the inequality of opportunities.

Advocates of the latter stress that inequality of outcomes (typically income) depend not only on circumstances that are beyond the control of the individual, such as parental background, but also on aspects that are under his or her control, such as effort, choices and so on. Moreover, some authors have suggested that from a moral standpoint, public policies should focus on equalizing opportunities instead of outcomes (incomes). In a seminal paper, Bourguignon, Ferreira and Melendez (2003) have developed a methodology to measure the proportion of the income distribution that is explained by inequality of circumstances of origin such as parental schooling, parents' occupation and race in Brazil. The methodology considers not only the direct effect that circumstances have on earnings, but also the indirect effect of circumstances on earnings through the accumulation of human capital. They found that the Gini coefficient is reduced in up to 10 percentage points (about 20 per cent of the Gini) after equalizing the set of circumstances mentioned earlier.

Table 1: Effect on Gini Coefficient of Equalizing Circumstances in Chile (Greater Santiago)

Cohort	24-37	38-51	52-65	24-65
Gini Coefficient	0.45	0.51	0.52	0.50
	Gini after Equalizing Circumstances			
Partial Effect	0.38	0.46	0.43	0.43
Total Effect	0.36	0.44	0.4	0.41

Source: Núñez and Tartakowsky (2006)

Núñez and Tartakowsky (2006) apply Bourguignon et. al. (2003) methodology to Chilean data and they find similar results to their study for Brazil. Table 1 shows the effect on the Gini coefficient for Greater Santiago of equalizing parental schooling, head of households age, household size, household composition (single versus biparental), and parents' job vulnerability. Even though there can be many relevant unobserved circumstances, these results do suggest that important circumstances, such as those mentioned above, play an important yet limited role in shaping income distribution. Hence, income distribution indicators would only reflect in part a society's degree and evolution of equality of opportunities, as they would be affected by other factors. In this context, perhaps a closer way of studying equality of opportunities is to examine the intergenerational income mobility, issue that we address next.

2. Theoretical framework

Following the previous literature, this paper analyzes intergenerational income transmission using a simplified version of model suggested by Becker and Tomes (1979). This model assumes that a family only consists of one individual at each generation. Consider two generations within a given family, father and child. Individual permanent income *Y* is assumed to derive from two components: individual endowment of human capital and individual ability denoted by *A*. Becker and Tomes assumes that a child's endowment of human capital is a result of his father's optimal allocation of his permanent income, where the father's utility depends of his own consumption and the child's permanent income. This framework yields the following relationship between father's and child's permanent income:

$$Y^{child} = \phi Y^{father} + \theta A^{child} \qquad (1)$$

This equation implies that father's permanent income has a positive causal influence on child's earnings or income captured by ϕ parameter. Equation (1) would also imply a second source of earnings correlation if child's ability is correlated to father's ability. Parameter θ can be interpreted as a causal effect of previous generations to the next that can be independent of father's investment decisions and budget constraints. This parameter will encompass all aspects of earnings determinants that money cannot buy, such as innate cognitive abilities, preferences or social networks, among others.

It is important to note that a simple regression of child's income on father's income will capture both transmission mechanisms. Hence, standard OLS estimates of intergenerational earnings transmission coefficient will provide an upward biased estimate of the "pure" causal effect of parental income on child's income. In this paper, we do not separate both effects[2]. Instead, we are interested in the estimation of reduced-form of intergenerational earnings regression. It constitutes, however, an important descriptive measure of the extent of intergenerational earnings mobility.

3. Empirical strategy and data

From the previous framework, if long-rung economic status were directly observed, the following log-linear relationship between the permanent income of father and child can be estimated by OLS:

$$Y_i^{child} = \beta_0 + \beta_1 Y_i^{father} + \varepsilon_i \tag{2}$$

where Y_i^{child} denotes the log of child's permanent income in family i and Y_i^{father} the log of his father's permanent income, and ε_i is an error term independent of Y_i^{father} usually assumed to be distributed as $N(0,\sigma^2)$. Our parameter of interest β_1 represents the elasticity of a child's long-run income with respect to his father's long-run income. There are two extreme cases. First if $\beta_1=0$ there is complete mobility. The income of the child shows no statistical association with the father's income. At the other extreme, if $\beta_1=1$ there is complete immobility, as a child born from a parent with an income *x* per cent above the mean will have an income exactly *x* per cent above the mean of his own cohort.

However, long-run incomes are not directly observed. Instead, most data sets only provide measures of current incomes or earnings. Solon (1992) and Zimmerman (1992) have shown that the use of income in a single year can seriously underestimate the true intergenerational transmission coefficient due to the presence of transitory components in current income, especially in combination with the use of a homogeneous sample. A solution to reduce this bias relies on panel data on fathers' income to obtain an average of father's current income over several years as a proxy for permanent income. Solon (1992)

[2] However, see Contreras, Fuenzalida and Núñez (2006) for an attempt to separate both effects using Chilean data.

shows that the inconsistency of the transmission coefficient estimator diminishes with the number of years over which incomes are averaged.

Another problem emerges when, as in this paper, there is no actual income information of father-child pairs. In this context, a solution proposed by Arellano and Meghir (1992) and Angrist and Krueger (1992) and followed by Bjorklund and Jantti (1999) for mobility studies is to use information from two samples: First, earnings equations can be estimated on an older sample of men in order to obtain coefficients of some earnings determinants, such as schooling, experience and occupation, for example. Then, these coefficients can be used to predict the income of the fathers of a sample of sons, employing the socio-demographic characteristics of the fathers reported by their sons. This technique is often known as two-sample instrumental variables estimation (TSIV). [3]

Assume that the log of father and son's current income at date t can be written as:

$$Y_{it}^{father} = Y_i^{father} + \mu_{it}^{father} \quad (3)$$

$$Y_{it}^{child} = Y_i^{child} + \mu_{it}^{child} \quad (4)$$

where μ_{it}^{father} and μ_{it}^{child} incorporates transitory fluctuations in father and child's current income and measurement error. Let Z_i^{father} denote a set of socio-demographic characteristics (like age, education, occupation, among others) of fathers from a sample of families i $\in$ I and assume that Y_{it}^{father} can be written as:

$$Y_{it}^{father} = Z_i^{father}\gamma + v_i^{father} + \mu_{it}^{father} \quad (5)$$

where v_i^{father} is independent of Z_i^{father}. Y_{it}^{father} it is not observed in sample I. Yet, if there exists a sample J from the same population as I, it can be used to provide an estimate of γ, $\hat{\gamma}$, derived from estimation of following equation:

[3] Although Dunn (2004) refers to this method as "two samples two stage least squares".

$$Y_{jt}^{father} = Z_j^{father}\gamma + v_j^{father} + \mu_{jt}^{father} \quad (6)$$

for j $\in$ J. From an OLS estimation of (6) one can obtain predictions of father's earnings in sample I: $\hat{Y}_{it}^{father} = Z_i^{father}\hat{\gamma}$. This prediction can in turn be used to estimate β_1 since equations (2), (3), (4) and (6) imply:

$$Y_{it}^{child} = \beta_0 + \beta_1\left(Z_i^{father}\hat{\gamma}\right) + \eta_{it} \quad (7)$$

where $\eta_{it} = \varepsilon_i + \mu_{it}^{child} + \beta_1 v_i^{father} + \beta_1\left(Z_i^{father}\left(\gamma - \hat{\gamma}\right)\right)$.

In this paper, the estimates of β_1 are based on the estimation of equations (6) and (7) on separate samples as we describe in the following section. In particular, in a first stage we estimate a Mincer version of equation (6) that allows for different schooling returns by educational level[4]:

$$Y_{js}^{father} = \gamma_0 + \gamma_1 S_{js} + \gamma_2 d_1(S_{js} - 8) + \gamma_3 d_2(S_{js} - 12) + \gamma_4 Exper_{js} + \gamma_5 Exper_{js}^2 + \varepsilon_{js} \quad (8)$$

where S_{js} represents the years of schooling of father in year s, [5] $Exper_{js}$ stands for father's potential experience[6] and ε_{js} is a random error term. In addition, dummy variables are defined as:

$$d_1 = \begin{cases} 1 & if \quad S > 8 \\ 0 & otherwise \end{cases} \qquad d_2 = \begin{cases} 1 & if \quad S > 12 \\ 0 & otherwise \end{cases}$$

In another specification we also use information of fathers' occupation that comes from a new survey realized in the middle of year 2006 to a sub-sample of the June 2004 version of Employment and Unemployment Survey of the Universidad de Chile, under the assumption that occupation is a good instrument to estimate the father's permanent income.

[4] In Chile, elementary education consists of the first eight years and secondary school consists of four additional years.
[5] The s year corresponds to time when father were taken investment decisions on his child's human capital.
[6] Potential experience is defined as: age minus years of schooling minus 6.

In a second stage, we use the estimated parameters in (8) and fathers' information reported by the sons to predict the fathers' income, as follows:

$$\hat{Y}_{is}^{father} = \hat{\gamma}_0 + \hat{\gamma}_1 S_{is} + \hat{\gamma}_2 d_1 (S_{is} - 8) + \hat{\gamma}_3 d_2 (S_{is} - 12) + \hat{\gamma}_4 Exper_{is} + \hat{\gamma}_5 Exper_{is}^2 \qquad (9)$$

Hence, we obtain the intergenerational income elasticity β_1 from:

$$Y_{it}^{child} = \beta_0 + \beta_1 \hat{Y}_{is}^{father} + \beta_2 age_{it} + \beta_3 age_{it}^2 + \eta_{it} \qquad (10)$$

where age_{it} stands for child's age and controls for life-cycle profiles in child's earnings.

Data

At the outset it must be stressed that unfortunately Chile does not have a nationally-representative survey of sons with data including fathers' characteristics that are likely to be correlated with long run earnings[7]. Our dataset comes from the Employment and Unemployment Survey for the Greater Santiago conducted annually by Universidad de Chile since 1957 and it is applied to approximately 4,000 households. This is important because the intergenerational income elasticities are likely to underestimate the nationwide elasticities, as the result of having a more homogeneous sample and leaving out parts of the population were intergenerational socioeconomic persistence is expected to be higher, such as the remote and rural areas as well as small urban areas, as we will discuss later.

In order to avoid selectivity issues associated with female participation in labor market, we focus only on father-son intergenerational income mobility. The analysis of intergenerational income mobility involving mothers and daughters remains as future research. However, we consider sons as well as daughters when we examine intergenerational educational mobility in section 5.2.

[7] The CASEN survey (Encuesta de Caracterización Socioeconómica) may be a possible database with this information but only for a sample of sons that live with their parents. Hence, problems arise from potential selection bias. Currently, we are working with this data and we are trying different solutions to address this problem.

The Employment and Unemployment Survey provides information on gender, age, educational level, employment status, occupational position, economic sectors and monthly income from wages, salaries and self-employment. All this information is relevant to estimate coefficients of Mincer equations that are employed to predict the unobserved income of fathers. Mincer equations like (8) were estimated for the male labor force in 15-65 age range with positive income and working at least 30 weekly hours.

Our sons' sample comes from the June 2004 version of the survey. In this year, additional to demographic and economic data, respondents were required to provide information about education and individual characteristics of their parents. We consider sons in the 23-65 age range to control for selectivity problems in samples outside this range. We eliminate unemployed and inactive individuals, with no positive incomes and missing parental information. Our sample was composed by 649 father-son pairs in the relevant age range.

Fathers' predicted income was estimated dividing the sons' sample in three sub-samples by age cohort: 23-34, 35-44, 45-54 and 55-65. We select our fathers' samples by assuming that the father's relevant investment decisions in human capital, which are expected to be a major source of socioeconomic transmission across generations, were taken approximately when the child was between 6 and 18 years old. These years correspond to 1987, 1977, 1967 and 1958 versions of the Employment and Unemployment Survey for the 23-34, 35-44, 45-54 and 55-65 cohorts, respectively. Those are periods of relative economic stability; hence, estimated coefficients of Mincer equations for mentioned years are similar to those of adjacent years.

A second data source comes from a new survey realized in the middle of year 2006 to a sub-sample of males previously surveyed in the Employment and Unemployment Survey of June 2004. In this new survey, individuals were required to provide additional information about specific occupation and other individual characteristics of their parents, as well as diverse family background information corresponding to period when they were about 15 years old. We use additional information about father's occupation[8] to estimate another specification of Mincer equations under the assumption that occupation is a good instrument to estimate the father's permanent income, in addition to fathers' schooling and potential experience.

4. Results

4.1 Estimates of intergenerational income mobility for Chile

Table 2 reports intergenerational regression coefficients for labor incomes[9]. Results are reported for the full sample in 23-65 age range. Estimates in this table are obtained using father's education, potential experience and occupation as predictors for father's income. First-step income regressions are provided in Table A.1 and A.2 in the appendix.

Table 2: Estimates of Intergenerational Labor Income Elasticity by TSIV (Greater Santiago)

	Father's income predicted from:	
Cohort	Schooling and Experience	Schooling, Experience and Occupation
24-65	0.54	0.52

Table 2 indicates that fathers' predicted log income has a significant positive effect on their sons' log income. For the whole sample, the estimated elasticity is around 0.52-0.54[10] depending on the predictors employed.

Table 3 reports the results of various recent intergenerational income mobility studies. Note that many of them employ the Greater Santiago sample, excepting those that employ the SIALS database (Contreras et. al.) It is interesting to note that employing only the SIALS data for the Metropolitan Region (slightly larger than Greater Santiago) yield

[8] The 5-level occupational categories are: employer (1); self-employed (2); employee (3); blue-collar worker (4) (reference) and domestic (household) workers (5).
[9] The estimates using personal incomes yield the same global elasticity.
[10] This number is a weighted average of elasticities of each age group.

fairly similar results that the Greater Santiago studies. In this context, the 0.67 elasticity obtained from the national urban SIALS database seems closer (but perhaps still an underestimate) of the country's intergenerational income elasticity. Another piece of evidence that reinforces this idea is the pattern of income mobility in Brazil, according to Ferreira and Veloso (2004): The more prosperous and more urban Brazilian Southwest has lower income elasticity than the rest of the country, and much lower than the poorer Northeast region.

Table 3: Evidence on Intergenerational Income Mobility for Chile

Study	Database	Father's Income Predictors	Population	Son's cohort	Elasticity
Nuñez and Risco (2004)	Employment and Unemployment Survey	Potential experience, schooling	Greater Santiago	23-55	0.55
Nuñez and Risco (2004)	Employment and Unemployment Survey and CASEN (Fathers' sample)	Potential experience, schooling	Greater Santiago	23-35	0.43
Contreras, Fuenzalida and Núñez (2006)	SIALS	Schooling	Greater Santiago	23-55	0.58
Contreras, Fuenzalida and Núñez (2006)	SIALS	Schooling	National Urban	23-55	0.67
This study	Employment and Unemployment Survey	Potential experience, schooling	Greater Santiago	23-65	0.54
This study	Employment and Unemployment Survey	Potential experience, schooling, occupation	Greater Santiago	23-65	0.52

Table 4 summarizes some selected evidence on intergenerational mobility. As can be seen, Chile presents relatively low intergenerational income mobility as compared to other developing and developed countries. Levels of intergenerational mobility in Chile are somewhat similar to Brazil's.[11] Some authors have suggested and provided some evidence of an overall positive relationship between cross-sectional income inequality and intergenerational persistence of inequality. [12] From this perspective, the evidence for Chile seems consistent with this hypothesis considering that Chile has a particularly unequal distribution of income.

[11] In this context, for comparison purposes it would be useful to obtain intergenerational incomes elasticities for large urban areas in Brazil.
[12] See Dunn (2004).

Table 4: International Evidence on Intergenerational Income Mobility

Study	Country	Son's cohort	Elasticity OLS	Elasticity IV
Osterbacka (2001)	Finland	25-45	0.13	
Corak and Heisz (1999)	Canada	29-32	0.23	
Lillard and Kilburn (1995)	Malaysia	>18	0.26	
Grawe (2001)	Malaysia	not reported		0.54
Björklund and Jänti (1997)	Sweden	29-38		0.28
Wiegand (1997)	Germany	27-33	0.34	
Lefrane and Trannoy (2004)	France	30-40		0.36-0.43
Solon (1992)	U.S.	25-33	0.29-0.39	
Solon (1992)	U.S.	25-33		0.45-0.53
Dearden, Machin and Reed (1997)	U.K.	33		0.39-0.59
Grawe (2001)	Nepal			0.44
Grawe (2001)	Pakistan			0.46
Dunn (2004)	Brazil	25-34	0.53	0.69
Ferreira and Veloso (2004)	Brazil	25-64		0.58
Ferreira and Veloso (2004)	Brazil (Southeast)	25-64		0.54
Ferreira and Veloso (2004)	Brazil (Northeast)	25-64		0.73
Ferreira and Veloso (2004)	Brazil (South)	25-64		0.62
Ferreira and Veloso (2004)	Brazil (Midwest)	25-64		0.55

4.2 Evolution of Intergenerational Mobility in Chile

Table 5 presents the intergenerational income elasticities by cohort employing the fathers' predicted income from schooling and experience only.

Table 5: Estimates of the Coefficient of Intergenerational Income Mobility

Cohort	Personal Income	Labor Income
23-34	0.46	0.46
35-44	0.54	0.52
45-54	0.63	0.65
55-65	0.59	0.58
All sample	**0.54**	**0.54**

As can be seen, the elasticity coefficient is monotonically decreasing for the three younger cohorts. A possible explanation of this phenomenon is that intergenerational mobility could have increased in the last decades. However, this pattern can also be associated with life-cycle effects on earnings, and therefore whether income mobility has increased in time is unclear.[13] Yet, in order to shed some light on intergenerational mobility in Chile, in this section we make a detour to examine whether there has been an increase in

[13] See Dunn (2004) for a discussion on the role of life-cycle effects on intergenerational income mobility elasticities.

intergenerational *educational* mobility. For this purpose, we employ the argument that the accumulation of schooling for the majority of the population in Chile happens before the early twenties, such that schooling levels remain fixed throughout the life cycle after the mid twenties or so. This would imply that life-cycle effects in years of schooling would not be important for the population in 23-65 age group. Accordingly, the existence of lower intergenerational mobility coefficients for the younger cohorts would be an indication of a change in the levels of educational mobility in time.

In order to substantiate this claim we explore the accumulation of schooling in Chile by age groups using the 1996-2001 CASEN Panel. Figure 1 shows the average individual accumulation of schooling in Chile by age groups between 1996 and 2001. Figure 1 shows that while schooling accumulation is significant before 20, it decreases thereafter, and in fact, it remains negligible after age 25. Hence, we claim that life-cycle effects in years of schooling are not important for the age group 23-65 considered in this study, and therefore finding evidence of higher intergenerational educational mobility for the younger cohorts would be suggestive of increasing *educational* mobility in time.

Figure 1: 1996-2001 Average Individual Differential in Years of Schooling by Age

Source: Panel CASEN 1996-2001

Table 6 presents the schooling intergenerational elasticities by cohorts. The evidence indicates lower values for the younger cohorts, although some stability in the last two cohorts. Using the argument presented above, this would suggest an overall increase in intergenerational educational mobility in the last decades.

Table 6: Schooling Elasticity by cohort

Cohort	Sons and Daughters	Sons	Daughters
23-34	0.15	0.15	0.14
	(7.48)**	(7.70)**	(4.14)**
35-44	0.15	0.15	0.15
	(6.48)**	(4.12)**	(5.55)**
45-54	0.29	0.24	0.37
	(9.75)**	(6.59)**	(7.57)**
55-65	0.37	0.41	0.32
	(7.97)**	(5.34)**	(6.82)**
All sample	**0.21**	**0.21**	**0.23**

Robust t statistics in parentheses

* significant at 5%; ** significant at 1%

Tables 7 and 8 provide more evidence of this pattern. While Table 7 reports the regression results with sons' schooling as dependent variable, Table 8 reports the intergenerational schooling elasticity. Tables 7 and 8 provide evidence of a positive association between fathers' and sons' schooling levels, as well as strong evidence of an expansion in schooling in time, particularly for daughters, as indicated by the coefficients of year of birth (*cohort* variable). But Tables 7 and 8 also provide strong evidence of a significant lower association between fathers and sons' schooling the younger the cohort, as indicated by the coefficients of the father's schooling-cohort interactive term in both specifications.

Table 7: Cohort Effects in years of schooling.

Dependent Variable: Sons' Schooling

Variables	All	Sons	Daughters
Father's schooling	17.183	14.440	19.945
	(5.21)**	(3.06)**	(4.29)**
Father's schooling*Cohort	-0.009	-0.007	-0.010
	(5.09)**	(2.98)**	(4.21)**
Cohort	0.119	0.022	0.089
	(7.10)**	(7.00)**	(3.55)**
Constant	-223.716	-40.527	-165.007
	(6.83)**	(6.65)**	(3.36)**
Observations	1197	649	548
Adj. R-squared	0.33	0.29	0.38

Note: Cohort is defined as offspring's year of birth.

Robust t statistics in parentheses

* significant at 5%; ** significant at 1%

Table 8: Cohort Effects in schooling intergenerational elasticity.

Dependent variable: Sons' log schooling

Variables	All	Sons	Daughters
Father's log schooling	15.156	16.548	13.554
	(5.67)**	(3.65)**	(4.66)**
Father's log schooling*Cohort	-0.008	-0.008	-0.007
	(5.60)**	(3.62)**	(4.59)**
Cohort	0.022	0.153	0.022
	(4.11)**	(7.11)**	(6.68)**
Constant	-40.567	-292.412	-41.082
	(3.90)**	(6.91)**	(6.35)**
Observations	1197	649	548
Adj. R-squared	0.32	0.29	0.35

Note: Cohort is defined as offspring's year of birth.
Robust t statistics in parentheses
* significant at 5%; ** significant at 1%

To what extent this lower statistical association between fathers' and sons' education in the younger cohorts is related or causes the lower intergenerational income elasticities of the younger cohorts reported earlier remains as a topic for future research. In particular, it must be noted that there can be several factors, such as the quality of education and social networks, or the existence of class-discrimination in the labor market, for example, that may yield different returns to schooling to individuals from different social backgrounds.[14] This would limit the capacity of transforming the higher degrees of educational mobility of the younger cohorts reported above into higher intergenerational income mobility. This, however, remains as a rich and open avenue for future research.

4.3 Patterns of Intergenerational Income Mobility in Chile

We finally examine some patterns of intergenerational income mobility. In particular, in this section we study whether the intergenerational transmission of the socioeconomic status varies across different segments of the population in Chile. We begin by examining transition matrix results for analyzing some heterogeneity in intergenerational income mobility.

[14] See for example Nuñez and Gutierrez (2004), which reports a 50% gap in earnings of professionals from different social backgrounds of origin but controlling for academic performance, experience, quality of school education, postgraduate studies, command of a second language, among other controls.

The main advantage of quantile transition matrixes over log-linear regression models is that it permits an assessment of whether there is more or less mobility at the bottom or at the top of the income distribution and provide a first way of looking at potential non-linearities in the intergenerational transmission process.

Table 9 reports estimates of quintiles transition matrix for labor income using father's education and potential experience as predictors for father's income, and deriving fathers' quintiles from the real corresponding income distribution. For purpose of robustness check, Table 10 reports an equivalent transition matrix in which fathers' quintiles are obtained from the distribution of fathers' predicted incomes. Both transition matrices yield similar patterns. In both cases the bottom-to-bottom and the top-to-top transition probabilities are large (37-30 and 47-57, respectively), a pattern that is, in fact, also observed for other countries. In addition, the probabilities of transiting from the lowest to the highest quintiles and vice versa are low, around 0 to 8 per cent, which is also consistent with the international evidence.

Table 9: Quintile Transition Matrix
Father's quintiles obtained from real income distribution

	Son				
Father	Bottom	2nd	3rd	4th	Top
Bottom	0.37	0.35	0.14	0.14	0.00
2nd	0.23	0.29	0.15	0.26	0.07
3rd	0.21	0.31	0.14	0.24	0.11
4th	0.15	0.31	0.10	0.20	0.23
Top	0.08	0.14	0.07	0.24	0.47
Immobility Index	**0.30**				

Table 10: Quintile Transition Matrix
Father's quintile obtained from distribution of fathers' predicted incomes

	Son				
Father	Bottom	2nd	3rd	4th	Top
Bottom	0.30	0.34	0.10	0.22	0.04
2nd	0.17	0.29	0.22	0.22	0.10
3rd	0.21	0.33	0.10	0.24	0.12
4th	0.12	0.18	0.11	0.26	0.33
Top	0.06	0.13	0.04	0.20	0.57
Immobility Index	**0.30**				

The transition matrixes in Tables 9 and 10 indicate that there is an important disparity in the intergenerational transmission mechanism. In fact, two salient features can be pointed out. First, there is more persistence at both extremes of the fathers' income distribution, while there is a significant degree of intergenerational mobility in the intermediate fathers' quintiles. Second, the matrixes also suggest an asymmetric degree of intergenerational persistence at the extremes of the fathers' income distribution, in particular a higher degree of persistence at the top quintile versus the bottom quintile. In order to explore these issues further, regression equations of the fathers' centiles versus sons' centiles are reported in Table 11.

Table 11: Alternative Functional Forms for Intergenerational Mobility
(Centiles of labor income)

	Father's centile from real income distribution				Father's centile from predicted income distribution			
Variables	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
Father's centile	0.453	-0.236	1.622	1.62	0.379	-0.157	1.027	1.050
	(10.97)***	(1.18)	(2.99)***	(2.97)***	(11.00)***	(1.15)	(3.09)***	(3.12)***
Father's centile^2		0.006	-0.035	-0.035		0.005	-0.024	-0.025
		(3.55)***	(3.09)***	(3.09)***		(4.14)***	(3.08)***	(3.11)***
Father's centile^3			0.0003	0.0003			0.0002	0.0002
			(3.71)***	(3.73)***			(3.81)***	(3.83)***
Constant	1.607	16.473	-3.324	22.394	5.312	12.151	3.506	34.479
	(0.11)	(1.04)	(0.20)	(2.80)***	(0.34)	(0.79)	(0.23)	(9.01)***
Observations	649	649	649	649	649	649	649	649
R-squared	0.17	0.19	0.21	0.18	0.15	0.17	0.19	0.18

Note: Models (1), (2), (3), (5), (6) and (7) include controls for son's life-cycle effect (age and age^2)
Robust t statistics in parentheses
* significant at 10%,** significant at 5%; *** significant at 1%

Table 11 shows that the association between the father's and son's centile is increasing as expected but not linearly, in agreement with the evidence reported by the transition matrixes. In fact, specifications 2 and 6 of Table 11 show that when a quadratic functional form is imposed, it yields a robust convex pattern. This indicates that, overall, the intergenerational income persistence is asymmetric, being higher in the upper part of the fathers' income distribution than in the bottom part. Yet, the cubic specification in 3, 4 7 and 8 outperform the quadratic model, indicating that intergenerational persistence is higher at the two extremes of the income distribution than in the intermediate segments of the fathers' income distribution, consistent with the evidence suggested by the transition matrixes. A similar pattern is observed when the specification is regressed for the 23-44 and the 45-65 age groups separately (see figures A3 and A4 in the appendix).

This fact is confirmed by comparing the coefficient of father's versus son's centiles at the bottom versus top quartile and quintile of the fathers' income distribution. Table 12 indicates that the slope of fathers' vs. sons' centiles is steeper at the top than at the bottom of fathers' income distribution, difference that is statistically significant. This again confirms the higher intergenerational persistence at the top of the fathers' income distribution relative to the bottom.

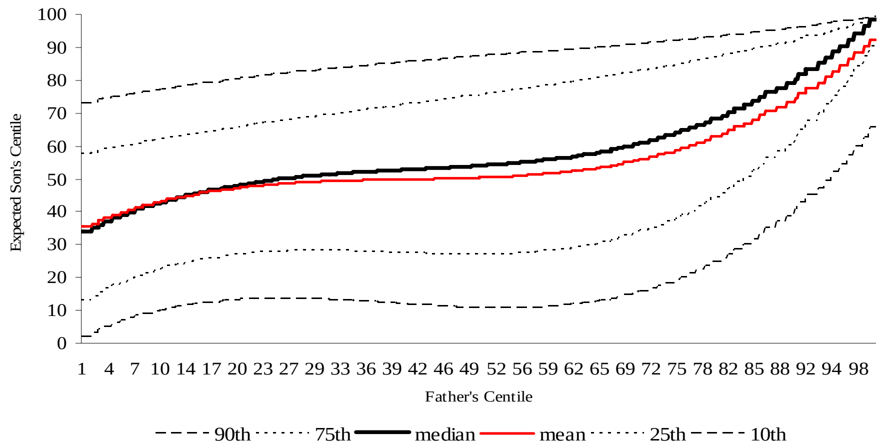
Table 12: Structural Change in OLS Coefficient of Father's vs. Son's Centiles in Upper and Lower Father's Quantile

		Estimates	F-test of structural change	P-value
Quartile	Top	1.46 (0.27)**	4.244	0.0023
	Bottom	0.7 (0.26)**		
Quintile	Top	1.88 (0.37)**	4.231	0.0025
	Bottom	0.91 (0.39)**		

Robust standard errors in parentheses
* significant at 5%; ** significant at 1%

This pattern is depicted in Figure 2, which shows the profile of cubic OLS and quantile regressions of fathers' vs. sons' centiles. Both the mean and the median show a clear cubic pattern, confirming the previous finding. But Figure 2 also illustrates how broad is the centile spectrum that sons are likely to end up in adulthood, given their fathers' centiles. Note that while in the most part of the fathers' income distribution the sons can end up in centiles often quite different from their parents, at the top of the fathers' income distribution, chances are that sons will occupy relative positions similar to those of their fathers, yet another indication of high intergenerational persistence at the top of the fathers' income distribution.

Figure 2: Quantile Cubic Regressions



In order to provide another view of the asymmetry between the degrees of persistence at the tails of the fathers' income distribution, we have employed the coefficients of specification 4 of Table 11 in order to obtain the expected son's income centile in adulthood, given a level of father's income centile. From these centiles we have constructed the expected son's centile for father's deciles, which are reported in Table 13.

Table 13: Son's Expected Centile in Adulthood given Father's Decile

Fathers' Decile	Son's Expected Centile (Father's Centile from Real Distribution)
1	30.04
2	39.90
3	45.20
4	47.49
5	48.35
6	49.32
7	51.98
8	57.88
9	68.59
10	85.66

Table 13 suggests a significant asymmetry in the patterns of mobility in Chile, which were already suggested by the previous analysis. In fact, the expected centiles of sons of fathers in the bottom decile is 30, that is, at the border of the 3rd and 4th decile. Although this shows an important degree of persistence, it also shows an important degree of the "regression-to-the-mean" effect for this group. Yet, sons of fathers in the top decile can expect to be, on average, on the 86th centile, that is well into the 9th decile and close to the top centile: the "regression-to-the-mean" effect does not seem to be very important in this case. This pattern repeats in a milder version when comparing the expected centiles of sons of fathers in the 2nd and 9th deciles, 40 and 69, respectively: the former sons are only 10 centiles below the median, while the latter are 19 centiles above it. As for the sons of fathers in deciles 3 to 8, their expected centile is only a few percentage points away from the median, suggesting an important degree of intergenerational income mobility at the centre of the father's income distribution. This evidence confirms an important asymmetry in the patterns of mobility at the extremes of fathers' income distribution in Chile, with more *relative* persistence at the very top than at the very bottom of the fathers' income distribution.

Finally, and considering the evidence of a significant amount of intergenerational persistence in the upper part of the father's income distribution, we examine how it compares with the available international evidence. The transition probabilities associated to the top quintile in Tables 9 and 10 do indeed seem high in comparison with the international evidence, as shown in Table 14. A similar pattern arises also when comparing the intergenerational persistence at the top quartile.

Table 14: Comparative evidence on persistence in top quantile

Study	Country	Top Quartile	Top Quintile
Veloso and Ferreira (2004)	Brazil		0.43
Österberg (2000)	Sweden	0.25	
Peters (1992)	U.S.	0.36-0.40	
Ng (2004)	Finland	0.40-0.52	
Fortin and Lefebvre (1998)	Canada	0.32-0.33	
Nuñez and Risco (2004)	Chile		0.50
This study	Chile	0.55-0.56	0.47-0.57

To recap, the evidence suggests three salient features of the patterns of mobility on Chile. First a higher degree of persistence at the extremes of the fathers' income distribution, while substantially more mobility in the intermediate segments of it. Second, there is evidence of an asymmetry between the degrees of persistence at the bottom versus the top of fathers' income distribution, the persistence being higher at the top quantiles.

Third, the persistence at the top seems relatively high in comparison with the available international evidence.

It is interesting to note that the latter two results seem quite consistent with recent evidence on intergenerational *occupational* mobility for Chile. In fact, Torche (2005) finds that Chile exhibits a high level of persistence in the occupations with highest social status in compassion with the international evidence, but a significant degree of mobility in the remaining occupations. It seems suggestive that the two investigations based on different methodologies and different conceptual frameworks reach somewhat converging conclusions. As a hypothesis awaiting research, this evidence may be associated with Chile's particular income distribution, namely the fact that Chile's income distribution is unequal for international standards basically due to the large share of the national income held by the top decile or so of the population, the remaining part of the population being particularly egalitarian. Put quite simply, perhaps the top decile or so of Chile's income distribution is largely responsible for Chile's unequal distribution of income as well as for shaping Chile's social mobility patterns. This hypothesis still awaits explicit investigation.

6. Conclusions

This paper describes some new findings on intergenerational income mobility in Chile. A shortcoming to study income mobility in Chile is the lack of income panels where both fathers and their offsprings' income can be observed. In this context, all the existing evidence is based on two-sample instrumental variables methodology, where fathers' income is predicted from income determinants of fathers reported by their sons, such as schooling, occupation and age.

Yet, by using this methodology, various salient features of intergenerational mobility can be obtained. First, the available intergenerational income mobility elasticities for Chile are in the range of 0.52 to 0.67. The figures derived for Greater Santiago are around 0.52-0.58. Yet these latter figures may underestimate the true nation-wide elasticities as some parts of the population are left out such as the remote and rural areas, and small urban areas, where intergenerational mobility can be expected to be lower. However, these figures stand as fairly high in comparison with the international evidence, mostly for developed countries, but similar to elasticities found for Brazil.

Second, intergenerational income elasticities are somewhat lower for the younger cohorts. This may suggest an increasing intergenerational mobility in time, although it can also be associated with life-cycle effects. We also find that educational mobility is lower for the younger cohorts. We claim that given the fact that schooling remains fixed for most individuals along the life cycle after the mid twenties, this finding strongly suggests an increase of intergenerational educational mobility in the last decades. This seems consistent with the significant expansion of school enrollment and of years of schooling in the last decades. Whether this has translated into higher intergenerational income mobility is open for future research.

Third, we also examine how intergenerational mobility varies along the fathers' relative income position. We find evidence of a high degree of intergenerational persistence at the bottom-to-bottom and top-to-top transition probabilities, while a significant degree of mobility in the intermediate segments of the income distribution. Moreover, we find evidence of an asymmetry in the degrees of persistence between the two tails of the fathers' income distribution: the intergenerational persistence is higher for the top fathers' income quintile or so than for the bottom quintile. The evidence suggests that this high degree of persistence at the top part of the fathers' income distribution is large for international standards, as suggested by the top-to-top transition probabilities. This evidence is consistent with recent findings of intergenerational occupational mobility for Chile by Torche (2005), who reports a high degree of persistence in occupations associated with high social status, but a significant degree of occupational mobility in the rest of the occupations spectrum. It is tempting to propose, as hypothesis, that this finding may be associated with Chile's particular income distribution, quite egalitarian for the 80 to 90 percent of the population, but rather unequal when the top quintile or decile is considered.

An interesting avenue for future research is to study the determinants of intergenerational and educational mobility in Chile. Although this paper has shed some light on some patters of Chilean intergenerational income mobility, there is a great deal of social mobility whose determinants need to be understood. In this respect, the dataset employed in this paper also includes various measures of circumstances of origin, such as family background, parental education and functional literacy, parental effort at home, quality of education, place of origin, ethnicity, among others. Examining the impact of these factors can shed light on the determinants that can promote more equality of opportunities.

Finally, more robust and conclusive evidence about intergenerational income mobility would require establishing proper panels where both parents' and offsprings' incomes can be observed. This is certainly a pending issue in Chile as in most developing countries.

References

Abul R. and F. Cowell (2002), "Intergenerational Mobility in Britain: Revisiting the Prediction Approach of Dearden, Machin and Reed", Ecole des Hautes Etudes Commerciales. Département D´Econométrie et d´Economie Politique. (disponible en: http://sticerd.lse.ac.uk/dps/darp/darp62.pdf).

Andersen, L. (2000), "Social Mobility in Latin America", Instituto de Investigaciones Socio-Económicas Universidad Católica Boliviana.

Angrist, J. and A Krueger (1992), "The Effect of Age at School Entry on Educational Attainment: An Application of Instrumental Variables with Moments from Two Samples", *Journal of American Statistical Association*, 87 (418), pp. 328-36.

Arellano, M. and C. Meghir (1992), "Female Labour Supply and On the Job Search: An Empirical Model Estimated Using Complementary Data Set", *Review of Economic Studies* 59(3), pp. 537-59.

Askew D., Brewington, J. and A. Touhey (2001), "An Examination of Intergenerational Income Mobility Using the Panel Study of Income Dynamics", *Econometrics* 375, (in: http://www.ups.edu/econ/pseje/articles/firstedition/atouhey.pdf)

Becker, G. and N. Tomes (1979), "An Equilibrium Theory of the Distribution of Income and Intergenerational Mobility", *Journal of Political Economy* 87, pp. 1153-1189.

Björklud, A. and M. Jäntti (1997), "Intergenerational Income Mobility in Sweden Compared to the United States", *American Economic Review* 87(5), pp. 1009-018.

Björklud, A. and M. Jäntti (1999), "Intergenerational Mobility of Socio-Economic Status in Comparative Perspective", mimeo Stockholm University.

Bourguignon, Ferreira and Menendez (2003), "Inequality of Outcomes, Inequality of Opportunities and Intergenerational Educational Mobility in Brazil". Washington DC: The World Bank.

Bowles, S. and H. Gintis (2002), "The Inheritance of Inequality", *Journal of Economic Perspectives* 16(3), pp. 3-30.

Contreras, D., M. Fuenzalida and J. Nuñez (2006), "Persistencia Intergeneracional del Ingreso en Chile y el Rol de la Habilidad de los Hijos", Masters thesis, Economics Department, Universidad de Chile.

Corak, M. and A. Heisz (1999), "The Intergenerational Earnings and Income Mobility of Canadian Men: Evidence from Longitudinal Income Tax Data", *Journal of Human Resources* 34(3), pp. 504-33.

Dahan, M. and A. Gaviria (1999), "Sibling Correlations and Social Mobility in Latin America", Inter-American Development Bank, Washington, D.C.

Dearden L., Machin S. and H. Reed (1997), "Intergenerational Mobility in Britain", *The Economic Journal* 107(440), pp. 47-66.

Dunn, C. (2003), "Intergenerational Earnings Mobility in Brazil" (Paper presented at 8th annual meeting LACEA).

Dunn, C. (2004), "Intergenerational Transmission of Lifetime Earnings: New Evidence from Brazil", mimeo Department of Economics and Population Studies Center, University of Michigan.

Ferreira, S. and F. Veloso (2004), "Intergenerational Mobility of Wages in Brazil". Unpublished paper.

Fields, G. (2001), "Distribution and Development. A New look at the developing world", Russell Sage Foundation and MIT Press.

Fortín, N. and S. Lefebvre (1998), "Intergenerational Income Mobility in Canada", Labor Markets, Social Institutions, and the Future of Canada's Children.

Hertz, T. (2001), "Education, Inequality and Economic Mobility in South Africa", Ph.D. thesis, University of Massachusetts.

Lefrane A. and A. Trannoy (2004), “Intergenerational Earnings mobility in France: Is France more mobile than the US?”, Accepté dans Annales d`Economie et de Statistiques.

Lillard, L. and M. Kilburn (1995), “Intergenerational Earnings Links: Sons and Daughters”, unpublished paper.

Nuñez, J. and C. Risco (2004), “Movilidad Intergeneracional de Ingresos en un País en Desarrollo: El Caso de Chile”, Working Paper N° 210, Department of Economics, Universidad de Chile.

Nuñez, J. and R. Gutiérrez (2004), “Classism, Meritocracy and Discrimination in the Labor Market: The case of Chile”, Working Paper N° 208, Department of Economics, Universidad de Chile.

Nuñez, J. and A. Tartakowsky (2006), “Inequality of Outcomes vs. Inequality of Opportunities in Chile”, mimeo, Department of Economics, Universidad de Chile.

Österbacka, E. (2001), “Family Background and Economic Status in Finland”, *Scand J. of Economics* 103(3), pp. 467-484.

Peters, E. (1992), “Patterns of Intergenerational Mobility in Income and Earnings”, *The Review of Economics and Statistics* 74, pp. 456-466.

Solon, G. (1992), “Intergenerational Income Mobility in the United States”, *The American Economic Review* 82(3), pp. 393-408.

Solon, G. (2002), “Cross-countries Differences in Intergenerational Earnings Mobility”, *The Journal of Economic Perspectives* 16(3), pp. 59-66.

Torche, F. (2005), “Unequal but fluid: social mobility in Chile in comparative perspective”, *American Sociological Review* 70(3), pp. 422-450.

Wiegand, J. (1997), “Intergenerational Earnings Mobility in Germany”, unpublished paper.

Zimmerman, D. (1992), “Regression toward Mediocrity in Economic Stature”, *The American Economic Review* 82, pp. 409-429.

Appendix

Table A.1: Estimates of Mincer equations using education and potential experience as regressors

	1958		1967		1977		1987	
Log Income	Personal Income	Labor Income	Personal Income	Labor Income	Personal Income	Labor Income	Personal Income	Labor Income
S	**0.1098**	**0.1043**	**0.0768**	**0.0725**	**0.0784**	**0.0808**	**0.0638**	**0.0626**
	(14.76)**	(14.20)**	(11.38)**	(10.78)**	(8.35)**	(8.60)**	(5.59)**	(5.43)**
(S-8)*d1	**0.0839**	**0.0873**	**0.1252**	**0.1288**	**0.1266**	**0.1240**	**0.1038**	**0.1066**
	(4.73)**	(4.99)**	(8.71)**	(8.98)**	(7.20)**	(7.04)**	(5.55)**	(5.65)**
(S-12)*d2	**-0.0675**	**-0.0722**	**-0.0249**	**-0.0329**	**-0.0322**	**-0.0343**	**0.1152**	**0.1112**
	(3.00)**	(3.25)**	(1.54)	(2.04)*	(1.88)	(2.01)*	(7.19)**	(6.89)**
Exper	**0.0604**	**0.0605**	**0.0620**	**0.0617**	**0.0636**	**0.0651**	**0.0539**	**0.0545**
	(14.20)**	(14.43)**	(18.73)**	(18.63)**	(16.60)**	(17.00)**	(13.11)**	(13.17)**
Exper2	**-0.0008**	**-0.0009**	**-0.0009**	**-0.0009**	**-0.0009**	**-0.0009**	**-0.0007**	**-0.0007**
	(10.47)**	(11.14)**	(13.79)**	(14.12)**	(11.52)**	(12.33)**	(7.87)**	(8.23)**
Constant	**8.9524**	**8.9958**	**11.4925**	**11.5125**	**6.0278**	**5.9930**	**8.3213**	**8.3199**
	(134.89)**	(137.61)**	(210.52)**	(211.27)**	(79.81)**	(79.34)**	(92.97)**	(92.11)**
Observations	1747	1736	2700	2691	2325	2321	2070	2068
Adj. R-squared	0.51	0.50	0.53	0.52	0.48	0.48	0.59	0.59

Absolute value of t statistics in parentheses

* significant at 5%; ** significant at 1%

Table A.2: Estimates of Mincer equations using education, potential experience and occupation as regressors

	1958		1967		1977		1987	
Log Income	Personal Income	Labor Income	Personal Income	Labor Income	Personal Income	Labor Income	Personal Income	Labor Income
S	**0.0852**	**0.0805**	**0.0526**	**0.0478**	**0.0584**	**0.0606**	**0.0511**	**0.0499**
	(11.46)**	(10.97)**	(7.84)**	(7.16)**	(6.59)**	(6.87)**	(4.94)**	(4.77)**
(S-8)*d1	**0.0496**	**0.0530**	**0.1033**	**0.1087**	**0.0808**	**0.0759**	**0.0629**	**0.0646**
	(2.89)**	(3.12)**	(7.47)**	(7.89)**	(4.87)**	(4.58)**	(3.64)**	(3.71)**
(S-12)*d2	**-0.0083**	**-0.0127**	**0.0199**	**0.0100**	**0.0195**	**0.0196**	**0.1418**	**0.1383**
	(0.38)	(0.58)	(1.26)	(0.64)	(1.21)	(1.22)	(9.73)**	(9.42)**
Exper	**0.0548**	**0.0550**	**0.0550**	**0.0545**	**0.0527**	**0.0538**	**0.0427**	**0.0432**
	(13.37)**	(13.59)**	(17.16)**	(17.10)**	(14.57)**	(14.96)**	(11.32)**	(11.38)**
Exper2	**-0.0008**	**-0.0008**	**-0.0008**	**-0.0008**	**-0.0007**	**-0.0008**	**-0.0005**	**-0.0006**
	(10.30)**	(10.90)**	(13.17)**	(13.52)**	(10.48)**	(11.29)**	(6.94)**	(7.34)**
Employer=1	**0.7414**	**0.7137**	**1.0662**	**1.0558**	**1.3777**	**1.4163**	**1.5265**	**1.5306**
	(10.82)**	(10.49)**	(15.37)**	(15.23)**	(18.79)**	(19.29)**	(21.22)**	(21.12)**
Self-employed=1	**0.1677**	**0.1220**	**0.2197**	**0.2033**	**0.2582**	**0.2637**	**0.0983**	**0.0994**
	(4.44)**	(3.27)**	(6.69)**	(6.21)**	(7.16)**	(7.35)**	(2.71)**	(2.72)**
Employee=1	**0.3755**	**0.3734**	**0.2997**	**0.2962**	**0.3794**	**0.3891**	**0.3284**	**0.3389**
	(9.96)**	(10.05)**	(9.95)**	(9.87)**	(10.59)**	(10.89)**	(9.08)**	(9.30)**
Domestic servants=1	**-0.7535**	**-0.7570**	**-0.3714**	**-0.7523**	**0.0025**	**-0.3188**	-	-
	(1.97)*	(2.01)*	(1.89)	(3.85)**	(0.01)	(0.91)	-	-
Constant	**9.0698**	**9.1128**	**11.6337**	**11.6607**	**6.1962**	**6.1659**	**8.5123**	**8.5107**
	(141.07)**	(143.92)**	(219.63)**	(221.22)**	(87.39)**	(87.34)**	(104.32)**	(103.31)**
Observations	1747	1736	2700	2691	2325	2321	2070	2068
Adj. R-squared	0.55	0.54	0.58	0.57	0.55	0.56	0.67	0.66

Absolute value of t statistics in parentheses

* significant at 5%; ** significant at 1%

Table A3: Quartile Transition Matrix

Father's quartile from real income distribution

	Son			
Father	Bottom	2nd	3rd	Top
Bottom	0.50	0.27	0.19	0.04
2nd	0.30	0.22	0.33	0.15
3rd	0.29	0.24	0.28	0.19
Top	0.14	0.14	0.18	0.54
Immobility Index	**0.38**			

Table A4: Quartile Transition Matrix

Father’s quartile from distribution of predicted income

	Son			
Father	Bottom	2nd	3rd	Top
Bottom	0.39	0.24	0.30	0.08
2nd	0.32	0.23	0.30	0.15
3rd	0.26	0.26	0.22	0.26
Top	0.12	0.12	0.20	0.55
Immobility Index	**0.35**			

Figure A1: Intergenerational Income Transition Probabilities (Quintiles)

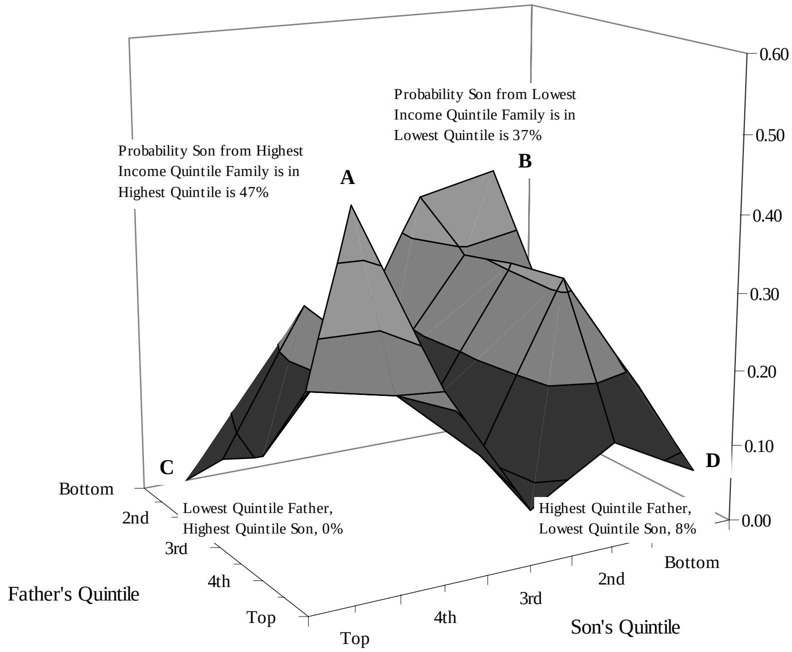


Figure A2: Intergenerational Income Transition Probabilities (Quartiles)

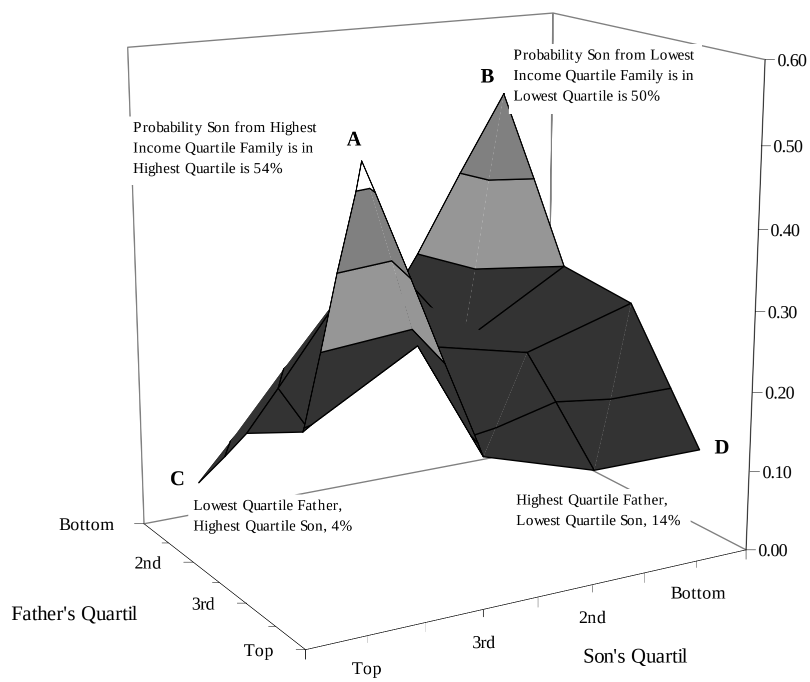


Figure A3: Quantile Cubic Regressions for 23-44 cohort

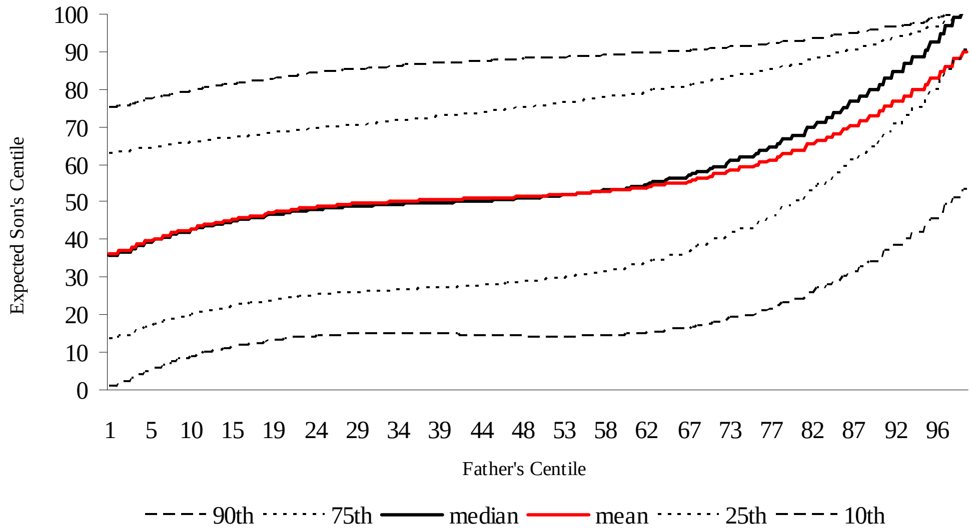


Figure A4: Quantile Cubic Regressions for 45-65 cohort