

Núñez, Javier; Miranda, Leslie

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Intergenerational income and educational mobility in urban Chile*

Movilidad intergeneracional del ingreso y la educación en zonas urbanas de Chile

JAVIER NÚÑEZ**
LESLIE MIRANDA***

Abstract

This paper provides evidence on the degree and patterns of intergenerational income and educational mobility in urban Chile. We find intergenerational income elasticities for Greater Santiago in Chile in the range of 0.52 to 0.54. This is lower than recent nation-wide elasticities for Chile of about 0.6-0.7, but still stands as fairly high in comparison with the comparable international evidence. We also find that intergenerational educational mobility is lower for the younger cohorts, which however does not necessarily imply an increase of intergenerational educational mobility in the last decades, as life-cycle effects may be at work. Finally, we find evidence of a higher degree of intergenerational persistence of income at the two extremes of the income distribution, which is more accentuated at the top centiles of the distribution. We suggest that this may mirror the unusually high concentration of income at the top of the income distribution in Chile, a hypothesis that requires further research.

Key words: *Intergenerational mobility, Schooling, Mobility patterns.*

Resumen

Este paper proporciona evidencia sobre el grado y los patrones de movilidad intergeneracional del ingreso y la educación en zonas urbanas de Chile. Encontramos elasticidades intergeneracionales del ingreso en el Gran Santiago

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** Department of Economics, Universidad de Chile.

*** Centro de Microdatos, Universidad de Chile.

del orden de 0.52-0.54, las que son menores que valores reportados para Chile en el rango de 0.6-0.7, pero son aún elevadas en relación a la evidencia internacional. También encontramos mayor movilidad en los cohortes más jóvenes, lo que no debe interpretarse necesariamente como aumentos en la movilidad del ingreso, pues efectos de ciclo de vida pueden estar presentes. Finalmente, encontramos evidencia de mayor persistencia intergeneracional del ingreso en los dos extremos de la distribución, la cual es particularmente acentuada en los centiles de ingreso más altos de la población. Sugerimos que este hallazgo puede reflejar la gran concentración de ingresos en los centiles altos que caracteriza a la distribución del ingreso en Chile, hipótesis que amerita investigación a futuro.

Palabras clave: *Movilidad intergeneracional, Escolaridad, Patrones de movilidad.*

JEL Classification: *D3, I2, J6.*

1. INTRODUCTION

Much of the literature on inequality in Chile has focused on the inequality of outcomes, typically the distribution of incomes, and little is still known about the country's levels of inequality of "opportunity", its assessment in comparative perspective and its evolution in recent times. A common approach to empirically assess a country's degree of equality of opportunity is the notion of intergenerational economic mobility: a higher level of equality of opportunity is expected to decrease the effect of an individual's early socioeconomic background on his economic achievement in adulthood, yielding therefore a higher level of intergenerational economic mobility.

This paper examines the degree of intergenerational income and educational mobility in urban Chile in comparative perspective, and some of its salient patterns. Since the research in intergenerational income mobility is fairly recent and limited in Chile, this work also proposes and discusses some exploratory hypothesis and some avenues for future research.

The study of intergenerational income mobility requires income data for pairs of fathers (or parents) and their offspring. As it is common in most of the developing world, such data is often limited or unavailable. In this work we follow the methodology often known as Two-Stage Two-Sample Least Squares (TSTSLS) developed originally by Björklud and Jäntti (1997) and widely applied, whereby father's incomes are predicted from data on income determinants such as father's schooling and education provided by their sons. We apply this methodology to Greater Santiago, Chile's main urban center, and analyze its results in comparative perspective. We find a lower intergenerational mobility in Greater Santiago compared with estimates using nationwide data, which may indicate that rural and small urban areas may exhibit lower educational and occupational opportunities than a large urban area such as Greater Santiago, an issue open for future research. Yet, the results for Greater Santiago are still relatively high compared to international evidence. This paper also explores

how intergenerational income mobility varies along parent's relative position in the income distribution in urban Chile.

This paper is structured as follows. 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 opportunity, and by discussing the need to increase the understanding of inequality of opportunity, a task for which the notion of intergenerational mobility is well suited. Section 3 presents the theoretical framework and the empirical strategy followed. Section 4 describes the dataset employed, and Section 5 presents and discusses the main results. Finally Section 6 concludes.

2. INEQUALITY OF OUTCOMES VS. INEQUALITY OF OPPORTUNITY IN CHILE

There has been a long debate about whether inequality, and the policies designed to deal with it, must tackle the inequality of outcomes or alternatively the inequality of opportunities. Advocates of the latter stress that inequality of outcomes (typically incomes) depend not only on the circumstances that are beyond the control of individuals, such as parental background, but also on factors that are (presumably) under their control, such as effort and choices. Accordingly, some authors have suggested that from a moral standpoint, public policies should focus on equalizing "opportunities" instead of outcomes (incomes)¹. In a seminal contribution, Bourguignon, Ferreira and Menéndez (2005, 2007) have developed a methodology to attempt to measure the proportion of the observed income distribution that is associated with inequality of (uncontrolled) circumstances of origin such as parental education and occupation and race in Brazil. The methodology distinguishes the direct effect that circumstances have on earnings in adulthood (the "partial effect") from the indirect effect of circumstances on earnings through the accumulation of human capital (schooling). They found that the Gini coefficient is reduced in up to about 10 percentage points (about 20 per cent) after equalizing the set of circumstances mentioned earlier.

Table 1 shows the effect on the Gini coefficient for Greater Santiago and Chile of equalizing across individuals various circumstances of origin including parental schooling, size and composition of the household during infancy (single versus biparental), the age of the household head, and a measure of parents' job vulnerability. Even though many relevant circumstances may certainly remain unobserved, these results do suggest that the important circumstances mentioned above play an important yet limited role in determining the income distribution, results that are similar to those reported for Brazil. These findings suggest that income distribution indicators seem only partially associated to a society's degree of "equality of opportunity", and that income distribution may be also affected by other factor presumably unrelated to exogenous circumstances. In this context, a closer and perhaps more direct way of assessing a country's degree

¹ For a discussion of these issues, see for example Bourguignon, Ferreira and Menéndez (2007).

of equality of opportunity is to examine its level of intergenerational income and educational mobility, issue that we address next.

TABLE 1
EFFECT OF EQUALIZING OBSERVED CIRCUMSTANCES OF ORIGIN ON GINI
COEFFICIENT, CHILE, MEN AGED 23-65

	Greater Santiago employment survey, 2004	Nationwide Casen survey 2006
Gini Coefficient	0.503	0.535
Gini after equalizing observed circumstances of origin		
Partial Effect	0.433	0.491
Total Effect	0.420	0.455

Source: Núñez and Tartakowsky (2007, 2009).

3. THEORETICAL FRAMEWORK AND EMPIRICAL STRATEGY

This paper motivates and illustrates the intergenerational transmission of the socioeconomic status and the related concept of intergenerational income mobility by means of a simplified version of the 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, that is, father and child. Individual permanent income Y is assumed to derive from two components: the individual endowment of human capital and individual ability denoted by A . Becker and Tomes (1979) assume 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 the father's and his child's permanent incomes:

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

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

It is important to note that a regression of the child's income on father's income would capture both transmission mechanisms. Hence, a standard OLS

estimate of the intergenerational earnings transmission coefficient would provide an upward biased estimate of the “pure” causal effect of parental income on child’s income. In this work, we do not attempt to separate both effects. Instead, and in line with the related literature, we are interested in the estimation of a reduced-form intergenerational earnings regression to measure the degree of intergenerational income mobility.

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

$$(2) \quad Y_i^{child} = \beta_0 + \beta_1 Y_i^{father} + \varepsilon_i$$

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, $\beta_1 = 0$ would depict a situation involving full intergenerational mobility, as the permanent income of the child in adulthood would show no statistical association whatsoever with the father’s permanent income. At the other extreme, $\beta_1 = 1$ would indicate a situation of very low mobility, whereby a child born from a parent with an income, say, x per cent above the mean of parent’s incomes will have an expected income located x per cent above the mean of his own generation. Alternatively, $1 - \beta_1$ can be interpreted as a “regression-to-the-mean” effect: If $\beta_1 = 0$, for example, the regression-to-the-mean effect is maximum, as sons will have the same expected value of income of adulthood regardless of their respective father’s relative socioeconomic status.

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 incomes in a single year can 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. An alternative to address this bias involves using panel data on father’s income to obtain an average of father’s current income over several years as a proxy for their permanent income. Solon (1992) shows that the inconsistency of the transmission coefficient diminishes with the number of years over which incomes are averaged.

Another problem emerges when, as in this paper, income information of father-child pairs is not available. In this context, a solution proposed by Arellano and Meghir (1992) and Angrist and Krueger (1992) and followed by Björklud and Jäntti (1997) for intergenerational mobility studies is to use information from two separate samples: first, earnings equations can be estimated using an older sample of men in order to obtain coefficients of some key earnings determinants, such as schooling, experience and occupation, for example. Then, these coefficients can be employed to predict the income of the fathers of a

sample of sons who have reported the required information about their fathers. This technique is often known as two-sample instrumental variables estimation (TSIV) or “two-sample two-stage least squares” (TSTSLs)².

More formally, assume that the log of the father’s and son’s current income at date t can be written as:

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

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

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

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

where v_i^{father} is independent of Z_i^{father} . Term Y_{it}^{father} is not observed in sample I. Yet, if there exists a separate sample J from the same population as I, it can be used to provide an estimate of γ , namely $\hat{\gamma}$, which would be derived from estimation of equation (6) using the sample of adult men J:

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

with $j \in J$. An OLS estimation of (6) would provide predictions of the father’s earnings in sample I: $\hat{Y}_{it}^{father} = Z_i^{father} \hat{\gamma}$. This prediction can in turn be used to estimate the intergenerational income elasticity coefficient β_1 since equations (2), (3), (4) and (6) imply:

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

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

In this work, 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 level³:

² See for example Dunn (2004).

³ In Chile, elementary education consists of the first eight years and secondary school consists of four additional years.

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

where S_{js} represents the years of schooling of father in year s ⁴, $Exper_{js}$ stands for father's potential experience⁵ and ε_{js} is a random error term. In addition, dummy variables for educational levels are defined as:

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

In another specification we also include information of father's occupation under the assumption that occupation is a good instrument to estimate the father's permanent income. This information comes from a second survey undertaken in June 2006 to a sub-sample of the June 2004 version of Employment and Unemployment Survey of Greater Santiago.

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

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

Hence, we obtain the intergenerational income elasticity β_1 from:

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

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

The methodology described above is subject to some well-known biases that have been identified in the related literature, which are worth pointing out. 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 separate explanatory variables of the son's earnings, the resulting omitted-variable problem would yield an upward bias in the intergenerational income elasticity.

⁴ The s year corresponds to time when father were taken investment decisions on his child's human capital.

⁵ Potential experience is defined as: age minus years of schooling minus 6.

Another source of bias is related to the ages of sons being considered for estimating equation (7). In particular, various studies have found that the estimated intergenerational elasticities increase substantially as son's earnings are observed further on in their careers. Accordingly, studies that use earnings data of sons in the early stages of their life-cycle 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 Due to the existence of these potential biases, in this paper we compare the results obtained for urban Chile with the results of international studies that follow a similar methodology, and are accordingly subject to the same kind of biases. We also report the age brackets of son's ages employed in all studies.

4. DATA

Our dataset comes from the Employment and Unemployment Survey for the Greater Santiago conducted annually by Universidad de Chile since 1957, which is applied to approximately 4,000 households. Greater Santiago is the largest urban centre in Chile, home to approximately 40 per cent of the country's population.

In order to avoid selectivity issues associated with female participation in the labor market, we focus only on fathers-sons 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.

The Greater Santiago 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 employed to estimate the 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 hours per week. Our sample of sons comes from the June 2004 version of the survey. In this year, additional to demographic and economic data, respondents were asked to provide information about educational and individual characteristics of their parents. We consider sons in the 23-65 age range to control for potential selectivity problems in individuals outside this age range. We eliminate unemployed and inactive individuals, those with no positive incomes and those missing parental information. Our sample was composed by 649 pairs of fathers and sons in the corresponding age range.

Father's predicted incomes were estimated dividing the sons sample in three sub-samples by age groups: 23-34, 35-44, 45-54 and 55-65 years old. We select the corresponding father's samples by assuming that the father's investment decisions in his son's 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 the 1987,

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

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 obtained from adjacent years.

A second data source comes from a new survey realized in June 2006 to a sub-sample of males previously surveyed in the Employment and Unemployment Survey of June 2004. In this new survey, individuals were asked 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 the information about the father's occupations⁷ to estimate a second Mincer equation specification under the assumption that occupation is a good instrument to estimate the father's permanent income, in addition to father's schooling.

5. RESULTS

5.1. Estimates of intergenerational income elasticity for urban Chile

Table 2 reports intergenerational regression coefficients for labor incomes⁸. Results are reported for the full sample in the 23-65 age range. Estimates in this table are obtained using father's education, potential experience and occupation as predictors of father's income, as described earlier. First-step income regressions are provided in Tables A.1 and A.2 in the appendix.

TABLE 2
ESTIMATES OF INTERGENERATIONAL LABOR INCOME ELASTICITY
BY TWO-SAMPLE TWO-STAGE LEAST SQUARES (TSTLS)
Greater Santiago, Chile

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

Table 2 indicates that father's predicted log income has a significant positive effect on their son's log income. For the whole sample, the estimated elasticity is around 0.52-0.54 depending on the predictors employed⁹.

Table 3 reports the results of other intergenerational income mobility studies for urban Chile. It is interesting to note that employing only the SIALS¹⁰ data

⁷ The 5-level occupational categories are: employer (1); self-employed (2); employee (3); blue-collar worker (4) (reference) and domestic (household) workers (5).

⁸ The estimates using personal incomes yield the same global elasticity.

⁹ This number is a weighted average of elasticities of each age group.

¹⁰ Second International Adult Literacy Survey.

TABLE 3
 TSTSLS INTERGENERATIONAL INCOME MOBILITY IN CHILE, USING URBAN AND NATIONWIDE DATA

Study	Database	Population	Key father's income predictors	Sons' ages	Elasticity
Núñez and Risco (2004)	Employment and Unemployment Survey (U. of Chile)	Greater Santiago	Potential experience, schooling	23-55	0.55
Contreras, Fuenzalida and Núñez (2006)	SIALS	Greater Santiago	Schooling	23-55	0.58
Contreras, Fuenzalida and Núñez (2006)	SIALS	Urban Population	Schooling	23-55	0.67
Núñez and Miranda (2010)	CASEN	Nationwide Data	Schooling	25-40	0.74
Núñez and Miranda (2010)	CASEN	Nationwide Data	Schooling and Occupation	25-40	0.57
This study	Employment/Unemployment Survey (U. of Chile)	Greater Santiago	Schooling	23-65	0.54
This study	Employment/Unemployment Survey (U. of Chile)	Greater Santiago	Schooling and Occupation	23-65	0.52

for the Metropolitan Region (similar but slightly larger than Greater Santiago) yields fairly similar results than the Greater Santiago studies although lower than the 0.67 elasticity obtained from the national urban SIALS database, and lower as well than the nationwide CASEN studies, which yield elasticities in the 0.6-0.7 range. This evidence suggests a lower intergenerational income elasticity for the large, more prosperous Greater Santiago urban area in comparison with the rest of the country.¹¹ This may be an indication of a lower intergenerational mobility in rural, semi-rural and small urban areas, where educational and occupational opportunities might be lower than in a large urban centre like Greater Santiago.

Table 4 summarizes some selected international evidence on (nationwide) intergenerational income mobility. As can be seen, urban Chile presents relatively low intergenerational income mobility in comparison with other developing and developed countries. This could be even more so for nationwide mobility in Chile if the previous discussion on the lower elasticity values for urban areas is considered. Levels of intergenerational mobility in Chile seem somewhat

TABLE 4
INTERNATIONAL EVIDENCE ON 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 (2004)	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	Lefrane and Trannoy (2004)	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	Björklud and Jäntii (1997)	28-36		0.52
Sweden	Björklud and Jäntii (1997)	29-38		0.28
Nepal	Grawe (2001)	–		0.44
Pakistan	Grawe (2001)	–		0.46

¹¹ This seems consistent with of evidence on the pattern of intergenerational income mobility in Brazil, according to Ferreira and Veloso (2004): The more prosperous and more urban Brazilian Southwest has an intergenerational income elasticity of 0.54, which is lower than the rest of the country, and much lower than the poorer Northeast region (0.73).

similar to Brazil's¹². Some authors have suggested and provided theoretical arguments and some evidence in favor of an overall positive relationship between cross-sectional income inequality and intergenerational persistence of inequality¹³. From this perspective, the evidence for Chile would be consistent with this hypothesis, considering that Chile has a particularly unequal distribution of income in compassion with the international evidence¹⁴.

5.2. Intergenerational income and educational mobility across cohorts

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

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, as mentioned earlier, this pattern can also be the result of life-cycle effects on earnings, and therefore whether income mobility has increased in time cannot be concluded¹⁵. However, a way round this problem is that unlike earnings, schooling levels usually exhibit little or negligible life-cycle effects after the time when most individuals stop studying and enter the labor market, usually around the mid twenties, after which schooling levels remain largely fixed for most individuals (as we shall see below). Following this reasoning we attempt to examine whether intergenerational *educational* mobility has exhibited changes in the recent decades by comparing educational mobility values across cohorts.

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

In order to substantiate this claim we explore the accumulation of schooling in Chile by age groups using the 1996-2001 CASEN Panel data. Figure 1 shows the average individual accumulation of schooling in years in Chile by age groups between 1996 and 2001. Figure 1 shows that while schooling accumulation is important before age 20, it decreases thereafter, and in fact, it

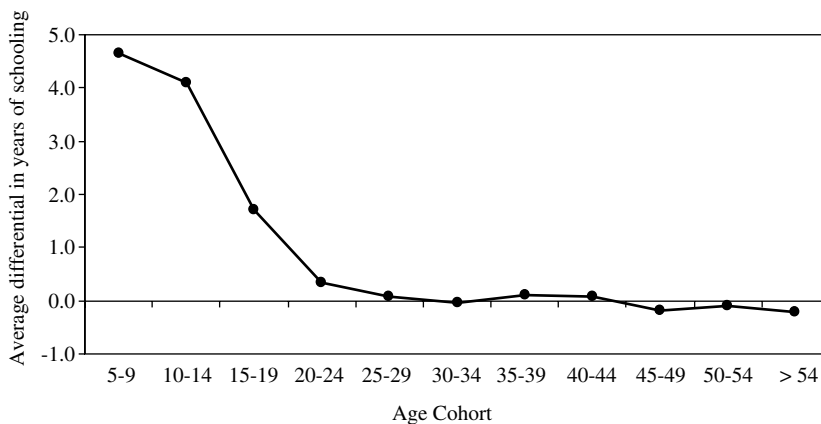
¹² In this context, for comparison purposes it would be useful to obtain intergenerational incomes elasticities for large urban areas in Brazil.

¹³ See for example Dunn (2004) and Andrews and Leigh (2008).

¹⁴ See for example De Ferranti, D., Perry, G., Ferreira F., and M. Walton (2003).

¹⁵ See Dunn (2004) for a discussion on the role of life-cycle effects on intergenerational income mobility elasticities.

FIGURE 1
1996-2001 AVERAGE INDIVIDUAL DIFFERENTIAL IN YEARS
OF SCHOOLING BY AGE



Source: Panel CASEN 1996-2001.

TABLE 6
SCHOOLING ELASTICITY BY COHORT

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

Robust t statistics in parentheses.

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

remains negligible after age 25 for the population as a whole. This suggests that life-cycle effects in years of schooling are unimportant for the age group 23-65 considered in this study. Therefore, finding evidence of higher intergenerational educational mobility for the younger cohorts would be suggestive of an increase of *educational* mobility in time.

Table 6 presents intergenerational schooling elasticities by cohorts. The evidence indicates lower values for the younger cohorts. Based on the argument presented above, this would be suggestive of an overall increase in intergenerational educational mobility in the last decades in Greater Santiago, for both men and women.

TABLE 7
COHORT EFFECTS IN YEARS OF SCHOOLING.
DEPENDENT VARIABLE: SONS' AND DAUGHTERS SCHOOLING

Variables	All	Sons	Daughters
Father's schooling	17.183 (5.21)**	14.440 (3.06)**	19.945 (4.29)**
Father's schooling*Cohort	-0.009 (5.09)**	-0.007 (2.98)**	-0.010 (4.21)**
Cohort	0.119 (7.10)**	0.022 (7.00)**	0.089 (3.55)**
Constant	-223.716 (6.83)**	-40.527 (6.65)**	-165.007 (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 INTERGENERATIONAL SCHOOLING ELASTICITY
Dependent variable: Sons' log schooling

Variables	All	Sons	Daughters
Father's log schooling	15.156 (5.67)**	16.548 (3.65)**	13.554 (4.66)**
Father's log schooling*Cohort	-0.008 (5.60)**	-0.008 (3.62)**	-0.007 (4.59)**
Cohort	0.022 (4.11)**	0.153 (7.11)**	0.022 (6.68)**
Constant	-40.567 (3.90)**	-292.412 (6.91)**	-41.082 (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%.

This is further confirmed in Tables 7 and 8. While these tables indicate a positive association between father's and son's schooling levels, as well as strong evidence of an expansion in schooling in time, particularly for daughters (as indicated by the coefficients of the *cohort* variable), they also provide evidence of a lower association between father's and son's schooling for the younger cohorts, as indicated by the coefficients of the interactive term in both specifications.

To what extent this evidence of lower statistical association between father's and son's education in the younger cohorts is related or causes the lower intergenerational income elasticities of the younger cohorts reported earlier

in Table 5 remains open for future research. In particular, it seems plausible that factors associated with socioeconomic background such as the quality of education, social capital and access to social networks, or the existence of class-discrimination in the labor market, for example, may yield different returns to schooling for individuals according to their social background¹⁶. Likewise, these factors may limit the transformation of the higher degrees of educational mobility of the younger cohorts reported above into higher levels of intergenerational income mobility. This hypothesis, however, remains open for future research.

5.3. Patterns of intergenerational income mobility in urban Chile

We finally examine some patterns of intergenerational income mobility in urban Chile. In particular, in this section we study whether the degree of the intergenerational transmission of the socioeconomic status varies across different population segments of the income distribution. Table 9 reports an estimates of quintile transition matrix for labor income using father's education and potential experience as predictors of father's income, and deriving father's quintiles from the real corresponding income distribution. As a robustness check, Table 10 reports an equivalent transition matrix in which father's quintiles are obtained from the distribution of the father's predicted incomes instead.¹⁷ Yet, 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 fairly low, around 0 to 8 per cent, also consistent with the international evidence.

TABLE 9
QUINTILE TRANSITION MATRIX
Father's quintiles obtained from real income distribution

Father	Son				
	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				

¹⁶ For example Núñez and Gutiérrez (2004) report a 50 per cent gap in earnings of Chilean professionals from different socioeconomic background of origin after controlling for academic performance, experience, school's academic performance, postgraduate studies, command of a second language, among other controls.

¹⁷ Quintile transition matrices are provided in Tables A3 and A4 in the appendix, also illustrated in Figures A1 and A2.

TABLE 10
QUINTILE TRANSITION MATRIX
 Father's quintile obtained from distribution of father's predicted incomes

Father	Son				
	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 suggest two important disparities in the intergenerational income transmission mechanism. First, they show more income persistence at both extremes of the father's income distribution, together with a substantial level of intergenerational mobility in the intermediate father's quintiles. Second, they suggest an asymmetry in the degree of intergenerational persistence at the two extremes of the father's 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 father's centiles versus son's centiles are reported in Table 11.

Table 11 shows that the association between father's and their son's income centiles is increasing (as expected) but not linearly. In fact, specifications 2 and 6 of Table 11 show that when a quadratic functional form is imposed it yields a robust overall convex pattern, indicating that the intergenerational income persistence is asymmetric, being higher in the upper part of the father's income distribution. Yet, the cubic specifications in 3, 4, 7 and 8 of Table 11 outperform the quadratic models, indicating that intergenerational persistence is higher at the two extremes of the income distribution than in the intermediate segments of the father's income distribution, consistent with the transition matrixes¹⁸. Table 12 also indicates by means of a structural change test that the slope of father's vs. son's centiles is statistically steeper at the top than at the bottom of the father's income distribution.

This pattern is depicted in Figure 2, which shows the profile of cubic OLS and quantile regressions of father's vs. son's income centiles. But Figure 2 also illustrates how broad is the spectrum of income centiles where sons are likely to end up in adulthood, conditional on their father's income centiles. Note that while for the most part of the father's income distribution the sons can end up as adults in income centiles often quite different from their parent's, at the top of the father's income distribution chances are that sons will occupy relative income positions similar to those of their fathers.

¹⁸ A similar pattern is observed when the specification is regressed separately for the 23-44 and the 45-65 age groups (see figures A3 and A4 in the appendix).

TABLE 11
ALTERNATIVE FUNCTIONAL FORMS FOR INTERGENERATIONAL INCOME MOBILITY IN URBAN CHILE
(Centiles of labor income)

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

FIGURE 2
QUANTILE CUBIC REGRESSIONS

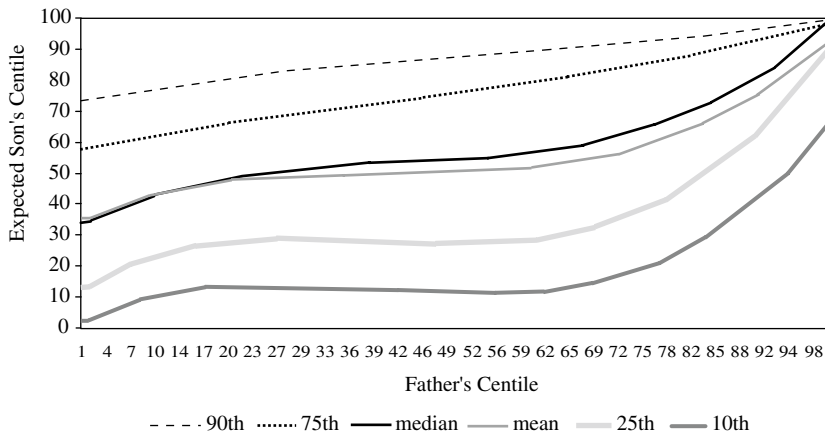


Table 13 indicates that the expected income centiles of sons of fathers in the bottom decile is 30, that is, at the border of the 3rd and 4th deciles. Although this shows an important degree of intergenerational persistence, it also shows an important degree of the “regression-to-the-mean” effect for this group. However, sons of fathers in the top income 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 itself in a milder version when comparing the expected

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

income centiles of sons of fathers in the 2nd and 9th deciles, namely 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 income centile is only a few percentage points away from the median, confirming an important degree of intergenerational income mobility at the centre of the father's income distribution.

It is interesting to note that the latter result is consistent with evidence on intergenerational *occupational* mobility for Chile. In particular, 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. As a hypothesis waiting for research, this converging evidence may be associated with Chile's particular income distribution, characterized by an unusually large share of the national income being held by the top decile or so of the population, the remaining part of the population being fairly egalitarian¹⁹. Put quite simply, it seems plausible that the top decile or so of Chile's income distribution may be largely responsible for Chile's unequal distribution of income as well as for shaping Chile's social mobility patterns, a hypothesis that is open for future research.

6. CONCLUSIONS

This paper has presented some findings on intergenerational income and educational mobility in urban Chile. A constraint to study intergenerational mobility in Chile, as in most of the developing world, is the lack of income panels where both fathers and their offspring's income can be observed. In this context, we follow the widely employed methodology known as two-sample two-stage

¹⁹ See for example De Ferranti *et al.* (2003).

least squares, whereby father's incomes are predicted from income determinants reported by their sons, such as father's schooling, occupation and age.

We find intergenerational income elasticities for Greater Santiago in Chile in the range of 0.52 to 0.54. These figures seem lower than nationwide intergenerational income elasticities in the range of 0.6-0.7, but still high in comparison with the international evidence. The lower values for Greater Santiago may reflect the higher educational and occupational opportunities than can be expected in a large urban centre, compared to rural and small urban areas of the country, yet this hypothesis requires explicit investigation.

The intergenerational income elasticities are somewhat lower for the younger cohorts. While this may suggest an increasing intergenerational mobility in time, it can also be associated with life-cycle effects on earnings, as indeed suggested by the literature. However, we find that intergenerational educational mobility is also lower for the younger cohorts for both sons and daughters. Based on the fact that life-cycle effects are unimportant as schooling remains fixed for most individuals along the life cycle after the mid twenties, this finding is suggestive of and increase of intergenerational educational mobility in urban Chile in the last decades. This seems consistent with the significant expansion of school and tertiary education observed in the last decades in Chile. Whether this phenomenon has indeed translated into higher intergenerational *income* mobility in recent decades is open for future research.

We also examine some patterns of intergenerational income mobility by studying how the degree of mobility varies along the father's relative income position. We find evidence of a higher degree of intergenerational income persistence (lower mobility) at the bottom-to-bottom and top-to-top transition probabilities, and a higher degree of intergenerational mobility in the intermediate segments of the father's income distribution a pattern that is consistent with the international evidence. However, we find also that intergenerational mobility is substantially lower at the top of fathers' income distribution. While this may be an expected pattern in other countries, this evidence is consistent with recent findings of intergenerational occupational mobility for Chile by Torche (2005), who reports a high degree of intergenerational persistence in occupations associated with high social status compared with a variety of other countries. It is tempting to propose, as hypothesis awaiting explicit research that this findings may be associated with Chile's particular income distribution, known to be quite egalitarian within the bottom 80 to 90 percent of the population, but relatively unequal for the country as a whole as a consequence of a high concentration of income in the top centiles of the distribution.

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APPENDIX
TABLE A.1
ESTIMATES OF MINCER EQUATIONS USING EDUCATION AND POTENTIAL EXPERIENCE AS REGRESSORS

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

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

Father	Son			
	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

Father	Son			
	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)

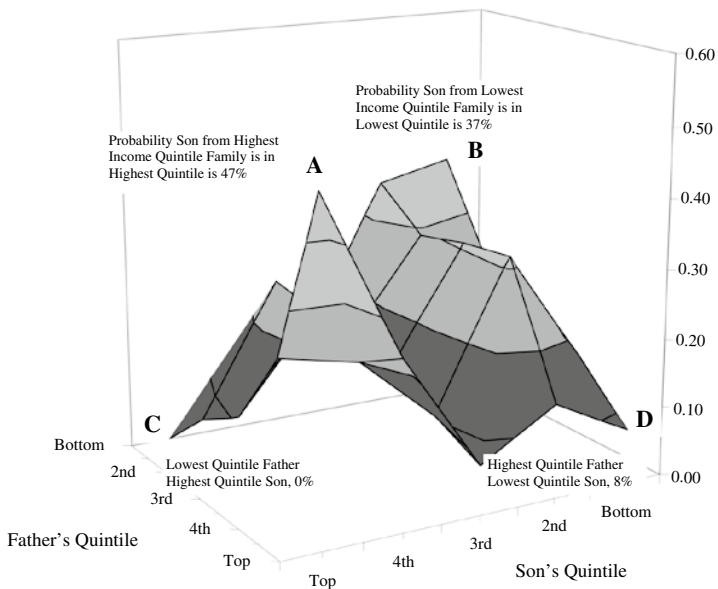


FIGURE A2
INTERGENERATIONAL INCOME TRANSITION PROBABILITIES
(Quartiles)

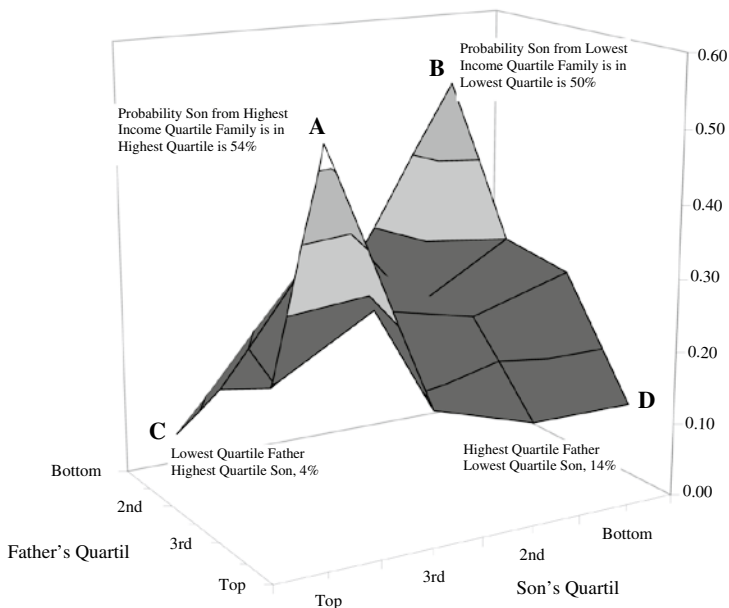


FIGURE A3
QUANTILE CUBIC REGRESSIONS FOR 23-44 COHORT

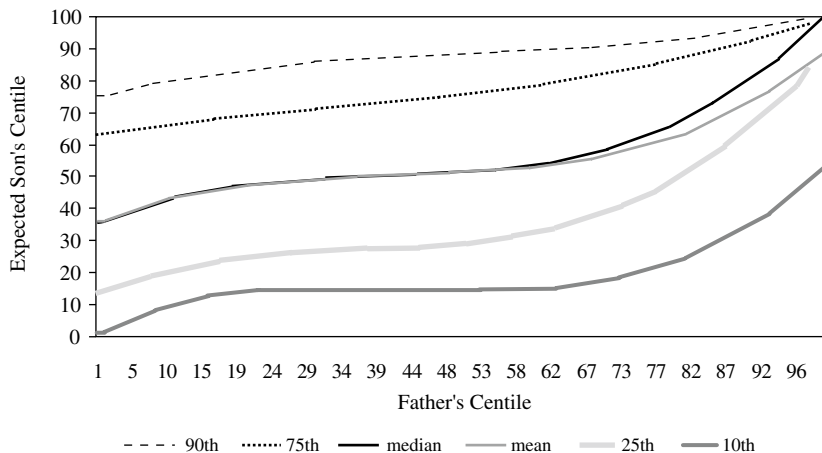


FIGURE A4
QUANTILE CUBIC REGRESSIONS FOR 45-65 COHORT

