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Facultad de Ciencias Sociales
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Intergenerational Educational Mobility: evidence from three approaches for Brazil, Chile, Uruguay and the USA (1995-2006)

Graciela Sanroman

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Intergenerational Educational Mobility: evidence from three approaches for Brazil, Chile, Uruguay and the USA (1995-2006)[1]

Graciela Sanroman[2]

Preliminar and incomplete

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[1]Guzman Ourens prepared the data on Brazil, Chile and the USA and collaborated with routines for Indexes estimation. I also appreciate his collaboration in the revision of this manuscript. Cecilia Gonzalez constructed the Uruguayan datasets.

[2]Departamento de Economía, FCS, UDELAR, e-mail: gsanroman@decon.edu.uy

Abstract

In this paper we estimate intergenerational educational mobility for Brazil, Chile, Uruguay and the USA along the period 1995-2006. We propose an index of intergenerational mobility based on the variance decomposition in an error-components model. We estimate three indexes of mobility: one based on the autoregressive Markov-chain regression, the second is the Dahan-Gaviria index and the last one is the index proposed in this work. We use data of teenagers and parents education and address the issue of top-censoring. We distinguish between relative and absolute mobility. We draw empirical conclusions for each country and compare results. We analyze theoretical and empirical attributes of indexes.

Keywords: Intergenerational Educational Mobility, Error-components Model, Censored data

Resumen

En este trabajo se estima la movilidad educativa intergeneracional para Brasil, Chile, Uruguay y los EEUU a lo largo del período 1995-2006. Se propone un índice de mobilidad intergeneracional basado en la descomposición de la varianza en un modelo de componentes de error. Se estiman tres índices de movilidad: el primero basado en una regresion autoregresiva del tipo cadenas de Markov, el segundo es el índice Dahan-Gaviria y el tercero el índice propuesto en este trabajo. Se utilizan datos de los años de estudio de los adolescentes de sus padres y se trata econométricamente el problema de la censura superior. Indices de movilidad relativa y absoluta son obtenidos, se derivan conclusiones empíricas para cada país y se comparan los resultados. Se analizan las propiedades de los índices desde el punto de vista teórico y empírico.

Palabras clave: Movilidad educativa intergeneracional, Modelo de componentes de error, Observaciones censuradas.

JEL: C24, C51, J62, O54

1 Introduction

Latin America is known as the most unequal region in the world. There are many empirical studies of inequality for the Latin American countries, but the vast majority of them rely on the static dimension of inequality, by measuring some index of inequality using the cross-section distribution of some variable of interest (e.g. income, education, consumption, occupation). The problem with this approach is that, if what really matters is the equal access to opportunities, these indexes are misleading: two societies with the same static inequality could be very different in terms of equality of opportunities depending on the processes of transference of the socioeconomic status from parents to their children. Therefore the study of intergenerational social mobility can help to understand and measure inequality.

There is not an unique definition of intergenerational mobility, but it can be thought as the process through which parental status is transferred to children (Becker and Tomes, 1979; Goldberger, 1989), and thus intergenerational mobility can be naturally measured as some index of the correlation between some parental and children socioeconomic outcome.

Although it is usually based on the same variables than in the analysis of static inequality, the study of intergenerational mobility is much more complicated. In particular, it is very difficult to use income and occupation if the aim is to measure correlation among different generations. It is well-known that observed income is influenced by time trends and cycles, aggregated and idiosyncratic temporary shocks; moreover, income varies significantly over the life cycle and the life-cycle patterns of income could change from one generation to the following. Thus, it is very difficult to measure the covariance between income of parents and those of their children that could be considered permanent using panel data, and it is almost impossible if only cross-sectional data is available. It is also very difficult to deal with the variable occupation within the study of intergenerational mobility. It also varies along the life cycle, but even more important is that both the set of occupations and the relative status of each occupation have been varying from one generation to the following since the industrial revolution. Besides, it is very difficult to measure the occupation status through a quantitative variable.

In this paper we focus on intergenerational educational mobility, hence our analysis relies on the educational attainments of children and parents. To focus on education has some advantages and disadvantages. On one hand, the education stays invariant from certain age and that is very convenient in comparison with the use of earnings, income or occupation. Moreover, measurements of education attainments are available for almost all countries. On the other hand, the main disadvantage in using education is that it could be poorly correlated with income and earnings, and in general with the welfare of people. Machin (2004) finds evidence for Britain which shows that the persistence of income has increased over time following the rapid expansion of post-compulsory education. However, he also shows that education attainments can help to explain income persistence within families across generations. Moreover, it is well known that better-educated people have better jobs, experience less unemployment and earn higher wages. That is particularly true in Latin America where returns to schooling are very high, and there is evidence showing that they have been increasing during last years.

One important limitation to study intergenerational mobility of Latin American countries is that surveys that collect information about the characteristics of the parents of interviewed adults are rare and non periodic; moreover, there are only few longitudinal surveys that could be use to measure intergenerational mobility. Previous comparative studies on intergenerational mobility for the Latin American countries are found in the works of Andersen (2001), Behrman et al (2001), Dahan and Gaviria (1999) and Fields et al (2007).

Dahan and Gaviria (1991) and Behrman et al (2001) use standard household surveys to measure intergenerational educational mobility through the correlation between the schooling attainments of teenager children still living with their parents. These surveys are commonly found for almost all Latin American countries during the last two decades, and that is an important advantage of relying the analysis on these data. Moreover, this information allows to analyze what type of society is being currently developed, in terms of intergenerational mobility, and therefore it allows to obtain useful policy implications, opposed to when the used data are about adults who already finished their education. However, a major econometric issue arises when using years of schooling of teenagers: many of them are still attending formal education, thus some non-negligible proportion of the observations about the years of schooling is top-censored.

In this work we use three different approaches to measure intergenerational educational mobility. First, we estimate mobility as is usual through the coefficient of the education of the parents in a first-order autoregressive Markov-chain regression where the dependent variable is a measure of the education of the teenager. Second, we calculate the index proposed by Dahan and Gaviria (1999). Finally, we perform a variance decomposition analysis using an error-components model (Arellano, 2003). We consider non-censored and censored versions of the Markov-chain regression and the error-components models. We also distinguish between relative and absolute mobility.

We use these three approaches to measure intergenerational educational mobility in three Latin American countries (Brazil, Chile and Uruguay), and compare results across these countries and with those of the United States of America (USA). We also analyze the evolution of intergenerational mobility within each country over the period 1995-2006 and compare time trends. The surveys we use are the Brazilian "Pesquisa Nacional por Amostra de Domicílios" (PNAD), the Chilean "Encuesta de Caracterización Socioeconómica" (CASEN), the Uruguayan "Encuesta Continua de Hogares" (ECH) and the "Current Population Survey" (IPUMS-CPS[1]) from the USA. Data are available for every year in the period 1995-2006 in the cases of Uruguay and the USA but for Brazil year 2000 is not available. Finally, the Chilean CASEN is available only for 1996, 1998, 2000, 2003 and 2006.

The selection of these countries has been made by taking into account that previous studies suggest that Uruguay and Chile exhibit lower inequality and a greater degree of intergenerational mobility than other countries of the region, while Brazil is considered the most unequal Latin American country. Moreover, evidence shows that during the last decades Chile and Brazil have experienced improvements in terms of equality, while the situation in Uruguay has been deteriorated, or at best stayed stagnant. In addition, the standard household surveys of these three countries includes reliable data on our variables of interest (age, years of completed formal schooling, attendance to education and relationships within the family). On the other hand, we include the USA in the analysis because it offers an extra-region point of comparison, commonly used and interesting by itself.

The main contribution of this work is to propose a flexible framework, based on the variance decomposition in an error-components model (Arellano, 2003), in order to obtain an index of intergenerational mobility. This approach is actually a simplification of the model estimated by Lillard and Willis (1994). We demonstrate that the indexes that are usually performed with the goal of measuring intergenerational mobility can be thought as par-

[1] King et al (2004)

ticular applications of the model that we propose. Also we show how different issues and features can be fitted in our framework, but they can not in the traditional approaches. Moreover, we contribute to the empirical analysis by estimating confidence intervals of the Dahan-Gaviria Index using the bootstrap method.

In addition, our empirical results are themselves interesting. We find that relative intergenerational educational mobility in the studied Latin American countries is lower than in the USA, but Brazil and Chile experienced substantial improvements in this features throughout the period of analysis according to all indexes. In terms of absolute mobility, we surprising find that at the end of the period Index 3 shows that Brazil and Chile were more mobile than the USA, but this result does not hold if we use Index 1. Uruguay has exhibited a bad performance and, although its position was not so unfavorable compared with the other countries of the region at the beginning, it was finally at the worst place both in term of absolute and relative mobility at the end.

The rest of this paper is organized as follows. In Sections 2 and 3 we review the characteristics of the indexes that were most commonly used in the literature of intergenerational mobility. In Section 4 we show how a new index of mobility can be obtained in the flexible framework of the error-components model. Section 5 analyzes some econometric issues that arise when studying intergenerational mobility. In Section 6 we briefly introduce the three indexes we use to measure mobility and analyze the empirical results. We analyze in detail the results for each country, afterwards we summarize across countries comparison and some evidence about the empirical performance of the three indexes. Finally, in Section 7 we present our conclusions and suggest some extensions of this study.

2 First-order autoregressive Markov-chain

Intergenerational mobility has been traditionally measured by estimating a simple linear first-order autoregressive Markov-chain regression.[2] The model can be written as follows,

$$S_{it} = \alpha + \beta S_{it-1} + u_{it} \quad i = 1, ...N \tag{1}$$

where S_{it} is a variable that measure some socioeconomic status of children, and S_{it-1} is the same variable for their parents (father/mother/average of

[2]Becker and Tomes (1979), Behrman (2000), Goldberger (1989), de Haan and Plug (2006), Solon (2003)

both). The index it here corresponds to each children in the sample, thus for example a family with two children will generates two non-related observations.

In this approach β becomes an index of intergenerational mobility, but the lower the index the higher is mobility. Estimates of equation (1) may be used to characterize intergenerational social mobility with socioeconomic indicators such as education, income, earnings, or occupational status.

Notice that $p\lim \widehat{\beta}_{MCO} = \frac{cov(S_{it}, S_{it-1})}{Var(S_{it-1})}$ and thus the point estimation of this index depends on the correlation between child and parent's outcomes and the variance of parents education. Therefore, it reflects how strongly children status is associated with parental status in comparison with the variability of this status within the cohort of their parents.

This model can be applied to measure absolute as well as relative intergenerational mobility. Absolute mobility is given by S measured in levels, while relative mobility is given by S in deviation from age-specific means. One alternative way to measure relative mobility is to consider S in levels and include age-dummies for children and parents. That alternative has the advantage that is more suitable to estimate the top-censored version of equation (1) as we explain later.

The first-order autoregressive Markov-chain approach has various shortcomings. First, the conditional variance of children education plays no role on the index (although determines its confidence interval). Second, the model deals with each child and his parent in an isolated way, thus some factors that are crucial to determine intergenerational mobility like assortative matings or the correlation between siblings are not taking into account. Finally, results could be different depending on whether S_{it-1} is referred to the mother, to the father or to some average of both.

Another standard way to characterize intergenerational mobility using Markov-chains is to use transition probability matrices for movements between generations among segments of the distribution (e.g. deciles). This allows to address asymmetries and other non-linearities, but it brings up the issue of how to reduce such a probability matrix to a scalar that characterize the extent of mobility.

3 Dahan-Gaviria Index

Dahan and Gaviria (1999) and Behrman et al (2001) use standard household surveys to measure intergenerational educational mobility through the correlation between the schooling attainment of teenagers still living with their parents. Behrman et al (2001) restrict the sample to teenagers aged sixteen to twenty. They argue that in Latin America a high proportion of young adults in this age range still live in the parental household and that going above this age group would imply substantial losses of information and probably biases, while to include children under sixteen would be non informative because schooling differences start becoming apparent precisely around this age.

Behrman et al (2001) measure intergenerational educational mobility using the Dahan-Gaviria Index (DG Index). The DG Index estimates the intergenerational mobility as the correlation of the educational attainments of siblings.

Notice that is expected that a considerable proportion of children aged 16-20 have not finished their educational cycle and thus the dependent variable of equation (1) is top-censored (Haan and Plug, 2006). In order to deal with top-censoring Dahan-Gaviria (1999) define a binary variable d that equals 1 if the individual years of schooling is above the median of his cohort.

$$d_{ij} = \begin{cases} 1 & S_{ij} > S_{aj} \\ 0 & \text{otherwise} \end{cases} \quad j = 1, 2..F_i \tag{2}$$

where S_{ij} and a_j are the years of schooling and the age of the j sibling of family i, and S_{a_j} is the median years of education of individuals aged a_j in the sample, F_i is the number of teenagers siblings of family i living together.

They include only families with $F_i \geq 2$. Afterwards, they calculate:

$$\rho_a = 1 - (1 - \rho_g)\frac{B-1}{B-F} \tag{3}$$

where $\rho_g = \frac{\sum_{i=1}^{F}(\overline{d}_i - \overline{d})^2}{\overline{d}(1-\overline{d})}$, $\overline{d}_i = \frac{1}{F_i}\sum_{j=1}^{F_i} d_{ij}$ $(i = 1, ...F)$, F is the number of the families in the sample, $\overline{d} = \frac{1}{F}\sum_{i=1}^{F}\overline{d}_i$, and $B\left(= \sum_{i=1}^{F} F_i\right)$ is the total number of teenagers in the sample. The authors indicate that the ρ_g also corresponds to the R–squared obtained by regressing d_{ij} on a set of dummy variables for each families i in the sample. Here ρ_a is the DG Index, it belongs to the unit interval and the higher the index the lower the mobility.

Dahan and Gaviria (1999) argue that ρ_a measures the extent to which schooling outcomes can be explained by family background. If there were perfect social mobility, family background would not matter, siblings would not be more alike than two people taken at random, and the DG Index would be close to zero. If there were little mobility, family background would matter very much, siblings would be very similar and the DG Index would be close to one.

We believe that the main contribution of the Dahan-Gaviria approach is that it relies on in the data of teenagers education. This alternative has two main advantages. First, periodic information about teenagers and their parents educational attainments are available for Latin American countries. Second, we consider that useful policy implications can be suggested when the study of intergenerational mobility is based on data about teenagers.

However, from our point of view the Dahan-Gaviria approach has some shortcomings. The main one is that parental attributes play no role in the DG Index: two equivalent families in terms of their observational teenagers characteristics could be very dissimilar in terms of parental characteristics, and this constitutes a major feature that affects intergenerational mobility.

Furthermore, to deal with censoring by defining a dummy variable could be an inefficient alternative, because useful information is lost in comparison with other available econometric techniques. Another limitation of the DG Index is its potential non-random sample-selection. As Behrman et al (2001) pointed out, low fertility households are more likely to be excluded than are high fertility households. If there is a trade-off between quantity and quality, the excluded low fertility households are likely to have relatively high child schooling. However, the authors conclude that it is not clear that this exclusion biases the estimates of intergenerational schooling mobility or affects cross-country comparisons. We think that the bias that is introduced, e.g. when not incorporating only children in the sample, could be more important than authors consider. There are great differences on fertility behavior between the countries under analysis. For example, Uruguay early processed the demographic transition, while Brazil began late and is still processing it. Thus, we suspect that the sample bias could affect the cross country analysis and also (at least for Brazil) over time comparisons.

Finally, another weakness of the DG Index, in comparison with the index β, is the calculus of its confidence interval. Actually, neither Dahan and Gaviria (1999) nor Behrman et al (2001) have calculated it. One contribution of this paper is to obtain the confidence interval for the DG Index. The 95 percent confidence interval of ρ_a is obtained as the 2.5^{th} and 97.5^{th} percentiles of the empirical distribution of ρ_a obtained through the bootstrap method. In order to apply the bootstrap method to the DG Index it is necessary to be careful. We proceed in this way: re-sample families not individuals (to this end the sample is shaped in a wide-form, i.e. we consider a row for each family). Afterwards, the variable that indexes families is changed in each resulting sample by simply re-index every row with a unique number, and then we calculate the DG Index as usual.

4 The error-components model framework

4.1 The basic framework

In this paper we propose an index to measure intergenerational mobility based on the variance decomposition of error-components in a "Random Effect model" (Arellano, 2003). The most simple specification of the model is as follows,

$$S_{ij} = \mu + \eta_i + v_{ij} \quad i = 1, ...N, \quad j = 1, ...N_i \tag{4}$$

where N_i is the number of members in family i, we assume

$$v_{ij} \sim iid(0, \sigma_v^2) \qquad \text{(Assumption 1)}$$

$$\eta_i \sim iid(0, \sigma_\eta^2) \qquad \text{(Assumption 2)}$$

$$v_{ij} \text{ and } \eta_i \text{ independent of each other} \qquad \text{(Assumption 3)}$$

The parameter of interest is,

$$\rho = \frac{\sigma_\eta^2}{\sigma_\eta^2 + \sigma_v^2} \tag{5}$$

This approach is commonly used for panel data analysis where i indexes an individual and j time, like in the study of earnings inequality and mobility over time. Arellano (2003) points out that in this model the parameter ρ indicates the fraction of the total variance that remain constant over time while the rest are differences that vary randomly over time and units. Notice that it is straightforward to consider a model where i indexes families and j each member of the family. In that context ρ tells us the fraction of the total variance of S that corresponds to variability across families that remains unchanged from one generation to the following.

Therefore using the parameter ρ with the goal of measure intergenerational mobility seems a natural application of this model, and that is precisely what we propose in this paper. This approach can be found in the work of Lillard and Willis (1994). Lillard and Willis explore evidence concerning the relationship among the educational attainments in Malaysia of four generations within a given family, using a sequential probit model which allows for correlations among unmeasured family and individual-specific components.

It is interesting to notice that parameters β and ρ_a of equations (1) and (3) can be thought like particular applications of the error-components model.

First, let us consider that the index j takes only two values 1 and 2, with S_{i1} denoting the parent's outcomes and S_{i2} the children's outcomes. In this specification, and under Assumptions 1, 2 and 3, $p\lim \widehat{\beta}_{MCO}$ of equation (1) equals ρ in (5).

Second, if we define $S_{ij} = d_{ij}$ of (2) and consider $j = 1, 2, ...F_i$ as in the DG Index, then the Balestra-Nerlove estimator (and also the MLE) of ρ in (5) is numerically almost identical to ρ_a in (3). The ML approach requires to assume that v_{ij} and η_i are normally distributed, but under these assumptions it is possible to obtain a confidence interval for the DG Index that does not rely on the bootstrap method.[3]

4.2 Error-components

The basic error-components model incorporates two unobservable variables: η_i and v_{ij}. In general, η_i captures in a very flexible way different types of unobservable or omitted attributes which characterize families or are common for their members (e.g., assortative matings, human capital, inherited ability or intelligence, income and wealth, borrowing constraints, social networks). On the other hand, v_{ij} is related with idiosyncratic characteristics of the j member of family i that would be not explained by his or her family background.

Notice that causalities can be interpreted in various ways within the model. For example, if η_i is intelligence, and intelligence is inherited, there will be a causal influence from η_i to the educational attainments of the members of the i family. Besides, if η_i capture income the process of transference can be though in the following way: parental education affects family income, which in turn influences investments in the children education (assuming non-perfect capital markets), and consequently teenagers educational attainments. A very convenient feature of the error-components model is that it is able to capture all these types of factors that are in the core of intergenerational transferences in a very parsimonious way, and separates them from factors that also influence the teenagers attainments but have no relationship with their family background.

Let us briefly present some intuitions about the effect of η_i on the variation of the index across countries or over time. It is obvious that changes in

[3] Nevertheless, to obtain the confidence interval of the DG Index in that way seems not to be appropiate because the dependent variable is a binary outcome.

the index are only related with varying unobservable or omitted factors. Although the incentives to use and develop intelligence could vary, the process of transference of genetic intelligence from one generation to the next will be invariant over time (at least when two generations are considered) and across countries, and thus we do no expect any effect of this unobservable characteristic in changes of the index. However, notice that we expect that the higher the degree of positive assortative matings the higher the index, and the intensity of assortative matings can change across societies and over time. On the other hand, the returns to schooling and the cost of education can vary significantly both across countries and over time. Thus, to distinguish the factors that cause changes on mobility could be interesting, and part of this work could be easily done by including observables into the analysis.

In addition, we can enrich the structure of the unobservables within the error-components model to obtain further evidence about the processes of intergenerational transferences. As we already pointed-out, the error-components model takes into account and relates the information about all members in the family and distinguish correlated and no correlated factors. Moreover, it is also able to isolate the correlations among group-specific components within families, like in Lillard and Willis (1994). The following equations are alternative specifications to (4) that illustrate this feature,

$$\begin{aligned} S_{ij} &= \mu + \eta_i + \eta_{ip} + v_{ij} \; i = 1, ...N, \; j = 1, ...N_i && (6) \\ S_{ij} &= \mu + \eta_i + \eta_{ic} + v_{ij} \; i = 1, ...N, \; j = 1, ...N_i && (7) \\ S_{ij} &= \mu + \eta_i + \eta_{ib} + \eta_{is} + v_{ij} \; i = 1, ...N, \; j = 1, ...N_i && (8) \\ S_{ij} &= \mu + \eta_i + \eta_{ip} + \eta_{ib} + \eta_{is} + v_{ij} \; i = 1, ...N, \; j = 1, ...N_i && (9) \end{aligned}$$

where η_{ih} is an error-component that captures the correlation between father and mother if $h = p$ (assortative matings), siblings if $h = c$ (allowing to obtain the DG Index in a more general framework), brothers if $h = b$ and sisters if $h = s$ which could be interesting if we suspect that gender differences are of interest for the country under analysis (Lillard and Willis, 1994).

It is not obvious for us which of the previous specification is the best in order to measure intergenerational mobility. However, to start by estimating equation (4) and only afterwards deal with richer decompositions seems to be a proper way to proceed with our research. Therefore, in this paper we concentrate in the simpler specification and leave the others to future research.

4.3 Observable characteristics and causal effects

The model can easily accommodate observable characteristics of the members of the family and their environment. However, it is not obvious which characteristics should be included to obtain an index of mobility. Including different sets of regressors would yield different measures of mobility. Natural candidates are cohort and age. For example, if we consider the model like in (4) without including regressors we will measure absolute mobility, while if we include dummies for the age of each individual we will measure relative mobility. To measure relative mobility by including age-dummies instead of considering the dependent variable through deviations from age-specific means has the main advantage that it is easier to deal with censored data, as we have already argued and will become clear below.

We believe that it is not appropriate to add more regressors to the model, if the goal is to measure intergenerational mobility. To include other observable characteristics (e.g. family income and size, the age of the parents when children was born, race, fertility, measurements of the quality and spread of the educational system) will be useful to analyze the driving forces that determine mobility, to test theoretical hypothesis about parents investments on children education and to relate differences across countries (ethnic groups, over time) with their characteristics.

Another easy way to analyze the evidence on the determinants of mobility is to estimate a macro panel-data model where the units are countries, the dependent variable is the index of mobility and the regressors are relevant characteristics of the countries at each point in time.

In addition, it would be interest to estimate the causal effect of parental education on children education. However, in this paper we focus on measuring mobility, and therefore we analize neither the determinants of mobility nor the causal effect of parental education.

5 Econometric issues in the error-components model framework

5.1 Heteroskedasticity

If ideal data about children and parents outcomes were available, the main econometric issue in our framework is heteroskedasticity. This is because assumption 1 seems to be very restrictive in the context of our analysis, it implies that,

$$Var(v_{ij}) = \sigma_v^2 \;\; \forall i \, \forall j$$

It would be interesting to relax Assumption 1 by allowing that the variance of v varies across cohorts. This can be done, for example, by allowing the unconditional variance of S for the parents generation to be different from that of their children. Let us consider a model that includes each child ($j = 2$) and one of his or her parents ($j = 1$) without considering any other members of the family, so the model becomes,

$$S_{ij} = \mu_j + \eta_i + v_{ij} \; i = 1, ...N, \; j = 1, 2$$

and let us assume,

$$\begin{aligned} Var(S_{i1}) &= \sigma_1^2 \;\; \forall i \\ Var(S_{i2}) &= \sigma_2^2 = \sigma_1^2 + a \quad \forall i \qquad a > -\sigma_1^2 \end{aligned}$$

$$\begin{aligned} v_{ij} &\sim iid(0, \sigma_{vj}^2) \;\; j = 1, 2 \\ \eta_i &\sim iid(0, \sigma_\eta^2) \end{aligned}$$

$v_{i1,}v_{i2}$ and η_i independent of each other.

Three type of measures can be derived in this context:

$$\rho_p = \frac{\sigma_\eta^2}{\sigma_1^2} = \frac{\sigma_\eta^2}{\sigma_\eta^2 + \sigma_{v1}^2} \tag{10}$$

$$\rho_c = \frac{\sigma_\eta^2}{\sigma_2^2} = \frac{\sigma_\eta^2}{\sigma_\eta^2 + \sigma_{v2}^2} = \frac{\sigma_\eta^2}{\sigma_\eta^2 + \sigma_{v1}^2 + a} \tag{11}$$

$$\rho = \frac{\sigma_\eta^2}{\sigma^2} = \frac{\sigma_\eta^2}{\sigma_\eta^2 + \sigma_v^2} = \frac{\sigma_\eta^2}{\sigma_\eta^2 + \sigma_{v1}^2 + \frac{a}{2} + \frac{b^2}{4}} \tag{12}$$

where μ_j=$E(S_{ij} \mid j)$, $\sigma_j^2 = V\,(S_{ij} \mid j)$ with $j = 1, 2;.b = \mu_2 - \mu_1$.

Parameter ρ_p is equivalent to β in equation (1) since it could be equivalent to the coefficient associated with parents outcomes in a linear regression where the dependent variable is given by children outcomes. Parameters defined in (10) and (11) can be interpreted in terms of intergenerational mobility, ρ_p measures the fraction of inequality within the parents cohort that is transferred to the next generation, while ρ_c measures inequality within the children cohort that is inheriting that of the previous generation.[4] On the

[4]Notice that although it seems not very natural to regress parent outcomes on children outcomes, the parameter that is obtained has in our context a clear and interesting interpretation.

other hand, ρ is based on a "mixture" of the variance of both generations. Notice that $\rho_p < \rho < \rho_c$ if $a < 0$, i.e. the variance within the children cohort is lower than that within the parents cohort, while $\rho_p > \rho > \rho_c$ in the opposite case. Moreover, the situation where $\rho_p = \rho = \rho_c$ can be thought like a "steady state" situation in terms of static inequality (characterized by σ) and ρ captures the fraction of inequality that is inherited from the previous generation.

It could be also interesting to allow for heteroskedasticity of unknown form (Arellano, 2003) or to relax assumption about the independence between v_{it} and η_i considering that the variance of the former might be affected by the latter. However, these issues are not analyzed in this study.

5.2 Censored data

In order to apply our methodology to the available information, censored data become a major issue. In particular if the outcomes of interest are educational attainments. This issue is also present in much of the recent work on intergenerational educational mobility which rely on samples for which information on children completed schooling is not always available.

De Haan and Plug (2006) focus on the issue of censoring. They first estimate the impact of parents schooling on children schooling using censored and uncensored samples of own-birth children and adoptees, and then investigate the consequences of three different methods that deal with censored observations: maximum likelihood approach, elimination of all school-aged children and replacement of observed with expected years of schooling. They found that the best alternative to address censoring is the latter. However, there is no data on expectations about teenagers definitive schooling in the available surveys for Latin American countries that we analyze in this paper. Thus, to address censoring we use the likelihood approach under the assumption that errors are normally distributed.

In the case of Index 1, the model can be re-written as follows

$$S^*_{ij} = \alpha + \beta_2 S_{i1} + u_{ij} \quad i = 1,..,N \; j = 1,...F_i \tag{13}$$

$$S_{ij} = \begin{cases} S^*_{ij} & \text{if } A_{ij} = 0 \\ C_{ij} & \text{if } A_{ij} = 1 \end{cases} \tag{14}$$

where S^*_{ij} is a latent variable that indicates the definitive educational attainment of child j in family i; A_{ij} is a dummy variable that equals 1 if he or she is attending school, college or university; S_{ij} is the observed completed years of schooling of children. We assume that S_{ij} is equal to

S^*_{ij} if A_{ij} equals 0, but is top-censored at value C_{ij} if A_{ij} equals 1. Notice that C_{ij} varies across individuals, and thus the model may be thought as a generalized Tobit model. Besides, S_{i1} indicates the years of schooling of his or her parent.[5]

In order to estimate the model we add the assumption that $u_{ij} \mid S_{i1} \sim Niid(0, \sigma_u^2)$, under this assumption the contribution to the log-likelihood of each individual is:

$$l_{ij} = (1-A_{ij}) \log \left[\frac{1}{\sigma_u} \phi \left(\frac{S_{ij} - \alpha - \beta S_{i1}}{\sigma_u} \right) \right] + A_{ij} \log \left[\Phi \left(-\frac{S_{ij} - \alpha - \beta S_{i1}}{\sigma_u} \right) \right]$$

where ϕ and Φ denotes the density and the cfd of the standard normal distribution. The log-likelihood of the sample is

$$L(S, A; \alpha, \beta, \sigma_u) = \sum_{i=1}^{N} \sum_{j=1}^{F_i} l_{ij}$$

Notice that in this model $\beta_2 = \frac{cov(S_{ij}, S_{i1})}{Var(S_{i1})}$ and thus censoring affects only the numerator of the index. Moreover, $cov(S_{ij}, S_{i1})$ will be higher in the censored version of the model (in comparison with the non-censored version) if the probability that a teenager is attending education is positively affected by the education of his parents (as is observed in practice). Censoring will also affect the estimation of σ_u and thus the confidence interval of β_2, but σ_u will not affect the point estimation of β_2. In consequence, we expect that the index of intergenerational educational mobility will be higher in the censored version than in the non-censored version of Index 1.

For the censored version of the error-components model[6], we consider the following specification,

$$\begin{aligned} S^*_{ij} &= \mu + \eta_i + v_{ij} \quad i = 1, ...N, \quad j = 1, ...N_i && (15) \\ S_{ij} &= \begin{cases} S^*_{ij} & \text{if } A_{ij} = 0 \\ C_{ij} & \text{if } A_{ij} = 1 \end{cases} && (16) \end{aligned}$$

where variables are as defined above. This model were introduced by Heckman and MaCurdy (1980) in the context of the female labor supply.

To estimate the model we need to add assumptions about the distribution functions of η_i and v_{ij}. We assume,

$$v_{ij} \sim Niid(0, \sigma_v^2) \qquad \text{(Assumption 1')}$$

$$\eta_i \sim Niid(0, \sigma_\eta^2) \qquad \text{(Assumption 2')}$$

$$v_{ij} \text{ and } \eta_i \text{ independent of each other.} \qquad \text{(Assumption 3')}$$

[5] We neglect censoring on parental education. Notice that top-censoring on parental education affects this model as measurement errors.

[6] Arellano and Honoré (2001), Honoré (1992)

The index of mobility is given by $\rho = \frac{\sigma_\eta^2}{\sigma_\eta^2+\sigma_v^2}$ as in (5). Here the estimation of both σ_η^2 and σ_v^2 are affected by censoring. In particular, following the same arguments than above we expect the estimation of σ_η^2 to be higher in comparison with the non-censored version of the variance component model. In addition, we also expect (by definition) that the estimation of σ_v^2 will be higher in this specification. Thus, we do not have any hypothesis about the result of the ratio between censored and non-censored versions of Index 3. However, we can conclude that if this ratio is greater than 1, it will be smaller than in Index 1.

The estimation of this model is based on the ML approach. As suggested in Lillard and Willis (1994), the likelihood can be written as the product of independent conditional on η densities (probabilities) integrated over the distribution of η. The conditional likelihood of any member of the family is given by

$$\mathcal{L}(S_{ij}, A_{ij}, \eta_i; \mu, \sigma_v\ ) = \left[\frac{1}{\sigma_v}\phi\left(\frac{S_{ij}-\mu-\eta_i}{\sigma_v}\right)\right]^{(1-A_{ij})}\left[\Phi\left(-\frac{S_{ij}-\mu-\eta_i}{\sigma_v}\right)\right]^{A_{ij}}$$

while the conditional contribution of each family is

$$\mathcal{L}(S_i, A_i, \eta_i; \mu, \sigma_v\ ) = \prod_{j=1}^{N_i}\mathcal{L}(S_{ij}, A_{ij}, \eta_i; \mu, \sigma_v\ )$$

and the unconditional likelihood is obtained integrating over the distribution of η as follows,

$$\mathcal{L}(S, A; \mu, \sigma_v, \sigma_\eta\ ) = \int_{-\infty}^{+\infty}\mathcal{L}(S_i, A_i, \eta; \mu, \sigma_v\ )\frac{1}{\sigma_\eta}\phi\left(\frac{\eta}{\sigma_\eta}\right)d\eta$$

where ϕ is the standard normal density function. The integral is computed by using some efficient numerical algorithm. Finally, the likelihood of the sample is given by

$$\mathcal{L}(S, A; \mu, \sigma_v, \sigma_\eta\ ) = \prod_{i=1}^{N}\int_{-\infty}^{+\infty}\mathcal{L}(S_i, A_i, \eta; \mu, \sigma\ )\frac{1}{\sigma_\eta}\phi\left(\frac{\eta}{\sigma_\eta}\right)d\eta$$

This model can be estimated through the xttobit comand of STATA, the xttobit relies in the use of the Gauss-Hermite quadrature. By default it uses 12 points of evaluation, but we use 25 in order to perform our estimates.

5.3 Other econometric issues

There are some additional econometric issues not adressed in this paper. First of all, both the censored versions of the first-order autoregressive Markov-chain and the error-components model are based in this work on the assumption of normally distributed errors.[7] These assumptions could be too restrictive in the context of our work where the dependent variable is years of schooling. To relax the assumption is not so difficult in the first type of model but is is nontrivial in the latter. Horowitz and Markatou (1996) propose to non-parametrically estimate the error-component model using deconvolution techniques (Arellano, 2003). Moreover, Honoré (1992) propose the Trimmed least absolute deviations (LAD) estimator to deal with censored regression models with fixed effects in panel data. He proved that this estimator is consistent and asymptotically normal under suitable regularity conditions, and that it is not necessary to maintain parametric assumptions on the error terms to obtain this result. However, the focus in Honoré (1992) is to demonstrate that, based on accurate orthogonality conditions, it is possible to obtain consistent estimates for the coefficients of the model (Arellano and Honore, 2001).

Another issue is regarding the measurement errors. As it is well-known every time we obtain estimates using the variable years of schooling we must be concerned about the potential measurement error bias. The concern about the effects of measurement errors in the study of intergenerational educational mobility has been studied by Solon (1989, 1999). Let us make a simple reasoning about how measurement errors could affect the indexes estimated in this paper. Let us consider two leading cases: uncorrelated and correlated measurement errors. First, if measurement errors vary randomly across families and individuals, we know that the estimates of almost all the indexes used in this paper will be biased downwards, being the exception the DG Index where it is not possible to conclude the direction of the bias. Second, if measurement errors are correlated within families the indexes from the first-order autoregressive regression will be also biased downwards (except if the correlation among measurement errors within families is perfect). Finally, in the error-components model the direction and the magnitude of the bias will be determined by the relationship between the correlation of measurement errors within families and the total variance of measurement errors.

At last, there are two additional econometric issues that are not addresed in this paper. On one hand, the question of whether to use sample weights

[7] In the first-order autoregressive model u_{ij} and in the error-components model v_{ij} and η_i.

could have some interest in our context.[8] On the other, the concern about no linear relationships between parental and children education is not analyzed.

6 Empirical results

6.1 Definition of Indexes

In this paper we analyze the evidence about intergenerational educational mobility on Brazil, Chile, Uruguay and the USA over the period 1995-2006. Attending the previous discussion we consider three diferent indexes to measure relative mobility and two indexes to measure absolute mobility.

Indexes of relative mobility are based on the censored version of the linear first-order autoregressive Markov-chain, the DG Index and the censored version of the error-components model and are denominated Index 1, Index 2 and Index 3, respectively.

Index 1 of relative mobility is given by β_2 in the following specification,

$$S_{ij}^{*} = \alpha + \beta_2 S_{i1} + \delta_j + \delta_1 + u_{ij} \quad i = 1, 2.....N \quad j = 1, 2..F_i \tag{17}$$

$$S_{ij} = \begin{cases} S_{ij}^{*} & \text{if } A_{ij} = 0 \\ C_{ij} & \text{if } A_{ij} = 1 \end{cases} \tag{18}$$

S_{ij}, S_{i1} and δ_j, δ_1 are observed completed years of schooling and age-dummies of teenager j of family i and their parents respectively. In the case of teenagers we consider five dummies for ages 16, 17, 18, 19 and 20, for parents we also consider five categories (to avoid saturation), categories are given by the quantiles of the empirical distribution of mothers/fathers ages in each year-sample. A_{ij} is a dummy variable that equals 1 if the teenager is attending school, college or university. We assume $u_{ij} \mid S_{i1}, \delta \sim Niid(0, \sigma_u^2)$.

We also estimate a non-censored version of this model in order to compare it with the censored version. Index 1 is calculated using children-mothers and children-fathers samples separetely. In the case of the children-mothers sample we include all the children aged 16-20 which live with his or her mother (similar for children-fathers). The 95^{th} confidence interval of β_2 is obtained through the MLE estimation of its variance.

Index 2 of relative mobility is the Dahan-Gaviria Index and it is given by ρ_a,

$$\rho_a = 1 - (1 - \rho_g)\frac{B-1}{B-F} \tag{19}$$

[8] Actually, we made some comparison (using Uruguayan data) between the results obtained with and without considering sample weights for almost all index (except for the censored version of Index 3), and results and conclusions didn't changed.

where ρ_g measure the correlation among the educational attainments of teenager siblings still living with their parents, B is the number of teenagers in the selected sample and F the number of families, respectively. The 95 percent confidence interval of ρ_a is obtained through the 2.5^{th} and 97.5^{th} percentiles of the empirical distribution of ρ_a, where its estimations are obtained through the bootstrap method.[9]

Finally, Index 3 of mobility corresponds to the parameter ρ_2,

$$\rho_2 = \frac{\sigma_\eta^2}{\sigma_\eta^2 + \sigma_v^2} \tag{20}$$

in the following error-components model,

$$S_{ij}^* = \mu + \eta_i + \delta_{ij} + v_{ij} \quad i = 1, ...N, \quad j = 1, ...N_i \tag{21}$$

$$S_{ij} = \begin{cases} S_{ij}^* & \text{if } A_{ij} = 0 \\ C_{ij} & \text{if } A_{ij} = 1 \end{cases} \tag{22}$$

we assume

$$v_{ij} \sim Niid(0, \sigma_v^2)$$
$$\eta_i \sim Niid(0, \sigma_\eta^2)$$
$$v_{ij} \text{ and } \eta_i \text{ independent of each other}$$

i denotes each family and j each member of the family, S and A are as defined above. δ_{ij} are age-dummies for individual j of family i, in the case of teenagers we consider five dummies for ages 16, 17, 18, 19 and 20, while for parents we consider separately five categories for mothers and five categories for fathers given by the quantiles of the empirical distribution of mothers/fathers ages in each year-sample. We select all families with at least one teenager living with at least one parent, and then consider the N_i members of the family that satisfy the condition of being teenagers aged 16 to 20 or being parents. As done with Index 1 we also estimate the model ignoring censoring. The 95 percent confidence interval of ρ_2 relies on the assumption that σ_η^2 and σ_v^2 are jointly normally distributed, and it is computed using the MLE of the variance-covariance of σ_η and σ_v and applying the Delta method.

We also estimate two indexes of absolute intergenerational educational mobility, that correspond to Index 1 and Index 3. They are obtained by simply not including age-dummies into the respective specification.[10]

[9] See Section 3 for a detailed description about the Dahan-Gaviria Index.

[10] The Dahan-Gaviria Index cannot be applied to obtain a measure of absolute mobility due to it is intrinsically constructed to measure relative mobility.

6.2 Results

In this section we start by analyzing the results by countries, in term of the magnitude of the indexes and their evolution over time. Afterwards, we summarize and enrich across countries comparisons. Finally, we derive some conclusions about the relative performances of the indexes.

6.2.1 Brazil

The Brazilian data come from the "Pesquisa Nacional por Amostra de Domicílios" (PNAD) for every year in the period 1995-2006, with the exception of the year 2000 when the survey was not realised. The sample of the survey changes in 2004, therefore some observations needed to be removed from that year on, in order to achieve homogeneity in the within-country analysis. Our results are then representative for all the country except for the six rural areas that make up the former Northern Region. Information about school attainment and years of schooling are available for every year.

As illustrated in previous work (Guimarães and Veloso, 2003; PNUD, 2009) our descriptive statistics show that Brazil is experiencing a rapid expansion of education among children, and in particular teenagers (Table 1). The fraction of teenagers aged 16-20 that were attending education rose from 0.58 in 1996 to 0.66 in 2006, and the average of observed years of schooling (YOS) jumped from 6.5 to 9.4. This allows Brazilian teenagers of recent cohorts to overcome the huge gap that was present by the teenage-cohort of 1996 (on average YOS were 10.2, 9.3 and 10.9 in Chile, Uruguay and the USA, respectively). Taking into account that the YOS is top-censored for those individuals still attending education, and under the assumption that latent years of education is normally distributed, we also calculate the expectation of YOS: 9.4 in 1996 and 12.6 in 2006, this last figure is 1 year above the same statistic for Uruguay, and less than and approximately 1 year below Chile and the USA, respectively. We also calculate the standard deviation (SD) of YOS and its expectation and observe that while the simple SD decreased 0.39, and the SD considering censoring was reduced by 0.83, between 1996 and 2006. However, the observed reduction of the variance of YOS was not sufficient to eliminate the gap with respect to the other countries: by 2006 Brazil continued as the most unequal country in terms of the variability of its teenagers education.

We also found that the YOS of the mothers (and fathers) of the teenagers included in the samples increased during the period under analysis. However, the standard deviation did not exhibit subtantial changes (Table 2). Finally, the covariance between children and mothers has decreased 1.4 in the simple calculus and 3.5 taking into account censoring (1 and 2.9 respectively in the case of children-fathers).[11] However, in Table 4 we observe that covariances at the end of the period (e.g. 5.3 for children-mothers) are substantially greater than those of Chile (2.0), Uruguay (3.7) and the USA (0.5).

In Tables 5, 6, 7 and 11, and in Figures 1, 2 and 3 we present the estimates of relative mobility indexes. All of them show that Brazil was the most immobile country in terms of relative intergenerational educational mobility at the beginning of the period under analysis. However, Brazil exhibited constant and significant improvements in term of relative mobility within the period under analysis. As a consequence, Brazil had displaced Uruguay to the worst situation by the end of the period, and it also reduced substantially its gap with respect to Chile and the USA.

The previous conclusions hold for every index; however, there are some interesting features that emerge when comparing the estimates of different indexes of relative mobility. First, the differences of magnitude among indexes are notorious; furthermore, the Brazilian gaps with respect to the other countries under Index 3 are much more narrow than those obtained when comparison is based on indexes 1 and 2. Second, all indexes significantly decreased over time in the Brazilian case; however the average size of the reduction vary among indexes. The annual size of decline is 0.029 on average according to Index 1 using children-mothers samples (0.027 in the children-fathers); while it falls from this level to 0.005 and 0.009 when we use Index 2 and 3, respectively.

In addition, we compare the estimates of indexes 1 and 3 with respect to the non-censored versions of the corresponding models. We observe that Index 1 almost doubles (it is on average 1.8 times) the estimations of the non-censored version. This ratio remains relatively stable within the period. On the other hand, Index 3 is on average 1.18 times the estimates of the non-censored version of the error-components model; moreover, this relationship increases steadily over the period, at an annual magnitude of 0.0003 (the relation is 1.15 at the beginning and 1.21 at the end). Notice that these results the staments presented above. First, Index 1 is expected to be higher in comparison with the non-censored version of the model. Second, we expect that the difference between the censored and non-censored version of Index 1

[11] Notice that, the variables are expresed in term of years of schooling, thus the covariances can also be interpreted in years.

will be larger than those of Index 3. Finally, it is also interesting to point-out that when we control on age-dummies Index 3 is larger that its non-censored version and thus we can conclude that censoring affects more variance of η_i than that of v_{ij} in this Brazilian case.

Let us move to the analysis of absolute intergenerational educational mobility in Brazil (see Tables 8, 9, 10 and 12; Figures 4 and 5). It is important to recall that we only estimate two indexes of absolute mobility (indexes 1 and 3), and that the only difference with the indexes of relative mobility is that we do not include age-dummies in the corresponding models.

The differences with respect to the relative measure in the case of Index 1 are really subtle; in the period 1996-1999 Index 1 of absolute mobility is in the order of 0.98 times Index 1 of relative mobility, while in the period 2001-2006 this relationship is of 0.96. Moreover, the average annual size of reduction are on average almost identical. The results of Index 3 of absolute mobility; however, show huge difference in comparison with Index 3 of relative mobility: on average the former is 0.58 times the latter, the relationship is 0.7 at the beginning and 0.5 at the end. These results are interesting and intuitive because it is well-known that average years of schooling of Brazilian people were very low until almost the end of the 20th century, but Brazil has been experienced an amazing expansion of its educational system during the last decades (see Tables 1, 2 and 3). Therefore, we expect absolute mobility to be greater than relative mobility, like Index 3 captures.

Results of comparison with other countries is unaffected if we measure mobility in absolute rather then relative terms when we use Index 1. However, surprising results arise in the case of Index 3. First, at the beginning of the period the estimation of Index 3 of absolute mobility for Brazil is very similar than those of Uruguay and the USA (0.46, 0.44 and 0.45, respectively), but greater than that of Chile (0.33). Second, the annual size of reduction of this index doubles that of Index 3 of relative mobility (0.018 vs 0.009). Finally, at the end of the period the result for Brazil is 0.29, markedly below those of Uruguay (0.48) and the USA (0.46), and a bit greater than that of Chile (0.23).

The relationship between Index 1 of absolute mobility with respect to its non-censored version is on average 1.77, very similar than that of Index 1 of relative mobility, but it has decreased continuously over the period. On the other hand, an interesting result is that every year Index 3 is lower than the estimates of the non-censored version of the error-components model (the relationship is 0.9 at the beginning and 0.7 at the end), indicating that censoring affects more the variance of v_{ij} than that of η, and that the variance of v_{ij} has increased compared with that of η_i during the period. These results differ from those analyzed when age-dummies are included.

Summarizing, the Brazilian results show an amazing process of improvements in terms of both teenager educational attainments and intergenerational educational mobility. At the beginning of the period it was in the worst situation in all these terms, according to all statistics. However, by the end of the period it has displaced Uruguay to the worst situation in accordance with almost all indicators, with the exception of the variance of teenagers education. Furthermore, the most surprising result is that Index 3 indicates that at the end of the period Brazil was placed in the second place (behind Chile) in terms of absolute mobility.

6.2.2 Chile

We use the Chilean "Encuesta de Caracterización Socioeconómica" (CASEN), which is available only for some years 1996, 1998, 2000, 2003 and 2006 of the period we are considering. This survey is representative for almost all the country (the exceptions being only some remote and inaccesible areas). Information about school attainment and years of schooling are available for each of these years.

At the beginning of the period, the average YOS of Chilean teenagers were higher than that of Brazil and Uruguay and lower in comparison with the USA (Table 1). In Chile and Brazil the average YOS increased during the period, but this happened faster in Brazil, while in Uruguay and the USA it remained stable, therefore the gap with respect to Brazil has fallen, with respect to Uruguay it has risen, and in comparison with the USA it has vanished. The proportion of teenagers attending education also increased and at the end of the period it was 0.66, similar than that of Brazil, higher than that of Uruguay (0.60), but substantially lower than that of the USA (0.85). The SD of YOS decreased and at the end of the period was 2.0, this figure is lower than those of Uruguay (2.2) and Brazil (2.8), but it is higher than the 1.56 of the USA. If censoring is taken into account we find that expected YOS has increased by 0.6 while its SD has decreased by 1.2 over the period.

The average YOS of the parents included in the samples increased around 1 year during the period under analysis, while it SD decreased by 0.3 for mothers and by 0.4 for fathers (Table 2). Finally, the covariance between children and both parents, falling to less than half in all cases (see Table 4).

These covariances at the end of the period are markedly lower than those of Brasil and Uruguay, but higher than that of the USA.

In terms of relative mobility, Chile was in second place at the end of the period according to all the indexes only behind the USA (Tables 5, 6, 7 and 11; Figures 1, 2 and 3). Its position at the beginning depends on the index we use: indexes 1 and 2 place Chile in the second place, but Index 3 does it in third place behind the USA and Uruguay (although the difference with respect to Uruguay is negligible). On the other hand, all indexes significantly decreased over time in the Chilean case: the average annual size of decline is 0.024 in Index 1 using children-mothers samples (0.022 in the children-fathers); and 0.013 and 0.007 when we use Index 2 and 3, respectively.

We compare the estimates of indexes 1 and 3 with those of the non-censored versions of the corresponding models. This shows that Index 1 more than doubles (it is on average 2.3 times) the estimation of the non-censored model, and the magnitude of the difference remains relatively stable over the period. On the other hand, Index 3 is on average 1.3 times the estimates of the non-censored version of the error-components. These results are similar than those obtained and analyzed above in the Brazilian case.

In terms of absolute intergenerational educational mobility Chile is the most mobile of the four countries according to Index 3. this conclusion holds in every year when Chilean data is available. Index 2, however, yields different results, at the beginning of the period Chile is in the third place behind the USA and Uruguay, and by the end of the period in the second place behind the USA.

The comparison between absolute and relative measures are very similar than in the Brazilian case in almost all features, but Index 1 of absolute mobility is approximately 0.92 times Index 1 of relative mobility, while in the case of Index 3 this ratio is 0.53 at the beginning and 0.42 at the end. Moreover, the linear trends of absolute mobility are very similar than those of relative mobility.

The ratios of censored versus non-censored versions of absolute mobility are on average 2.2 and 0.8 for indexes 1 and 3 respectively. Therefore, as in Brazil Index 3 is lower when censoring is considered and this ratio has fallen over the period.

The analyzed evidence about Chile indicates that this country was in a relative good position in terms of teenagers education and intergenerational mobility at the beginning, and in addition it continued improving its performance on these features along the period. The average education of Chilean teenagers were similar than that for the USA at the end of the period, which is particularly remarkable since the USA configures a high standard of comparison. Even more, Chile is in the second place in terms of relative mobility but Index 3 indicates that at the end of the period it was the country with the highest absolute mobility

6.2.3 Uruguay

The data for Uruguay come from the "Encuesta Continua de Hogares" (ECH) for every year of the period 1995-2006. To obtain homogeneous samples, we consider only observations from urban areas of more than 5,000 inhabitants. We also take into account the changes in the way information is collected and harmonize the measures of the variable YOS along the period.

At the beginning of the period average YOS of Uruguayan teenagers was almost 3 years greater than those of the Brazilian, but 1 and 1.6 years lower than those of the Chilean and the teenagers of the USA, respectively (Table 1). Within the cohort of parents, the Uruguayan average YOS of mothers was the highest within the included Latin American countries: attained 8.5 in comparison with 4.8 for Brazil, 7.6 for Chile and 12.6 for the USA, while for fathers these averages were 7.9, 4.8, 7.9 and 12.9, respectively (Table 2). At the end of the period the Uruguayan averages of the parents cohort exhibited a moderate increase, but that of the teenagers remained unchanged, the latter diverges from the expansion of education among teenagers of Brazil and Chile, pointed out above. On the other hand, although the proportion of teenagers attending education went up from 0.54 to 0.60 between the extremes of the period, it was comparatively very low in the whole period. As a consequence, the Uruguayan teenagers were the poor educated by the end of the period. This conclusion is reinforced if we take into account censoring: we observe that expected years of education is 11.4 in comparison with 12.6 of Brazil, 13.4 of Chile and 13.8 of the USA. On the other hand, the SD of teenagers YOS is lower than that of Brazil, and higher than those of Chile and the USA, considering and not considering censoring.

The covariances between children and parents were almost identical at the beginning than at the end; but they exhibited a hump-shaped behavior within the period (Table 4). In comparison with Brazil covariances are low, but they are high next to those of Chile and very high compared to those of the USA.

According to indexes 1 and 3, Uruguay had the second place behind the USA in terms of relative mobility at the beginning of the period, but it was in third place in accordance with Index 2. Nevertheless, its bad performance determined that by the end of the period it was displaced to the worst position according to all indexes (Tables 5, 6, 7 and 11; Figures 1, 2 and 3). In addition, the linear trends of the indexes are not significant, except for Index 1 in the children-mothers samples, where we find a significant increase at an average annual size of 0.001. However, in the Uruguayan case, a second-order polynomial better fits observed time trends of indexes 1 and 3, with significant positive and negative coefficients associated with trend and trend square, respectively. Finally, the time trend is not significantly different from 0 in the case of Index 2 in both specifications.

The ratio between Index 1 and its non-censored version is 2.4 on average and presents a hump-shaped pattern over the period. In the case of Index 3 this ratio is on average 1.3 but is somewhat U-shaped. These numbers are very similar to those of Chile.

Essentially, the same facts are indicated by the indexes of absolute mobility. However, in particular Index 1 exhibit a deep deterioration during the period. In the linear specification, the trend coefficients are both around 0.01, and are significant at the 0.01 and 0.06 levels, in the children-mothers and children-fathers samples respectively; while in the second-order specification the coefficient of the linear trend is 0.05 and that of the quadratic term -0.004, and they are all significant at the 0.05 levels, in both samples.

The relationship between censored and non-censored version of Index 1 of absolute mobility is similar than for Brasil and Chile, although the ratio is a bit greater (2.4). However, in the case of Index 3 the estimates considering the censoring are greater than those obtained not considering it, which differs from results found for Chile and Brazil. Nevertheless, the ratio is small: it is 1.06 on average. That feature indicates that censoring affects a bit more the variance of η_i than that of v_{ij} in the Uruguayan case.

At last, we can conclude that Uruguay exhibited a very bad performance in terms of their teenagers schooling and intergenerational educational mobility, in comparison with the other three countries, during the period under analysis. Although its situation was not so unfavourable at the beginning of the period, all the statistics and estimates show that its situation got worse in almost all analyzed dimensions. Evidence also shows that this is the result of a deterioration in the first part of the period and a moderate improvement in the later years of the analysis. It is pertinent to add that Uruguayan public education budget has been increased substantially since 2005 (47 percent in real terms between 2004 and 2008); however no subtantial change was introduced in the educational system, except for the "Plan Ceibal", which provides one laptop free of charge (with costless connectivity) per child and teacher in public schools. This plan began to be implemented by year 2007.[12] Therefore, we expect that Uruguay will improve in the analyzed features, at least in the long run.

6.2.4 The United States of America

The "Current Population Survey", which has information for every year in our period (IPUMS-CPS[13]) is used for the USA. This is a monthly survey of about 50,000 households and is representative of all U.S. non-institutional civilian. Information about school attainment and attendance to school are available for every year. However, the questionaries gather YOS when it is either one to four or five to eight, for these cases we have used the imputation suggested in Jaeger (1997).[14]

It is well-known that the average YOS is very high for people of the USA, moreover the SD of YOS is very low, in particular in comparison with the Latin American countries. The USA attained its high standard during the eighties, and since then YOS continued increasing slowly, alternating periods of augment and stagnation; from 1995 it has increased but at a very small rate (Tables 1, 2 and 3). On the other hand, the SD of YOS decreased for both cohorts of teenagers and parents along the period. Besides, we find that expected YOS obtained for teenagers by considering censoring seems reliable because it is only somewhat above than observed YOS of parents (13.8 against 13.2 and 13.3 for mothers and fathers respectively); however, the SD of expected YOS is very well below those of their parents (2.1 against 2.8 and 2.9 respectively).

We find interesting results when analyzing covariances between children and parents. First, the simple covariance between teenagers and mothers was only 0.8 at the beginning (0.9 for fathers), but it also decreased and attained a value very close to 0.5 (for both mothers and fathers) at the end. Second, censoring affects these covariances much more than it does for Brazil, Chile and Uruguay, suggesting that parental education influences strongly the probability of teenagers attendance to education.

The USA exhibit the greater degree of relative intergenerational educa-

[12] Taken from the project One Laptop per Child (OLPC) of the scientist Nicholas Negroponte.

[13] King et al (2004)

[14] The proportion of teenagers in these categories are 5 and 4 percent in 1996 and 2006 respectively, while these figures for the population aged 30-55 are 7 and 5.4.

tional mobility in every year according to all the indexes. However, the magnitude of the gap depends on the index: indexes 1 and 2 show wider differences than Index 3 (Tables 5, 6, 7, 10, Figures 1, 2, 3). The conclusion about time trends depends also on the index: Index 1 within the children-mother samples and Index 3 shows a significant decrease (at annual sizes of 0.004 and 0.003, respectively), while time trend is not significant in the cases of Index 2 and Index 1 (within the children-father samples). The ratio between indexes 1 and 3 and their non-censored versions attained at the end 3.5 for the Index 1 children-mothers (3.9 for children-fathers) and 1.6 for Index 3, these figures are substantially higher than those of Brazil, Chile and Uruguay. This is not surprising given the results analized above about the relationship between simple covariances and their censored counterparts.

Furthermore, when comparing censored and non-censored versions of relative mobility indexes we find that for Index 1 the ratio oscillate in the range of 2.8 and 3.9, becoming stable since the year 2001 around the value of 3.8. The ratio in the case of Index 3 is much more stable and it is around 1.55 over the whole period. All these figures are larger than those obtained for Brazil, Chile and Uruguay, in particular in the case of Index 1.

In terms of absolute mobility, conclusions are only subtly different if we consider Index 1; except for the fact that time trends are not significant either in children-mother or children-fathers samples (Tables 8 and 10, Figure 4). We also find that the sizes of Index 3 of relative and absolute mobility are very similar. However, when considering Index 3 some unexpected features appear. First, at the beginning of the period absolute mobility in the USA (0.45) is below that of Chile (0.33) and similar than that of Uruguay (0.44) and Brazil (0.46). Second, Brazil and Chile were much more mobiles than the USA and Uruguay by the end of the period, according to Index 3 (0.22, 0.29, 0.46 and 0.48, respectively). Finally, time trend of absolute mobility is not significant in accordance with this index. (Tables 9 and 10, Figure 5).

The ratio between the estimates of censored and non-censored versions of indexes 1 and 3 of absolute mobility exhibit an increasing pattern over the period: it is 2.9 at the beginning and 3.7 at the end for Index 1, and 2 and 3.2 respectively for Index 3. Therefore, we can conclude that in the USA censoring affects much more the variance of η_i than that of v_{ij}. It is interesting to recall that in the case of Index 3 the ratio is lower than 1 for Brazil and Chile, and only a bit greater than 1 for Uruguay.

Summarizing, the USA reached a very high worldwide standard in terms of its population educational attainments by the eighties, and since then kept this situation and also improved it moderately. In this study, the USA appears as the most mobile country in terms of relative intergenerational educational mobility. We also find that in the USA, the degree of relative mobility is similar than that of absolute mobility in accordance with both indexes 1 and 3, which might be expected because of the stability of the average YOS of teenagers and parents over the period. Finally, in terms of absolute mobility the USA is also in first place according to Index 1, but it is in third place in accordance with Index 3.

6.2.5 Summary of across countries comparison

We find that the USA is by large the country with most relative intergenerational educational mobility, while Chile is in the second place; it does not matter which index we use. However, Brazil and Uruguay are alternatively in the third or the fourth place depending on the index we use (Table 10).

We also find robust evidence that relative intergenerational educational mobility has increased over the period under analysis in the cases of Brazil and Chile, while different indexes indicate different time trends in the USA with some indexes showing a very slow increases of mobility and others stagnation. Finally, in the case of Uruguay almost all indexes indicate stagnation (Table 11).

In terms of absolute intergenerational mobility the USA also occupies the first place except when using the censored version of the variance decomposition index, where is displaced to the third position behind Chile and Brazil at the end of the period. Uruguay is in fourth position in all cases, except once when the index of Uruguay is almost identical to that of Brazil (Table 10). All indexes indicates that absolute mobility has significantly increased in Brazil and Chile, but it remained stable in the USA and it decreased in Uruguay (Table 12). We also find the magnitud of Index 3 of absolute mobility being very similar to that of relative mobility in the case of the USA, contrary to results for Brazil and Chile where the values of Index 3 of relative mobility are well below those of absolute mobility.

On the other hand, the third place of the USA in terms of absolute mobility is surprising, but it can be rationalized by the fact that this country exhibit a very high YOS standard at the beginning of the period which increased very slowly from then on; as a consequence we expect absolute mobility to be similar than relative mobility, as both indexes show.

Descriptive statistics can also help to explain the fact that Index 3 positions Chile and Brasil as the most mobile countries in absolute terms. In particular, in the case of Brazil because its educative system experimented an "actual revolution": the average YOS of Brazilian teenagers increased as much as 3 years along only one decade. Furthermore, the good Chilean performance in terms of absolute mobility is the result of combining a good starting point and non-negligible improvements over the period under analysis.

Some interesting conclusions arise from the comparison between censored and non-censored versions of indexes 1 and 3. Let us firstly consider Index 1 of relative mobility, in this case the lowest ratio is found for Brazil where it is on average 1.8, while for Chile it is 2.3, for Uruguay it is 2.4 and for the USA equals 3.3. In the case of Index 3 of relative mobility these figures are 1.2, 1.3, 1.3 and 1.6. Therefore, in the case of relative indexes the estimates are pushed up by censoring. This is as expected in the case of Index 1, because we know that the probability of attending education is positively affected by parents schooling. However, the ratios are informative because they show that the effect of the censoring is substantially greater for the USA. On the other hand, we expect that both the numerator and the denominator of Index 3 are augmented by top-censoring, and thus no hypothesis about the final result can be inferred from intuition. Thus, results indicating that the censored version is greater than the non-censored one is useful, and show that the censoring affects more the variance of the error that is invariant within families, than that of the error that randomly varies across families and individuals. Finally, these figures for Index 1 of absolute mobility are 1.8, 2.2, 2.4 and 3.4, allowing us to derive similar conclusions than when analyzing Index 1 of relative mobility. Nevertheless, the comparison between censored and non-censored models in the case of Index 3 adds interesting results, the ratio is 0.8, 0.8, 1.06 and 2.5 for Brazil, Chile, Uruguay and the USA respectively, and thus shows that the censoring affects less the variance of the error that is invariant within families than that of the error that randomly varies across families and individuals in Brazil and Chile, but the opposite holds for the USA, while the effects are balanced in the case of Uruguay.

6.2.6 Comparison of indexes empirical performance

Evidence we obtain allows us to gain some insights about the comparative performance of the indexes of intergenerational educational mobility estimated in this paper.

First of all, the SD of the indexes varies notoriously among them (Tables

5 to 9, Figures 1 to 5). The SD of Index 3 is low in absolute magnitudes, and also when compared with indexes 1 and 2. If we focus on indexes of relative mobility, the ratio between SD of Index 3 and that of Index 2 is on average 0.3 in all countries, while if we compare SD of indexes 3 and 1 the ratio is around 0.5 for Brazil, Chile and Uruguay and it is 0.67 for the USA (Tables 5 to 7). These figures are 0.64, 0.71, 0.55 and 0.70 respectively, when comparing the SD of indexes 3 and 1 of absolute mobility (Tables 8 and 9).

In addition, Index 3 is smoother than the others, which seems to be a positive attribute for an index of intergenerational mobility, since this is a structural characteristic of countries, and therefore we should expect that it changes very slowly over time (Figures 1 to 5).

Furthermore, the conclusions that arise from the estimates of Index 3 are in accordance to the stylized facts we extract from descriptive statistics and, although some results were unexpected (e.g. the first place of Brazil in terms of absolute mobility), all of them can be rationalized in the context of our knowledge about recent performance of the educational systems of the countries under analysis.

Index 1 failures to capture the expected differences between relative and absolute mobility in the cases of Brazil and Chile, and for the relationship between censored and non-censored it cannot account for the trade-off between the effect of censoring on covariances between children and parents against the effect on the variance of children education.

Index 2 is unable to take into account some evidence about changes of fertility behavior during the period: e.g. in Table 3 we observe that the proportion of households with at least two teenagers has fallen from 0.073 to 0.041 in Brazil, and increased from 0.027 to 0.034 in the USA. Moreover, we find that the proportion of households (with a woman as head or wife) with at least one teenager is more than four times that with two or more teenagers, in all countries and almost all years; therefore, we lost a lot of information when the study of mobility is based only on correlation among siblings.[15]

In Table 1 we find that the average YOS as well as its SD and the proportion of teenagers attending education do not change dramatically when we restrict the sample to those households with at least two children; however in some cases resulting ratios vary over the period. In addition, we observe that the simple covariance between children and mothers education is greater in

[15] The exception is Brazil where this ratio is 3.3 at the beginning and 4.6 at the end of the period.

the sample with at least two teenagers per household than in the sample with one or more teenagers per household. The opposite is true when we analyze the calculus of covariances considering the censoring. Moreover, both ratios vary over time in all countries. Finally, we believe that although Index 2 is able to capture the decreasing time trend in the cases of Chile and Brasil, the width of its confidence interval makes us doubt about the reliability of time trends results, at least when the period is short and the size of the changes relatively small.[16]

It is important to notice that except for the analysis about the SD of the indexes, the other conclusions are conjectural and can be controversial. This is because we actually do not know for certain the real processes which generated the data. An interesting extension of this paper would be to use Monte Carlo experiments in order to compare the performance of the indexes under different types of controlled stylized facts.

7 Concluding remarks

In this paper we study the evolution of intergenerational educational mobility for Brazil, Chile, Uruguay and the USA over the period 1995-2006.

We propose a flexible econometric approach given by the variance decomposition in an error-components model (Arellano, 2003). We consider a model where families are the units of analysis and each member constitutes a particular observation. We distinguish between two types of error-components (unobservables): one that captures characteristics or attributes which are shared for all members of the family, and another that varies randomly across individuals and families. Within this framework we calculate an index of intergenerational mobility, given by the fraction of the variance of these two error-components that can be asociated to the error that capture common factors within families. This approach can be thought as one simplification of the model estimated by Lillard and Willis (1994).

We estimate three different indexes of intergenerational mobility and compare their results. Index 1, as in the traditional approach, is based on a first-order autoregressive Markov-chain that relates children outcomes with parental outcomes, Index 2 is the Dahan-Gaviria Index, and Index 3 is the new index proposed in this paper. The three indexes are used to measure relative mobility, while we also estimate absolute mobility through indexes 1

[16]Notice that the width of the confidence interval of Index 2 is on average 0.053 for Brazil and 0.087 for Chile.

and 3.

We use data from standard household surveys, and follow Dahan-Gaviria's idea of relying the analysis on data about teenagers, although we also use information on parents education. Many teenagers are still attending education, and thus we do not observe the definitive educational attainment of all teenagers; as a consequence a non-negligible portion of the observations of teenagers schooling is top-censored. We address this issue by estimating indexes 1 and 3 considering the censoring (under the assumption that errors are normally distributed). On the other hand, Index 2 considers censoring by defining a dummy variable that indicates if the teenager's years of schooling is above the median of his or her cohort.

We show that indexes 1 and 2 can be thought as particular applications of the error-components model. In addition we argue that Index 3 has some advantages in comparison with them, from both theoretical and empirical points of view.

Results show that education has been expanding quickly in Brazil and it reached huge improvements in term of teenagers schooling and intergenerational mobility over the period. In the case of Chile, at the beginning its situation was already good compared with the other Latin American countries, and in addition it also improved over the period under analysis. On the other hand, the performance of Uruguay was really bad and at the end of the period was the worst one in almost all dimensions. Finally, as previous evidence demonstrate teenagers are very well educated in the USA and in this country the degree of intergenerational mobility is high; this situation is found at the beginning of the period and remains stable (with some moderate improvements) along the period. These conclusions hold for the results of all indexes.

However, there are some other features that depend on the index we use. The most salient differences arise when comparing indexes of absolute mobility. First, values of Index 1 of relative mobility are similar than those of Index 1 of absolute mobility, but we find substantial differences between relative and absolute measures of mobility when we use Index 3, in the cases of Brazil and Chile. On the other hand, the censored version of Index 1 is greater that the estimates without considering the censoring in all cases, as might be expected. However, censoring affects in different manners results of Index 3: estimates are lower than when censoring is not addresed for Brazil and Chile, but for the USA is higher and for Uruguay both estimates are similar.

At the beginning the ranking of relative mobility in descending order was: the USA, Uruguay, Chile and Brasil according to indexes 1 and 3, and the USA, Chile, Uruguay and Brasil in accordance with Index 2. At the end of the period the ranking was: the USA, Chile, Brazil and Uruguay in accordance with all indexes, except for Index 3 which indicates that relative mobility is similar in Brazil and Uruguay.

In terms of absolute mobility the ranking was: the USA, Uruguay, Chile and Brazil at the beginning and the USA, Chile, Brazil and Uruguay at the end according to Index 1. Surprising results arise from estimates of Index 3: at the beginning Chile was in the first place and the estimates of the index for Brazil, Uruguay and the USA were similar; while at the end of the period the ranking was: Chile, Brazil, the USA and Uruguay. It seems unexpected that the USA appeared less mobile than Brazil and Chile; however this feature can be explained by some stylized facts. The main explanation is that along the period the USA educational system was essentially in steady state; while it experienced a huge expansion in Brazil and a significant improvement in Chile, at least in terms of teenagers attendance to education and their average years of schooling.

Results indicates that the efforts that Chilean and Brazilian governments have done during the past decade are being successful. However, they show that Uruguay needs to introduce substantial changes in its educational system in order to avoid the risk of impoverishing the human capital of the new cohorts next to those of other countries in the region.

There are some extensions of this paper in our long run research agenda. The most interesting for us are the following. First, to relax the assumption that error-components are normally distributed (Horowitz and Markaton 1996, Arellano and Honoré, 2001; Honoré, 1992). Second, to perform Monte Carlo experiments to compare the performance of the indexes. Third, to enrich the structure of error-components (Lillard and Willis, 1994) and to include observable characteristics with the goal of studying the determinants of intergenerational mobility. Finally, to analyze which of those determinants can be considered causal effects using the IV approach (Arellano and Bover, 1995). Furthermore, there are other interesting extensions like to address the issue of heteroskedasticity, the use of sample weights, to obtain the indexes for other countries or for larger periods of time.[17]

[17] Notice that Index 3 can be also applied to Census data, and thus to perform long run analysis for some countries where these type of data is available since the beginning of the 20^{th} Century.

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Table 1: Descriptive statistics on teenagers (aged 16-20) education. Selected years

	No. obs.	Proportion attending education (1)	Average YOS (2)	Expected YOS (t-cen) (3)	Std. Dev. YOS (4)	Std. Dev. Exp. YOS (t-cen) (5)
			All teeenagers			
Brazil						
1996	52478	0.584	6.48	9.41	3.16	4.97
2001	28701	0.690	8.76	12.95	3.17	5.17
2006	28321	0.659	9.44	12.65	2.77	4.14
Chile						
1996	18784	0.602	10.18	12.78	2.57	4.39
2000	17643	0.602	10.31	12.74	2.44	4.00
2006	14994	0.664	10.97	13.38	2.01	3.19
Uruguay						
1996	8786	0.544	9.27	11.23	2.48	3.80
2000	3710	0.597	9.45	11.98	2.59	4.39
2006	12013	0.603	9.24	11.36	2.24	3.73
USA						
1996	14080	0.813	10.86	13.73	1.64	2.29
2000	7151	0.817	10.87	13.70	1.61	2.21
2006	12799	0.852	10.84	13.84	1.56	2.07
			Households with at least two siblings			
Brazil						
1996	25798	0.574	6.24	9.08	3.18	4.94
2001	12445	0.689	8.53	12.85	3.25	5.36
2006	10478	0.648	9.16	12.35	2.84	4.27
Chile						
1996	7650	0.564	9.94	12.29	2.70	4.38
2000	7074	0.573	10.14	12.38	2.50	4.00
2006	5284	0.646	10.92	13.18	2.01	3.13
Uruguay						
1996	3418	0.506	9.11	10.88	2.62	3.84
2000	1403	0.561	9.25	11.52	2.65	4.40
2006	4116	0.577	9.07	11.05	2.31	3.77
USA						
1996	4970	0.803	10.89	13.88	1.74	2.50
2000	2556	0.805	10.91	13.71	1.68	2.28
2006	4733	0.839	10.94	13.94	1.64	2.18

(1) Column 1 is the proportion of teenagers aged 16-20 in the sample that is attending school, college or university.
(2) Column 2 is the simple average of completed years of schooling of the teenagers in the sample.
(3) Column 3 is the expected years of schooling of teenagers taking into account top-censoring, normal distribution assumed.
(4) Column 4 is the simple standard deviation of completed years of schooling of the teenagers in the sample.
(5) Column 5 is the standard deviation of expected years of schooling of teenagers taking into account top-censoring, normal distribution assumed.

Table 2: Descriptive statistics on parents education. Selected years

	No. obs.	Average Age	Age-Percentil 5	Age-Percentil 95	Average number of teenagers at home	Average YOS	Std Dev YOS
				Mothers			
Brazil							
1996	37400	45.2	35	58	1.369	4.80	4.25
2001	21355	44.7	35	57	1.306	6.76	4.47
2006	22127	44.4	35	57	1.245	7.53	4.54
Chile							
1996	14216	44.3	35	57	1.276	7.62	4.23
2000	13411	44.6	35	58	1.271	7.56	4.05
2006	12176	44.8	36	56	1.226	8.66	3.86
Uruguay							
1996	6720	46.1	36	58	1.258	8.45	3.90
2000	2835	45.9	36	58	1.256	8.92	3.92
2006	9495	45.6	35	57	1.221	8.99	3.69
USA							
1996	10968	43.6	35	53	1.229	12.61	2.97
2000	5577	44.0	35	53	1.235	12.68	3.00
2006	9800	44.9	35	55	1.246	13.21	2.77
				Fathers			
Brazil							
1996	30622	48.9	37	65	1.376	4.81	4.43
2001	16934	48.2	37	64	1.322	6.49	4.54
2006	17418	47.9	36	64	1.253	7.08	4.57
Chile							
1996	12458	47.6	36	64	1.284	7.93	4.42
2000	11890	48.2	37	64	1.277	7.66	4.24
2006	12034	48.4	38	63	1.226	8.63	4.03
Uruguay							
1996	5764	49.5	38	64	1.255	7.95	3.88
2000	2408	49.0	38	64	1.259	8.60	4.01
2006	7520	48.8	37	63	1.223	8.69	3.73
USA							
1996	9008	46.4	36	58	1.234	12.93	3.30
2000	4625	46.4	36	58	1.236	13.04	3.16
2006	8108	47.3	37	59	1.249	13.35	2.95

(1) These statistics corresponds to head and spouse in households with at least one child aged 16-20.

Table 3: Statistics about the Heads or Spouses of Households

	No. obs.	Average YOS 1 (1)	Average YOS 2 (2)	Average YOS 3 (3)	Proportion 1 (4)	Proportion 2 (5)	Proportion 3 (6)
Women Aged 32-60							
Brazil							
1996	37400	5.68	4.90	4.56	0.762	0.238	0.073
2001	21355	7.27	6.86	6.46	0.793	0.207	0.054
2006	22127	8.03	7.71	6.90	0.811	0.189	0.041
Chile							
1996	14216	7.76	7.71	7.34	0.767	0.233	0.057
2000	13411	7.49	7.65	7.30	0.769	0.231	0.056
2006	12176	7.75	8.68	8.60	0.815	0.185	0.038
Uruguay							
1996	6740	7.88	8.40	8.63	0.805	0.195	0.045
2000	2841	8.51	8.90	8.95	0.829	0.171	0.038
2006	9495	8.66	8.99	9.00	0.818	0.182	0.036
USA							
1996	10968	12.54	12.71	12.25	0.870	0.130	0.027
2000	5577	12.80	12.72	12.55	0.870	0.130	0.028
2006	9800	13.15	13.24	13.09	0.848	0.152	0.034
Men Aged 35-65							
Brazil							
1996	30622	5.72	4.94	4.51	0.772	0.228	0.071
2001	16934	7.07	6.63	6.10	0.798	0.202	0.055
2006	17418	7.69	7.23	6.57	0.816	0.184	0.041
Chile							
1996	12458	8.13	8.01	7.70	0.770	0.230	0.058
2000	11890	7.68	7.73	7.43	0.773	0.227	0.056
2006	12034	7.93	8.63	8.61	0.791	0.209	0.043
Uruguay							
1996	5786	7.81	7.86	8.26	0.794	0.206	0.047
2000	2419	8.39	8.59	8.58	0.819	0.181	0.040
2006	7520	8.45	8.66	8.80	0.816	0.184	0.037
USA							
1996	9008	12.80	13.00	12.66	0.874	0.126	0.026
2000	4625	13.00	13.03	13.07	0.876	0.124	0.027
2006	8108	13.22	13.38	13.25	0.855	0.145	0.032

(1) Columns 1, 2 and 3 correspond to the years of schooling of head or spouse in households with no child, one child and more than one child respectively.
(2) Columns 4, 5 and 6 are sample proportions of head or spouse in households with no child, one child and more than one child respectively.
(3) We restrict the sample to women aged 32-62 and men aged 35-65 taking into account the 2.5th and 97.5th percentiles of the observed distribution of mothers/fathers ages.

Table 4: Covariances between YOS of Teeenagers and Parents

	Mothers				***Fathers***			
	All teenagers		Teenagers with at least one sibling		All teenagers		Teenagers with at least one sibling	
	Simple (1)	Censored (2)	Simple (1)	Censored (2)	Simple (1)	Censored (2)	Simple (1)	Censored (2)
Brazil								
1996	6.69	12.30	6.99	12.29	6.97	13.31	7.24	13.10
2001	6.10	10.91	6.74	11.62	6.23	11.39	6.76	12.10
2006	5.31	8.80	5.98	9.41	5.21	8.64	5.70	9.05
Chile								
1996	4.40	9.88	5.09	10.41	4.41	10.27	5.16	11.00
2000	3.55	7.06	3.78	7.12	3.57	7.25	3.74	7.15
2006	2.02	4.72	2.17	4.62	2.05	4.90	2.30	5.07
Uruguay								
1996	3.83	8.46	4.65	10.11	3.48	8.26	4.27	10.20
2000	4.34	10.80	5.14	12.75	4.56	10.96	5.59	13.07
2006	3.69	8.41	4.34	9.61	3.26	7.90	3.92	9.18
USA								
1996	0.80	2.59	1.08	3.08	0.92	2.85	1.41	3.72
2000	0.75	2.14	1.04	2.63	0.73	2.36	1.05	2.99
2006	0.48	1.78	0.62	1.99	0.51	2.05	0.74	2.43

(1) Is the simple the simple covariance between years of schooling of children and their parents.
(2) Is the estimation of the covariance between children and parents years of education taking into account top-censoring on the YOS of children, normal distribution is assumed.

Table 5: Index 1 of Relative Mobility. Children-Mothers and Children-Fathers (1)

a. Children-Mothers

	Censored				Non-censored			
	Brasil	Chile	Uruguay	USA	Brasil	Chile	Uruguay	USA
1995	0.740	--	0.510	0.264	0.399	--	0.236	0.089
	[0.009]	--	[0.017]	[0.013]	[0.004]	--	[0.009]	[0.004]
1996	0.697	0.587	0.550	0.262	0.378	0.258	0.253	0.082
	[0.009]	[0.013]	[0.018]	[0.013]	[0.004]	[0.006]	[0.009]	[0.005]
1997	0.706	--	0.591	0.261	0.381	--	0.279	0.093
	[0.009]	--	[0.020]	[0.013]	[0.004]	--	[0.010]	[0.005]
1998	0.676	0.548	0.584	0.261	0.361	0.240	0.278	0.081
	[0.009]	[0.011]	[0.020]	[0.014]	[0.004]	[0.005]	[0.010]	[0.005]
1999	0.636	--	0.663	0.255	0.344	--	0.274	0.075
	[0.009]	--	[0.023]	[0.014]	[0.004]	--	[0.009]	[0.005]
2000	--	0.468	0.711	0.220	--	0.226	0.275	0.076
	--	[0.009]	[0.024]	[0.013]	--	[0.004]	[0.010]	[0.005]
2001	0.568	--	0.648	0.211	0.303	--	0.268	0.066
	[0.009]	--	[0.021]	[0.013]	[0.004]	--	[0.008]	[0.005]
2002	0.546	--	0.694	0.239	0.299	--	0.262	0.070
	[0.008]	--	[0.026]	[0.010]	[0.004]	--	[0.009]	[0.004]
2003	0.512	0.437	0.677	0.260	0.279	0.192	0.235	0.073
	[0.008]	[0.009]	[0.028]	[0.011]	[0.003]	[0.004]	[0.009]	[0.004]
2004	0.484	--	0.665	0.246	0.273	--	0.235	0.071
	[0.008]	--	[0.027]	[0.011]	[0.003]	--	[0.009]	[0.004]
2005	0.455	--	0.660	0.235	0.258	--	0.239	0.060
	[0.007]	--	[0.026]	[0.011]	[0.003]	--	[0.009]	[0.004]
2006	0.441	0.338	0.605	0.207	0.256	0.143	0.262	0.059
	[0.007]	[0.009]	[0.012]	[0.011]	[0.003]	[0.004]	[0.005]	[0.003]

b. Children-Fathers

	Censored				Non-censored			
	Brasil	Chile	Uruguay	USA	Brasil	Chile	Uruguay	USA
1995	0.727	--	0.528	0.254	0.383	--	0.250	0.077
	[0.010]	--	[0.019]	[0.013]	[0.004]	--	[0.009]	[0.004]
1996	0.690	0.553	0.548	0.250	0.358	0.232	0.230	0.075
	[0.010]	[0.014]	[0.021]	[0.013]	[0.004]	[0.006]	[0.010]	[0.004]
1997	0.680	--	0.607	0.239	0.355	--	0.274	0.079
	[0.010]	--	[0.023]	[0.013]	[0.004]	--	[0.011]	[0.004]
1998	0.648	0.524	0.580	0.218	0.341	0.218	0.262	0.067
	[0.010]	[0.012]	[0.023]	[0.013]	[0.004]	[0.005]	[0.011]	[0.004]
1999	0.622	--	0.682	0.231	0.330	--	0.282	0.060
	[0.010]	--	[0.026]	[0.014]	[0.004]	--	[0.010]	[0.005]
2000	--	0.442	0.698	0.227	--	0.207	0.276	0.066
	--	[0.010]	[0.027]	[0.014]	--	[0.004]	[0.011]	[0.005]
2001	0.572	--	0.626	0.188	0.294	--	0.245	0.051
	[0.010]	--	[0.025]	[0.014]	[0.004]	--	[0.009]	[0.005]
2002	0.537	--	0.680	0.240	0.294	--	0.249	0.061
	[0.010]	--	[0.029]	[0.011]	[0.004]	--	[0.009]	[0.003]
2003	0.511	0.416	0.644	0.239	0.272	0.174	0.231	0.064
	[0.009]	[0.010]	[0.031]	[0.011]	[0.004]	[0.004]	[0.010]	[0.003]
2004	0.482	--	0.686	0.245	0.266	--	0.231	0.065
	[0.009]	--	[0.032]	[0.011]	[0.004]	--	[0.010]	[0.003]
2005	0.449	--	0.637	0.249	0.253	--	0.212	0.063
	[0.008]	--	[0.031]	[0.012]	[0.004]	--	[0.010]	[0.004]
2006	0.423	0.327	0.560	0.220	0.244	0.133	0.226	0.056
	[0.008]	[0.009]	[0.014]	[0.011]	[0.004]	[0.004]	[0.006]	[0.003]

(1) The Index 1 of Relative Mobility is obtained as the coefficient of the years of schooling of the parent in a first-order autoregressive Markov-chain regression where the dependent variable is the years of schooling of the child and dummies for the age of children and parents are included. In the case of teenagers we consider five dummies for ages 16 to 20, for parents we consider five categories given by the quantiles of the empirical distribution of mother/fathers ages in each year-sample.

(2) The estimation of the index in the censored version is based on the assumption that errors are normally distributed.

(3) Standard deviation in brackets.

Table 6: Index 2 of Relative Mobility: Dahan-Gaviria Index (1)

	Brasil	Chile	Uruguay	USA
1995	0.533	--	0.476	0.140
	[0.011]	--	[0.036]	[0.043]
1996	0.524	0.337	0.444	0.172
	[0.011]	[0.035]	[0.031]	[0.028]
1997	0.531	--	0.479	0.116
	[0.012]	--	[0.034]	[0.033]
1998	0.526	0.297	0.453	0.106
	[0.011]	[0.018]	[0.034]	[0.036]
1999	0.523	--	0.519	0.173
	[0.010]	--	[0.033]	[0.029]
2000	--	0.274	0.543	0.160
	--	[0.017]	[0.035]	[0.027]
2001	0.476	--	0.506	0.111
	[0.011]	--	[0.038]	[0.034]
2002	0.493	--	0.494	0.175
	[0.011]	--	[0.039]	[0.020]
2003	0.505	0.226	0.491	0.133
	[0.011]	[0.019]	[0.037]	[0.032]
2004	0.488	--	0.474	0.139
	[0.012]	--	[0.042]	[0.023]
2005	0.491	--	0.517	0.155
	[0.013]	--	[0.037]	[0.037]
2006	0.466	0.206	0.487	0.133
	[0.034]	[0.020]	[0.022]	[0.021]

(1) The Index 2 of Relative Mobility (the Dahan-Gaviria Index) captures the correlation between the educational attainments of teenager siblings that are living in the same household. The educational attainment is measure through a binary variable that equals one if children years of schooling is greater that the median of his age-specific group.

(2) Standard deviation in brackets. Standard deviation is given by the range between 97.5th and 2.5th percentiles of the empirical distribution of the index using the bootstrap method (1000 reps.) divided by four.

Table 7: Index 3 of Relative Mobility. Variance decomposition of error-components (1)

	Censored				Non-censored			
	Brasil	Chile	Uruguay	USA	Brasil	Chile	Uruguay	USA
1995	0.659	--	0.587	0.500	0.575	--	0.450	0.322
	[0.004]	--	[0.010]	[0.009]	[0.004]	--	[0.010]	[0.008]
1996	0.641	0.612	0.576	0.507	0.554	0.479	0.428	0.326
	[0.004]	[0.006]	[0.010]	[0.009]	[0.004]	[0.007]	[0.010]	[0.008]
1997	0.655	--	0.572	0.506	0.564	--	0.454	0.338
	[0.004]	--	[0.010]	[0.009]	[0.004]	--	[0.010]	[0.008]
1998	0.643	0.591	0.580	0.514	0.554	0.461	0.452	0.332
	[0.004]	[0.006]	[0.011]	[0.009]	[0.004]	[0.006]	[0.011]	[0.008]
1999	0.633	--	0.609	0.499	0.544	--	0.475	0.322
	[0.004]	--	[0.010]	[0.009]	[0.004]	--	[0.011]	[0.008]
2000	--	0.557	0.625	0.490	--	0.449	0.480	0.318
	--	[0.005]	[0.010]	[0.009]	--	[0.005]	[0.011]	[0.008]
2001	0.583	--	0.615	0.479	0.490	--	0.465	0.305
	[0.004]	--	[0.010]	[0.010]	[0.004]	--	[0.011]	[0.009]
2002	0.585	--	0.608	0.483	0.494	--	0.462	0.304
	[0.004]	--	[0.011]	[0.007]	[0.004]	--	[0.011]	[0.006]
2003	0.584	0.561	0.591	0.485	0.487	0.425	0.439	0.302
	[0.004]	[0.005]	[0.011]	[0.007]	[0.004]	[0.005]	[0.012]	[0.006]
2004	0.581	--	0.596	0.485	0.483	--	0.443	0.308
	[0.004]	--	[0.011]	[0.007]	[0.004]	--	[0.011]	[0.006]
2005	0.567	--	0.597	0.479	0.468	--	0.433	0.299
	[0.004]	--	[0.011]	[0.007]	[0.004]	--	[0.012]	[0.006]
2006	0.569	0.537	0.594	0.474	0.469	0.398	0.451	0.292
	[0.004]	[0.006]	[0.006]	[0.007]	[0.004]	[0.005]	[0.006]	[0.006]

(1) The Index 3 of Relative Mobility is obtained as the fraction of the variance of the years of schooling within a sample of teenagers, mothers and fathers living together that can be interpreted as "common factors" within each family. Index 3 is estimated through ML in a model like the "Random Effects" model commonly used in panel data analysis (Arellano, 2003). The difference is that here we consider that a given family constitutes the "unit", and include one observation for each member of the family (i.e. we have not data over time but across family members). Thus, we consider two types of errors (unobservables) one that is invariant within each family and another one that varies randomly across families and individuals. Dummies for the age of children and parents are included. In the case of teenagers we consider five dummies for ages 16 to 20, for parents we consider five categories for mothers and five for fathers given by the quantiles of the empirical distribution of ages in each year-sample. The point-estimation in the censored version relies on the assumption that both errors are normally distributed and are independent of each other, and is obtained by the xttobit STATA routine, which uses the Gauss-Hermite quadrature (25 points are used to perform the estimation).

(2) Standard deviation in brackets. Standard deviation of the index is obtained by applying the Delta method and the MLE of the variance-covariance of errors, and relies on the assumption that both errors are jointly normally distributed.

Table 8: Index 1 of Absolute Mobility. Children-Mothers and Children-Fathers

a. Children-Mothers

	Censored				Non-censored			
	Brasil	Chile	Uruguay	USA	Brasil	Chile	Uruguay	USA
1995	0.731	--	0.494	0.265	0.397	--	0.231	0.091
	[0.009]	--	[0.017]	[0.012]	[0.004]	--	[0.009]	[0.006]
1996	0.688	0.547	0.542	0.272	0.374	0.244	0.246	0.084
	[0.009]	[0.012]	[0.017]	[0.013]	[0.004]	[0.006]	[0.009]	[0.006]
1997	0.692	--	0.581	0.267	0.377	--	0.276	0.096
	[0.009]	--	[0.019]	[0.013]	[0.004]	--	[0.010]	[0.007]
1998	0.661	0.511	0.579	0.264	0.358	0.230	0.279	0.093
	[0.008]	[0.010]	[0.019]	[0.013]	[0.004]	[0.005]	[0.010]	[0.007]
1999	0.619	--	0.647	0.262	0.345	--	0.278	0.078
	[0.008]	--	[0.022]	[0.014]	[0.004]	--	[0.009]	[0.007]
2000	--	0.433	0.694	0.227	--	0.218	0.279	0.079
	[1.000]	[0.009]	[0.023]	[0.012]	--	[0.004]	[0.010]	[0.006]
2001	0.544	--	0.639	0.221	0.304	--	0.275	0.061
	[0.008]	--	[0.021]	[0.013]	[0.004]	--	[0.008]	[0.007]
2002	0.528	--	0.683	0.250	0.302	--	0.267	0.065
	[0.008]	--	[0.025]	[0.010]	[0.004]	--	[0.009]	[0.005]
2003	0.491	0.402	0.671	0.270	0.278	0.185	0.235	0.071
	[0.008]	[0.008]	[0.027]	[0.010]	[0.004]	[0.004]	[0.009]	[0.005]
2004	0.466	--	0.666	0.256	0.272	--	0.242	0.071
	[0.007]	--	[0.026]	[0.011]	[0.003]	--	[0.009]	[0.005]
2005	0.438	--	0.657	0.246	0.260	--	0.249	0.063
	[0.007]	--	[0.025]	[0.011]	[0.003]	--	[0.009]	[0.005]
2006	0.425	0.312	0.612	0.225	0.256	0.133	0.268	0.061
	[0.006]	[0.008]	[0.012]	[0.010]	[0.003]	[0.004]	[0.005]	[0.005]

b. Children-Fathers

	Brasil	Chile	Uruguay	USA	Brasil	Chile	Uruguay	USA
1995	0.717	--	0.503	0.249	0.383	--	0.243	0.079
	[0.010]	--	[0.018]	[0.012]	[0.004]	--	[0.009]	[0.006]
1996	0.683	0.521	0.528	0.250	0.358	0.223	0.222	0.081
	[0.010]	[0.013]	[0.020]	[0.013]	[0.004]	[0.006]	[0.010]	[0.006]
1997	0.668	--	0.584	0.246	0.355	--	0.267	0.089
	[0.009]	--	[0.022]	[0.013]	[0.004]	--	[0.011]	[0.006]
1998	0.630	0.485	0.567	0.225	0.343	0.211	0.263	0.075
	[0.009]	[0.011]	[0.022]	[0.012]	[0.004]	[0.005]	[0.011]	[0.007]
1999	0.609	--	0.656	0.236	0.335	--	0.280	0.061
	[0.009]	--	[0.025]	[0.014]	[0.004]	--	[0.010]	[0.007]
2000	--	0.404	0.678	0.228	--	0.199	0.282	0.070
	[1.00]	[0.009]	[0.025]	[0.013]	--	[0.004]	[0.011]	[0.007]
2001	0.550	--	0.611	0.200	0.301	--	0.249	0.053
	[0.010]	--	[0.024]	[0.013]	[0.004]	--	[0.009]	[0.007]
2002	0.517	--	0.662	0.247	0.300	--	0.248	0.062
	[0.009]	--	[0.027]	[0.010]	[0.004]	--	[0.009]	[0.005]
2003	0.492	0.376	0.633	0.245	0.278	0.170	0.230	0.065
	[0.009]	[0.009]	[0.030]	[0.010]	[0.004]	[0.004]	[0.010]	[0.005]
2004	0.466	--	0.676	0.249	0.268	--	0.231	0.066
	[0.008]	--	[0.031]	[0.011]	[0.004]	--	[0.010]	[0.005]
2005	0.431	--	0.646	0.252	0.258	--	0.224	0.062
	[0.008]	--	[0.030]	[0.011]	[0.004]	--	[0.010]	[0.005]
2006	0.411	0.298	0.558	0.228	0.248	0.125	0.230	0.057
	[0.007]	[0.008]	[0.014]	[0.011]	[0.004]	[0.004]	[0.006]	[0.005]

(1) The Index 1 of Absolute Mobility is obtained as the coefficient of the years of schooling of the parents in a first-order autoregressive Markov-chain regression where the dependent variable is the years of schooling of the children and no other regressor is included. The point estimation of the index in the censored version is based on the assumption that the errors are normally distributed.

(2) Standard deviation in brackets.

Table 9: Index 3 of Absolute Mobility. Variance decomposition of error-components (1)

	Censored				Non-censored			
	Brasil	Chile	Uruguay	USA	Brasil	Chile	Uruguay	USA
1995	0.461	--	0.440	0.450	0.514	--	0.422	0.227
	[0.005]	--	[0.011]	[0.009]	[0.004]	--	[0.010]	[0.008]
1996	0.442	0.326	0.424	0.459	0.499	0.379	0.401	0.236
	[0.005]	[0.008]	[0.011]	[0.009]	[0.004]	[0.007]	[0.010]	[0.008]
1997	0.439	--	0.428	0.470	0.506	--	0.424	0.241
	[0.005]	--	[0.011]	[0.009]	[0.004]	--	[0.011]	[0.008]
1998	0.416	0.315	0.457	0.471	0.498	0.374	0.437	0.234
	[0.005]	[0.007]	[0.012]	[0.009]	[0.004]	[0.006]	[0.011]	[0.009]
1999	0.377	--	0.474	0.468	0.480	--	0.462	0.209
	[0.005]	--	[0.012]	[0.009]	[0.004]	--	[0.011]	[0.008]
2000	--	0.252	0.487	0.457	--	0.332	0.470	0.219
	--	[0.006]	[0.012]	[0.009]	--	[0.005]	[0.011]	[0.008]
2001	0.301	--	0.501	0.451	0.422	--	0.462	0.199
	[0.005]	--	[0.012]	[0.010]	[0.004]	--	[0.011]	[0.008]
2002	0.303	--	0.489	0.474	0.424	--	0.462	0.149
	[0.005]	--	[0.012]	[0.007]	[0.004]	--	[0.011]	[0.006]
2003	0.290	0.247	0.471	0.474	0.410	0.331	0.436	0.154
	[0.005]	[0.006]	[0.013]	[0.007]	[0.004]	[0.005]	[0.011]	[0.006]
2004	0.296	--	0.479	0.477	0.407	--	0.442	0.157
	[0.005]	--	[0.012]	[0.007]	[0.004]	--	[0.011]	[0.006]
2005	0.286	--	0.483	0.467	0.395	--	0.435	0.155
	[0.005]	--	[0.012]	[0.007]	[0.004]	--	[0.011]	[0.006]
2006	0.290	0.225	0.484	0.462	0.396	0.305	0.444	0.142
	[0.005]	[0.006]	[0.007]	[0.007]	[0.004]	[0.006]	[0.006]	[0.006]

(1) The Index 3 of absolute Mobility is obtained as the fraction of the variance of the years of schooling within a sample of teenagers, mothers and fathers living together that can be interpreted as "common factors" within each family. Index 3 is estimated through ML in a model like the "Random Effects" model commonly used in panel data analysis (Arellano, 2003). The difference is that here we consider that a given family constitutes the "unit", and include one observation for each member of the family (i.e. we have not data over time but across family members). Thus, we consider two types of errors (unobservables) one that is invariant within each family and another one that varies randomly across families and individuals. Only a common intercept is included in the model. The point-estimation in the censored version relies on the assumption that both errors are normally distributed and are independent of each other, and is obtained by the xttobit STATA routine, which uses the Gauss-Hermite quadrature (25 points are used to perform the estimation).

(2) Standard deviation in brackets. Standard deviation of the index is obtained by applying the Delta method and the MLE of the variance-covariance of errors, and relies on the assumption that both errors are jointly normally distributed.

Table 10: Summary of results at the beginning and at the end of the period

	Brazil	Chile (1)	Uruguay	USA
a. Indexes of Relative Mobility				
Index 1 Children-Mothers (Markov chain, Censoring considered)				
1995	0.740	0.587	0.510	0.264
2006	0.441	0.338	0.605	0.207
Index 1 Children-Fathers (Markov chain, Censoring considered)				
1995	0.727	0.553	0.528	0.254
2006	0.423	0.327	0.560	0.220
Index 2 (DG-Index)				
1995	0.533	0.337	0.476	0.140
2006	0.466	0.206	0.487	0.133
Index 3 (Variance decomposition, Censoring considered)				
1995	0.659	0.612	0.587	0.500
2006	0.569	0.537	0.594	0.474
b. Indexes of Absolute Mobility				
Index 1 Children-Mothers (Markov chain, Censoring considered)				
1995	0.731	0.547	0.494	0.265
2006	0.425	0.312	0.612	0.225
Index 1 Children-Fathers (Markov chain, Censoring considered)				
1995	0.717	0.521	0.503	0.249
2006	0.411	0.298	0.558	0.228
Index 3 (Variance decomposition, Censoring considered)				
1995	0.461	0.326	0.440	0.450
2006	0.290	0.225	0.484	0.462

(1) In the case of Chile the year 1995 is not available, therefore the results actually correspond to 1996.

Table 11: Linear trend of Indexes of Relative Mobility over the period 1995-2006

	Brazil	Chile	Uruguay	USA
Index 1 Children-Mothers (Markov chain, Censoring considered)				
Trend	**-0.029****	**-0.024****	**0.011***	**-0.004***
	[-0.031 - -0.026]	[-0.031 - -0.017]	[0.001 - 0.020]	[-0.007 - -0.000]
Constant	0.745**	0.611**	0.571**	0.263**
	[0.731 - 0.760]	[0.563 - 0.659]	[0.510 - 0.632]	[0.242 - 0.284]
Index 1 Children-Mothers (Markov chain, Censoring not considered)				
Trend	**-0.014****	**-0.011****	**-0.001**	**-0.003****
	[-0.015 - -0.013]	[-0.015 - -0.008]	[-0.005 - 0.002]	[-0.004 - -0.002]
Constant	0.398**	0.275**	0.266**	0.089**
	[0.371 - 0.384]	[0.228 - 0.269]	[0.244 - 0.290]	[0.067 - 0.082]
Index 1 Children-Fathers (Markov chain, Censoring considered)				
Trend	**-0.027****	**-0.022****	**0.007**	**-0.001**
	[-0.029 - -0.026]	[-0.030 - -0.015]	[-0.004 - 0.017]	[-0.004 - 0.003]
Constant	0.729**	0.577**	0.586**	0.238**
	[0.722 - 0.736]	[0.526 - 0.628]	[0.518 - 0.653]	[0.215 - 0.261]
Index 1 Children-Fathers (Markov chain, Censoring not considered)				
Trend	**-0.013****	**-0.010****	**-0.004**	**-0.002***
	[-0.013 - -0.012]	[-0.013 - -0.007]	[-0.007 - 0.000]	[-0.003 - -0.000]
Constant	0.378**	0.248**	0.267**	0.074**
	[0.371 - 0.384]	[0.228 - 0.269]	[0.244 - 0.290]	[0.067 - 0.082]
Index 2 (DG-Index)				
Trend	**-0.005****	**-0.013****	**0.003**	**0.000**
	[-0.008 - -0.003]	[-0.017 - -0.009]	[-0.002 - 0.008]	[-0.005 - 0.005]
Constant	0.535**	0.341**	0.475**	0.142**
	[0.520 - 0.550]	[0.314 - 0.369]	[0.441 - 0.508]	[0.111 - 0.173]
Index 3 (Variance decomposition, Censoring considered)				
Trend	**-0.009****	**-0.007***	**0.002**	**-0.003****
	[-0.011 - -0.007]	[-0.012 - -0.002]	[-0.001 - 0.005]	[-0.004 - -0.002]
Constant	0.660**	0.610**	0.587**	0.509**
	[0.646 - 0.675]	[0.578 - 0.643]	[0.568 - 0.606]	[0.500 - 0.517]
Index 3 (Variance decomposition, Censoring not considered)				
Trend	**-0.011****	**-0.008****	**0.000**	**-0.003****
	[-0.013 - -0.008]	[-0.009 - -0.007]	[-0.004 - 0.003]	[-0.005 - -0.002]
Constant	0.575**	0.487**	0.455**	0.333**
	[0.561 - 0.589]	[0.482 - 0.493]	[0.435 - 0.475]	[0.325 - 0.342]

(1) Trends are obtained by regressing the corresponding indexes in a linear trend over time (time trend value is 0 for 1995). Data are available for every year in the period 1995-2006 in the cases of Uruguay and the USA, from Brazil 2000 is absent and from Chile there are only data for years 1996, 1998, 2000, 2003 and 2006.

(2) Confidence intervals at the 5 percent level of significance in brackets.

(3) * significant at 5%; ** significant at 1%

Table 12: Linear trend of Indexes of Absolute Mobility over the period 1995-2006

	Brazil	Chile	Uruguay	USA
Index 1 Children-Mothers (Markov chain, Censoring considered)				
Trend	**-0.029****	**-0.023****	**0.012***	**-0.003**
	[-0.032 - -0.027]	[-0.030 - -0.016]	[0.003 - 0.021]	[-0.006 - 0.000]
Constant	0.733**	0.570**	0.556**	0.267**
	[0.718 - 0.748]	[0.525 - 0.614]	[0.500 - 0.611]	[0.247 - 0.287]
Index 1 Children-Mothers (Markov chain, Censoring not considered)				
Trend	**-0.014****	**-0.011****	**0.000**	**-0.003****
	[-0.015 - -0.012]	[-0.015 - -0.006]	[-0.004 - 0.003]	[-0.004 - -0.002]
Constant	0.395**	0.263**	0.262**	0.093**
	[0.386 - 0.404]	[0.234 - 0.291]	[0.238 - 0.285]	[0.085 - 0.101]
Index 1 Children-Fathers (Markov chain, Censoring considered)				
Trend	**-0.028****	**-0.022****	**0.009**	**0.000**
	[-0.029 - -0.027]	[-0.029 - -0.015]	[-0.001 - 0.019]	[-0.003 - 0.003]
Constant	0.717**	0.540**	0.561**	0.240**
	[0.712 - 0.723]	[0.493 - 0.587]	[0.497 - 0.626]	[0.220 - 0.259]
Index 1 Children-Fathers (Markov chain, Censoring not considered)				
Trend	**-0.012****	**-0.010****	**-0.002**	**-0.002****
	[-0.013 - -0.011]	[-0.013 - -0.006]	[-0.006 - 0.001]	[-0.004 - -0.001]
Constant	0.378**	0.240**	0.260**	0.081**
	[0.372 - 0.384]	[0.217 - 0.263]	[0.236 - 0.285]	[0.072 - 0.090]
Index 3 (Variance decomposition, Censoring considered)				
Trend	**-0.018****	**-0.010***	**0.005****	**0.001**
	[-0.023 - -0.014]	[-0.018 - -0.003]	[0.002 - 0.008]	[-0.001 - 0.003]
Constant	0.455**	0.332**	0.439**	0.459**
	[0.425 - 0.485]	[0.282 - 0.381]	[0.418 - 0.460]	[0.449 - 0.470]
Index 3 (Variance decomposition, Censoring not considered)				
Trend	**-0.012****	**-0.008***	**0.002**	**-0.010****
	[-0.015 - -0.010]	[-0.012 - -0.003]	[-0.001 - 0.006]	[-0.013 - -0.007]
Constant	0.519**	0.386**	0.429**	0.248**
	[0.504 - 0.535]	[0.356 - 0.417]	[0.406 - 0.452]	[0.229 - 0.267]

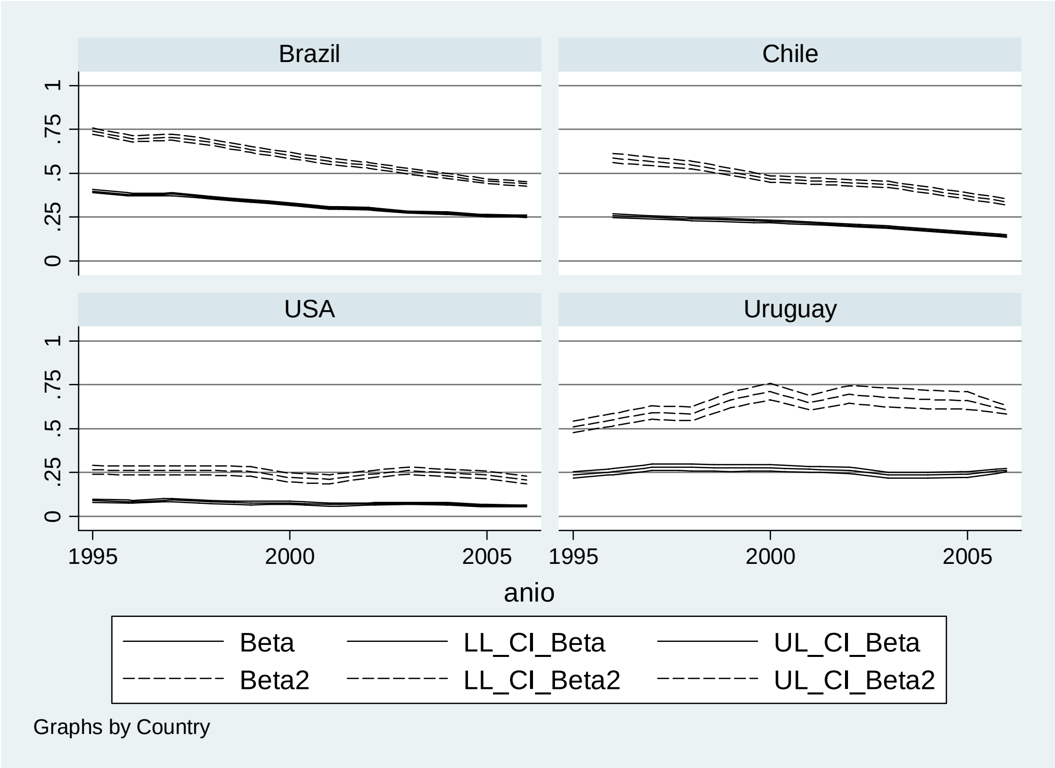
(1) Trends are obtained by regressing the corresponding indexes in a linear trend over time (time trend value is 0 for 1995). Data are available for every year in the period 1995-2006 in the cases of Uruguay and the USA, from Brazil 2000 is absent and from Chile there are only data for years 1996, 1998, 2000, 2003 and 2006.

(2) Confidence intervals at the 5 percent level of significance in brackets.

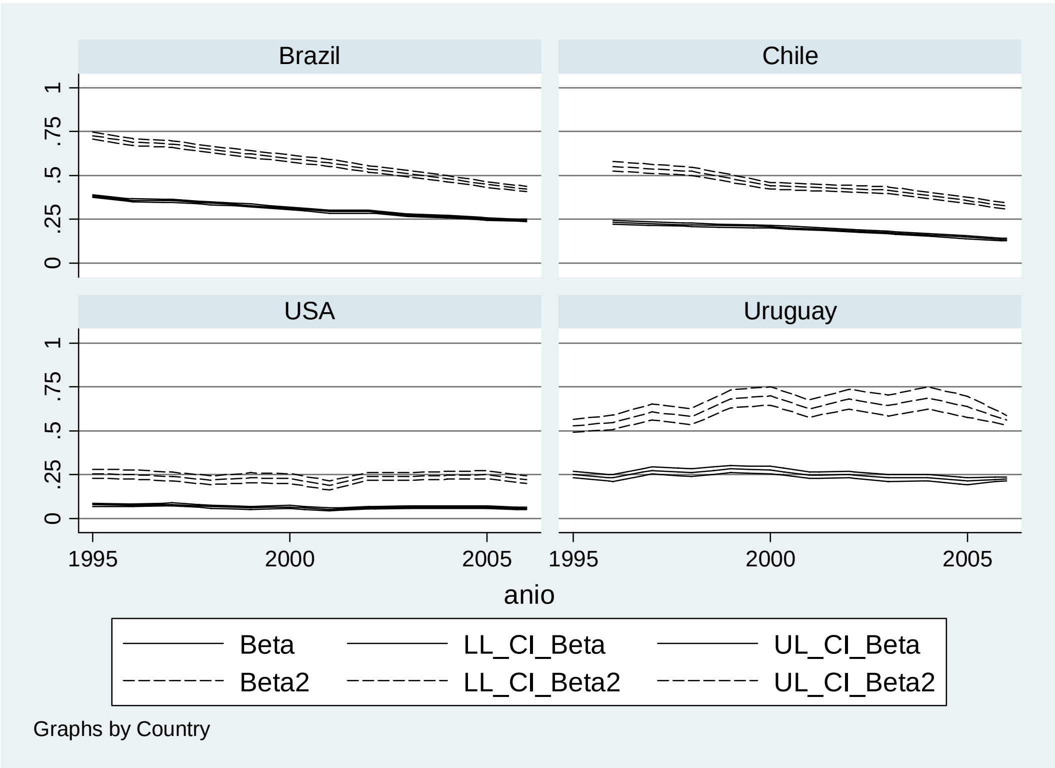
(3) * significant at 5%; ** significant at 1%

Figure 1: Index 1 of Relative Mobility. Children-Mothers and Children-Fathers

a. Children-Mothers



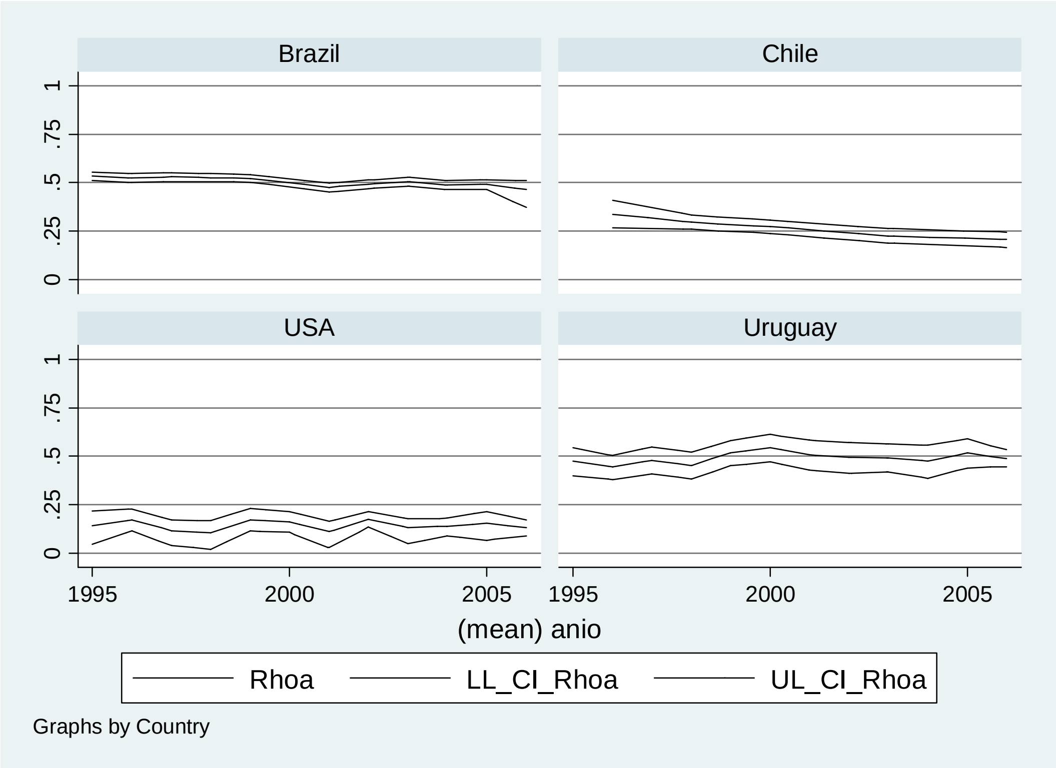
b. Children-Fathers



(1) Beta is the Index when censoring is not considered; Beta2 is the Index when censoring is considered.

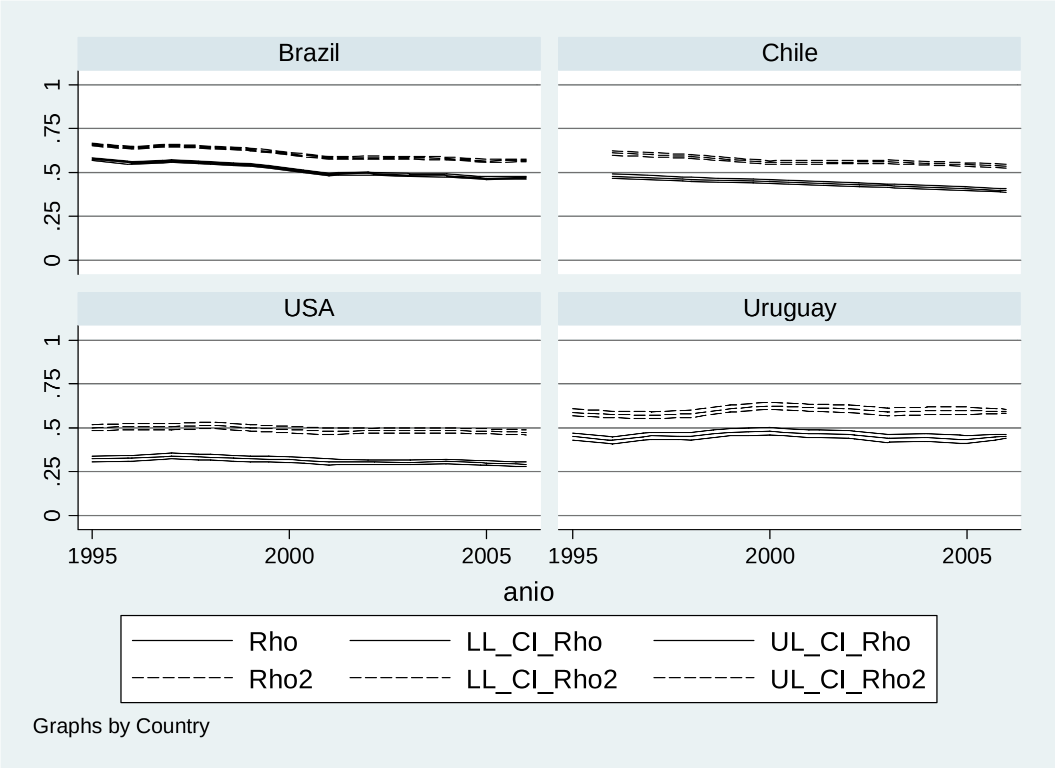
(2) LL_CI and UL_CI are the lower and upper limits of the 95[th] confidence interval, obtained under normal distribution.

Figure 2: Index 2 of Relative Mobility. Dahan-Gaviria Index



(1) LL_CI and UL_CI are the 2.5th and 97.5th percentiles of the empirical distribution of Rhoa obtained through the bootstrap method using 1000 repositions.

Figure 3: Index 3 of Relative Mobility. Variance decomposition of error-components

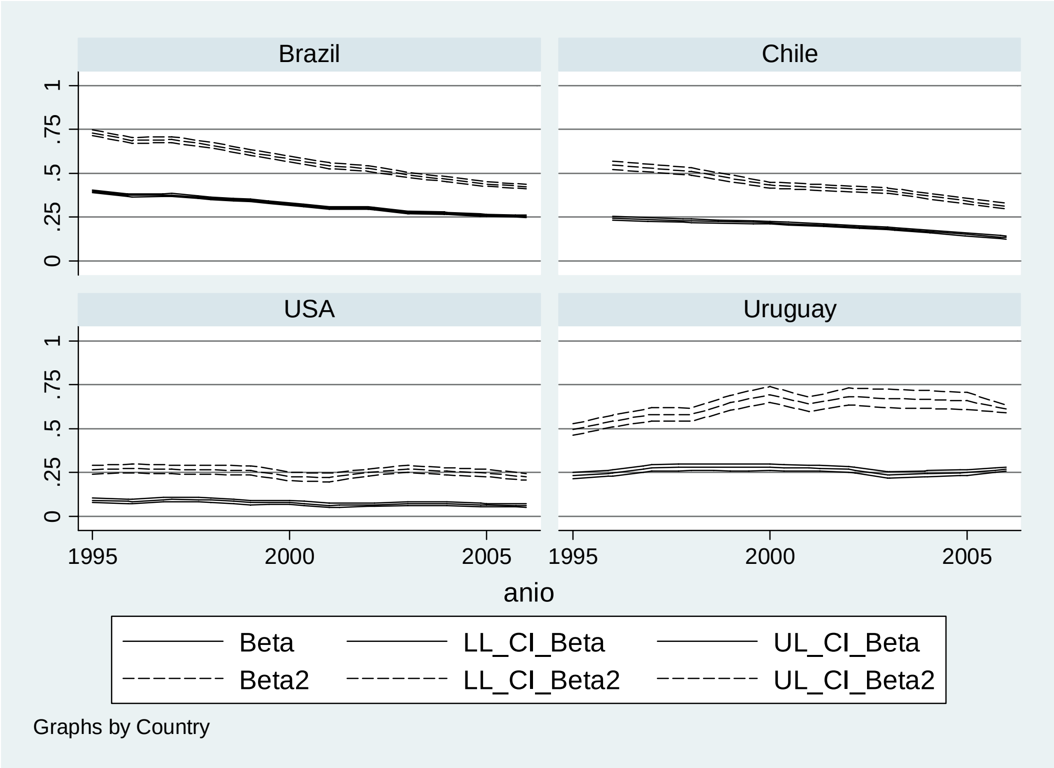


(1) Rho is the Index when censoring is not considered; Rho2 is the Index when censoring is considered.

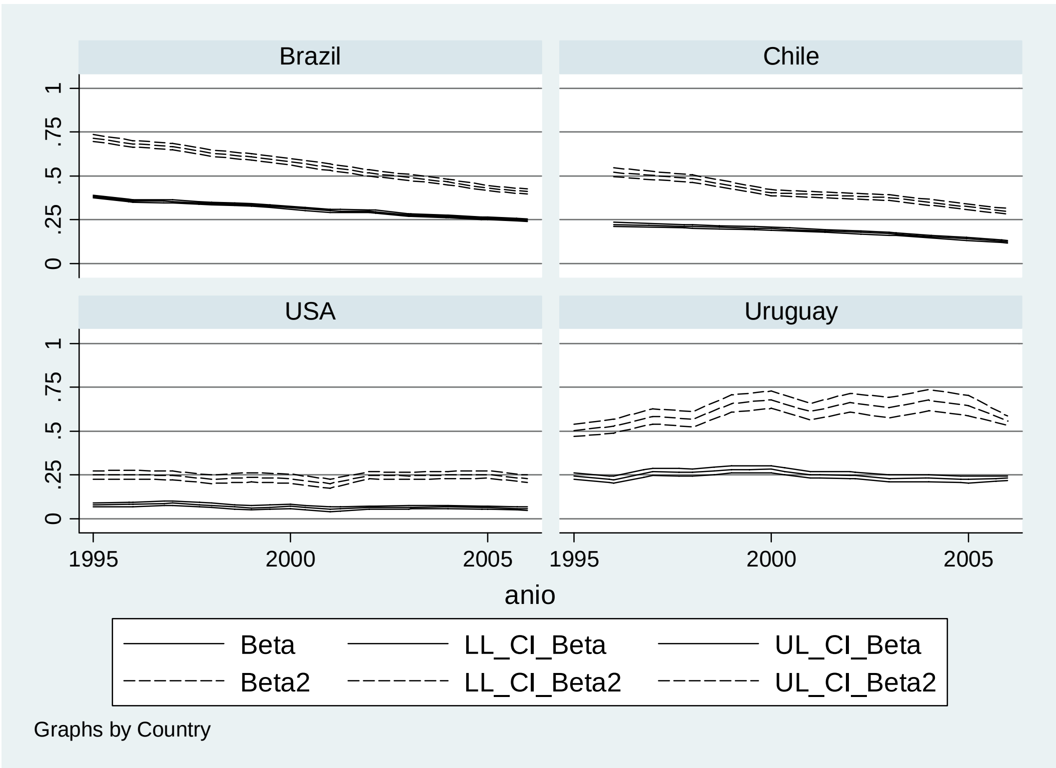
(2) LL_CI and UL_CI are the lower and upper limits of the 95^{th} confidence interval, obtained under jointly normally distributed error-components.

Figure 4: Index 1 of Absolute Mobility. Children-Mothers and Children-Fathers

a. Children-Mothers



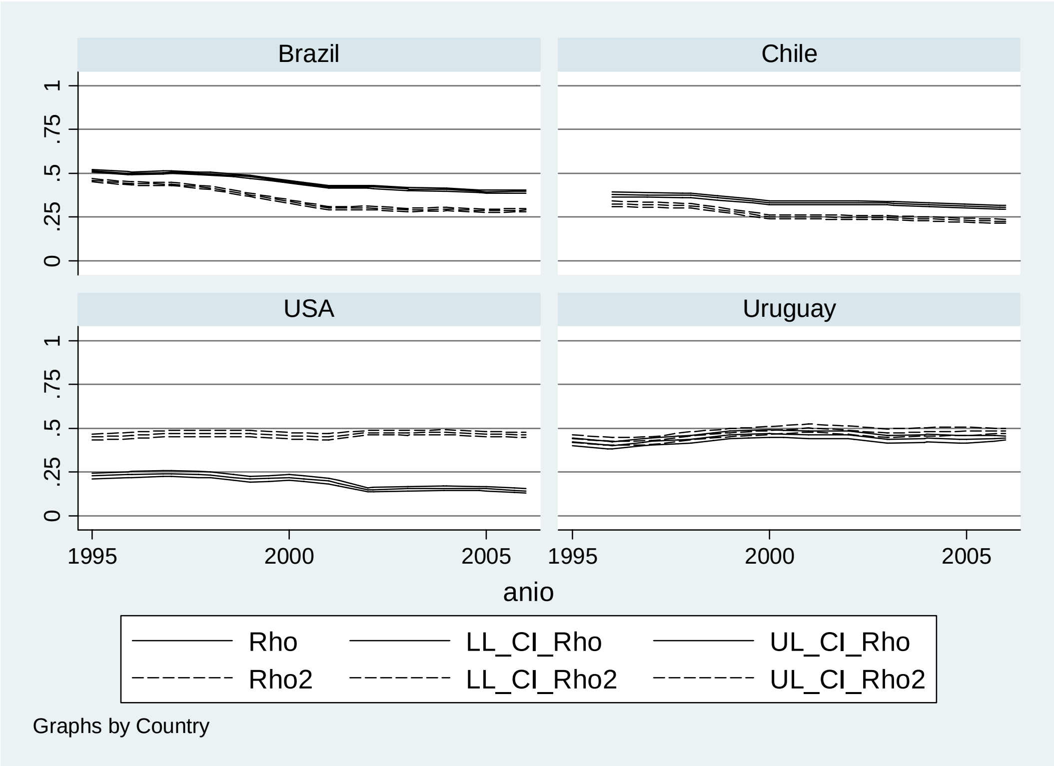
b. Children-Fathers



(1) Beta is the Index when censoring is not considered; Beta2 is the Index when censoring is considered.

(2) LL_CI and UL_CI are the lower and upper limits of the 95th confidence interval, obtained under normal distribution.

Figure 5: Index 3 Absolute Mobility. Variance decomposition of error-components



(1) Rho is the Index when censoring is not considered; Rho2 is the Index when censoring is considered.

(2) LL_CI and UL_CI are the lower and upper limits of the 95th confidence interval, obtained under jointly normally distributed error-components.