# Educational expansion, earnings compression and changes in intergenerational economic mobility : Evidence from French cohorts, 1931-1976

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**Abstract**

This paper analyzes long-term trends in intergenerational earnings mobility in France. I estimate intergenerational earnings elasticities for male cohorts born between 1931 and 1975. This time period has witnessed important changes in the French labor market and educational system, in particular a large expansion in access to secondary and higher education as well as an important compression of earnings differentials. Intergenerational mobility is estimated using a two-sample instrumental variables approach. I pay special attention to the bias that may arise when assessing trends in mobility from cohorts observed at different stages of their life-cycle. Over the period, intergenerational earnings mobility exhibits a V-shaped pattern. Mobility falls between cohorts born in the mid 1930s and those born in the mid 1950s, but subsequently rises. For cohorts born in the first half of the 1970s, age-adjusted intergenerational earnings elasticity amount to around .55. This value is statistically significantly higher than the elasticity estimated for the baby boom cohorts. It is also slightly lower than the elasticity estimated for cohorts born in the 1930s but the difference is not statistically significant. Changes in the extent of mobility mostly reflects the evolution of cross-section earnings inequality, rather than variations in positional mobility.

**JEL Codes:** D1, D3, J3

**Keywords:** Intergenerational mobility, earnings, inequality, trends, elasticity, correlation, education, France.

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# 1 Introduction

Over the last fifteen years, an important body of research has investigated the extent of the intergenerational transmission of income inequality. Two main results have emerged from this literature. First, individual economic well-being, in developed societies, is much more strongly influenced by family background than was thought two decades ago : on average, between 20 and 60% of economic advantage is transmitted, within families, from one generation to the next (Solon 1999, Black & Devereux 2010). Second, the transmission of economic inequality varies considerably across countries and countries where inequality is lower generally tend to exhibit higher intergenerational mobility (Björklund & Jäntti 2009).

From a theoretical perspective, the determinants of intergenerational economic mobility are now well established (Becker & Tomes 1979, Solon 2004). However, beyond the above-mentioned stylized facts, the factors that shape intergenerational economic mobility empirically have not been much explored. Why does the degree of intergenerational mobility vary across countries ? To what extent does it change over time ? How does the level of economic inequality relate to the persistence of inequality across generations ? Have recent changes in the wage structure affected the degree of economic mobility ? What policy intervention in general, and what features of the educational system in particular, may help foster equality of opportunity ? Such important questions remain largely unanswered.

The objective of this paper is to analyze changes over time in the extent of intergenerational earnings mobility in France over the second half of the twentieth century. This period appears particularly interesting for the study of economic mobility, since it witnessed a considerable expansion of access to secondary and higher education, as well as an important reduction in the degree of earnings inequality. In particular, given the large reduction in earnings inequality that occurred throughout the 1960s and early 1970s, older cohort were exposed to a much larger degree of inequality of family environment than the more recent ones. In this context, looking at changes across cohorts in economic mobility may help us improve our understanding of how the intergenerational transmission of inequality is influenced by the overall economic and social environment.

In this paper, intergenerational mobility is measured by the now standard intergenerational earnings elasticity (IGE), which can be obtained by regressing the log of individual earnings on the log of their father's earnings. I estimate cohort-specific IGEs for male cohorts born between 1931 and 1976. In the absence of linked parent-child data sets measuring earnings over such a long period, I use a two-sample instrumental variables approach (Arellano & Meghir 1992, Angrist & Krueger 1995) as first applied to the estimation of the IGE by Björklund & Jäntti (1997). The estimation exploits a labor force survey covering the period 1964-2003 that contains information on both individual earnings and several parental characteristics, including father's education which is used to form a prediction of father's earnings.

As is now well understood, the estimation of the intergenerational earnings elasticity, in particular the assessment of trends, is vulnerable to what has been referred to as the life-cycle bias. This bias arises from the fact that current earnings measured early (resp. late) in the life-cycle tend to underestimate (resp. overestimate) the extent of permanent earnings inequality among fathers or sons (Jenkins 1987, Grawe 2006, Haider & Solon 2006). In this paper, I use the specification of Lee & Solon (2009) to provide estimates of the average and cohort-specific IGE in France that correct for life-cycle bias.

Changes in earnings inequality may affect the extent of intergenerational mobility in various ways. For instance, a more compressed earnings structure in the father's generation may weaken the link between family income and child's human capital investment. Next, a compression of earnings differentials among children is expected, other things equal, to mechanically decrease the IGE. To account for the changes across cohorts in the IGE, I compare its evolution to the evolution of cross-sectional earnings inequality among sons and among fathers. I also analyze changes over time in the intergenerational earning correlation in order to assess changes in positional mobility. Lastly, I try to isolate the contribution of educational expansion to changes in economic mobility using a decomposition approach, that allows me to disentangle two factors : changes in the association between family income and child's human capital, on the one hand, and changes in the returns to human capital, on the other hand.

This paper relates to a series of recent papers that have looked at changes over time in the IGE in various countries. The most extensively studied country is by far the United States. Several studies have estimated trends in the IGE using the Panel Study of Income

Dynamics (PSID) data and reached mixed conclusions. Early PSID studies include in particular Fertig (2003) and Mayer & Lopoo (2005) . One of the limitations of these studies, as shown in subsequent work (Hertz 2007, Lee & Solon 2009) is that life-cycle bias leads to underestimate the IGE for the most recent and youngest cohort. One of the limitation of the PSID for the study of changes over time in the IGE is that it offers a relatively limited cohort span and a rather small sample. The current conclusion that arises from the PSID data is that between the late 1970s and the early 2000s, the IGE has remained roughly constant for males. The longer-run perspective adopted in Aaronson & Mazumder (2008) is probably closer to the perspective of the present paper. They estimate changes in the IGE between 1940 and 2000, using census data and relying, as I do here, on a two-sample instrumental variables approach. Their conclusion is that the IGE exhibits a large fall between 1950 and 1980 and a sharp rise in the recent period. The assessment of trends in economic mobility has attracted researchers' attention in several other countries, including Britain - where non consensus has been reached on trends at work (Ermisch & Francesconi 2004, Blanden, Goodman, Gregg & Machin 2004, Nicoletti & Ermisch 2008, Erikson & Goldthorpe 2010)-, Finland (Pekkala & Lucas 2007), Italy (Piraino 2007), Norway (Bratberg, Nilsen & Vaage 2003) and Sweden (Björklund, Jäntti & Lindquist 2009). With respect to the existing literature on trends in the IGE, the contribution present paper is twofold. First, I analyze of a country that has not been studied so far, over a relatively long time period. Second, I am able to provide a more detailed account of the sources of change in the IGE than what is usually offered in existing papers, using an original decomposition.

Four main results emerge from this paper. First, taking into account life-cycle biases and using an estimation procedure comparable to state-of-the-art estimates reveals that the average IGE in France is around .5, a value higher than what was originally found in Lefranc & Trannoy (2005). Second the IGE has fallen from a high of value of .6 for cohorts born in the 1930s to around .45 for those born in the 1950s, but has subsequently risen to a level close to the beginning of the period. Third, the initial fall in the IGE results from the joint effect of a more equal labor market and a more opened educational system. Fourth, the recent rise in the IGE partly reflects a rise in the association between parental income and child's education. The rest of the paper is organized as follows. I first discuss the estimation procedure and the data used in the analysis (section 2). Then I present the results of the first-step estimation (section 3) and analyze the main trends in the IGE across cohorts (section 4). Finally, I examine long-term changes in earnings inequality and its contribution to changes in the IGE in section 5 and examine the role the educational expansion in section 6. Section 7 concludes.

# 2 Estimation method and data

## 2.1 Estimation method

Most of the economic analysis of intergenerational mobility focuses on estimating the IGE in permanent (or long-term) earnings. This elasticity is given by the coefficient $\beta$ in the following intergenerational earnings regression model :

$$Y_i = \beta_0 + \beta X_i + \epsilon_i \tag{1}$$

where $Y_i$ denotes the log of individual $i$'s long-term earnings and $X_i$ denotes his father's long-term earnings. As already discussed in the literature, $\beta$ should not be seen as a structural parameter measuring the causal effect of parental resources on child's earnings, but rather as a "catch-all" descriptive measure of the intergenerational association in earnings, capturing all possible channels of transmission.

To assess trends in the IGE, one can rely on the following extension of the intergenerational regression model, that allows for cohort heterogeneity in the parameters :

$$Y_{ic} = \beta_{0c} + \beta_c X_{ic} + \epsilon_{ic} \tag{2}$$

where $c$ is an index of the birth cohort of the children and $\beta_c$ is the IGE for cohort $c$. The main objective of this paper is to assess changes in $\beta_c$ for the widest possible range of cohorts.

The direct estimation of equation 2 for a large interval of cohorts requires a considerable wealth of information. In fact, not only does it call for a linked data set in which both father and child's earnings are observed, but for each generation one needs to observe a time-series of individual earnings in order to measure long-term earnings. Very few data sets satisfy this data requisite although there are some exception like the PSID. But even this fairly rich and long panel fails to cover a wide range of children's cohorts. In France, I am not aware of any linked father-child data set that conveys information on long-term earnings.

In this paper, I estimate the $\beta_c$s using a two-sample instrumental variables (TSIV) approach as originally derived in Arellano & Meghir (1992) and Angrist & Krueger (1995). This method was first applied to the estimation of the IGE by Björklund & Jäntti (1997). The basic principle behind TSIV estimation is to replace $X_{ic}$ in equation 2 by a prediction $\hat{X}_{ic}$ formed on the basis of some observable father's characteristics, $Z_{ic}$. Here, I use father's education to predict father's earnings. In the rest of this section, I discuss the properties of TSIV estimation and present the details of the specification used in the paper.

The data requirements for TSIV estimation are significantly less stringent than for the direct estimation. The prediction is derived from a first-step equation which is estimated on a sample that is representative of the fathers' population, and in which one observes both earnings and the characteristics $Z_{ic}$. Given the estimation of the first-step, the data requirement for the estimation of $\beta_c$ is to observe both child's income and father's characteristics.

TSIV has been extensively used for the estimation of the IGE and its properties are discussed in several papers including Solon (1999) and Nicoletti & Ermisch (2008). These properties depend on the choice of the instrument. If the instrument only affects child's earnings through its effect on father's earnings, TSIV estimates of the $\beta_c$s are consistent. Indeed, in this case TSIV estimation offers the significant advantage of over-riding the *attenuation bias* that typically arises, because of classical measurement errors, when estimating equation 2 with long-term earnings replaced by current earnings (Solon 1992, Zimmerman 1992, Mazumder 2001)). However, if the instrument has a direct effect on the child's outcome, than the TSIV estimates is biased and the direction of the bias depends on the sign of the direct effect. When using father's education as an instrument, the expectation is that the direct effect will be positive, hence resulting in an overestimation of the IGE. However, in practice, the order of magnitude of this overestimation turns out to be small, as discussed in Björklund & Jäntti (1997).

Another important source of bias in the estimation of the IGE is what has been recently referred to as the *life-cycle bias* (Jenkins 1987, Grawe 2006, Haider & Solon 2006). This bias arises when using current (usually annual) earnings instead of permanent earnings in the estimation of the IGE. In the presence of individual heterogeneity in earnings growth over the life-cycle, current earnings measures permanent earnings with error. Furthermore, it can be shown that the error is not of the classical type and is correlated with both true permanent earnings and individual age.[1] As a result, differences in current earnings across individuals will in general provide a biased estimate of permanent income differentials. Since age-earnings profiles are steeper for high income individuals, current income differentials, measured at an early stage of the life-cycle, will underestimate permanent income differentials; current income at the end of the life-cycle will over-estimate permanent income differentials.

This form of measurement error will introduce an asymmetric bias in the estimation of $\beta$, depending on whether child or father's earnings are affected by this bias. Using current earnings early (resp. late) in the life-cycle, as a proxy for *child's* permanent earnings will lead to underestimate (resp. overestimate) $\beta$. Conversely, using current earnings early (resp. late) in the life-cycle, as a proxy for *father's* permanent earnings will lead to overestimate (resp. underestimate) the IGE.

Accounting for life-cycle biases is of paramount importance when assessing trends over time in the IGE. Mechanically, younger cohorts will be observed at an earlier stage of their life-cycle then older cohorts, resulting in a lower IGE. In this case, inadequate treatment of life-cycle bias will induce a spurious downward trend across cohorts in the value of the IGE (Hertz 2007, Lee & Solon 2009, Nicoletti & Ermisch 2008).[2] To account for this bias, my specification follows the one of Hertz and Lee & Solon and allows the IGE to vary with child's age by introducing an interaction term between child's age and father's predicted

[1] The classical measurement error case refer to the situation where measurement error is independent of the true value

[2] As a result, the use of different sample selection criteria for fathers' and children's ages across studies of intergenerational mobility jeopardizes the comparability of IGE estimates across countries, as discussed in Grawe (2006).

earnings. By focusing on the main effect of father's earnings one can wipe out the effect of child's age on the cross-cohort comparison. The reference age for children used in the estimation of the IGE is the age of 40, as suggested by the rule of thumb of Haider & Solon (2006). Similarly, for all cohorts I predict father's earnings at the age of 40, hence eliminating life-cycle bias on the dependant variable side as well.

In the end, I use the following specification for the second-step equation :

$$Y_{ict} = \alpha_t + \beta_c \hat{X}_{ic} + g(age_{ict}) \times \hat{X}_{ic} + f_C(age_{ict}) + e_{ict} \tag{3}$$

where $i$ and $t$ are indices for individual and time. $c$ denotes the five-year birth cohort of individual $i$. The $\alpha_t$s denote time dummies and $f$ and $g$ are fourth order polynomial functions in individual age. I allow the age profile to vary with year birth and consider four "super cohorts" indexed by $C$. The birth cohorts of these four groups are the following : 1933-1942, 1943-1949, 1950-1959, 1960-1973.[3] In principle, the use of polynomial functions for age would allow to simultaneously include time and cohort dummies. Cohort dummies however turn out to be insignificant when added to this specification and their inclusion does not affect the results. $\hat{X}_{ic}$ is predicted father's earnings at age 40; the variable $age$ is normalized to zero at age 40. Consequently $\beta_c$ denotes the IGE for cohort $c$ if, as suggested in Haider & Solon (2006) the life-cycle bias is zero at age 40.

Let us now turn to the specification of the first-step equation. Its purpose is to predict father's income at the age of 40. The prediction is based on information on father's education. One of the difficulties is that for some of the children's cohort, in particular the oldest ones, their father's cohort is observed fairly late in its work career. For these cohorts, earnings differentials in mid-career has to be predicted on the basis of end of career wage differentials by education group. Hence, one needs to take away the wage growth that occurred in between. Furthermore, for the prediction of wage differentials by education to be consistent, one needs to account for heterogeneity by wage growth by education. This is done by estimating parametric, yet flexible, education-specific age-earnings profiles.

[3] The cutoff years are chosen to balance group size.

Formally, the second-step model I estimate is the following :

$$X_{ict} = \alpha_t + \sum_j \gamma_c^j Educ_{ic}^j + f(age_{ict}, Educ_{ic}) + e_i \quad (4)$$

where $Educ_{ic}^j$ is a set of education dummies; $f_c(age_{ict}, Educ_{ic})$ is a fourth polynomial in age, specific to each level of education[4]; $age_{ict}$ is centered at age 40. This equation is used to predict father's earnings at age 40 as :

$$\hat{X}_{ict} = \sum_j \hat{\gamma}_c^j Educ_{ic}^j$$

### 2.2 Data

**Data sets and sample selection** The data are taken from the first five waves of the FQP (*Formation, Qualification, Profession*, i.e. Education, Training and Occupation) surveys conducted by INSEE in 1964, 1970, 1977, 1985 et 1993. A new sample is drawn for each wave, so that the data do not have a panel structure. The number of individuals surveyed varies across waves : 25 000 individuals in 1964, 38 000 from 1970 to 1985 and in 2003, and 19 000 in 1993. For all waves but 1993 and 2003, individuals surveyed are taken from a stratified sample of the French population of working age, with different sampling probabilities for each stratus. The FQP surveys focus on the description of individual labor market outcomes, education as well. As discussed below, it also includes information on several parental characteristics that may be used in TSIV estimation.

In the analysis, I use two distinct samples. The main sample is the sample of children, on which the second-step equation (equation 3) is estimated. For this sample, I use waves 1970 to 2003 of the survey. In each wave, the sample is restricted to male heads of household, born between 1931 and 1975 and aged 28 to 50 years old at the date of the survey. Since income is not reported for these categories, I exclude self-employed children as well as children whose father was self-employed from the sample. However, I test for the sensitivity of the results to this exclusion.

The second sample used in the analysis is the sample of "pseudo-fathers" on which the

[4]In variants of this model, I also allowed for cohort heterogeneity in the function $f$, without any significant impact on the results.

first-step equation (equation 4) is estimated. This sample should be representative of the population of the fathers of the individual sampled in our children sample. For this sample, I use all waves of the survey, from 1964 to 2003, and restrict the sample to male heads of household, aged 25 to 60 years old as of the survey date, who report at least one child, and are not self-employed.

As previously discussed,the equations estimated on both samples allow for heterogeneity by cohort in the effect of the explanatory variables. For the estimation of the first-step equation, I use three-year cohorts to warrant large enough groups in each cohorts.[5] For the estimation of the second-step equation, where the sample relies on a smaller number of survey waves, I use five-year cohorts.

The matching of individuals from the children and the pseudo-fathers samples is based on the father's characteristics used in the prediction of father's earnings (as discussed below), as well as on reports, provided in the children sample, of the year of birth of the father. Given the age restriction imposed in the children and pseudo-fathers samples, the oldest children cohort observed in the sample was born in 1931 and the oldest cohort of pseudo-fathers from which to predict fathers' earnings was born in 1904. This 27 years gap is reasonable given that the mean age of the fathers at the birth of their children was a bit above 30 in 1933.[6] For children whose father was born before 1904, we assign the predicted father's wage of the cohort born in 1904. When information on father's birth year is missing the prediction of father's earnings is based on the distribution of birth age computed from non-missing observations.

**Main variables** For all individuals surveyed, the data contain detailed information on education, as well as training, labor market experience, 4-digits occupation and industry when relevant. Individual annual earnings (excluding unemployment benefits) in the previous year and number of months worked full- and part-time are also collected in all waves except 1964. In 1964, annual earnings are recorded in interval form, using 9 intervals. Hence, all estimations results reported for wave 1964 are based on interval regression. In all waves earnings refer to labor earnings and are only recorded for salaried workers.

[5] I group the first two cohorts, 1903-1905 and 1906-1908.
[6] Daguet (2002).

All surveys provide information about the respondent's current family (marital status, number of children) and family of origin (number of siblings, respondent's birth rank). Waves 1977 through 2003 also contain a detailed description of the educational attainment and 2-digits occupation of the father of the respondent, and information about the geographical location of the respondent's parents. This information is reported *a posteriori* by survey respondents and refer to the time when the respondent left the schooling system.

While other characteristics of the father are available in the data set, in particular occupation, only education is used in the first step to predict father's earnings. The reason for this is the lack of synchronicity between the father's age at which father's occupation is reported by the child and the age at which occupation is observed in the fathers' sample. On the one hand, children are asked to report the occupation of their father at the time they finished school. On the other hand, the various father's cohorts are observed at different points of their work career. For instance the oldest cohorts are typically observed in their fifties. Hence, using child's report of occupation would be misleading as it would amount to assume that occupation stayed constant between the middle and the end of the career. The same problem could arise for other time-varying characteristics such as industry. On the contrary, it is reasonable to assume that father's education stayed constant over the work career.

In all waves, education is recorded using a 10 levels education classification that distinguishes between general and vocational education but the categories changed several time over the five waves. I recoded education using a consistent classification across survey waves. The classification is based on the highest degree achieved by the individual and distinguishes between six different categories that reflect key stages in the French educational system. The first one gathers individuals with no degree. The second one corresponds to individuals who passed the certification exam organized at the end of primary education (*certificat d'études primaires.*) This was the major degree taken in older cohorts, among children of the lower and middle class.[7] Next, we consider intermediate secondary education degrees, for the general and vocational tracks. The last two groups considered are individuals who hold un upper secondary degree (*baccalauréat*) or a higher education

[7] Starting in 1972, the *certificat d'études primaires* was only taken by adults, in the context of adult education programs. It was abandoned in 1989.

degree.

The main summary statistics are given in table 1. We now turn to the discussion of the results of the estimation of the first-step equation and the analysis of trends in the earnings structure and the distribution of education.

# 3 First-step estimates and trends in educational attainment and returns to education

Before presenting the results of the first-step equation, it is useful to briefly document the main historical trends in educational enrollment in France over the twentieth century. These trends are described in figure 1. The major evolution is the large rise in access to secondary and higher education. Among cohorts born at the beginning of the century, a very large share of about 70% of the population exhibits a very low level of education, with at most a primary education degree. At this period, mass-education is confined to primary schools. Secondary education is to a large extent a privilege of the upper class. The degree of tracking is extremely high at this time. At the level of primary education, two tracks co-exist. The first one offers regular primary education, as well as the possibility of two extra-years of advanced primary training (*classes primaires supérieures*). The second track is integrated into high-schools (*lycées)*, that at the time concurrently offer primary education from the age of six. The two tracks are entirely disconnected and only the children who attended the second one are offered the chance to reach secondary education degrees.

The opening up of access to secondary education takes place gradually after 1930 and leads to a steady rise in the share of individuals with lower and upper secondary degrees. This results from several policy reforms occurring between 1936 and 1975, but also reflects the development of schooling infrastructure to accommodate the rise in number of pupils. Among the key stages of educational reform in France in the twentieth century, one should mention the extension of compulsory education from 13 to 14 (Zay, 1936)[8] then 16 years old (Berthoin, 1959) and two key reforms undertaken to abolish the strong tracking at work

[8]The name is the name of minister of education responsible for the reform mentionned, the date is the date when the reform was enacted.

in secondary education (Fouquet,1966; Haby, 1977). The rise in access to higher education starts in the 1950s, for cohorts born in before WWII and develops throughout the 1960s and the 1970s among the baby-boom cohorts. It further accelerates at the end of the period for cohorts born between the late 1960s and the 1970s. Lastly, it is worth emphasizing that while trends are somewhat similar to other developed countries, educational attainment in France is, throughout the period, markedly lower than in comparable industrialized countries. For instance, at the end of the period, only about 35% of the population obtain a higher education degree and 20% reach the level of upper secondary education.

Let us now turn to the analysis of earnings differentials by level of education. The analysis is based on the estimation of equation 4. Recall that this equation allows for heterogeneity by cohort in the effect of education and heterogeneity by education-groups in age-earnings profiles. The detailed estimation results are provided in the appendix table 4 and summarized in figures 2 and 3.

Figure 2 presents the evolution over time of the earnings structure, by level of education. The earnings premia attached to each of the six levels of education correspond to the coefficients $\gamma_c$s in equation 4. These premia are expressed in deviation from the mean income in each cohort and are predicted at age 40, using estimated age-earnings profiles. The major result that emerges from the figure is that France experiences a marked decline in the returns to education over the twentieth century. The largest fall occurs between cohorts born at the beginning of the century and early baby-boomers born around 1940.

Whether this compression of education earnings premia led to a reduction of the overall degree of earnings inequality cannot be deducted directly from figure 2. This is because the evolutions of earnings dispersion also depends on changes in the distribution of education in the population over time. In fact, at beginning of the century, a very small share of the total population was earning the high wage premia attached to tertiary and upper secondary degrees, as already discussed. I now directly examine trends in earnings inequality over time, as measured by the Gini coefficient. With our data, the major difficulty arising in assessing the evolution of earnings inequality across cohorts, is that cohorts are observed at different points of their life-cycle. To account for that I subtract education and cohort specific age effects to predict mid-career wages for each cohort and compare earnings inequality by cohort. Results are displayed in figure 4. Three main findings emerge from this figure. First, cohorts born in the 1920s and the 1930s experienced a high degree of wage inequality. Second, wage inequality fell markedly in the post WWII period throughout between the 1940 and the 1955 birth cohorts. Lastly, the within-cohort level of wage inequality stayed approximately constant across cohorts born after 1955.

These results on long-term trends in earnings inequality are consistent with results derived from alternative data sources and methodologies. Selz & Thélot (2004) estimate standard Mincer equations for the period 1964 to 1998 and show that the returns to education has fallen over time in France. Based on fiscal data, the results reported by Piketty indicate a significant fall, throughout the 1970s and early 1980s, in the ratio between the average wage of higher-grade professionals and the average wage of manual workers in the manufacturing sector (Piketty (2001), figure 3-7). Piketty (2003) also reports that the 1930s, 1950s and the 1960s where periods of historically high earnings inequality.

Two main factors may account for the fall in earnings inequality displayed in figure 4. The first one is the massive wage compression that occurred at the end of the 1960 (in particular in 1968, after the 1/3 rise in the minimum wage) and in the early 1970's. The second one is the competitive wage adjustment that followed the rise in the supply of highly educated workers, as discussed in Goux & Maurin (2000).

Lastly, figure 3 presents age-earnings profiles by level of education, estimated in equation 4. Following Murphy & Welch (1990), age profile is captured by a fourth-degree polynomial. The results are consistent with evidence reported elsewhere of a fanning out of wage profiles level of education.[9] In particular, the age-earnings profiles of workers with degrees equal to lower secondary vocational degrees or lower are fairly similar among themselves but also flatter than the age profiles of individuals with other secondary degrees (upper secondary or lower secondary general degree). The steepest profile corresponds, as expected, to individuals with higher education.

To summarize, the extent of wage and educational inequality has varied considerably across cohorts over the last century. We now investigate how much of this inequality has been transmitted across generations and the extent to which this intergenerational

[9]See for instance Lillard (1977) for early evidence.

transmission has varied over time.

# 4 Changes in the intergenerational earnings elasticity

## 4.1 Main results

Table 2 reports estimates of the IGE for various specifications of the intergenerational regression model. The first coefficient in column 1 is the average IGE across all cohorts included in the analysis. Over the full sample, the average intergenerational elasticity amounts to .53. This value is consistent with the estimate of .4 reported in Lefranc & Trannoy (2005). The fact that this latter estimate is based on a younger sample of children and does not correct for life-cycle bias may account for the discrepancy. As shown below, the rest of the gap may also be explained by the inclusion of both older and younger children cohorts, for which the IGE is higher. However, this value of the IGE appears high compared to estimates obtained for other developed countries and surveyed for instance in Björklund & Jäntti (2009). IGEs of comparable magnitude are only found in low mobility countries such as the United States (Mazumder 2005), Italy (Mocetti 2007), and the United Kingdom (Dearden, Machin & Reed 1997). This confirms that a large fraction of inequality is transmitted across generations in France.

Column 2 of table 2 and figure 5 show our main estimates of the IGE for each of the nine five-year cohorts. The evolution of the IGE exhibits a V-shaped pattern over the period. The degree of intergenerational transmission decreases across cohorts until cohorts born in the late 1950s but rises at the end of the period. Furthermore, as summarized in figure 6, the IGE is significantly higher for the early and late cohorts than for the middle ones.

The value estimated for the two cohorts born in the 1930s is around .6, which is very high, compared to values reported elsewhere. By comparison, Aaronson & Mazumder (2008) report a value of the IGE around .35 for US cohorts born in the same decade and Pekkala & Lucas (2007) report a similar figure for Finland. Two factors are likely to account for this relatively low degree of economic mobility among the oldest cohorts. First, historical sociological evidence indicate that the degree of educational mobility was very low in France at the beginning of the twentieth century (Thélot & Vallet 2000). Second, these cohorts also experienced the high degree of labor market inequality at work in France in the 1950s and the 1960s.

The decline in the IGE occurs for the first part of the baby-boom cohorts, i.e. individuals born in the second half of the 1940s and in the 1950s. In these three cohorts, the IGE reaches a low value of .45, which is still relatively high by international standards. The IGE subsequently rises for cohorts born in the 1960s and in the early 1970s to reach .55.

In the rest of the paper, we investigate the determinants of this rise and fall of economic mobility in France, after performing some sensitivity analysis.

## 4.2 Sensitivity analysis

As suggested by Hertz (2007) and Lee & Solon (2009), it is important to examine the role played life-cycle effects in the estimated mobility trends. To this end, figure 7 compares the main estimates discussed in the previous section with the results one would obtain without controlling for life-cycle biases. Omitting the interaction between father's earnings and age leads on average to underestimate the IGE . Three main points need to be emphasized. First, for all cohorts, the bias is negative. The finding of a negative bias, despite having imposed an age restriction centered on the mid-career (28-50 years old) can be easily explained by the concavity of the interaction effect, as shown in figure 8. Second, the bias is for most cohorts relatively small and at most equal to -.05, with the notable exception of the last cohort. For individuals born in the early 1970s, who are surveyed earlier in their life-cycle, the bias is more important and close to -.1. Third, the V-shaped trend in the intergenerational persistence of inequality is largely present even without controlling for life-cycle effects. Including the interaction term between child's age and father's wage simply slightly reinforces the estimated upward trend in the IGE in the most recent period.

The second robustness check I perform amounts to assess to the influence of excluding self-employed workers. Since labor income is not reported by self-employed workers, I excluded from my samples both self-employed children and the children of self-employed fathers. The latter category represents about 30% of the children's sample while the former one amounts to 13%. There is no way to satisfactorily predict the incidence of excluding self-employed workers from the analysis. To explore this question, I simulated father's earnings for the children of self-employed workers on the basis of the Mincer equation estimated for salaried fathers. This allows me to include the children of self-employed fathers in the estimation, as long as they are themselves salaried.[10] Results are given in table 2, column 4 and in figure 9. The level and time trends are very similar to those previously discussed. Of course, the validity of this robustness check hinges upon stringent restrictions. It requires that the relationship between schooling and earnings is similar for salaried and self-employed workers. Or at least that the bias induced by the use of a Mincer equation estimated on the sole sample of salaried workers stays constant over time. Both hypothesis are of course open to discussion. However, at the very least, figure 9 indicates that the relationship between father's *education* and child's wage is very similar for the children of self-employed and salaried fathers and suggests that the trends in figure 5 may capture a general pattern of changes in economic mobility in the French society.

## 5 Changes in earnings inequality and the intergenerational correlation coefficient

Changes in economic mobility, as captured by the intergenerational earnings elasticity, are of course deeply connected to the evolution of earnings inequality in society. Since the IGE is a regression coefficient its value is sensitive to the variance of both fathers' and children's earnings. In particular, for a given distribution of parental characteristics, any reduction in earnings inequality among children, would "mechanically" lead to a fall in the IGE.

This should by no means suggest that the IGE provides an inadequate measure of economic mobility. In fact, if some policy change brings more equality in the children's generation, for a given degree of inequality among parents, the intergenerational transmission of inequality unambiguously decreases and most people would probably agree that intergenerational mobility has increased. One should however note that a reduction in earnings inequality in the children's generation leaves unaffected the chances that a child from a disadvantaged background succeeds better than a child from a more privileged one,

[10] In my sample, this is the case for 75% of the children of self-employed fathers.

or *vice versa*. This specific form of mobility is usually referred to as positional mobility.[11] It corresponds to the sociological notion of exchange (as opposed to structural) mobility.

Positional mobility can be measured by the intergenerational correlation coefficient (IGC), $\rho$. This coefficient is given by the usual formula and is by construction unaffected by changes in the variance of earnings in the children's or father's generation. The link between the IGC and the IGE is given by:

$$\beta = \rho \frac{\sigma_Y}{\sigma_X} \tag{5}$$

In the steady state, the variance of earnings is constant across generations and the IGE and IGC are identical. This is no longer the case whenever $\sigma_Y$ and $\sigma_X$ differ. Furthermore, this formula makes clear that changes in the IGE will reflect both changes in positional mobility and the evolution across generations of earnings inequality. I now examine the historical evolution of these two components.

The main challenge for assessing trends in the ratio $\frac{\sigma_Y}{\sigma_X}$ is that, as already discussed, I do not observe permanent earnings for children and fathers. For children, I observe current earnings. Furthermore, the point in individual life-cycles where current earnings are observed varies across cohorts. This problem is addressed by removing age effects around age 40. Removing life-cycle effects, however, does not eliminate transitory earnings components. As a consequence, $\sigma_Y$ will be overestimated. On the contrary, for fathers, my earnings measure is predicted on the basis of observable characteristics. Since this leaves out unobserved permanent characteristics, $\sigma_X$ will be underestimated. However, if the proportional bias in the estimation of $\sigma_Y$ and $\sigma_X$ stays constant over time, the estimation of the trends in $\frac{\sigma_Y}{\sigma_X}$ will be consistent, although the estimation of the level will be biased.[12] Of course, given an estimate of $\frac{\sigma_Y}{\sigma_X}$, equation 5 implies that the IGC can be estimated by the product of the ratio $\frac{\sigma_Y}{\sigma_X}$ with our estimated of the IGE. Again, only the trends in the IGE are likely to be consistently estimated.

Trends in the IGC and in the father-son ratio of log-earnings standard-deviations are

[11] For a recent discussion of the definition and measurement of mobility, see for instance Cowell & Flachaire (2011).

[12] The results reported in Moffitt & Gottschalk (1995) are supportive of this assumption. This assumption is implicit in Aaronson & Mazumder (2008)

given in figure 10. Until the late 1950s, the evolution of the ratio $\frac{\sigma_X}{\sigma_Y}$ is broadly similar to the trends in the IGE. It is relatively stable over the 1930s and the first half of the 1940s. It strongly rises between the second half of the 1940s and the 1950s, under the impulse of the fall of earnings inequality in the children's generation, already noted in figure 4. After that date, it slowly decreases since, with a lag, earnings inequality starts declining among parents. The IGC on the contrary follows a different evolution. It is roughly stable and if anything slowly declining over cohorts born in the 1930s and 1940s. It rises in the 1950s to reach a level slightly above the starting value. It keeps mildly increasing over the latest cohorts.

Several conclusions emerge from these discrepant trends. The first one is that the large fall in the IGE occurring between cohorts born in the early 1930s and those born in the late 1950s is largely driven by the fall in cross-section earnings inequality among children. In other words, children born after World War II (WWII) inherit a smaller share of their parents' economic advantage because society has become more equal. Not because the degree of positional mobility in the French society has increased. Second, the later rise in the IGE is, to some extent, the delayed effect of this fall in cross-section inequality. With time, the ratio $\frac{\sigma_X}{\sigma_Y}$ decreases and the IGE converges back to its steady-state value, which depends on the degree of positional mobility, i.e. the IGC. Lastly, the recent evolution is made even worse by the slow rise in the IGC, which suggests declining positional mobility.

## 6 The contribution of educational expansion

Educational policy is often seen as the means *par excellence* of equalizing life-chances and fostering social mobility. On the opposite side, educational investment is also often considered as one of the main channels of the intergenerational transmission of ability. In this perspective, the rise in access to upper secondary and higher education that occurred after World War II may have contributed to the evolution of intergenerational mobility. I investigate this contribution in this section.

The contribution of education acquisition to the intergenerational earnings elasticity can be summarized by the following system :

$$
\begin{aligned}
H_{ic} &= \gamma_c X_{ic} + u_{ic} && (6)\\
Y_{ic} &= \beta_c^1 H_{ic} + \beta_c^2 X_{ic} + e_{ic} && (7)
\end{aligned}
$$

where $H$ denotes the human capital of the child. Equation 6 captures the relationship between parental income and human capital accumulation, as discussed for instance in Becker & Tomes (1979) and Solon (2004). In equation 7, child's earnings are determined by child's human capital and, residually, parental earnings. Using this system, the IGE can be expressed as :

$$\beta_c = \beta_c^1 \gamma_c + \beta_c^2 \tag{8}$$

As this equation makes clear, the intergenerational earnings elasticity is a function of three key parameters : the effect of parental income on the human capital of the child ($\beta_c^1$), the returns to human capital ($\gamma_c$) and the residual effect of parental earnings on child's earnings, conditional on human capital ($\beta_c^2$).

To implement this decomposition, I first estimate equation 7 by regressing child's income on child's number of years of education and predicted father's income. As for the main IGE estimates, the coefficient on father's income should not be interpreted in a causal sense but as a catch-all measure of the residual impact of all family attributes related to income, once educational attainment has been taken into account. The results are given in column 2 of table 3 and in figure 11.[13] Several results emerge from this estimation. First, similar to the results found for the US (Aaronson & Mazumder 2008) and Sweden (Björklund et al. 2009), I find that the residual effect of father's earnings $\beta_c^2$ accounts for 50 to 60% of the total IGE. In other terms, education acquisition accounts for at most half of the intergenerational transmission of earnings inequality. Second, the time trend followed by the residual elasticity is roughly similar to that of the base IGE : the IGE falls for intermediate cohorts and rises again for the most recent ones. Interestingly, while the overall IGE at the end of the period is lower than for cohorts born in the 1930s, it is

[13]The econometric model also allows for year and age effects, as well as age $\times$ father's earnings interactions. Table 3, column 1, re-estimates the main IGE model on the sub-sample with non-missing education data.

no longer the case for the residual elasticity. For cohorts born in the 1970s, the residual transmission of earnings inequality is higher than for cohorts born in the 1930s.

The contribution of education acquisition to the intergenerational transmission of earnings can be computed as the gap between the overall IGE, $\beta_c$ and the residual elasticity $\beta_c^2$. This contribution is represented in figure 11. It falls almost continuously over the period, from a high value of about .26 for cohorts born in the 1930s to a low value of about .16 for cohorts born in the 1970s. As emphasized by equation 8, this evolution results from changes in two parameters : the semi-elasticity of human capital to parental income and the earnings returns to human capital. The evolution of the latter component is given in table 3, column 2. The returns to education fall over time from .055 to about .03. This fall occurs in two steps, with a first drop between the early 1930s and the early 1940s cohorts and a second one for cohorts born in the late 1960s and early 1970s. These values of the returns to education lay on the low end of estimates reported in other studies, probably owing to the inclusion of father's earnings as an additional regressor. However trends are consistent with previous results (Selz & Thélot 2004).

This fall in the returns to education should induce, other things equal, a decrease in the IGE. But the overall evolution of the contribution of education to the intergenerational transmission of earnings also depends on the semi-elasticity $\gamma$ of years of schooling with respect to father's earnings. This semi-elasticity is reported in column 3, table 3 and in figure 12. The statistical association between years of education and parental earnings falls over time until cohorts born in the first half of the 1960s but subsequently rises to reach an even higher level than for the early 1930s cohorts.

These results indicate that educational mobility has risen over time over most cohorts. They are consistent with studies that have focused on the association between parents' and child's education (e.g. Thélot & Vallet 2000) and who report a upward mobility trend throughout the twentieth century. However, one should note that these studies usually fail to report the most recent fall in educational mobility noted here, with the exception of Vallet & Selz (2007).

This increase in educational mobility, combined with the effect of the fall in the returns to education, leads to a fall in the intergenerational transmission that occurs through educational acquisition. To disentangle the two contributions, I plot in figure 11 the evolution of $\beta_c^1\bar{\gamma}$, i.e. the component of the IGE arising through education that would have occurred if the returns to education had stayed constant over time at their mean value $\bar{\gamma}$. Since it amounts to hold constant any evolution that might have occurred on the labor market, this comes closer to measuring the contribution of changes that have taken place in the educational system. This simulation casts a different light on the contribution of education to the intergenerational transmission of inequality, as it indicates a fall in mobility, even beyond the initially low levels of the 1930s.

# 7 Conclusion

In this paper, I have estimated trends in intergenerational mobility in France for cohorts of children born between the 1930s and the mid-1970s. The first result arising from this analysis pertains to the overall level of intergenerational mobility in France. Once life-cycle effects are taken into account, the intergenerational earnings elasticity in France amounts to an average value of .53. This value is much higher (although consistent) with estimates previously reported for France and indicates a very low degree of mobility by international standards. Second, I show that the intergenerational earnings elasticity followed a V-shaped patterns across birth cohorts. From a high value of .6 for cohorts born in the 1930s, the intergenerational elasticity falls to a low value of .45 for cohorts born in the late 1940s and the 1950s. It subsequently rises to reach a value of .55 for cohorts born in the early 1970s.

The fall in the intergenerational transmission of earnings inequality experienced by the cohorts born after World War II seems largely related to the decrease in earnings inequality and the returns to education that occurred for the most part in the 1970s and early 1980s. This reduction in inequality has brought closer together the earnings prospects of individuals whose parents had faced very different earnings levels. At the same time, over this period, the degree of positional mobility, as captured by the intergenerational earnings correlation has remained roughly unchanged.

In the end, this rise in intergenerational mobility turns out to be short-lived and limited to the generational transition between two societies : the unequal society experienced by pre-baby boom cohorts and the more equal one enjoyed by the baby-boomers. The early baby-boomers appeared more mobile than their parents only because they enjoyed more *intra*-generational equality. As the early baby-boomers became parents themselves, their children also experienced this this less unequal society. Yet, since positional mobility did not improve, this led the intergenerational elasticity to fall back toward its initial level.

All in all, these results empirically demonstrates that the intergenerational elasticity is very sensitive to inequality dynamics and that its evolution, outside the steady-state, could provide a misleading characterization of the long-run evolution of intergenerational mobility. This contrasts with the greater stability of the intergenerational earnings correlation.

Whether intergenerational mobility in France will get back, in the near future, to the low level experienced by the cohorts born in the 1930s is an open question. There are reasons, however to be pessimistic. First, positional mobility seems to have slightly decreased. Second, while intergenerational mobility initially benefited from the large educational expansion of the post-WWII era, the association between parents' earnings and child's educational achievement has also risen recently. If earnings inequality was to rise among recent cohorts, the degree to which economic inequality is transmitted across generations could well rise to unprecedented levels.

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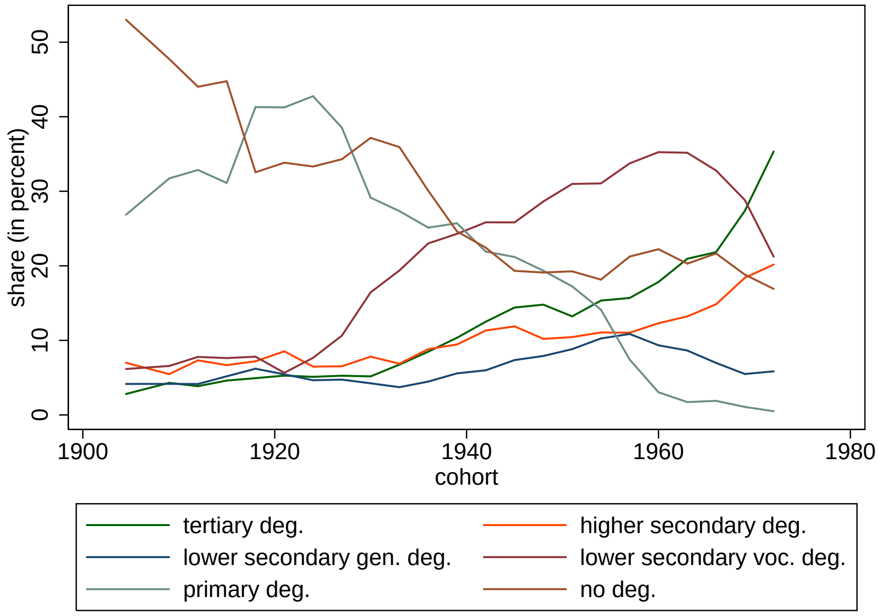
Table 1: Summary statistics

| | | cohort | | | | | | | | |
|---|---|---|---|---|---|---|---|---|---|---|
| | all | 1931-35 | 1936-41 | 1941-45 | 1946-51 | 1951-56 | 1956-61 | 1961-65 | 1966-71 | 1971-75 |
| Observations | 29415 | 2484 | 4817 | 4338 | 4935 | 3972 | 3056 | 2563 | 1771 | 1479 |
| 1970 survey (share) | 0.128 | 0.503 | 0.380 | 0.165 | 0 | 0 | 0 | 0 | 0 | 0 |
| 1977 survey (share) | 0.216 | 0.378 | 0.336 | 0.428 | 0.398 | 0 | 0 | 0 | 0 | 0 |
| 1985 survey (share) | 0.282 | 0.118 | 0.284 | 0.326 | 0.435 | 0.588 | 0.236 | 0 | 0 | 0 |
| 1993 survey (share) | 0.119 | 0 | 0 | 0.0814 | 0.167 | 0.195 | 0.256 | 0.295 | 0 | 0 |
| 2003 survey (share) | 0.254 | 0 | 0 | 0 | 0 | 0.217 | 0.508 | 0.705 | 1 | 1 |
| labor earnings (current Francs) | 88875.9 | 39480.4 | 52859.8 | 70068.0 | 80734.4 | 109236.8 | 121605.2 | 130779.9 | 133049.3 | 120713.1 |
| | (63013.9) | (37909.4) | (49627.7) | (59149.0) | (54220.1) | (55242.8) | (61111.0) | (63688.0) | (56964.8) | (49038.5) |
| father's predicted earnings (normalized) | 1.254 | 1.083 | 1.131 | 1.205 | 1.240 | 1.310 | 1.339 | 1.359 | 1.391 | 1.450 |
| | (0.534) | (0.470) | (0.505) | (0.547) | (0.548) | (0.566) | (0.559) | (0.505) | (0.456) | (0.436) |
| father's year of birth | 1930.5 | 1912.6 | 1916.2 | 1922.5 | 1927.1 | 1932.6 | 1938.2 | 1944.0 | | |
| | (11.05) | (7.714) | (7.618) | (7.108) | (6.869) | (6.902) | (7.078) | (6.797) | | |
| father self-employed (share) | 0.274 | 0.344 | 0.344 | 0.311 | 0.272 | 0.251 | 0.217 | 0.225 | 0.205 | 0.185 |
| | (0.446) | (0.475) | (0.475) | (0.463) | (0.445) | (0.434) | (0.412) | (0.418) | (0.404) | (0.389) |
| father's highest degree (distribution) | | | | | | | | | | |
| tertiary degree | 0.203 | 0.128 | 0.166 | 0.203 | 0.215 | 0.210 | 0.186 | 0.200 | 0.262 | 0.360 |
| higher secondary degree | 0.133 | 0.106 | 0.120 | 0.147 | 0.132 | 0.130 | 0.125 | 0.127 | 0.157 | 0.192 |
| lower sec. gen. degree | 0.304 | 0.244 | 0.272 | 0.277 | 0.320 | 0.333 | 0.349 | 0.362 | 0.333 | 0.230 |
| lower sec. voc. degree | 0.0621 | 0.0360 | 0.0387 | 0.0507 | 0.0610 | 0.0762 | 0.0947 | 0.0916 | 0.0582 | 0.0647 |
| primary degree | 0.125 | 0.199 | 0.197 | 0.173 | 0.147 | 0.124 | 0.0714 | 0.0152 | 0.00944 | 0.00680 |
| no degree | 0.172 | 0.287 | 0.206 | 0.149 | 0.125 | 0.126 | 0.174 | 0.204 | 0.181 | 0.147 |
| individual's highest degree (distribution) | | | | | | | | | | |
| tertiary degree | 0.0680 | 0.0416 | 0.0511 | 0.0646 | 0.0582 | 0.0676 | 0.0777 | 0.0842 | 0.0929 | 0.133 |
| higher secondary degree | 0.0597 | 0.0501 | 0.0548 | 0.0630 | 0.0635 | 0.0616 | 0.0550 | 0.0604 | 0.0523 | 0.0802 |
| lower sec. gen. degree | 0.120 | 0.0354 | 0.0523 | 0.0796 | 0.111 | 0.128 | 0.150 | 0.188 | 0.224 | 0.291 |
| lower sec. voc. degree | 0.0371 | 0.0320 | 0.0358 | 0.0437 | 0.0431 | 0.0373 | 0.0273 | 0.0263 | 0.0515 | 0.0315 |
| primary degree | 0.331 | 0.314 | 0.325 | 0.338 | 0.349 | 0.354 | 0.343 | 0.323 | 0.313 | 0.239 |
| no degree | 0.385 | 0.527 | 0.481 | 0.411 | 0.375 | 0.352 | 0.347 | 0.318 | 0.267 | 0.226 |

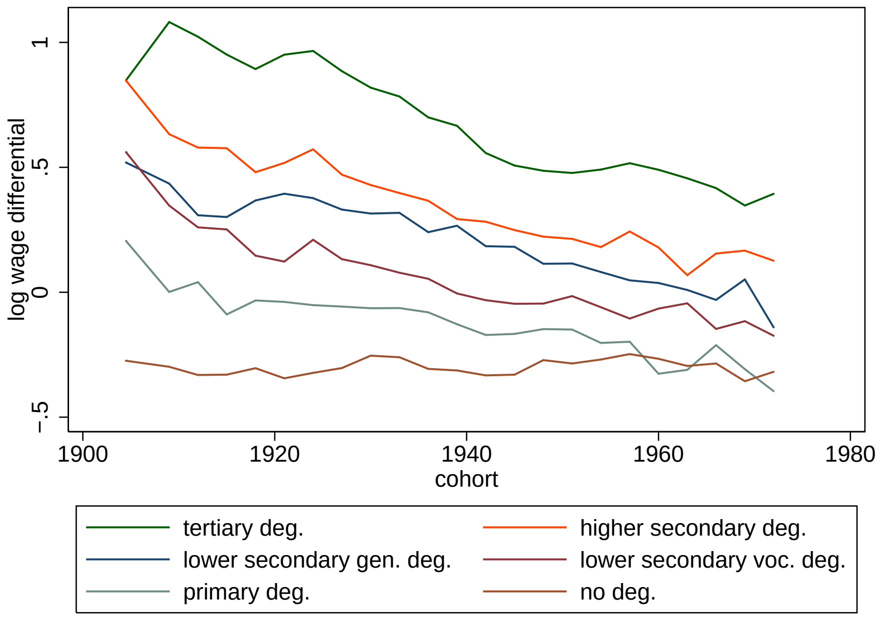
Notes : standard-deviations in parenthesis. Computations are based on the main sample, as described on page 8.

Figure 1: Distribution of education by cohort



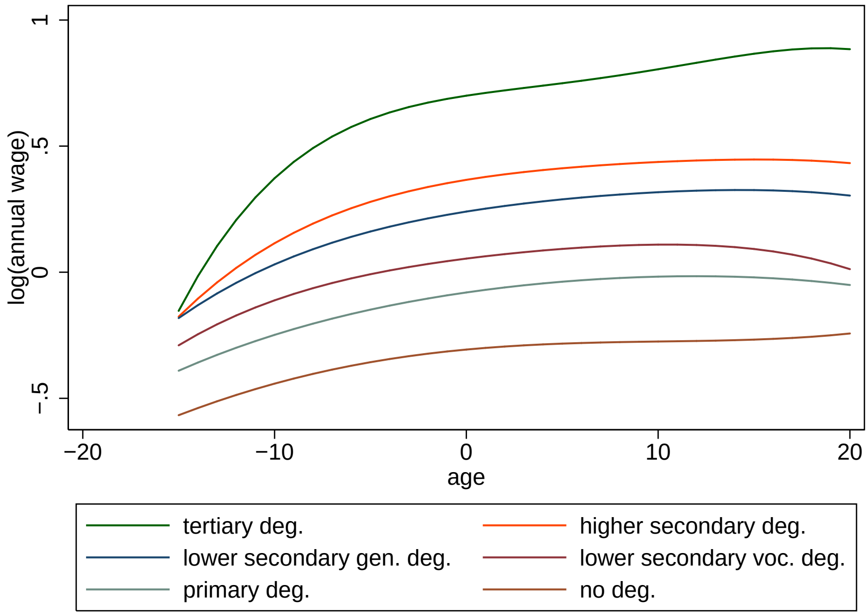

Notes : The figure gives the distribution of highest degree obtained, by cohort for three-year cohorts.

Figure 2: Education earnings differentials by cohort - predicted at age 40



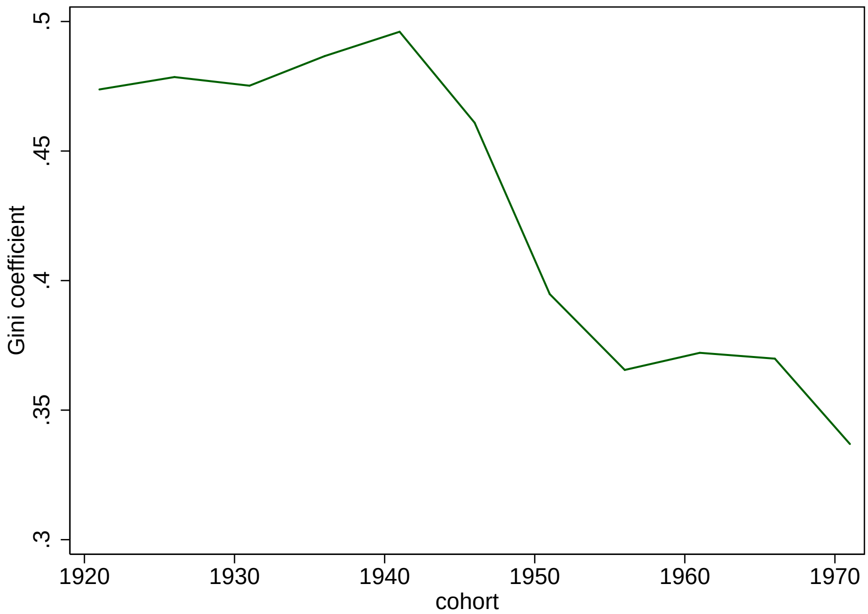

Notes : The figure gives the log annual earnings differential between each educational group and the mean annual earnings, for each three-years cohorts, based on the estimates of equation 4, reported in table 4. Earnings differentials are predicted at age 40.

Figure 3: Age-earnings profile by education




Notes : The figure gives the age-earnings profile for each educational group based on the estimates of equation 4, reported in table 4. The earnings variable on the $y$-axis is the log of annual earnings.

Figure 4: Labor earnings inequality by cohort




Notes : to account for life-cycle effects, age-effects are subtracted in order so as to predict earnings at age 40 for each cohort. This is done using an earnings equation that allows for education- and cohort-specific age-effects.

Table 2: Intergenerational earnings elasticity, by cohort

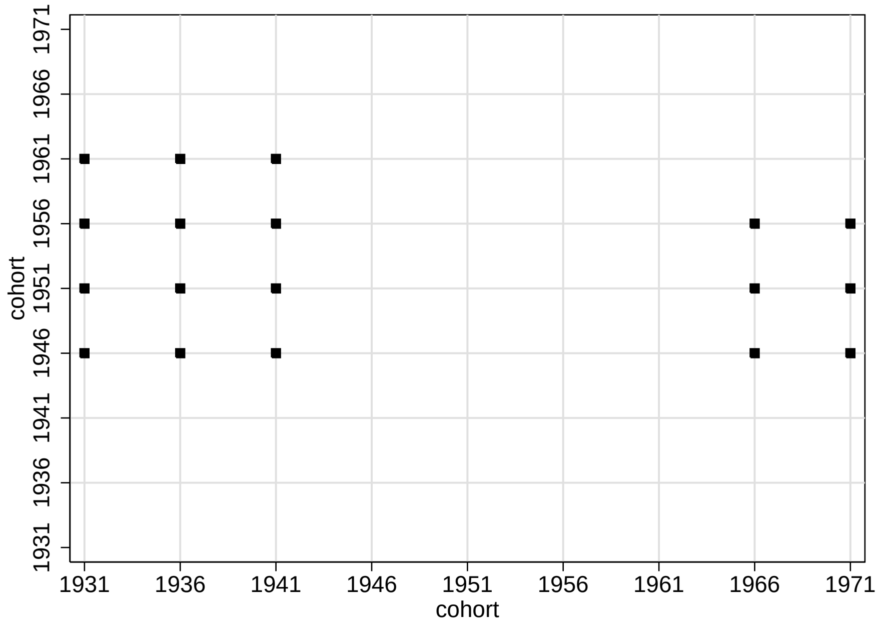
| | (1) | (2) | (3) | (4) |
|---|---|---|---|---|
| all cohorts | .530 | | | |
| | (0.0104) | | | |
| 1931-1935 | | 0.626 | 0.614 | 0.650 |
| | | (0.0245) | (0.0240) | (0.0207) |
| 1936-1940 | | 0.593 | 0.565 | 0.605 |
| | | (0.0184) | (0.0169) | (0.0155) |
| 1941-1945 | | 0.561 | 0.522 | 0.588 |
| | | (0.0184) | (0.0167) | (0.0159) |
| 1946-1950 | | 0.459 | 0.411 | 0.473 |
| | | (0.0185) | (0.0161) | (0.0164) |
| 1951-1955 | | 0.441 | 0.396 | 0.466 |
| | | (0.0204) | (0.0179) | (0.0183) |
| 1956-1960 | | 0.441 | 0.403 | 0.442 |
| | | (0.0223) | (0.0202) | (0.0201) |
| 1961-1965 | | 0.492 | 0.456 | 0.520 |
| | | (0.0234) | (0.0225) | (0.0211) |
| 1966-1970 | | 0.543 | 0.503 | 0.542 |
| | | (0.0289) | (0.0280) | (0.0267) |
| 1971-1975 | | 0.559 | 0.477 | 0.566 |
| | | (0.0357) | (0.0325) | (0.0323) |
| father's earnings×(age-40) | 0.00474 | 0.00414 | | 0.00494 |
| | (0.00132) | (0.00137) | | (0.00119) |
| father's earnings×(age-40)$^2$ | -0.000782 | -0.000574 | | -0.000570 |
| | (0.000191) | (0.000198) | | (0.000172) |
| Observations | 21317 | 21317 | 21317 | 29489 |

Notes : Standard errors in parentheses. The dependent variable is the log of annual earnings. Reported estimates are based on equation 3. Columns 1-3 exclude self-employed children and the children of self-employed fathers, as discussed in page 8. Column 4 includes the children of self-employed fathers, whose earnings are predicted on the basis of the first-step equation estimated on non self-employed fathers. All equations include year and age effects.

Figure 5: IGE by cohort

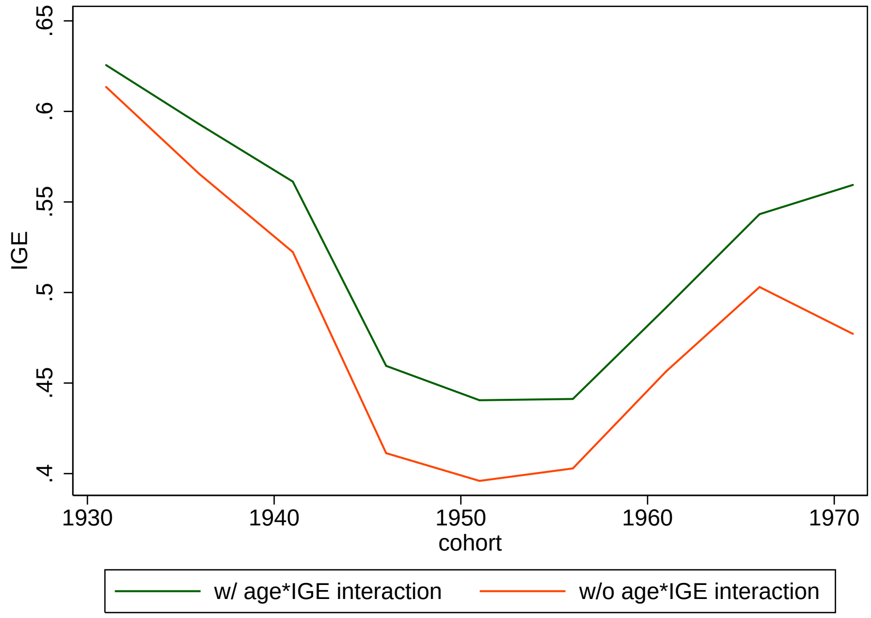

<u>Notes :</u> reported IGEs are based on the estimates in table 2, column 2.

Figure 6: Statistical significance of IGE differences between cohorts



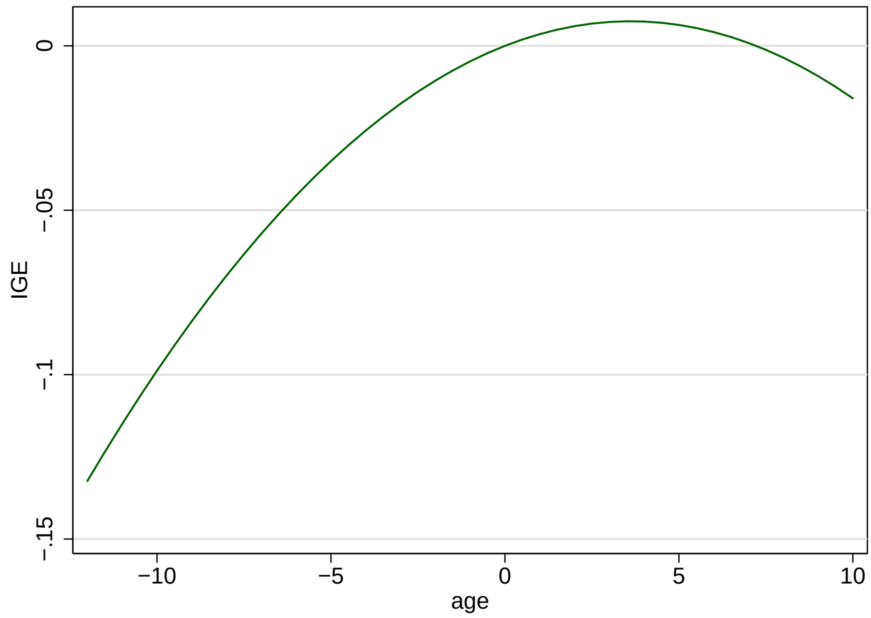

<u>Notes :</u> a square indicates that the IGE for the cohort on the horizontal axis is significantly higher than the IGE for the cohort on the vertical axis, at the 1% level; tests are based on the estimates in table 2, column 2.

Figure 7: IGE by cohort - Influence of age× father's wage interactions




Notes : reported IGEs are based on the estimates in table 2, columns 2 and 3.

Figure 8: IGE - age× father's wage profile

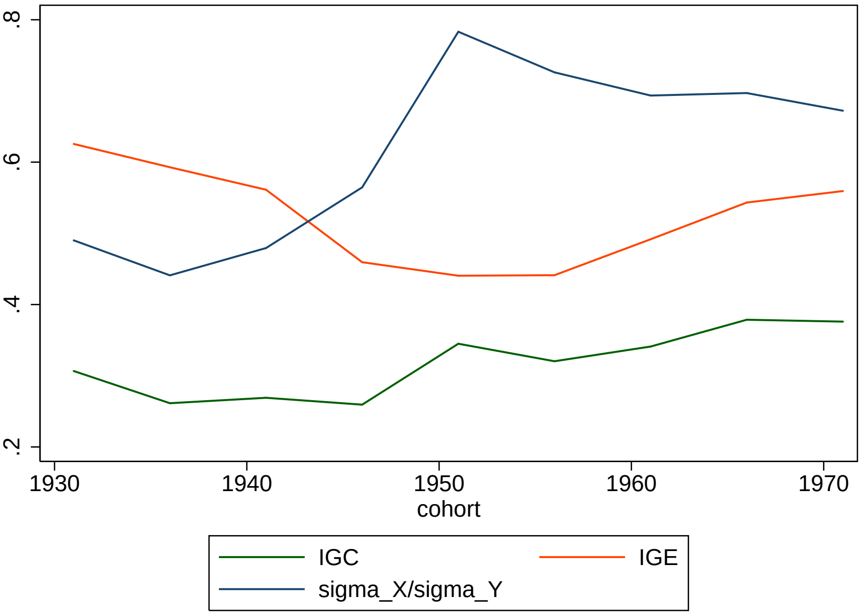



Notes : age profiles are based on the estimates in table 2, columns 2

Figure 9: IGE by cohort - Influence of the inclusion the children of self-employed


Notes : reported IGEs are based on the estimates in table 2, columns 2 and 4.

Figure 10: IGE and IGC by cohort



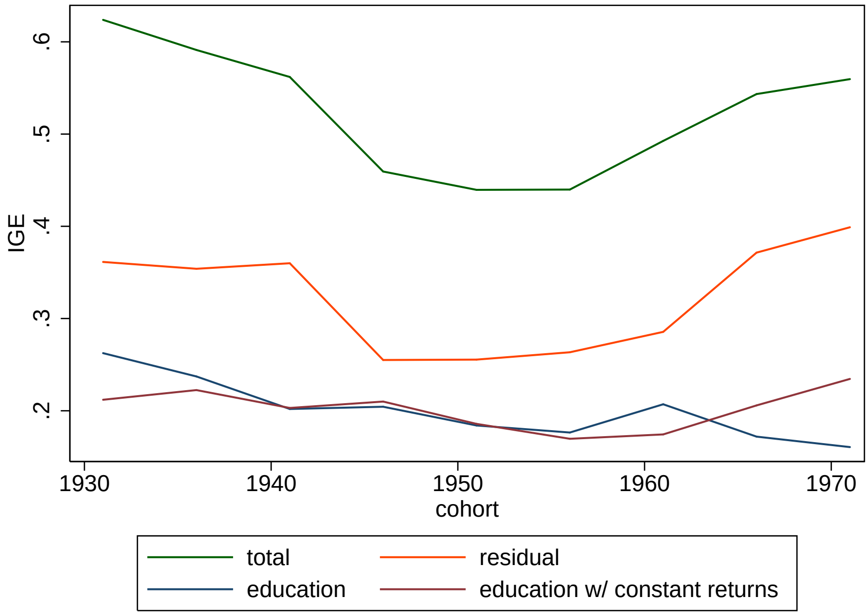

<u>Notes :</u> see details on page 17. The ratio $\frac{\sigma_X}{\sigma_Y}$ is computed using earnings net of age effects for children and predicted earnings for fathers. The estimated IGC is derived from estimates of $\frac{\sigma_X}{\sigma_Y}$ and the IGE

Table 3: Decomposition of the intergenerational earnings elasticity and analysis of the contribution of education, by cohort

| | | (1) | (2) | (3) |
|---|---|---|---|---|
| father's earnings | | | | |
| × cohort | 1931-1935 | 0.624 | 0.361 | 4.678 |
| | | (0.0245) | (0.0259) | (0.195) |
| | 1936-1940 | 0.591 | 0.354 | 4.909 |
| | | (0.0184) | (0.0193) | (0.146) |
| | 1941-1945 | 0.562 | 0.360 | 4.481 |
| | | (0.0184) | (0.0190) | (0.146) |
| | 1946-1950 | 0.459 | 0.255 | 4.634 |
| | | (0.0185) | (0.0192) | (0.147) |
| | 1951-1955 | 0.440 | 0.256 | 4.101 |
| | | (0.0204) | (0.0208) | (0.162) |
| | 1956-1960 | 0.440 | 0.263 | 3.744 |
| | | (0.0223) | (0.0228) | (0.178) |
| | 1961-1965 | 0.493 | 0.286 | 3.848 |
| | | (0.0234) | (0.0247) | (0.186) |
| | 1966-1970 | 0.543 | 0.371 | 4.542 |
| | | (0.0289) | (0.0308) | (0.230) |
| | 1971-1975 | 0.560 | 0.399 | 5.176 |
| | | (0.0357) | (0.0384) | (0.284) |
| number of years of education | | | | |
| × cohort | 1931-1935 | | 0.0564 | |
| | | | (0.00268) | |
| | 1936-1940 | | 0.0485 | |
| | | | (0.00191) | |
| | 1941-1945 | | 0.0452 | |
| | | | (0.00196) | |
| | 1946-1950 | | 0.0445 | |
| | | | (0.00190) | |
| | 1951-1955 | | 0.0453 | |
| | | | (0.00224) | |
| | 1956-1960 | | 0.0477 | |
| | | | (0.00275) | |
| | 1961-1965 | | 0.0522 | |
| | | | (0.00292) | |
| | 1966-1970 | | 0.0365 | |
| | | | (0.00314) | |
| | 1971-1975 | | 0.0315 | |
| | | | (0.00341) | |
| Observations | | 21285 | 21285 | 21285 |

Notes : Standard errors in parentheses. Dependant variables : columns (1) and (2), log annual earnings; columns (3), number of years of education.

Figure 11: Contributions to changes in the IGE by cohort




Notes : total denotes the overall IGE; residual denotes the residual elasticity of child's earnings w.r.t. father's earnings, conditional on child's education; education denotes the gap between total and residual and represents the component of the intergenerational transmission that occurs through education acquisition; education w/ constant returns assumes that the wage returns are constant. See page 18 for details.

Figure 12: Effect of parental income on child's education


Notes : see table 3, column 3.

# A First-step estimation results

Table 4: First-step equation estimation- Dependant variable : log(annual earning

| | | Coefficient | Standard-error |
|---|---|---|---|
| Intercept | | 11.07849 | 0.02317 |
| Survey wave 1964 | | -2.02079 | 0.05239 |
| Survey wave 1970 | | -1.54555 | 0.03763 |
| Survey wave 1977 | | -0.77385 | 0.02083 |
| Survey wave 1985 | | REF | REF |
| Survey wave 1993 | | 0.195 | 0.02162 |
| Survey wave 2003 | | 0.36777 | 0.0452 |
| Higher education × | cohort 1903-1908 | 0.81475 | 0.12866 |
| | cohort 1909-1911 | 1.16551 | 0.10197 |
| | cohort 1912-1914 | 1.10396 | 0.08439 |
| | cohort 1915-1917 | 1.08556 | 0.08202 |
| | cohort 1918-1920 | 1.06448 | 0.06438 |
| | cohort 1921-1923 | 1.15622 | 0.05626 |
| | cohort 1924-1926 | 1.1459 | 0.04904 |
| | cohort 1927-1929 | 1.13798 | 0.04263 |
| | cohort 1930-1932 | 1.06117 | 0.03787 |
| | cohort 1933-1935 | 1.02424 | 0.03331 |
| | cohort 1936-1938 | 0.97107 | 0.03055 |
| | cohort 1939-1941 | 0.9793 | 0.03024 |
| | cohort 1942-1944 | 0.93336 | 0.02902 |
| | cohort 1945-1947 | 0.88995 | 0.03228 |
| | cohort 1948-1950 | 0.83229 | 0.03663 |
| | cohort 1951-1953 | 0.79944 | 0.0432 |
| | cohort 1954-1956 | 0.82623 | 0.04941 |
| | cohort 1957-1959 | 0.85314 | 0.05731 |
| | cohort 1960-1962 | 0.81899 | 0.06513 |
| | cohort 1963-1965 | 0.78852 | 0.07122 |
| | cohort 1966-1968 | 0.76642 | 0.07735 |
| | cohort 1969-1971 | 0.68211 | 0.08634 |
| | cohort 1972-1974 | 0.79002 | 0.09634 |
| | cohort 1975-1977 | 0.89199 | 0.12369 |
| Upper secondary education × | cohort 1903-1908 | 0.81341 | 0.11635 |
| | cohort 1909-1911 | 0.71667 | 0.09527 |
| | cohort 1912-1914 | 0.66017 | 0.08053 |
| | cohort 1915-1917 | 0.71072 | 0.07883 |
| | cohort 1918-1920 | 0.65198 | 0.06351 |
| | cohort 1921-1923 | 0.72269 | 0.05454 |

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| | | Coefficient | Standard-error |
|---|---|---|---|
| | cohort 1924-1926 | 0.75209 | 0.04972 |
| | cohort 1927-1929 | 0.72382 | 0.04221 |
| | cohort 1930-1932 | 0.67149 | 0.03715 |
| | cohort 1933-1935 | 0.63794 | 0.03384 |
| | cohort 1936-1938 | 0.63742 | 0.03057 |
| | cohort 1939-1941 | 0.60632 | 0.03114 |
| | cohort 1942-1944 | 0.65838 | 0.03028 |
| | cohort 1945-1947 | 0.63164 | 0.03356 |
| | cohort 1948-1950 | 0.56864 | 0.03854 |
| | cohort 1951-1953 | 0.53562 | 0.04506 |
| | cohort 1954-1956 | 0.51598 | 0.05174 |
| | cohort 1957-1959 | 0.5798 | 0.05963 |
| | cohort 1960-1962 | 0.50812 | 0.06751 |
| | cohort 1963-1965 | 0.40081 | 0.07498 |
| | cohort 1966-1968 | 0.50465 | 0.08356 |
| | cohort 1969-1971 | 0.50156 | 0.09001 |
| | cohort 1972-1974 | 0.52222 | 0.10059 |
| | cohort 1975-1977 | 0.52277 | 0.12804 |
| Vocational lower secondary education × | cohort 1903-1908 | 0.52683 | 0.10706 |
| | cohort 1909-1911 | 0.43077 | 0.08902 |
| | cohort 1912-1914 | 0.34072 | 0.07847 |
| | cohort 1915-1917 | 0.38576 | 0.07435 |
| | cohort 1918-1920 | 0.31771 | 0.06465 |
| | cohort 1921-1923 | 0.32774 | 0.0567 |
| | cohort 1924-1926 | 0.39001 | 0.04807 |
| | cohort 1927-1929 | 0.38538 | 0.04041 |
| | cohort 1930-1932 | 0.35044 | 0.03377 |
| | cohort 1933-1935 | 0.31888 | 0.02919 |
| | cohort 1936-1938 | 0.32502 | 0.02602 |
| | cohort 1939-1941 | 0.30817 | 0.02576 |
| | cohort 1942-1944 | 0.34435 | 0.02664 |
| | cohort 1945-1947 | 0.33661 | 0.0295 |
| | cohort 1948-1950 | 0.30067 | 0.03372 |
| | cohort 1951-1953 | 0.30649 | 0.03925 |
| | cohort 1954-1956 | 0.27512 | 0.04615 |
| | cohort 1957-1959 | 0.23128 | 0.05271 |
| | cohort 1960-1962 | 0.26338 | 0.06003 |
| | cohort 1963-1965 | 0.28808 | 0.06708 |
| | cohort 1966-1968 | 0.20279 | 0.07467 |
| | cohort 1969-1971 | 0.21897 | 0.08381 |
| | cohort 1972-1974 | 0.2218 | 0.0964 |
| | cohort 1975-1977 | 0.13839 | 0.11209 |

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| | | Coefficient | Standard-error |
|---|---|---|---|
| General lower secondary education × | cohort 1903-1908 | 0.48551 | 0.16436 |
| | cohort 1909-1911 | 0.51844 | 0.113 |
| | cohort 1912-1914 | 0.38922 | 0.09497 |
| | cohort 1915-1917 | 0.43514 | 0.09527 |
| | cohort 1918-1920 | 0.53869 | 0.06946 |
| | cohort 1921-1923 | 0.59943 | 0.05973 |
| | cohort 1924-1926 | 0.55672 | 0.05514 |
| | cohort 1927-1929 | 0.58391 | 0.04942 |
| | cohort 1930-1932 | 0.55715 | 0.04569 |
| | cohort 1933-1935 | 0.55848 | 0.04114 |
| | cohort 1936-1938 | 0.51152 | 0.03915 |
| | cohort 1939-1941 | 0.57927 | 0.03974 |
| | cohort 1942-1944 | 0.56049 | 0.03729 |
| | cohort 1945-1947 | 0.56478 | 0.03746 |
| | cohort 1948-1950 | 0.45991 | 0.04103 |
| | cohort 1951-1953 | 0.43696 | 0.04703 |
| | cohort 1954-1956 | 0.41605 | 0.05376 |
| | cohort 1957-1959 | 0.38419 | 0.05995 |
| | cohort 1960-1962 | 0.36589 | 0.06886 |
| | cohort 1963-1965 | 0.34144 | 0.07735 |
| | cohort 1966-1968 | 0.31887 | 0.09141 |
| | cohort 1969-1971 | 0.3862 | 0.1149 |
| | cohort 1972-1974 | 0.25528 | 0.12127 |
| | cohort 1975-1977 | 0.18722 | 0.14538 |
| Primary education × | cohort 1903-1908 | 0.17138 | 0.09765 |
| | cohort 1909-1911 | 0.08483 | 0.08412 |
| | cohort 1912-1914 | 0.12143 | 0.07472 |
| | cohort 1915-1917 | 0.04544 | 0.07036 |
| | cohort 1918-1920 | 0.13836 | 0.05796 |
| | cohort 1921-1923 | 0.16632 | 0.0505 |
| | cohort 1924-1926 | 0.1284 | 0.0437 |
| | cohort 1927-1929 | 0.1956 | 0.03771 |
| | cohort 1930-1932 | 0.17811 | 0.03275 |
| | cohort 1933-1935 | 0.17726 | 0.02868 |
| | cohort 1936-1938 | 0.19038 | 0.02618 |
| | cohort 1939-1941 | 0.18506 | 0.02556 |
| | cohort 1942-1944 | 0.20507 | 0.02746 |
| | cohort 1945-1947 | 0.21596 | 0.03033 |
| | cohort 1948-1950 | 0.1985 | 0.03508 |
| | cohort 1951-1953 | 0.17256 | 0.04162 |
| | cohort 1954-1956 | 0.13222 | 0.04978 |

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| | | Coefficient | Standard-error |
|---|---|---|---|
| | cohort 1957-1959 | 0.13845 | 0.06296 |
| | cohort 1960-1962 | 0.002561 | 0.09492 |
| | cohort 1963-1965 | 0.02179 | 0.12145 |
| | cohort 1966-1968 | 0.13773 | 0.12587 |
| | cohort 1969-1971 | 0.027 | 0.15483 |
| | cohort 1972-1974 | REF | REF |
| | cohort 1975-1977 | 0.25003 | 0.27188 |
| No degree × | cohort 1903-1908 | -0.30789 | 0.09562 |
| | cohort 1909-1911 | -0.21478 | 0.082 |
| | cohort 1912-1914 | -0.25026 | 0.07362 |
| | cohort 1915-1917 | -0.1957 | 0.06831 |
| | cohort 1918-1920 | -0.1328 | 0.05839 |
| | cohort 1921-1923 | -0.13952 | 0.05116 |
| | cohort 1924-1926 | -0.14278 | 0.04413 |
| | cohort 1927-1929 | -0.05048 | 0.03753 |
| | cohort 1930-1932 | -0.01186 | 0.03143 |
| | cohort 1933-1935 | -0.0195 | 0.02644 |
| | cohort 1936-1938 | -0.03573 | 0.02393 |
| | cohort 1942-1944 | 0.04328 | 0.02649 |
| | cohort 1945-1947 | 0.05283 | 0.03011 |
| | cohort 1948-1950 | 0.07432 | 0.03414 |
| | cohort 1951-1953 | 0.03682 | 0.04068 |
| | cohort 1954-1956 | 0.06587 | 0.04801 |
| | cohort 1957-1959 | 0.08896 | 0.05413 |
| | cohort 1960-1962 | 0.06238 | 0.06132 |
| | cohort 1963-1965 | 0.03734 | 0.0688 |
| | cohort 1966-1968 | 0.06437 | 0.0777 |
| | cohort 1969-1971 | -0.02133 | 0.0863 |
| | cohort 1972-1974 | 0.0776 | 0.09626 |
| | cohort 1975-1977 | -0.06945 | 0.11295 |
| Higher education × | age | 0.01171 | 0.0030438 |
| | $age^2$ | -0.0007823 | 0.0002258 |
| | $age^3$ | 0.00009903 | 0.00001326 |
| | $age^4$ | -3.31E-06 | 1.01E-06 |
| Upper secondary education× | age | 0.01238 | 0.0030364 |
| | $age^2$ | -0.0008048 | 0.0002229 |
| | $age^3$ | 0.0000373 | 0.00001137 |
| | $age^4$ | -9.86E-07 | 9.17E-07 |
| Vocational lower secondary education× | age | 0.0096945 | 0.0027407 |
| | $age^2$ | -0.000448 | 0.000169 |
| | $age^3$ | 0.00001352 | 7.71E-06 |

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| | | Coefficient | Standard-error |
|---|---|---|---|
| | $age^4$ | -1.03E-06 | 6.80E-07 |
| General lower secondary education× | age | 0.01222 | 0.0033687 |
| | $age^2$ | -0.0005944 | 0.0002856 |
| | $age^3$ | 0.00002086 | 0.00001465 |
| | $age^4$ | -6.88E-07 | 1.18E-06 |
| Primary education× | age | 0.01091 | 0.0027908 |
| | $age^2$ | -0.0005035 | 0.0001539 |
| | $age^3$ | 6.33E-06 | 8.46E-06 |
| | $age^4$ | -2.36E-07 | 6.35E-07 |
| No degree × | age | 0.0070362 | 0.002777 |
| | $age^2$ | -0.0005369 | 0.0001511 |
| | $age^3$ | 0.00001314 | 8.43E-06 |
| | $age^4$ | 2.04E-07 | 6.25E-07 |
| $\sigma$ | Normal | 0.4423 | 0.0014413 |
| Number of Observations | | 48245 | |
| Noncensored Values | | 41394 | |
| Log Likelihood | | -36966.33378 | |