Projet ANR-14-CE22-0006

ELITISME

ELITISME WORKING PAPER SERIES

ANR-ELITISME-2015-006

Equity of European education policies

Sophie Dantan

PROJET FINANCÉ PAR L'ANR - PROJECT FUNDED BY THE ANR

ANR

Labex MME-DII
Modèles Mathématiques et Économiques de la Dynamique, de l'Incertitude et des Interactions

UNIVERSITÉ de Cergy-Pontoise

ÉCOLE POLYTECHNIQUE UNIVERSITÉ PARIS-SACLAY

ENS CACHAN

université PARIS-SACLAY

Equity of European education policies.

Sophie DANTAN[1]

Abstract

This article analyses and compares patterns of intergenerational education mobility across countries in order to determine which ones implement the most efficient policy against intergenerational persistence. To explain the differences between national patterns, variations of national policies across cohorts and countries are exploited to evaluate their effect on the transmission of educational inequalities. In particular, the effect policies devoted to early grades of education are distinguished from the effect of education policies devoted to higher grades. This strategy allows investigating the relevance of the timing of education policies. Results suggest that investments in early grade are more efficient in reducing educational inequalities.

Keywords: intergenerational educational mobility; educational policies.

[1] This research was achieved at *THEMA, Université de Cergy Pontoise*

Introduction

The intergenerational transmission of education can be thought as the result of the transmission of endowments in human capital and/or as the result of parental and public investments in children's human capital. If the inheritance of endowment in human capital is the main channel of transmission of education, then the educational policies should be inefficient in reducing the education persistence. If on the contrary, investments in education play the greater part in the education transmission, then educational policies can significantly enhance education mobility. In the latter case, the efficiency of the policy measures might however depend on the period of the life-cycle at which they are implemented. It is indeed likely that inequalities in cognitive skills acquired in primary and secondary educations translate into inequalities in achieving post-secondary education and into wages via the returns to education in the labour market. In the terminology of Cameron and Heckman (1998, 2001), the effect of parental background on the ability to acquire skills is called "long-term" effect of parental background, by opposition to a "short-term" effect which is linked to parents' financial capacities to pay for high education when credit market is imperfect. Assuming that the transmission of parents' endowment in human capital is not the main channel of education transmission, the efficiency of the public policy in mitigating this transmission strongly depends on the extent to which persistence in education is due to short-term or long-term effects.

This question is crucial to evaluate the efficiency of education policies. Indeed, if persistence in education is due to credit constraints on tertiary schooling investment, then public policies should aim at reducing them, for instance by subsidising able students from poor families. If on the contrary, the educational persistence is not due to this short-term effect but rather to the long-term effect of parental background, reducing credit constraints on college attendance is inefficient. In that case, implementing human capital-enhancing policies at the earliest stages of education would have a better and longer-lasting effect on the education mobility.

Evaluating the respective relevance of the long-term and short-term policies then constitute an important step on the path to improve intergenerational mobility but also a way for researchers to determine whether the observed correlation between parents' and offspring's correlation is due to the fact that poor parents transmit a lower endowment in human capital and have then no interest in investing in their children's – in that case, neither long-term nor short-term policies would impact the educational mobility- or to the fact that they do not invest enough in their children's education due to credit constraints – in that case, the timing of the efficient educational policy indicates at which step of the school life the constraints are the most binding.

From the European micro-survey EU-SILC, this paper measures and compares the persistence in education across some European countries and exploit variations (across time and countries) in national education policies to explain the observed differences.

The first section reviews the main findings and recommendations of the economic literature to enhance equity of educational opportunities. The second section describes the empirical strategy implemented to measure the effect of parental background on education. The third section presents and compares different measures of persistence in education across the considered European countries. The fourth section aims at explaining the observed differences by comparing the efficiency of national education policies. The last section concludes.

1. Empirical evidence on the impact on education policy

1.1. International comparison

The comparison of intergenerational mobility in education reveals strong discrepancies across countries and across time. Among the numerous authors who compare the educational mobility, Hertz *et al.* (2007) achieve a remarkable research by examining the intergenerational persistence in education across 42 countries from 1927 to 1972. They find a global decrease in the education immobility measured by the regression coefficients of offspring's education length on the parental one, whereas the correlation between both lengths has remained relatively stable. This suggests that the observed decrease in the global intergenerational persistence in education is mainly due to a decrease in the variance of schooling attainment from a generation to the next, rather than to a decrease in the intergenerational correlation of parent's and offspring's education. This is not however true in all countries: in most of Nordic countries, Ireland and Switzerland, both correlation and regression coefficient have fallen whereas they both rise in United Kingdom; in Italy, Norway and Sweden, the regression coefficient decreased whereas the correlation remain stable in Ireland and Eastern European countries. On average, the mobility is found to be higher in developed countries and, more particularly, in Nordic countries: their intergenerational correlation hardly reaches 30% against 41% in Eastern European countries.

Consistently, Couch and Dunn (1997) find a stronger correlation between father's and son's and between mother's and daughter's educational attainments in the US than in Germany. Gantzeboom and Nieuwbeerta (1999) find a decrease in education mobility (measured by the regression coefficient) in six Eastern European countries from 1940 to 1985 which appears to be stronger than in other European countries, probably because it was exacerbated during communism. By contrast, Blanden and Machin (2004) show that the effect of parental income on child's education has increased between 1950 and 1970 in UK but stabilized when a reform was introduced to enlarge the access to tertiary education. Checchi *et al.* (2008) regress the offspring's education length on their parents' from an Italian panel by controlling for various individual characteristics and find that persistence in education significantly decreased at the beginning of the 20th century.

Lefranc *et al.* (2008) find a lower intergenerational correlation of father's and offspring's education in Japan than in France which they attribute to a more elitist (although publicly-funded) tertiary educational system in France. Parental socio-economic background is also found to impact significantly the PISA tests scores, in particular in Austria, Czech Republic, France and United Kingdom (OECD 2010, Causa and Chapuis 2009).

Other authors compute indexes of educational mobility built on educational mobility matrices. Among them, authors generally distinguish indexes of movement - measure how far the offspring's education is from the parents' - from indexes of equity of opportunity - which measure to what extent the actual situation differ from equity of opportunities. Whatever the chosen type of index, Sweden, Denmark and Germany are found to be mobile countries in contrast to UK, Italy and Ireland (de Broucker and Underwood 1998, Foley 2006).

Movement and equality of opportunity indexes are not necessarily correlated. For instance, Chevalier (2004) examines indexes from the education duration surveyed in 20 countries and conclude that educational mobility is higher in countries with lower earning returns to education and higher public education expenditures. Chevalier *et al.* (2009) observe a cross-

country increase in the equity of opportunity but not of the movement. In the same way, by computing several educational mobility indexes, Checchi *et al.* (1999) find the intergenerational education persistence to be higher in Italy than in the US despite its more centralized and publicly-funded school system, except when considering a movement index. Their explanation is that this movement index mirrors the structural post-war rise in education which occurred in Italy but not in the US.

1.2. Effect of education at different stages

According to the seminal model of human capital developed by Becker and Tomes (1979), the educational attainment implicitly reflects the level of human capital and can then be modelled as a function of the private and public investments in education and of the child's own endowment in human capital. The own endowment in human capital represents the part of human capital inherited from parents but not transmitted through financial channels. It comprise genetic endowment but also non-cognitive skills which have strong effect on the child's social insertion. In their model, Becker and Tomes (1979) assume that parents can borrow to fund their children's education and so the persistence of educational inequalities can only be explained by the fact that poor children are less endowed with human capital and thus do not have to invest more in education. Their finding that public policy is inefficient in enhancing the intergenerational mobility, is challenged by recent advances in the modelling of human capital transmission.

Solon (2004) extends their model, relaxes the borrowing hypothesis and shows that redistributive policy can actually improve social mobility. The observed differences of educational persistence across countries can thus be explained by differences in the progressivity of the redistributive policies. Moreover, in Becker and Tomes (1979), the production of human capital is assumed to depend on the cumulated investment made during childhood, with no distinction made between investments in primary, secondary or tertiary schools: these subsequent investments are assumed to be perfect substitutes. Carneiro and Heckman (2002) propose an alternative modelling in which the human capital is produced following a self-productive and complementary process of formation of human capital: a higher level of human capital cumulated at the first stage of the life-cycle induces a higher acquired human capital and a higher return of the investment in education at the following stages.

Consistently with the theory of human capital formation developed by Carneiro and Heckman (2002) several empirical studies have emphasized the positive effect of pre-school attendance on the educational and social achievement of children, from national surveys, (see Dumas and Lefranc 2012 in France; Berlinski et al (2008) in Uruguay; Goodman and Sianesi 2005 in United Kingdom; Magnuson *et al.* 2007 for US). Havnes and Mogstad 2009 in Norway; Bauer and Riphahn 2009 in Switzerland and Leuven *et al.* 2006 for Netherlands). Except Goodman and Sianesi (2005), all authors found that this positive effect is stronger and longer-lasting for children of disadvantaged families. This effect of early education in reducing social inequalities was confirmed by cross-country evidence (Braga *et al.* 2011; Schütz *et al.* 2005; OECD 2010).

The positive effect of increasing the minimum school exit age on social achievement was also demonstrated by exploiting its historical variations in United States and Canada (Goldin and Katz 2003; Oreopoulos 2007) and in some continental European (see for instance Grenet (2013) for France; Pekkarinnen (2005) for Finland; Pischke and Watcher (2005) for Germany; Plug (2001) for Netherlands; Meghir and Palme (2005) for Sweden). Those national results

obtained from weak variations of the minimum exit age are confirmed by cross-country estimations which exploit variations of the exit age both within and across European countries (Murtin and Viarengo 2011, Brunello *et al.* 2009). Interestingly, Brunello *et al.* (2009) find that an increase in exit age benefits more to the less able individuals (those in the bottom deciles of the school duration distribution) in term of education duration than to the more able ones.

Reforms of the compulsory school may even have had an intergenerational effect on educational achievement by influencing parental education. Indeed, to determine whether the intergenerational transmission of human capital occur through a genetic channel, several authors use the reform of exit age as instrument in the regression of the offspring's education on the parental one, with contrasted results (Chevalier 2004, Chevalier *et al.* 2005 and Oreopoulos *et al.* 2006 find a positive effect of parental education on achievement while Black *et al.* 2005 find none). Nevertheless, Holmlund *et al.* (2008) and Pronzato (2009) demonstrate that the mother's education is a more important determinant of children's achievement in the culturally-poorer families whereas the father's one is more important for the achievement of children from culturally-richer families.

Turning to the effect of other policy, several empirical studies suggest a significant effect of reducing the class size on achievement and equity (Angrist and Lavy 1999 in Israel, Wössman 2005 from a cross-country study, Leuven *et al.* (2008) in Netherlands; Browning and Heinesen 2007 and Bingley *et al.* 2005 in Denmark; Bonesroning 2003 in Norway; Piketty 2004, Gary-Bobo and Mahjoub 2010, Bressoux *et al.* 2009).Empirical results differ by their significance, which can be attributed to national specificities (notably the already low number of pupils per class in Nordic countries), but also to the hypothesis that decreasing the class size can affect the educational achievement and enhance educational mobility but only at the lowest stages of schooling (consistently with Gary-Bobo and Mahjoub 2010).

The effect of easing the access to tertiary education is more controversial. While Kane (1995) finds that the enrolment gap between low-income and high-income children increased more in states with the strongest increases of tuition fees, other studies about reform of the tuition scheme in the US (in particular of the PELL grant) have led to contradictory conclusions (see Kane 2001 for a review). More generally, the idea that the lower enrolment of low-income high-school completers reflect credit constraint is challenged by studies which estimate the effect of parental income on college attendance by controlling for the students' ability. Results obtained by Cameron and Heckman (2002) Carneiro and Heckman (2002) Heckman and Cameron (1998, 2001) and Belzil and Hansen (2003) and Dearden *et al.* (2004) suggests that *the long-term effect* of parental background on the cognitive and non-cognitive skills explain better the poor students' lower college attendance than the short-term effect of parental background on the affordability of college.

Eckstein and Wolpin (1999) find a weak effect of working during attendance to college on the probability to drop out when including in their model a latent heterogeneity to capture unobservable variables. They demonstrate that the value of latent variable is positively correlated to motivation and ability. Consistently, Keane and Wolpin (2001) estimate a joint model of schooling, working and saving decisions and conclude that relaxing borrowing constaints woud have little effect on schooling decisions although it consistutes an important deteminant of working during school. However, as mentioned by Belley and Lochner (2007), those latter estimations (as those achieved by Heckman and coauthors) are obtained from the same old data (the NLSY79) so that the moderate effect of credit constraints on college entry might no longer be true. They notably underline a larger effect of the family income in 1997 than in 1979 and find that this increase cannot be explained without appealing to borrowing constraints. Cross-country comparisons of college reforms also suggest that implementing

reforms which aim at expanding the university access, increasing grant and enhancing the autonomy of universities tend to enhance equity (Braga *et al.* 2011; Oliveira *et al.* 2009).

2. Empirical strategy

2.1. An ordered model of educational choice

Following Cameron and Heckman (1998), this section presents an ordered discrete educational choice model. Assume that the utility provided by the choice of a grade l (l=1 to $L^{\max}$) is the difference between the return to investment R_{il} and a fixed cost of investment c_l at this stage:

$$U_{il} = R_{il} - c_l \tag{1}$$

Under the assumption that both the return and the cost increase with the grade l and that R_{il} is a concave and c_l is a convex function of this grade, they show that utility is maximized by only one level of education l^*. The chosen level of investment l^* verifies the properties:

$$R_{il^*} - c_{l^*} \geq R_{i,l^*-1} - c_{l^*-1} \text{ and } R_{il^*} - c_{l^*} \geq R_{i,l^*+1} - c_{l^*+1} \tag{2}$$

Moreover, assume that the benefit R_{il} can be written as the product of fixed cost R_l by a function φ_i of the family characteristics:

$$R_{i,l} = R_l \times \varphi_i \tag{3}$$

Thus, a level l of educational attainment is chosen if and only if :

$$R_l.\varphi_i - c_l \geq R_{l-1}.\varphi_i - c_{l-1} \text{ and } R_l.\varphi_i - c_l \geq R_{l+1}.\varphi_i - c_{l+1} \tag{4}$$

$$\Leftrightarrow \frac{c_l - c_{l-1}}{R_l - R_{l-1}} \leq \varphi_i \leq \frac{c_{l+1} - c_l}{R_{l+1} - R_l} \tag{5}$$

φ_i can be interpreted as the family-specific marginal return to achieving studies. Each family i chooses the optimal level l so that φ_i is higher than the net marginal cost of achieving the grade l but lower than the marginal net cost of achieving *l-1*.

Thus the probability of achieving a level of education l as

$$P_{il} = \Pr(T_l < \varphi_i \leq T_{l+1}) = \Pr(\varphi_i \leq T_{l+1}) - \Pr(\varphi_i \leq T_l) \tag{6}$$

$$\text{where } T_l = \frac{c_l - c_{l-1}}{R_l - R_{l-1}} \; for \; 1 < l < L^{\max}, \quad T_1 = -\infty$$

Assuming a normal distribution of the marginal return φ_i, Cameron and Heckman (1998) estimated an ordered probit model of educational attainment. To estimate the effects of parental education on a child's own education while controlling for other characteristic of family background, this paper follows Cameron and Heckman (1998) and estimate an ordered probit of educational choice. To do so, φ_i is assumed to be a measure of the propensity to achieve studies and to be split into a deterministic function of the father's education $E_{i,t-1}$ and other individual characteristics X_i and into a normally-distributed error-term ε_i.

$$\varphi_i = X_i.\eta + \lambda.E_{i,t-1} + \varepsilon_i \tag{7}$$

Under these assumptions, the probability that an individual achieves a certain level of education, given the conditioning variables ofhis/her propensity to achieve higher education, can be estimated by maximum likelihood as:

$$\Pr(E_i = low) = \Pr(\varphi_i \le T_{medium})$$
$$\Pr(E_i = medium) = \Pr(T_{medium} < \varphi_i \le T_{high}) \quad (8)$$
$$\Pr(E_i = high) = \Pr(T_{high} < \varphi_i),$$

where T are the cut off points at which the probability of achieving each educational outcome changes. Ordered probit estimation allows to compute the percentage change in the probability of achieving each educational level induced by a change in paternal education (or other conditioning variables subsumed in φ_i) at a country specific mean point, that is to say for an individual whose control variables take the country-specific average values.

The father's education is coded as a discrete variable taking 3 values "low", "medium" and "high". "medium" is normalized as the reference value so that the marginal effect corresponding to "low" and "high" can be interpreted as a penalty of having a low-educated father and a premium of having a high-educated father, respectively.

2.2. Effect of national policies

In order to assess the role of public expenditures devoted to education in explaining intergenerational education persistence, the ordered probit can be estimated by including interactions betwwen parental education and national policies among among the explanatory variables. The variation in policies across countries and cohorts, cohorts being defined as the set of individual of the same age insuch estimations. To explore whether certain policies or institutions mitigate or reinforce educational persistence, policy and institutional variables are interacted with parental background in a logit[2] model that allows heterogeneity across countries and over time (cohorts) to be taken into account. More specifically, the latent propensity to achieve education is defined as follow:

$$\varphi_i = \lambda \cdot E^f_{iac} + \delta \cdot P_{iac} \cdot E^f_{iac} + \phi \cdot P_{iac} + Q_{c\cdot} + Z_a + H_{ac} + \varepsilon_{iac} \quad (9)$$

where a, c, and i index time (cohort), country and individual respectively; P denotes the policy or institutional variable; Q denotes country fixed effects; Z denotes time fixed effects; H denotes interactions between country and the time (time being a continuous variable); φ is the propensity to achieve studies, E^f is the paternal education and the residuals ε_{iac} are logistically-distributed. Migrants are excluded from the sample and control variables (family type, number of siblings) are excluded from the estimation since they are missing in the Swedish sample. Moreover, migrants are more or less likely to be impacted by educational policies depending on their date of arrival, a variable which is not available in our dataset. Robust standard errors account for the clustering of individuals in country-cohort classes.

The analysis exploits family, country and cohort variations in educational achievement, parental background and variation in policies to determine whether a policy exacerbate or mitigate the persistence in education. Since the specification includes country-specific cohort effects, it cannot identify the direct impact of policies on the dependent variable but enable to

[2] Considering a logit rather than a probit model enables a more easy interpretation of the coefficient.

control for the various changes in national characteristics which influence the education attaiment.

In contrast to a standard regression, a logit estimation does not provide easily interpretable coefficients and the computation of a marginal effect is complicated by the fact that some variables are interacted. Actually, the obtained coefficients measure the effect of variables on the log odd ratios of achieving certain levels of education. For instance , in order to analyse the effect of having a high-educated father on the odd ratio of achieving a high rather than a lower education, one should consider the following equation:

$$\frac{\dfrac{P(E=high\,|\,E_{iac}^{f}=high)}{P(E\neq high\,|\,E_{iac}^{f}=high)}}{\dfrac{P(E=high\,|\,E_{iac}^{f}=medium)}{P(E\neq high\,|\,E_{iac}^{f}=medium)}}=\exp(\lambda_{high}+\delta^{high}\cdot P_{iac})$$

If having a high-educated father increase the propensity to achieve tertiary education ($\lambda_{high}+\delta^{high}\cdot P_{iac}>0$) then this ratio exceeds one, which means that when having a high-educated father, the odd of achieving tertiary education is greater than when having a medium-educated father. Another interpretation of this ratio can be obtained by rewriting it as follows:

$$\frac{\dfrac{P(E=high\,|\,E_{iac}^{f}=high)}{P(E=high\,|\,E_{iac}^{f}=medium)}}{\dfrac{P(E\neq high\,|\,E_{iac}^{f}=high)}{P(E\neq high\,|\,E_{iac}^{f}=medium)}}=\exp(\lambda_{high}+\delta^{high}\cdot P_{iac})$$

According to this equation, if having a high-educated father rather than a medium-educated one increase the propensity to achieve high education ($\exp(\lambda_{high}+\delta^{high}\cdot P_{iac})>1$), then the risk ratio of being graduated with a high- rather than medium-educated father exceeds the risk ratio of having no degree.

The formula of the odd ratio illustrates that a policy variable which mitigates the premium of having a high-educated father ($\delta^{high}<0$) will decrease the odd ratio and enhance the education mobility. On the contrary, policies which reinforce the premium are found to exacerbate the transmission of education inequalities. More precisely, a one-unit increase (decrease) in the variable policy will decrease (increase) the odd ratio by δ^{high} percentage point.

The odd ratio of achieving tertiary education with a low-educated father rather than a medium-educated one can be written in a similar fashion by replacing the effect of having a high-educated father $\lambda_{high}+\delta^{high}\cdot P_{iac}$ by the effect of having a low-educated one $\lambda_{low}+\delta^{low}\cdot P_{iac}$ in previous equations. If having a low-educated father decreases the propensity to achieve studies, then this ratio is lower than one but can be increased (decreased) by policies which reduce (exacerbate) the penalty of having alow-educated father. The magnitude of the increase (decrease) attain δ^{low} percentage point for one-unit increase of the policy variable.

The total effect of a policy on the education mobility can finally be measured by considering the odd ratio of achieving tertiary education with a high-educated father rather than a low-educated one:

$$\log\left(\frac{\frac{P(E=high\,|\,E_{iac}=high)}{P(E\neq high\,|\,E_{iac}=high)}}{\frac{P(E=high\,|\,E_{iac}=low)}{P(E\neq high\,|\,E_{iac}=low)}}\right)=(\lambda_{high}-\lambda_{low})+(\delta^{high}-\delta^{low})\cdot P_{iac} \qquad (10)$$

According to this formula, the effect of a policy variable on this odd ratio - and then on the education intergenerational mobility - corresponds to the gap $(\delta^{high}-\delta^{low})$ between its effect on the premium of having a high-educated father and its effect on the penalty of having a low-educated father. If this gap is positive then the policy exacerbates the peristence in education, if it is negative then the policy mitigate this persistence.

3. Persistence in education across Europe

3.1. Data

The empirical analysis relies on the recently released European Union Statistics on Income and Living Conditions (EU-SILC) household database. EU-SILC is the first international micro dataset that allows quantitative analysis of the transmission of both social advantage and disadvantage across European countries in a comparable way. The analysis focuses on the 2005 EU-SILC cross-section, which includes a specific module on intergenerational transmission of poverty. Exploiting the longitudinal dimension of the data would require a longer time span than is currently available in the EU-SILC data (2004-2006). The poverty module allows linking economic and educational outcomes of successive cohorts of individuals to their family's socio-economic background. Individuals aged between 24 and 64 from a sample of the households covered by the EU-SILC survey are asked retrospective information on their socio-economic situation as teenagers. Such retrospective information covers family composition, age, educational levels, activity status and occupation of parents. However, information on parents' wages is lacking.

The sample used in the analysis consists of individuals aged 25 to 54 and is divided into three successive age groups or cohorts: 25-34, 35-44 and 45-54 years of age. Most results of the country-by-country analysis are presented for the 35-44 year old cohort in order to reduce life-cycle measurement error in individuals' economic outcomes, as said before. The results of the youngest cohort (25-34 years old) have to be interpreted carefully because of the potential difficulties associated with measuring permanent wages/incomes at young ages and because of the inability to ensure that the individuals under consideration have reached their desired educational level. For the same reason, individuals that are selected to respond to the intergenerational module, but are reported to still be in education are dropped from the analysis. However, including them has little impact on the results. Conversely, the variation across cohorts is used to identify the association between policies and wage persistence in a panel data setting. The remaining sample consists in a little bit more than 85000 individuals unequally distributed across countries. Italian, Polish, Spanish and French people represent

more than a third of the sample whereas each other countries represent about 3% of the sample. This unbalance is accounted for in cross-country estimations by reweighting individuals. The distribution of age is more balanced across countries.

Table 1: Sample distribution by country and cohort

countries	#observations	% of the total sample	24-34	35-44	45-54
Austria	3,190	3,73	29,36	40,88	29,75
Belgium	3,006	3,52	32,74	37,34	29,92
Cyprus	2,878	3,37	36,33	34,87	28,81
Czech Republic	2,598	3,04	36,00	30,25	33,75
Denmark	2,199	2,57	28,96	39,17	31,88
Estonia	2,349	2,75	29,76	34,01	36,24
Finland	2,566	3,00	29,92	35,21	34,87
France	5,795	6,78	32,92	35,86	31,22
Greece	2,631	3,08	36,74	37,24	26,02
Hungary	3,599	4,21	34,21	30,90	34,89
Iceland	883	1,03	32,44	37,98	29,58
Ireland	1,867	2,19	37,82	28,15	34,03
Italy	11,389	13,33	32,20	38,40	29,39
Lithuania	2,520	2,95	29,50	37,55	32,95
Luxembourg	2,554	2,99	30,91	38,80	30,30
Netherlands	2,548	2,98	31,66	35,85	32,50
Norway	2,119	2,48	30,20	36,95	32,85
Poland	8,819	10,33	37,77	31,59	30,64
Portugal	2,516	2,95	39,07	35,39	25,54
Slovakia	4,297	5,03	31,37	32,81	35,82
Slovenia	2,381	2,79	30,07	36,90	33,02
Spain	7,162	8,39	39,98	34,64	25,38
Sweden	1,880	2,20	36,07	33,37	30,56
UK	3,668	4,29	28,91	36,50	34,58
Total	*85,414*	*100*	*33,87*	*35,55*	*30,58*

Source: Author's computations from SILC data

Education is defined, for both parents and offspring, as the highest International Standard Classification of Education (ISCED) level attained by the individual. For the purpose of estimation, the five ISCED categories may be aggregated into three modalities: (1) low education (pre-primary, primary, lower secondary, i.e. ISCED 0-2); (2) medium education (upper-secondary and post-secondary i.e. ISCED 3, 4); and (3) higher education (tertiary i.e. ISCED 5, 6).

The distribution of fathers' education varies significantly across countries and cohorts (see Table 2). Whatever, the considered cohort, the Eastern countries are characterized by a low proportion of low-educated father and a strong proportion of father with medium education. The proportion of low-educated fathers in Denmark, Finland, Iceland, and Norway seems higher than in Eastern countries but significantly lower than in Central and Southern Europe. The Nordic countries, Belgium, Estonia and UK exhibit particularly high proportion of high-educated fathers compared to other countries. Consistently with previous literature, the

educational attainment exhibits an increasing pattern as shown by the increase (decrease) in the proportion of high- (low-) educated fathers.

Table 2: Aggregated father's educational attainment by cohort

cohort country\father's education	25-34			35-44			45-54		
	low	medium	high	low	medium	high	low	medium	high
Austria	49,53	46,73	3,74	57,29	38,64	4,08	60,90	35,11	3,99
Belgium	49,08	26,23	24,70	61,21	23,09	15,70	68,68	18,63	12,69
Cyprus	56,74	28,55	14,71	75,35	18,20	6,45	88,01	7,88	4,11
Czech Republic	6,71	81,78	11,51	15,00	75,94	9,06	25,62	68,40	5,98
Denmark	30,70	46,68	22,62	42,54	42,32	15,14	47,48	42,48	10,04
Estonia	16,45	52,00	31,55	38,86	42,81	18,33	55,33	32,60	12,07
Finland	41,49	29,94	28,57	61,70	18,24	20,07	80,05	11,09	8,86
France	76,13	10,22	13,65	85,73	6,67	7,60	90,79	4,63	4,57
Greece	72,93	16,68	10,38	81,63	9,75	8,62	89,43	6,82	3,75
Hungary	20,42	65,88	13,70	42,31	48,97	8,72	55,27	37,14	7,60
Iceland	30,87	49,00	20,13	34,74	50,83	14,44	40,61	49,64	9,75
Ireland	69,75	13,40	16,85	78,35	13,10	8,55	87,95	7,89	4,17
Italy	77,82	17,41	4,78	85,09	10,89	4,02	86,09	10,17	3,74
Lithuania	28,55	51,14	20,31	60,92	27,64	11,44	77,84	16,04	6,12
Luxembourg	50,72	31,08	18,20	61,36	29,35	9,29	64,09	26,50	9,41
Netherlands	51,04	22,42	26,54	61,09	18,43	20,48	71,42	17,12	11,46
Norway	22,97	46,09	30,94	31,55	46,49	21,97	37,50	44,68	17,82
Poland	22,60	66,81	10,59	46,39	47,42	6,19	62,88	32,30	4,82
Portugal	89,42	5,19	5,40	94,07	2,53	3,39	94,58	3,22	2,21
Slovakia	11,57	76,33	12,10	29,19	61,01	9,80	51,98	40,75	7,26
Slovenia	40,16	54,72	5,12	57,21	39,83	2,96	60,48	35,67	3,85
Spain	80,43	7,72	11,85	84,35	5,28	10,37	88,69	3,82	7,49
Sweden	58,25	19,31	22,44	73,31	11,40	15,29	79,80	9,84	10,36
UK	50,98	24,07	24,95	59,52	24,72	15,76	69,37	19,47	11,16
Total	*59,1*	*27,47*	*13,44*	*70,23*	*20,51*	*9,26*	*75,97*	*17,47*	*6,56*

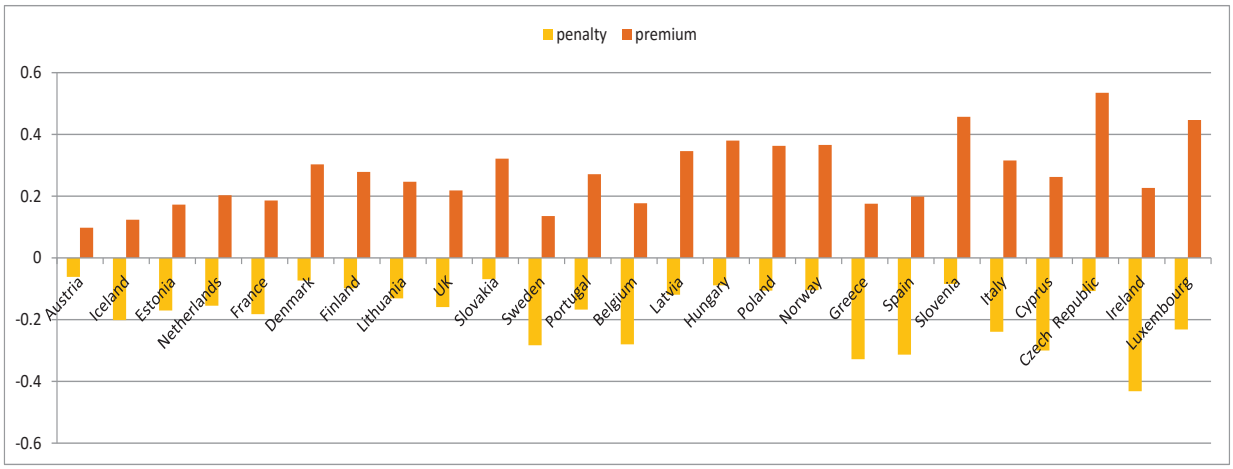
Source: Author's computations from SILC data

3.2. Country-by-country estimations

Figure 1 and Figure 2 report the marginal effects associated with changes in a father's education on the probability of the child to achieve tertiary education for the 35-44 year old cohort.

For pairs of fathers and sons, the estimated premium is particularly large in Luxembourg, Czech Republic, Slovenia, Latvia, Hungary, Poland where the probability of achieving tertiary education is more than 30 percentage points higher for a son whose father had tertiary education, compared to one whose father had upper-secondary education. The penalty of coming from a low-educated family is particularly high in Belgium, Ireland, Spain, Cyprus and Greece where the probability for a son of low-educated father to obtain a degree is more than 20 percentage point lower, compared to a son of medium-educated father. The cross-country pattern in the estimated probability premium for 35-44 year old women is quite similar from that of men. Nevertheless, women's probability premium is significantly lower in all countries except in Iceland, Estonia, Netherlands, Finland, UK, Norway, Ireland, and Luxembourg. Similarly, the estimated penalty of having a low-educated father is higher than for men in some countries, in particular Portugal, Greece, Czech Republic, Slovenia, Latvia, Estonia but also Norway and Denmark.

Figure 1: Probability premium and penalty of achieving tertiary education due to father's education levels, men 35-44[3]



Source: Author's estimations from SILC data

Figure 2: Probability premium and penalty of achieving tertiary education due to father's education levels, women 35-44

Source: Author's estimations from SILC data

Persistence in tertiary education and in below-secondary education can be summarized by measuring the difference in the estimated probabilities to achieve tertiary and below-secondary education respectively (Figure 3). A larger gap in either measure implies that

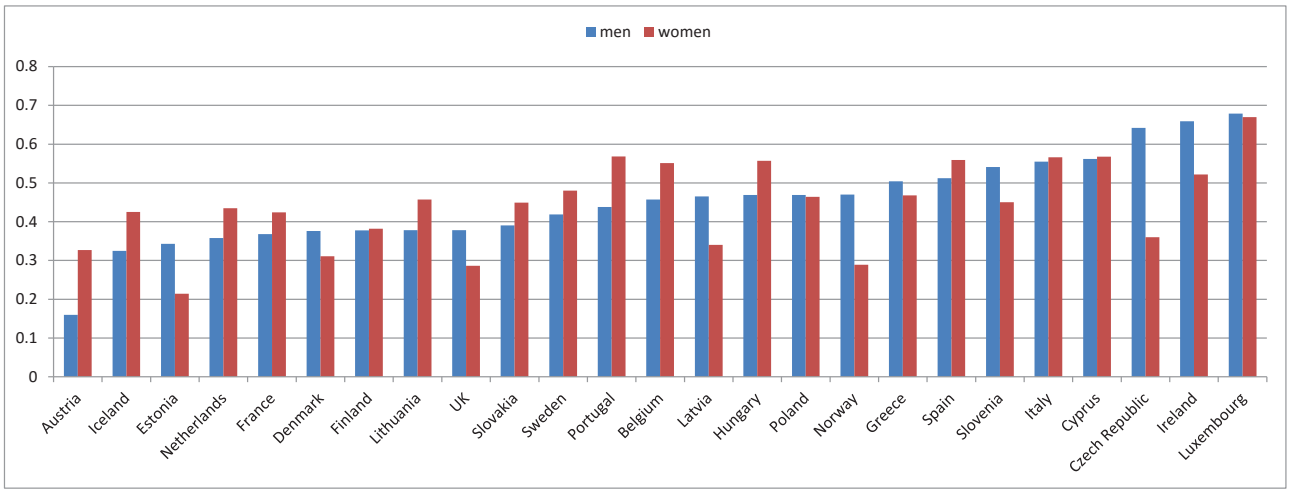
[3] The figures show the percentage point changes in the probability of an offspring to achieve tertiary education depending on the offspring's parental background. The probability premium is the increase of an offspring to achieve tertiary education given that his/her father had tertiary education relative to an offsring whose father had upper-secondary education. The probability penalty is the decrease in the probability of an offspring to achieve tertiary education given that his/her father had lower-secondary education relative to an offspring whose father had upper-secondary education.

Based on the ordered probit estimation of individuals' educational attainment. Marginal effects reported.

For Sweden, the number of siblings and household composition are missing so that they are not included in the regression.

fathers' education more strongly influences that of individuals and, therefore, indicates stronger persistence in education across generations.

Figure 3: Summary measure of wage persistence, 35-44



Source: Author's estimations from SILC data

1. Persistence in tertiary education is measured as the distance between the estimated probability premium and penalty. Thus, it measures the percentage point increase in the probability of an offspring having a father with tertiary education to achieve tertiary education relative to an offspring having a father with below-upper secondary education. A larger number implies a larger gap, thus stronger persistence in tertiary education or a higher degree of educational immobility across generations.
Based on ordered probit estimation of individuals' educational attainment. Marginal fixed effects reported.

A drawback of comparing the marginal effects (probability penalty and premium) across countries is that these marginal effects are computed at a country-specific "mean point" – as explained above, that is to say for an individual whose characteristics (number of siblings, family composition) take country-specific average values. To ensure that the summary measure does not reflect the differences in national mean points but the degree transmission of education, its values are re-estimated at a common "mean point" - that of the Denmark. This correction has no effect on the ranking of countries by the summary measure.

4. Educational mobility and national policies

4.1. School expenditures

The first specification of Equation 9 concerns the effect of the expenditures devoted to education per pupils in points of GDP per capita. Expenditures devoted to the pre-primary and primary educations are found to have no effect on the premium of having a high-educated father but seems to significantly mitigate the penalty of having a low-educated father (Table 3 and Table 4).

Consistently with prior findings that early intervention benefits more to worse-off sons, increasing the spending in each pupil's pre-primary education by one percentage point of GDP per capita is found to decrease the odd ratio of Equation 10 by 5.4% against 2.1 to 2.4% for a similar increase in primary education. This conclusion holds for girls: their odd ratio decreases by 5.4% when one more GDP point is invested in early education, against 1.8% when investing in primary education.

The effects of education expenditures in secondary and tertiary education are on the contrary very low or insignificant which confirms that investing in earlier stages of education enhance the education mobility more significantly than investing in the latter stages. Controlling for the enrolment rate at each stage slightly modifies the coefficients without however modifying the declining pattern of the policy effect on the education mobility along the school cycle.

Table 3: The effect of education spending per pupil in % of GDP per capita, men

	preprimary_habca	primary_habcary	young_habary	young_hatmary	old_habcmary	old_habcrimary_habca	rimary_habc	
X=	pre-primary education		primary education		primary education		primary education	
when the offspring was	*3 to 5*		*6 to 8*		*9 to 10*		*6 to 10*	
low-educated father	-1.862***	-2.373***	-1.395***	4.658***	-1.479***	2.649**	-1.459***	4.749***
	(0.204)	(0.324)	(0.119)	(1.161)	(0.113)	(1.123)	(0.116)	(1.252)
high-educated father	1.197***	1.526***	1.311***	2.311	0.926***	1.462	1.212***	2.680*
	(0.244)	(0.383)	(0.165)	(1.435)	(0.161)	(1.361)	(0.158)	(1.527)
low-educated father*expense per pupil devoted to X	0.074***	0.061***	0.021***	0.008*	0.024***	0.015***	0.022***	0.009*
	(0.018)	(0.017)	(0.005)	(0.005)	(0.004)	(0.005)	(0.005)	(0.005)
high-educated father*expense per pupil devoted to X	0.005	0.017	-0.002	-0.000	0.013**	0.014*	0.001	0.002
	(0.022)	(0.022)	(0.007)	(0.008)	(0.006)	(0.007)	(0.006)	(0.007)
low-educated father*enrollment rate in X		0.009**		-0.057***		-0.039***		-0.059***
		(0.004)		(0.011)		(0.011)		(0.012)
high-educated father*enrollment rate in X		-0.006		-0.010		-0.006		-0.015
		(0.004)		(0.013)		(0.013)		(0.014)
T(medium)	-1.921***	-1.735***	-0.094	-0.943***	0.548**	0.090	-0.117	-0.992***
	(0.287)	(0.287)	(0.277)	(0.310)	(0.272)	(0.301)	(0.277)	(0.310)
T(high)	0.747***	0.947***	2.489***	1.649***	3.100***	2.646***	2.438***	1.572***
	(0.270)	(0.268)	(0.279)	(0.315)	(0.277)	(0.310)	(0.281)	(0.316)
Observations	10,373	10,373	20,987	20,987	23,925	23,925	22,851	22,851
N_clust	98	98	240	240	271	271	252	252
r2_p	0.178	0.180	0.160	0.162	0.157	0.158	0.160	0.162

X=	secondary education		tertiary education		tertiary education	
when the offspring was	*12 to 15*		*19 to 22*		*23 to 25*	
low-educated father	-1.104***	-2.675***	-1.194***	-2.005***	-1.197***	-1.784***
	(0.127)	(0.366)	(0.090)	(0.180)	(0.090)	(0.186)
high-educated father	1.602***	3.790***	1.216***	2.226***	1.246***	2.116***
	(0.193)	(0.464)	(0.134)	(0.239)	(0.133)	(0.248)
low-educated father*expense per pupil devoted to X	0.006	0.002	0.004***	0.010***	0.005***	0.010***
	(0.006)	(0.006)	(0.002)	(0.002)	(0.002)	(0.002)
high-educated father*expense per pupil devoted to X	-0.016*	-0.006	0.002	-0.005*	0.001	-0.005
	(0.009)	(0.009)	(0.003)	(0.003)	(0.003)	(0.003)
low-educated father*enrollment rate in X		0.019***		0.017***		0.011***
		(0.004)		(0.003)		(0.003)
high-educated father*enrollment rate in X		-0.026***		-0.020***		-0.016***
		(0.005)		(0.004)		(0.004)
T(medium)	0.563*	0.359	-0.148	0.093	-0.330	-0.200
	(0.298)	(0.296)	(0.209)	(0.231)	(0.232)	(0.251)
T(high)	3.129***	2.939***	2.507***	2.758***	2.400***	2.536***
	(0.304)	(0.301)	(0.212)	(0.232)	(0.235)	(0.253)
Observations	31,871	31,871	47,393	47,338	45,176	45,176
N_clust	351	351	517	516	501	501
r2_p	0.152	0.156	0.149	0.151	0.147	0.148

Source: Author's estimations from SILC and Eurydice data

Table 4: The effect of education spending per pupil in % of GDP per capita, women

X=	pre-primary education		primary education		primary education		primary education	
when the offspring was	*3 to 5*		*6 to 8*		*9 to 10*		*6 to 10*	
low-educated father	-1.519***	-1.864***	-1.134***	2.653***	-1.223***	2.106**	-1.223***	3.059***
	(0.208)	(0.254)	(0.103)	(1.017)	(0.101)	(1.049)	(0.104)	(1.124)
high-educated father	0.792***	0.829*	1.256***	2.338*	1.225***	1.536	1.162***	2.419*
	(0.294)	(0.433)	(0.167)	(1.257)	(0.163)	(1.358)	(0.169)	(1.408)
low-educated father*expense per pupil devoted to X	0.054***	0.048***	0.014***	0.006	0.018***	0.011***	0.018***	0.009**
	(0.017)	(0.016)	(0.005)	(0.004)	(0.004)	(0.004)	(0.005)	(0.004)
high-educated father*expense per pupil devoted to X	0.029	0.036	-0.000	-0.001	0.002	0.003	0.005	0.004
	(0.024)	(0.023)	(0.007)	(0.007)	(0.007)	(0.007)	(0.007)	(0.008)
low-educated father*enrollment rate in X		0.006*		-0.036***		-0.031***		-0.040***
		(0.003)		(0.010)		(0.010)		(0.011)
high-educated father*enrollment rate in X		-0.001		-0.011		-0.003		-0.012
		(0.005)		(0.012)		(0.013)		(0.013)
T(medium)	-2.220***	-2.075***	-0.913***	-1.366***	-1.578***	-2.024***	-0.865***	-1.357***
	(0.266)	(0.312)	(0.280)	(0.302)	(0.244)	(0.269)	(0.257)	(0.283)
T(high)	0.186	0.335	1.417***	0.967***	0.766***	0.323	1.444***	0.956***
	(0.252)	(0.298)	(0.283)	(0.308)	(0.242)	(0.270)	(0.260)	(0.288)
Observations	10,747	10,747	21,955	21,955	24,884	24,884	23,824	23,824
N_clust	98	98	240	240	271	271	252	252
r2_p	0.140	0.141	0.136	0.137	0.140	0.141	0.139	0.140

X=	secondary education		tertiary education		tertiary education	
when the offspring was	*12 to 15*		*19 to 22*		*23 to 25*	
low-educated father	-1.223***	-2.831***	-1.031***	-1.870***	-0.915***	-1.549***
	(0.143)	(0.365)	(0.097)	(0.202)	(0.093)	(0.197)
high-educated father	1.181***	2.553***	1.055***	1.443***	1.302***	1.780***
	(0.156)	(0.488)	(0.142)	(0.232)	(0.136)	(0.271)
low-educated father*expense per pupil devoted to X	0.016**	0.012*	0.002	0.008***	0.000	0.005**
	(0.007)	(0.007)	(0.002)	(0.002)	(0.002)	(0.002)
high-educated father*expense per pupil devoted to X	0.007	0.015**	0.005*	0.002	0.000	-0.003
	(0.007)	(0.007)	(0.003)	(0.003)	(0.003)	(0.003)
low-educated father*enrollment rate in X		0.019***		0.017***		0.012***
		(0.004)		(0.003)		(0.003)
high-educated father*enrollment rate in X		-0.017***		-0.008*		-0.009**
		(0.005)		(0.004)		(0.004)
T(medium)	-1.794***	-1.935***	-1.194***	-0.823***	-1.230***	-1.002***
	(0.324)	(0.309)	(0.257)	(0.278)	(0.273)	(0.289)
T(high)	0.553*	0.427	1.244***	1.624***	1.266***	1.500***
	(0.325)	(0.310)	(0.257)	(0.277)	(0.273)	(0.288)
Observations	33,071	33,071	49,467	49,410	47,072	47,072

Source: Author's estimations from SILC and Eurydice data

Replacing the expenditures devoted to education by the number of pupils per teacher and the proportion of children in private schools in previous estimation enables to evaluate the extent to which these variables reinforce the persistence in education (Table 5 and Table 6). Surprisingly, the number of pupils per teacher is found to mitigate more significantly the persistence in education when increased in pre-primary and secondary classes than in primary classes. Indeed, an additional pupil in pre-primary and secondary classes is found to exacerbate the penalty for being the son of a low-educated father and then to increase by 4.4% the sons' odd ratio of being graduated when having a high- rather than low-educated father. By contrast, this odd ratio increases by only 3.2% when increasing the number of pupils in primary classes of children aged from 9 to 10.

For daughters, the effect of the number of pupils follow an even more surprising pattern: it is found to reinforce the penalty of having a low-educated father at pre-primary and secondary stages but also to decrease the premium of having a high-educated father at primary and secondary stages. An additional pupil could then increase the odd ratio by 1.6% at pre-primary stages, decrease it by 3% at primary stage and by 0.3% at the secondary stage (or even increase it by 0.9% if enrolment is controlled for).

Those ambiguous results at primary and secondary stages of education suggest that increasing the number of pupils in pre-primary class could reduce the propensity to achieve studies of "culturally-poor" children more strongly than that of other children but that, at later stages,

increasing the number of pupils would have the same negative effect on children's future social achievement independently of their social origin. As suggested by Schütz *et al.* (2005), this effect could be due to the fact that rich children are the "first" to benefit from an expansion of pre-primary schooling.

Table 5: The effect of the ratio of pupil per teacher, men

X=	pre-primary education		primary education		primary education		primary education		secondary education	
when the offspring was	*3 to 5*		*6 to 8*		*9 to 10*		*6 to 10*		*12 to 15*	
low-educated father	-0.116	-0.203	-0.848***	1.794**	-0.698***	0.629	-0.731***	1.784**	-0.647***	-2.560***
	(0.161)	(0.223)	(0.253)	(0.886)	(0.231)	(0.717)	(0.239)	(0.809)	(0.215)	(0.512)
high-educated father	1.252***	1.266***	0.656	0.698	0.955***	0.936	0.669*	0.909	1.171***	3.241***
	(0.180)	(0.203)	(0.410)	(0.950)	(0.350)	(0.929)	(0.376)	(0.902)	(0.280)	(0.694)
low-educated father* #pupils per teacher in	-0.044***	-0.045***	-0.012	0.002	-0.019*	-0.012	-0.016	-0.006	-0.032**	-0.033**
	(0.007)	(0.008)	(0.011)	(0.013)	(0.011)	(0.011)	(0.011)	(0.011)	(0.016)	(0.016)
high-educated father* #pupils per teacher in	0.005	0.005	0.029*	0.031*	0.017	0.017	0.030*	0.030*	0.011	-0.001
	(0.008)	(0.008)	(0.018)	(0.018)	(0.015)	(0.016)	(0.016)	(0.016)	(0.019)	(0.021)
low-educated father*enrollment rate in X		0.002		-0.029***		-0.014*		-0.027***		0.022***
		(0.003)		(0.009)		(0.007)		(0.008)		(0.005)
high-educated father*enrollment rate in X		-0.000		-0.001		0.000		-0.003		-0.021***
		(0.003)		(0.009)		(0.009)		(0.009)		(0.006)
T(medium)	-1.092***	-1.084***	-1.033	-1.046	-1.296**	-1.296**	-1.344**	-1.352**	-1.348	-1.384
	(0.350)	(0.348)	(0.971)	(0.973)	(0.652)	(0.652)	(0.659)	(0.659)	(1.360)	(1.256)
T(high)	1.762***	1.771***	1.607*	1.600*	1.304**	1.305**	1.294**	1.291*	1.264	1.245
	(0.352)	(0.350)	(0.971)	(0.973)	(0.652)	(0.652)	(0.659)	(0.659)	(1.358)	(1.255)
Observations	21,060	21,060	19,896	19,896	23,329	23,329	21,416	21,416	27,811	27,811
N_clust	198	198	214	214	243	243	233	233	281	281
r2_p	0.166	0.166	0.164	0.165	0.160	0.161	0.161	0.162	0.150	0.154

Source: Author's estimations from SILC and Eurydice data

Table 6: The effect of the ratio of pupils per teacher, women

X=	pre-primary education		primary education		primary education		primary education		secondary education	
when the offspring was	*3 to 5*		*6 to 8*		*9 to 10*		*6 to 10*		*12 to 15*	
low-educated father	-0.364**	-0.445**	-0.739***	1.290*	-0.736***	0.821	-0.726***	1.209*	-0.351	-2.259***
	(0.149)	(0.204)	(0.237)	(0.735)	(0.245)	(0.728)	(0.245)	(0.698)	(0.221)	(0.485)
high-educated father	1.269***	1.120***	1.977***	3.079***	1.871***	2.212**	1.929***	3.001***	1.922***	4.119***
	(0.204)	(0.260)	(0.364)	(0.806)	(0.358)	(0.931)	(0.355)	(0.896)	(0.299)	(0.662)
low-educated father* #pupils per teacher in	-0.030***	-0.031***	-0.012	-0.002	-0.013	-0.006	-0.012	-0.004	-0.055***	-0.052***
	(0.007)	(0.007)	(0.011)	(0.011)	(0.011)	(0.011)	(0.011)	(0.011)	(0.017)	(0.015)
high-educated father* #pupils per teacher in	0.003	0.001	-0.031**	-0.023	-0.026*	-0.024	-0.029*	-0.024	-0.048**	-0.061***
	(0.010)	(0.010)	(0.015)	(0.015)	(0.015)	(0.016)	(0.015)	(0.015)	(0.021)	(0.020)
low-educated father*enrollment rate in X		0.002		-0.022***		-0.017**		-0.021***		0.021***
		(0.002)		(0.007)		(0.007)		(0.007)		(0.004)
high-educated father*enrollment rate in X		0.003		-0.012*		-0.004		-0.012		-0.022***
		(0.003)		(0.008)		(0.009)		(0.009)		(0.006)
T(medium)	-1.704***	-1.660***	-1.354***	-1.373***	-1.544***	-1.544***	-1.522***	-1.532***	-1.649***	-1.524***
	(0.434)	(0.440)	(0.364)	(0.364)	(0.297)	(0.297)	(0.295)	(0.295)	(0.483)	(0.421)
T(high)	0.837*	0.882**	1.028***	1.011***	0.846***	0.848***	0.857***	0.849***	0.771	0.915**
	(0.435)	(0.440)	(0.361)	(0.360)	(0.293)	(0.293)	(0.291)	(0.291)	(0.481)	(0.418)
Observations	21,764	21,764	20,672	20,672	24,281	24,281	22,358	22,358	29,183	29,183
N_clust	198	198	214	214	243	243	233	233	281	281
r2_p	0.142	0.142	0.146	0.147	0.147	0.147	0.144	0.144	0.145	0.149

Source: Author's estimations from SILC and Eurydice data

The proportion of pupils in private schools is expected to increase social persistence but with a decreasing effect as the proportion of private schooling increases. To test this hypothesis, the father's education is interacted with a second-order polynomial of the proportion of pupils in private schools. Unfortunately, the lack of data does not enable to estimate the effect of the proportion of private schooling in secondary and tertiary education (Table 7 and Table 8).

At pre-primary and primary stages of education, the share of children in private schools has a weak effect on the premium for having a high-educated father but increase significantly the penalty for having a low-educated one at the primary stages. As shown by the coefficient of the polynomial terms, this negative effect decreases as the proportion increase which is consistent with the hypothesis that the more accessible the private schooling, the lower the social segregation. This might explain why Netherlands is a mobile country despite the

proportion of pupils in private schools exceeds 50%. More precisely the results suggest that increasing the proportion of pupils in private primary schools increase the odd ratio of Equation 10 and the persistence in education until the proportion reaches at least 20%.

Table 7: Effect of the proportion of pupils in private schools, men

X=	pre-primary education		primary education		primary education		primary education	
when the offspring was	*3 to 5*		*6 to 8*		*9 to 10*		*6 to 10*	
low-educated father	-0.839***	-1.187***	-0.900***	4.950***	-0.729***	4.378***	-0.790***	5.233***
	(0.114)	(0.213)	(0.120)	(1.718)	(0.118)	(1.436)	(0.108)	(1.550)
high-educated father	1.282***	1.251***	1.111***	1.600	0.996***	2.343	1.133***	2.201
	(0.196)	(0.244)	(0.159)	(1.905)	(0.151)	(1.850)	(0.141)	(1.715)
low-educated father* % pupils in private X	-0.005	-0.004	-0.037***	-0.004	-0.046***	-0.019	-0.045***	-0.008
	(0.008)	(0.008)	(0.013)	(0.016)	(0.013)	(0.015)	(0.012)	(0.015)
low-educated father*(% pupils in private X)²	-0.000	-0.000	0.001***	0.000	0.001***	0.000	0.001***	0.000
	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)
high-educated father* % pupils in private X	-0.019*	-0.019*	0.002	0.004	0.005	0.010	0.001	0.006
	(0.011)	(0.010)	(0.015)	(0.018)	(0.014)	(0.015)	(0.013)	(0.016)
high-educated father*(% pupils in private X)²	0.000**	0.000*	-0.000	-0.000	-0.000	-0.000	-0.000	-0.000
	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)
low-educated father*enrollment rate in X		0.005*		-0.059***		-0.051***		-0.061***
		(0.003)		(0.017)		(0.014)		(0.016)
high-educated father*enrollment rate in X		0.000		-0.005		-0.014		-0.011
		(0.003)		(0.019)		(0.018)		(0.017)
T(medium)	-1.063***	-1.036***	-2.093***	-1.905***	-0.572**	-0.512**	-0.410*	-0.338
	(0.275)	(0.276)	(0.250)	(0.254)	(0.256)	(0.253)	(0.245)	(0.245)
T(high)	1.248***	1.277***	0.350	0.545**	1.813***	1.878***	2.012***	2.091***
	(0.276)	(0.276)	(0.254)	(0.256)	(0.262)	(0.259)	(0.251)	(0.251)
Observations	13,331	13,331	17,706	17,706	19,849	19,849	20,912	20,912
N_clust	160	160	171	171	188	188	202	202
r2_p	0.156	0.156	0.157	0.159	0.146	0.147	0.155	0.156

Source: Author's estimations from SILC and Eurydice data

Table 8: Effect of the proportion of pupils in private schools, women

X=	pre-primary education		primary education		primary education		primary education	
when the offspring was	*3 to 5*		*6 to 8*		*9 to 10*		*6 to 10*	
low-educated father	-1.014***	-1.244***	-0.814***	2.781**	-0.760***	2.715	-0.792***	2.535*
	(0.167)	(0.215)	(0.126)	(1.284)	(0.128)	(1.736)	(0.116)	(1.355)
high-educated father	1.285***	1.071***	1.429***	3.906**	1.299***	3.680	1.294***	5.025***
	(0.210)	(0.260)	(0.188)	(1.811)	(0.211)	(2.240)	(0.184)	(1.834)
low-educated father* % pupils in private X	0.007	0.008	-0.041***	-0.021	-0.036***	-0.017	-0.038***	-0.019
	(0.009)	(0.008)	(0.013)	(0.013)	(0.014)	(0.014)	(0.012)	(0.013)
low-educated father*(% pupils in private X)²	-0.000	-0.000	0.001***	0.000*	0.001***	0.000	0.001***	0.000*
	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)
high-educated father* % pupils in private X	-0.010	-0.009	-0.032*	-0.019	-0.015	-0.003	-0.017	0.005
	(0.011)	(0.011)	(0.017)	(0.019)	(0.019)	(0.020)	(0.017)	(0.018)
high-educated father*(% pupils in private X)²	0.000	0.000	0.000*	0.000	0.000	0.000	0.000	-0.000
	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)
low-educated father*enrollment rate in X		0.003		-0.036***		-0.035**		-0.034**
		(0.003)		(0.012)		(0.017)		(0.013)
high-educated father*enrollment rate in X		0.003		-0.025		-0.024		-0.038**
		(0.003)		(0.018)		(0.022)		(0.018)
T(medium)	-1.423***	-1.368***	-1.236***	-1.156***	-1.665***	-1.551***	-1.789***	-1.665***
	(0.317)	(0.323)	(0.167)	(0.172)	(0.296)	(0.300)	(0.259)	(0.263)
T(high)	0.633**	0.688**	0.948***	1.030***	0.548*	0.664**	0.384	0.509*
	(0.318)	(0.323)	(0.168)	(0.172)	(0.300)	(0.303)	(0.263)	(0.266)
Observations	14,198	14,198	18,423	18,423	20,621	20,621	21,794	21,794
N_clust	160	160	171	171	188	188	202	202
r2_p	0.131	0.131	0.137	0.137	0.139	0.139	0.136	0.136

Source: Author's estimations from SILC and Eurydice data

4.2. Compulsory school

In order to determine the effect of the duration of the compulsory school on the education persistence, Equation 10 is estimated with the policy variable being the minimum school leaving age. To do so, the variations of this variable across a selection of 14 countries and across time are used. Variations across time are due to educational reforms which successively increase the minimum age in Western countries in appendix). The main difficulty encountered by most of authors for estimating the effect of such reform is to disentangle the effect of minimum age from the effect of other variables which induced an increase in the education across time. Equation 10 does not aim at estimating the direct effect of the minimum age but its interacted effect with paternal education so that the interactions of country and cohort dummies are included to control for the increasing pattern of education across countries.

Our estimations then indicate that increasing by one year the minimum school leaving age mitigate the penalty for having a low-educated father without altering the premium for having a high-educated father, so that it finally mitigates the odd ratio of Equation 10 by 22%.

In contrast to previous estimations of the effect of compulsory schooling, this latter one uses a large sample of individuals whose age vary from 25 to 85 years old whereas the minimum leaving age vary very punctually across time. One could then argue that the obtained effect of the school leaving age on the education mobility might mirror the increase of the education mobility within countries due to other factors than the minimum age. To ensure that the estimated effect actually measures the mobility-enhancing effect of the leaving age, the sample is then restricted to individuals born a few years before and after the cohort concerned by the most recent national reform of each country (as in Brunello *et al.* 2009). After this restriction, the estimated effect of the leaving age is found to be significant but lower: for men, one more year of compulsory school is found to decrease the odd ratio by 14% whereas for women it decreases it only by approximately 7%. Indeed, for daughters, although an additional year is found to decrease the penalty of having a low-educated father, it is also found to increase significantly the premium for having a tertiary-educated one.

Table 9: Effect of the minimum school leaving age

	Men				Women			
year birth compared to the concerned cohort	any	+/- 10 years before	+/- 7 years before	+/- 5 years before	any	+/- 10 years before	+/- 7 years before	- 5 years befo
low-educated father	-4.311***	-3.321***	-3.109***	-3.253***	-4.366***	-4.282***	-3.982***	-3.954***
	(0.468)	(0.591)	(0.691)	(0.852)	(0.537)	(0.518)	(0.598)	(0.727)
high-educated father	2.092***	0.929	0.794	0.317	2.126***	-0.456	-0.621	-0.914
	(0.533)	(0.649)	(0.743)	(0.932)	(0.644)	(0.680)	(0.784)	(0.915)
low-educated father*minimum leaving age	0.221***	0.155***	0.140***	0.147***	0.224***	0.218***	0.199***	0.194***
	(0.031)	(0.039)	(0.046)	(0.057)	(0.035)	(0.034)	(0.039)	(0.047)
high-educated father*minimum leaving age	-0.064*	0.019	0.031	0.065	-0.066	0.102**	0.116**	0.131**
	(0.035)	(0.043)	(0.049)	(0.062)	(0.043)	(0.045)	(0.052)	(0.060)
T(medium)	-2.518***	-1.981***	-2.077***	-1.998***	-1.568***	-2.144***	-1.993***	-2.443***
	(0.055)	(0.065)	(0.076)	(0.079)	(0.082)	(0.077)	(0.076)	(0.100)
T(high)	-0.324***	0.203***	0.082	0.160*	0.524***	-0.107	0.012	-0.486***
	(0.052)	(0.070)	(0.080)	(0.084)	(0.084)	(0.081)	(0.078)	(0.097)
Observations	52,735	26,035	19,021	14,231	55,354	27,544	19,958	14,805
r2_p	0.138	0.118	0.115	0.114	0.152	0.122	0.115	0.112

Source: Author's estimations from SILC data

4.3. Intergenerational transmission

The reforms of the minimum leaving age did not only concern the surveyed individuals but also their father. From the father's year of birth, it is then possible to determine the minimum age at which the parents were allowed to leave school and then to estimate its impact both on parents' and offspring's educational attainment. The parental minimum school leaving age is uncorrelated with a parent's ability but is likely to increase the parent's educational attainment and income. If the transmission of education is only due to the inheritance of human capital, such a variable should then have a null effect on the offspring's educational attainment, while it should significantly increase it if the transmission occurred through father's investment in his offspring's human capital.

As said previously, distinguishing the effect of the minimum leaving age from the effect of other national characteristics arises some issues because of its weak and monotonic variations within countries across time. Controlling for cohort dummies and country dummies in a cross-country estimation might enable to circumvent this issue but due to the weak variations of the minimum age across countries for a given cohort, another strategy is chosen in the present chapter. Rather than controlling for each parent's cohort/time dummies, in each country, the sample is restricted to individuals whose father or mother was born a few years before and after the cohort concerned by the reform (in countries where several reforms of the minimum leaving age have been implemented, the oldest one is now considered). By choosing an appropriate number of years and controlling for country dummies, it is then possible to estimate the effect of the parental minimum leaving age respectively on the parent's and offspring's education in a logit model whose propensity can be written:

$$\varphi_i = \phi \cdot ycomp_{ac} + Q_{c\cdot} + \varepsilon_{iac}$$

where $ycomp$ corresponds to the minimum school leaving age.

Since this variable is continuous, its obtained coefficient corresponds to the increase in the odd of having a tertiary education when increasing by one year the minimum age, consistently with the equation:

$$\log\left(\frac{P(E=high)}{P(E\neq high)}\right) = \phi \cdot ycomp_{ac} + Q_c \qquad (11)$$

The results a priori suggest that the rise in minimum leaving age which has occurred during parents' youth has increased their propensity to obtain a degree. But this effect is not robust to restricting the sample to parents born a few years before or after the cohort concerned by those reforms. This suggests that the estimated effect indeed mirrors the effect of other education-enhancing factors but could also be due to the decrease in coefficients' accuracy induced by the reduction of the sample size. Whatever the reason for this lack of significance, it mainly concerns the effect of the minimum age on the paternal educational attainment, the effect on the maternal education remaining significant at 10%.

The results also indicate that a one-year increase in the minimum age increased by at least 7% the offspring's odd of being graduated. The stronger effect of the maternal education on the children's future achievement is consistent with previous literature which documents the effect of changes in compulsory education on educational attainment by gender. As previously said, this effect might be due to the fact that mother's education has a stronger impact at the bottom of the parental education distribution, which is the part of the distribution the most affected by such reforms.

The transmission of this effect to offspring also seems to depend on the offspring's gender. Indeed, when a woman is affected by a one-year increase in the minimum exit age, her daughters' odd to be graduated increases by at least 8% whereas her sons' exhibit a lower (if not zero) increase[4]. Surprisingly, this pattern holds for the offspring of a man who is affected by an increase in the minimum leaving age, although the father's educational may be unchanged according to the estimations. One explanation could be that the father's and mother's age tend to be close so that a change in the leaving age which occur in a father's youth might have affected the mother's educational attainment and translated into an increase in the offspring's education. Another hypothesis could be that the reforms of the exit age have increased the father's human capital enough to increase his income and his offspring's outcome but not enough to permit the father reaching an upper level of education

Whatever the reason for this effect of parental minimum age on offspring's attainment, it underlines that the transmission of human capital might not only due to the inheritance of some genetic or ethic characteristics but also to unequal investment in children's education due to income disparities.

[4] Interestingly, the effect of the parents' minimum leaving age on the offspring's attainment is found to be more significant when the age difference between father and offspring is controlled for with a country-specific coefficient. This latter variable is found to increase very significantly the propensity to achieve high education, probably because the father's income tends to increase with his age, so that older father can afford paying more for their offspring's education. Since the increase of in the minimum leaving age might delay the age of first parenthood, the minimum age might be negatively correlated to the age difference between parents and offspring. Then, if the age difference is not controlled for, the estimated effect of the minimum age is likely to be biased downward.

Table 10: Effect of the paternal minimum school leaving age on father's and offspring's education

	Mother's education				Son's education								Daughter's education							
mother's year of birth compared to the concerned cohort:	any	+/- 10 yrs before	+/- 7 yrs before	+/- 5 yrs before	any	+/- 10 years before		+/- 7 years before		+/- 5 years before			any	+/- 10 years before		+/- 7 years before		+/- 5 years before		
maternal minimum leaving age	0.445***	0.161***	0.107***	0.071*	0.140***	0.178***	0.081***	0.122***	0.019	0.080**	0.010	0.056	0.279***	0.345***	0.081**	0.195***	0.020	0.131***	-0.014	0.084***
	(0.038)	(0.033)	(0.038)	(0.038)	(0.026)	(0.025)	(0.029)	(0.031)	(0.036)	(0.036)	(0.037)	(0.038)	(0.033)	(0.030)	(0.033)	(0.030)	(0.034)	(0.028)	(0.037)	(0.029)
T(medium)	6.125***	2.073***	1.293***	0.808*	0.688**	2.760***	-0.219	2.538***	-0.987**	1.887***	-1.124**	1.252*	2.195***	5.292***	-0.421	4.291***	-1.308***	4.038***	-1.733***	3.522***
	(0.535)	(0.422)	(0.462)	(0.468)	(0.325)	(0.385)	(0.365)	(0.637)	(0.443)	(0.682)	(0.457)	(0.699)	(0.396)	(0.427)	(0.420)	(0.634)	(0.445)	(0.639)	(0.485)	(0.717)
T(high)	7.390***	3.451***	2.722***	2.228***	2.588***	4.692***	1.928***	4.841***	1.379***	4.292***	1.261***	3.678***	3.931***	7.084***	1.549***	6.333***	0.763*	6.174***	0.337	5.667***
	(0.534)	(0.425)	(0.467)	(0.474)	(0.322)	(0.384)	(0.357)	(0.632)	(0.436)	(0.680)	(0.457)	(0.699)	(0.395)	(0.429)	(0.419)	(0.641)	(0.439)	(0.649)	(0.481)	(0.732)
Country dummies	YES	YES	YES	YES	YES	YES	YES	YES	YES	YES	YES	YES	YES	YES	YES	YES	YES	YES	YES	YES
*(offspring age-father's age)	NO	NO	NO	NO	NO	YES	NO	YES	NO	YES	NO	YES	NO	YES	NO	YES	NO	YES	NO	YES
Observations	100,164	26,015	16,933	12,453	48,385	48,383	18,661	12,592	8,288	8,288	6,128	6,128	51,779	51,776	13,421	13,418	8,645	8,644	6,325	6,324
r2_p	0.0731	0.0857	0.0879	0.0851	0.0570	0.0665	0.0669	0.0811	0.0839	0.0936	0.0765	0.0871	0.0467	0.0652	0.0374	0.0591	0.0410	0.0611	0.0412	0.0646

Source: Author's computations from SILC data

Table 11 : Effect the maternal minimum school leaving age on mother's and offspring's education

	Father's education				Son's education								Daughter's education							
father's year of birth compared to the concerned cohort:	any	+/- 10 yrs before	+/- 7 yrs before	+/- 5 yrs before	any	+/- 10 years before		+/- 7 years before		+/- 5 years before			any	+/- 10 years before		+/- 7 years before		+/- 5 years before		
paternal minimum leaving age	0.264***	0.089***	0.017	0.003	0.127***	0.154***	-0.010	0.058	-0.055	0.031	-0.047	0.003	0.320***	0.376***	0.080**	0.191***	-0.000	0.100**	-0.013	0.068
	(0.030)	(0.028)	(0.028)	(0.029)	(0.031)	(0.030)	(0.037)	(0.036)	(0.041)	(0.041)	(0.043)	(0.044)	(0.039)	(0.035)	(0.039)	(0.039)	(0.040)	(0.040)	(0.046)	(0.048)
T(medium)	2.968***	0.599*	-0.331	-0.585	0.539	1.917***	-1.445***	1.506**	-2.032***	1.472*	-1.967***	-0.176	2.670***	5.405***	-0.619	3.725***	-1.643***	2.680***	-1.856***	1.994*
	(0.353)	(0.348)	(0.352)	(0.357)	(0.370)	(0.457)	(0.455)	(0.714)	(0.516)	(0.818)	(0.535)	(1.029)	(0.473)	(0.453)	(0.493)	(0.717)	(0.545)	(0.702)	(0.644)	(1.045)
T(high)	4.331***	2.161***	1.310***	1.078***	2.458***	3.873***	0.876**	3.865***	0.420	3.961***	0.522	2.338**	4.432***	7.244***	1.408***	5.824***	0.517	4.890***	0.312	4.208***
	(0.353)	(0.344)	(0.354)	(0.363)	(0.367)	(0.456)	(0.447)	(0.713)	(0.503)	(0.819)	(0.530)	(1.023)	(0.472)	(0.455)	(0.495)	(0.723)	(0.551)	(0.715)	(0.642)	(1.056)
Country dummies	YES	YES	YES	YES	YES	YES	YES	YES	YES	YES	YES	YES	YES	YES	YES	YES	YES	YES	YES	YES
*(offspring age-father's age)	NO	NO	NO	NO	NO	YES	NO	YES	NO	YES	NO	YES	NO	YES	NO	YES	NO	YES	NO	YES
Observations	102,849	21,414	13,116	9,536	49,902	49,901	10,382	10,381	6,385	6,385	4,646	4,646	52,947	52,944	11,032	11,029	6,731	6,730	4,890	4,889
r2_p	0.0746	0.0695	0.0753	0.0768	0.0562	0.0672	0.0681	0.0777	0.0776	0.0878	0.0708	0.0789	0.0464	0.0713	0.0357	0.0568	0.0420	0.0578	0.0443	0.0593

Source: Author's estimations from SILC data

5. Concluding remarks

Comparing countries' relative positions in terms of educational and wage persistence reveals interesting patterns. Consistently with prior research, Nordic countries appear as more mobile than Southern and Eastern Europe which is essentially due to their very low mobility at the bottom of the paternal education distribution. The analysis of the evolution of education mobility reveals a general increase in all countries which corresponds to an increase in equality of opportunity, whereas the mobility as a movement between social classes has remained rather stable. Whereas the persistence in education has regularly decreased in Western countries and particularly in Nordic countries, thanks to the educational reforms undertaken to increase the accessibility of early education, it has slightly increased in Eastern countries under the effect of a more recent liberalization of the economy.

Estimating the effect of educational policies shows that a high investment in the early stages decreases more substantially the persistence in education than investment at the secondary and tertiary educations, in particular for boys when this high investment result in a lower ratio of pupils per teacher at pre-primary stage. This higher significance reinforces the relevance of Cunha and Heckman's (2007) theory of dynamic complementarity of human capital and pledges for a redistribution of the public resources devoted to education toward pre-primary and primary levels of education in countries where the educational mobility is low.

Increasing the duration of compulsory school by delaying the exit age also appears as a way to increase education mobility since it mitigates the inequality in the propensity to achieve studies between the offspring of low-educated and high-educated father. Moreover, increasing the exit age is found to increase a male's own propensity to achieve higher studies, but also that of his offspring, which proves that the inheritance of endowment in human capital is not the only channel of transmission of education but that unequal investment in education might also explain the intergenerational persistence in education.

Last, the social segregation induced by the existence of private schooling is also found to increase social persistence, except in countries where private schooling is an institution that enroll enough pupils to avoid segregation. Decreasing the segregation by pooling children from poor and rich families in the same schools might indeed increase the educational attainment of the socially disadvantaged pupils by preventing them with suffering from the negative peer effects induced by the concentration of poor pupils.

References

Ammermüller, A. (2005) "Educational Opportunities and the Role of Institutions", ZEW Discussion Paper, No. 05-44.

Angrist, J. Lavy, V. (1999) "Using Maimonides' Rule to Estimate the Effect of Class Size on Scholastic Achievement", Quarterly Journal of Economics, Vol. 114, No2, pp. 533-574

Bauer, P. Riphahn,R. (2006), "Timing of school tracking as a determinant of intergenerational transmission of education", Economics Letters 91: 90-97.

Bauer, P. Riphahn,R. (2009), "Kindergarten enrollment and the intergenerational transmission of education." Discussion paper no. 4466, Institute for the Study of Labor (IZA), Bonn.

Becker, G.S. and N. Tomes (1979), "An Equilibrium Theory of the Distribution of Income and Intergenerational Mobility", Journal of Political Economy, Vol. 87, pp. 1153-89.

Belley, P. Lochner, L. (2007) "The Changing Role of Family Income and Ability in Determining Educational Achievement," Journal of Human Capital, University of Chicago Press, vol. 1(1), pages 37-89.

Belzil, C. Hansen, J. (2003) "Structural Estimates of the Intergenerational Education Correlation," IZA Discussion Papers 973, Institute for the Study of Labor.

Berlinski, S. Galiani, S. Gertler, P. (2009) "The effect of pre-primary education on primary school performance," Journal of Public Economics, Elsevier, vol. 93(1-2), pages 219-234, February.

Berlinski, S. Galiani, S. Manacorda, M. (2008). "Giving children a better start: Preschool attendance and school-age profiles," Journal of Public Economics, Elsevier, vol. 92(5-6), pages 1416-1440, June.

Bingley, P. Jensen, V.Walker,I. (2005) "The Effects of School class Size on Length of Post-Compulsory Education: Some cost-benefit Analysis", IZA Discussion paper No. 1605, Bonn, Germany.

Black, S. Devereux, P. Salvanes,K. (2005) "Why the apple doesn't fall far: Understanding intergenerational transmission of human capital", American Economic Review 95: 437-449.

Blanden, J. Goodman,A. Gregg,P. Machin,S. (2004), "Changes in intergenerational mobility in Britain," in M. Corak (ed.) Generational Income Mobility in North America and Europe (Cambridge University Press).

Blanden, J. Gregg, P. Macmillan,L. (2006), "Accounting for Intergenerational Income Persistence: Non-Cognitive Skills, Ability and Education", CEE Working Paper.

Bonesrønning, H. (2003). "Class size effects on student achievement in norway: Patterns and explanations." Southern Economic Journal, 69(4):952–965.

Braga, M. Checchi, D. Meschi, E. (2011) "Institutional Reforms and Educational Attainment in Europe: A Long Run Perspective." IZA Discussion Papers 6190, Institute for the Study of Labor (IZA).

Bressoux, P. Kramarz, F. Prost,C. (2009) "Teachers' Training, Class Size, and Students' Outcomes: Learning from Administrative Forecasting Mistakes", Economic Journal, 119, 540-561.

Browning, M. Heinesen, E. (2007) "Class Size, Teacher Hours and Educational Attainment", Scandinavian Journal of Economics, 109, 415-438.

Brunello, G. Fort, M. Weber, G. (2009) "Changes in Compulsory Schooling, Education and the Distribution of Wages in Europe," Economic Journal, 119(536): 516-539.

Brunello,G. Checchi, D. (2007) "Does school tracking affect equality of opportunity? New international evidence," Economic Policy, CEPR & CES & MSH, vol. 22, pages 781-861.

Cameron, S. Heckman, J. (1998) "Life Cycle Schooling and Dynamic Selection Bias: Models and Evidence for Five Cohorts," NBER Working Papers 6385, National Bureau of Economic Research, Inc.

Cameron, S. Heckman, J. (2001) The Dynamics of Educational Attainment for Black, Hispanic, and White Males. Journal of Political Economy, June.

Carneiro, P. Heckman,J. (2002) "The Evidence on Credit Constraints in Post-secondary Schooling," Economic Journal, Royal Economic Society, vol. 112(482), pages 705-734, October.

Causa,O. Chapuis,C. (2009). "Equity in Student Achievement Across OECD Countries: An Investigation of the Role of Policies," OECD Economics Department Working Papers 708, OECD, Economics Department.

Checchi, D. Flabbi, L. (2005) "Intergenerational mobility and schooling decisions in Italy and Germany.: The Impact of Secondary School Tracks." IZA Discussion Papers 2876, Institute for the Study of Labor (IZA).

Checchi, D. Ichino, A. Rustichini, A. (1999) ”More equal but less mobile? Education financing and intergenerational mobility in Italy and in the U.S.” Journal of Public Economics, 74:35 1393.

Chevalier, A. (2004), “Parental Education and Child’s Education: A natural experiment”, Discussion paper no. 1153 (Institute for the Study of Labor (IZA), Bonn). 82

Chevalier, A. Denny,K. McMahon,D. (2009), “A Multi-country study of inter-generational educational mobility” in: Dolton,P. Asplundh,R and Barth,E. eds., Education and inequality Across Europe (Edward Elgar, London).

Chevalier, A. Harmon, C. O’Sullivan,V. Walker,I. (2005) “The Impact of parental income and education on the schooling of their children”, Discussion paper no. 1496 (Institute for the Study of Labor (IZA), Bonn).

Couch, K. Dunn, T. (1997). "Intergenerational Correlations in Labor Market Status: A Comparison of the United States and Germany," Journal of Human Resources, University of Wisconsin Press, vol. 32(1), pages 210-232.

Cunha, F. Heckman,J. (2007) ""The Technology of Skill Formation,"" American Economic Review, American Economic Association, vol. 97(2), pages 31-47, May.

de Broucker, P. Underwood,K. (1998) "Intergenerational education mobility: An international comparison with a focus on postsecondary education." Education Quarterly Review, 5(2):30-51.

Dearden, L., McGranahan, L. Sianesi B. (2004), "The Role of Credit Constraints in Educational Choices: Evidence from the NCDS and BCS70", CEE Working Paper, LSE, London.

Dobbelsteen, S. Levin, J. Oosterbeek, H. (2002) "The Causal Effect of Class Size on Scholastic Achievement: Distinguishing the Pure Class-Size Effect From the Effect of Changes in Class Composition", Oxford Bulletin of Economics and Statistics, 64, 17-38.

Dumas, C. Lefranc, A. (2012) "Early schooling and later outcomes: Evidence from pre-school extension in France," in J. Ermisch, M. Jantti and T.Smeeding (eds) From parents to children: The Intergenerational Transmission of Advantage. New York: Russell Sage Foundation, p. 164-189

Eckstein, Z. Wolpin,K. (1999). "Why Youths Drop Out of High School: The Impact of Preferences, Opportunities, and Abilities," Econometrica, Econometric Society, vol. 67(6),

Foley, K. (2004), "International differences in education mobility across generations", unpublished manuscript, Department of Economics, University of British Columbia, Vancouver.

Ganzeboom, H. Nieuwbeerta,P. (1999) "Access to education in six Eastern European countries between 1940 and 1985: Results of a crossnational survey." Communist and Post-Communist Studies, 32(4):339-357.

Gary-Bobo, R. Mahjoub, M-B. (2010), “Estimation of Class-Size Effects, using “Maimonides‟ Rule and Other Instruments: The Case of French Junior High Schools”, unpublished

Goldin, C. Katz, L. (2003) "Mass secondary schooling and the state: The role of state compulsion in the high school movement." National Bureau of Economic Research Working Paper no. 10075, Cambridge, MA.

Goodman, A. Sianesi,B. (2005) "Early education and children's outcomes: low long do the impacts last?," Fiscal Studies, Institute for Fiscal Studies, vol. 26(4), pages 513-548, December.

Grenet, J. (2013), Is Extending Compulsory Schooling Alone Enough to Raise Earnings? Evidence from French and British Compulsory Schooling Laws. The Scandinavian Journal of Economics.

Hanushek, E.A. Woessmann, L. (2005) "Does Educational Tracking Affect Performance and Inequality? Differences- in-Differences Evidence Across Countries," Economic Journal, vol. 116(510): C63-C76

Havnes, T. Mogstad, M. (2009) "No child left behind: Universal child care and children's long-run outcomes", IZA Discussion Papers 4561, Institute for the Study of Labor (IZA).

Hertz, T., T, Jayasundera, P. Piraino, S. Selcuk, N. Smith and A. Verashchagina (2007), "The Inheritance of Educational Inequality: International Comparisons and Fifty-Year Trends", The B.E. Journal of Economic Analysis and Policy, Vol 7.

Holmlund, H. Lindahl, M. Plug,E. (2008) "The Causal Effect of Parent's Schooling on Children's Schooling: A Comparison of Estimation Methods", IZA Discussion Paper No. 3630.

Kane, T (2001) "College Going and Inequality: A Literature Review" Russell Sage Foundation working paper.

Kane, T. (1995) "Rising Public College Tuition and College Entry: How Well Do Public Subsidies Promote Access to College?," NBER Working Papers 5164, National Bureau of Economic Research.

Keane, M. Wolpin, K. (2001) "The Effect of Parental Transfers and Borrowing Constraints on Educational Attainment" International Economic Review , Vol. 42, No. 4 November, pp. 1051-1103

Lefranc, A. Ojima, F. Yoshida, T. (2008). "The Intergenerational Transmission of Income and Education: A Comparison of Japan and France," RSCAS Working Papers 2008/25, European University Institute.

Leuven, E. Lindahl, M. Oosterbeek, H. Webbink,D. (2006) "Expanding Schooling Opportunities for 4-Year-Olds," IZA Discussion Papers 2434, Institute for the Study of Labor.

Leuven,E. Oosterbeek, H. Ronning, M. (2008) "Quasi-Experimental Estimates of the Effect of Class Size on Achievement in Norway" Scandinavian Journal of Economics, 110, 663-693.

Magnuson, K. Ruhm, C. Waldfogel, J. (2007). "Does prekindergarten improve school preparation and performance?," Economics of Education Review, Elsevier, vol. 26(1), pages 33-51, February.

Meghir C. Palme,M. (2005) "Educational Reform, Ability, and Family Background." American Economic Review, 95(1):414-424, March.

Mocetti, S. (2008) "Educational choices and the selection process before and after compulsory schooling," Temi di discussione (Economic working papers) 691, Bank of Italy, Economic Research Department.

Murtin, F. Viarengo, M. (2011) "The Expansion and Convergence of Compulsory Schooling in Western Europe, 1950–2000", Economica, Vol. 78(311): 501-522.

OECD (2010), "A Family Affair: Intergenerational Social Mobility across OECD Countries", in OECD, Economic Policy Reforms 2010: Going for Growth, OECD Publishing.

Oliveira-Martins, J. Boarini, R. Strauss, H. De La Maisonneuve, C. (2009) "The policy determinants of investment in tertiary education", OECD Journal: Economic Studies, OECD Publishing, vol. 2009(1), pages 1-37.

Oreopoulos, P. (2007) "Do Dropouts Drop Out Too Soon? Wealth, Health, and Happiness from Compulsory Schooling." Journal of Public Economics 91, (11-12): 2213-2229.

Oreopoulos, P. Page,M. Stevens,A. (2006) "The intergenerational effects of compulsory schooling", Journal of Labor Economics 24: 729-760.

Pekkarinen, T. Uusitalo, R. Kerr, S. (2009) "School Tracking and Development of Cognitive Skills." IZA Discussion Papers 4058, Institute for the Study of Labor (IZA).

Pekkarinen, T. Uusitalo, R. Pekkala,S. (2006) "Education Policy and Intergenerational Income Mobility: Evidence from the Finnish Comprehensive School Reform". IZA discussion paper 2204.

Pekkarinnen, T. (2005) "Gender Differences in Educational Attainment: Evidence on the Role of Tracking Age from a Finnish Quasi-Experiment." Discussion Paper 1897, Institute for the Study of Labor (IZA).

Piketty, T. (2004) "Should We Reduce Class Size or School Segregation? Theory and Evidence from France", PSE Jourdan, ENS, Paris, France. 27

Pischke, J. Watcher,T. (2005) "Zero Returns to Compulsory Schooling in Germany: Evidence and Interpretations." Working Paper 11414, National Bureau of Economic Research, June.

Plug, E. (2001) Season of Birth, Schooling and Earnings. Journal of Economic Psychology, 22:641-660.

Pronzato, C. (2009) "An examination of paternal and maternal intergenerational transmission of schooling", Journal of Population Economics, Springer, vol. 25(2), pages 591-608, January.

Schütz, G. Ursprung, H. Woessmann, L. (2005) "Education Policy and Equality of Opportunity," IZA Discussion Papers 1906, Institute for the Study of Labor (IZA).

Solon, G. (2004), "A Model of Intergenerational Mobility Variation over Time and Place." In Generational Income Mobility in North America and Europe, ed. Miles Corak. 38–47. Cambridge University Press.

Waldinger, F. (2006) "Does Tracking Affect the Importance of Family Background on Students. Test Scores?" mimeo, London School of Economics

Wössmann, L.(2005) "Educational Production in Europe", Economic Policy, 20, issue 43, 445-504.

Appendix

Table 12: Reforms of the minimum school leaving age which affect the sample

country	year of reform	concerned cohort by reform	minimal leaving age	% concerned offspring	% concerned fathers
Austria		before 1948	14	17,18	89,45
	1962	1948	15	82,82	10,55
Belgium		before 1969	14	73,59	100
	1983	1969	18	26,41	0
Denmark		before 1947	11	18,13	89,65
	1958	1947	14	24,35	10,15
	1971	1957	16	57,51	0,2
Finland		before 1964	13	60,04	100
	1977	1964	16	39,96	0
France		before 1953	14	30,37	96,86
	1967	1953	15	69,63	3,14
Greece		before 1964	12	56,05	100
	1976	1964	15	43,95	0
Ireland		before 1958	14	43,25	99,5
	1972	1958	15	56,75	0,5
Italy		before 1950	11	21,89	95,69
	1963	1950	14	78,11	4,31
Netherlands		before 1937	13	0	57,75
	1950	1937	15	38,81	24,99
	1971	1956	16	8,05	0,22
	1975	1959	17	53,14	17,04
Portugal*		before 1947	9	19,99	86,98
	1956	1947	10	11,69	10,63
	1964	1954	12	51,54	2,4
	1986	1974	15	16,49	0
Spain		before 1957	13	34,21	99,48
	1970	1957	14	54,81	0,52
	1990	1976	16	10,98	0
Sweden		before 1950	15	24,82	93,35
	1962	1950	16	75,18	6,65
UK		before 1933	14	0	56,72
	1947	1933	15	40,97	42,62
	1973	1958	16	59,03	0,66
Norway		before 1957	12	40,12	99,64
	1969	1957	15	59,88	0,36

* for women, the reform which increased the minimum leaving age from 9 to 10 was implemented in 1960 ; thus the considered cohort for women is that of 1951.