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# The Relative Effect of Family Characteristics and Financial Situation on Educational Achievement

ARNAUD CHEVALIER and GAUTHIER LANOT

ABSTRACT *Children from poorer backgrounds are generally observed to have lower educational outcomes than other youth. However, the mechanism through which household income affects the child's outcomes remains unclear. Either, poorer families are financially constrained or some characteristics of the family make the children less likely to participate in post-compulsory education. We propose a methodology that separates financial and familial effects. As in previous studies, we find that pupils from poorer families are less likely to invest in education. However, a financial transfer would not lead to a significant increase in schooling investment, which supports the view that the family characteristic effects dominate the financial constraint effects.*

## Introduction

Schooling attainment and other choices made during adolescence reflect to some extent the conditions in which children are growing up (see Haveman and Wolfe, 1995 for the UK, and Gregg *et al.*, 1999, for the UK, for recent reviews). Children from poorer backgrounds are generally observed to achieve lower outcomes later in life i.e. less schooling, more crime, higher risk of teenage pregnancy. However, the mechanism through which household income affects the child's outcomes is still unclear. This is a question of importance in order to adequately determine policies to reduce inequalities. Focusing on schooling achievement, two main theoretical strands can be distinguished.

First, as advocated by Becker and Tomes (1986), poorer families are financially constrained which prevents them from investing in the human capital of their

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The data is available after authorisation is obtained from the Economic and Research Council's Data Archive at the University of Essex.

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offspring. The effect of family income on child's attainment is direct, thus, policies of financial support could reduce the differences in achievement between children from different backgrounds.

Second, poorer parents may be endowed with characteristics that make them both less successful on the labour market and worse at parenting (Mayer, 1997). Family background characteristics might also affect the motivation, access to career information or the discount rate of the child (Card, 1999). Then, the usually observed income effect is an artifact due to the relationship between family income and some unobservable family characteristics. Therefore, direct support to the children, in the form of extra educational attention for example, would be more efficient than financial support at reducing inequality in schooling achievements.

Whether the income effect is causal, or merely reflects the correlation of income and some unobservable characteristics of the parents, remains unclear (see Mayer, 1997, for a review). The controversy that ensued from the publication of the book entitled the 'Bell Curve' (Herrnstein and Murray, 1994) is an example of the recent debate on the efficiency of financial support at reducing educational inequality (see Heckman, 1995 or Goldberger and Manski, 1995, for reviews and methodological criticisms). According to these authors, cognitive ability mainly determines success at school. The observed effect of income only reflects the correlation between ability and family wealth. As long-term improvement of cognitive ability is costly and 'of limited scope' (Herrnstein and Murray), the authors conclude that public interventions using financial incentives to reduce educational inequalities are bound to fail. Cameron and Heckman (1998, 1999) also support the idea that educational decisions do not stem from short-term financial constraints but have their origins in the long-term effects of family characteristics on ability/motivation and other unobserved characteristics (see also Cameron and Taber, 2000 or Shea, 2000 for the US).[1] Harmon and Walker (2000) for the UK rely on schooling contingent income to identify income effect (child support), but find it has no effect on the probability of staying in post-compulsory education.

On the other hand, Rice (1987) for the UK and more recently Acemoglu and Pischke (2000) for the US, find that family income has a strong effect on the child participation in post-compulsory education. Acemoglu and Pischke (2000) use change in the income distribution over time and across states to identify the effect of family income on college enrolment and estimate an income elasticity of 0.14. Furthermore, Dynarski (1999, 2000) for the US, provide some evidence that financial support can be efficient and cost effective. Dynarski (1999) for example, relies on a natural experiment, the suppression of the Social Security Student Benefit Program, to estimate that a $1000 aid increases the probability of attending college by 4% for 'poor' students. This would indicate that a policy of contingent financial support could be efficient at increasing post-compulsory education attendance.

Britain is characterized by a lower rate of post compulsory schooling than other European countries. In order to increase schooling, an education maintenance allowance (EMA) where 16- to 19-year olds would be given financial support to attend post-compulsory schooling if the family income falls below a threshold[2] is currently being piloted in the country, but this paper attempts at estimating first, the relative effect of parental characteristics and parental financial situation and then, the effect of financial transfers on the educational choice of children.

We use a model developed by Cameron and Heckman (1998) to distinguish between the direct and indirect effect of family income on schooling decisions.

More specifically, by simulating a financial support policy, we are able to maintain the observed and unobserved characteristics of the family constant, and to estimate the direct effect of financial transfers on schooling decisions. We assume that the parents' altruistic behaviour is not affected by the introduction of the benefit, i.e. parents keep on transferring to their children the same amount as before the introduction of the scheme, so that the allocation is a true increase in the education budget of the family. Our estimate is an over-estimate of the effect if parents reallocate towards themselves some of the money previously allocated to their children's education. Our estimate may therefore be substantially biased if this substitution effect is large.

To summarise our findings, as in previous studies, we find that pupils from poorer families are less likely to invest in education. A financial transfer would be effective if financial constraints dominate the family characteristic effects *i.e.* students from poorer background are financially constrained in their educational decision. We find that an education benefit would not lead to significant increase in schooling, which supports the view that the family characteristic effects dominate the financial constraint effects.

## 2. A Model of Education Decision

In this section, we review the basic model derived by Cameron and Heckman (1998). The optimal level of schooling is defined in terms of costs and returns, where the cost, $C(s|x)$, is defined to be convex in years of schooling and depends solely on time-invariant family or individual characteristics, $x$, and years of schooling, $s$. The discounted return to schooling, $R(s)$, is assumed to be a concave function of years of schooling independent of the individual characteristics. To ensure the existence of a unique optimal schooling level, the returns to zero years of schooling are assumed to be positive, whereas the costs are null. Formally, the assumptions are:

$$\begin{cases} \dfrac{\partial C(s|x)}{\partial s} > 0 \qquad \dfrac{\partial^2 C(s|x)}{\partial s^2} > 0 \qquad \text{and } C(0|x) = 0 \\ \dfrac{\partial R(s)}{\partial s} > 0 \qquad \dfrac{\partial^2 R(s)}{\partial s^2} < 0 \qquad \text{and } R(0) > 0 \end{cases} \tag{1}$$

The optimal amount of schooling $s^\star$ is then the unique solution to the maximization:

$$\max_{s \in \{0,\ldots,S\}} R(s) - C(s|x)$$

We allow for the presence of unobserved heterogeneity, and assume that the cost function has the following functional form:

$$C(s|x) = C(s)\varphi(x)\varepsilon \tag{2}$$

where $\varphi(x)$ is a function of family characteristics and the observed ability and $\varepsilon$ is a random variable accounting for the unobserved heterogeneity of each pupil. The heterogeneity may reflect differences in individual ability or any other

unobserved characteristics, which accounts for unobserved variations of the cost of reaching a certain level of schooling. Without loss of generality, we further assume that:

$$\varepsilon > 0, \qquad E[\varepsilon] = 1 \qquad \text{and } \varphi(x) > 0$$

The following system of inequalities guarantees that $s^\star$ is the optimal level of schooling.

$$\begin{cases} R(s^\star) - C(s^\star)\varphi(x)\varepsilon \geq 0 \\ R(s^\star) - C(s^\star)\varphi(x)\varepsilon \geq R(s^\star - 1) - C(s^\star - 1)\varphi(x)\varepsilon \\ R(s^\star) - C(s^\star)\varphi(x)\varepsilon \geq R(s^\star + 1) - C(s^\star + 1)\varphi(x)\varepsilon \end{cases} \tag{3}$$

Thus, for each individual, at the optimal educational level, $s^\star$, the unobserved component of the cost function, $\varepsilon$, is bounded[3]:

$$\frac{R(s^\star) - R(s^\star - 1)}{[C(s^\star) - C(s^\star - 1)]\varphi(x)} \geq \varepsilon \geq \frac{R(s^\star + 1) - R(s^\star)}{[C(s^\star + 1) - C(s^\star)]\varphi(x)} \tag{4}$$

Assuming that $\varepsilon$ is continuously distributed, the probability of choosing $s^\star$ years of schooling when growing up in a family with characteristics $x$ is:

$$\text{Prob}(s|x) = \Pr\left[\frac{R(s^\star + 1) - R(s^\star)}{[C(s^\star + 1) - C(s^\star)]^\star \varphi(x)} \leq \varepsilon \leq \frac{R(s^\star) - R(s^\star - 1)}{[C(s^\star) - C(s^\star - 1)]^\star \varphi(x)}\right] \tag{5}$$

This model will take the familiar form of an ordered probit model[4] where $\varphi(x) = \exp(-X\beta)$ and

$$l(s) = \ln\left(\frac{R(s^\star + 1) - R(s^\star - 1)}{[C(s^\star + 1) - C(s^\star - 1)]}\right)$$

and assuming that $\ln(\varepsilon)$ is normally distributed.

The ratio of marginal revenue over marginal cost can be calculated using the cut-off points deduced from the ordered probit estimation of the model.

$$\frac{mR(s = j)}{mC(s = j)} = \frac{\exp(\mu_j)}{exp(-\overline{X}\beta)} \tag{6}$$

where $\mu_j$ is the cut-off point defining the $j$th educational group *i.e.* $\Pr(s = j) = \Phi(\mu_j - \overline{X}\beta) - \Phi(\mu_{j-1} - \overline{X}\beta)$, and $\overline{X}\beta$ is measured at the average characteristics of the cohort.

Everything else equal, in particular unobservable characteristics, pupils who invest least in education experience either a reward function with a lower than average growth or a cost function with a higher than average growth, or both. Since the returns function is assumed to be positive and concave while the cost function

is assumed positive and convex (1), the ratio of marginal return over marginal cost is decreasing in years of education and tends towards zero[5].

## 3. Data

We use the National Child Development Study (NCDS) and the British Cohort Study (BCS). These two surveys were designed to observe the development of a cohort of children at different points in time. They also contain extensive information on schooling achievements and various ability measures and are therefore particularly appropriate for our analysis. Using two cohorts, we can also test the stability of our results. However, one of the drawbacks of these data sets concerns the quality of the income data. In particular, earnings are coded into a limited number of categories. In order to estimate the effect of a small change in paternal income, we attempt to transform this categorical earning variable into a continuous one using additional information extracted from the Family Expenditure Survey.

The NCDS is a continuous longitudinal survey of persons living in Great Britain who were born in the first week of March 1958. We use information collected when the respondents were aged 7, 11, 16 and 33. Respondents who are still in education at the last wave are dropped. The family background characteristics are collected when the child was 11. They include parental education, father's socio-economic group[6], number of siblings, and dummies for the presence of natural parents and race. A dummy for whether the child was brought up in a council estate captures some neighbourhood effects. Father's earnings (in grouped category) were reported in 1974 when the child was 16; this measure is used as a proxy for family income. Information on a single year is only a crude proxy for the financial situation of the household as the child was growing up (see Wolfe *et al.*, 1996, for an exposition of this problem). However, it can be argued that this is the constraint faced by the adolescent while making his/her educational choice. Additionally, many interviews in 1974 were conducted during the 'three-day week'[7]. It is unclear whether adjusted earnings were reported (Mickewright, 1986) thus the earnings variable is likely to be noisy. At age 7, all children's abilities in reading and mathematics were measured in a series of tests. As these tests were conducted at a young age, they are moderately affected by schooling already attained. These measures reflect not only the 'natural ability' of the child, but also the support, material and emotional provided by the parents.

The design of the BCS is similar to the NCDS; all children born in Great Britain in the first week of April 1970 were surveyed. Children and parents were interviewed when the child was 5, 10, 16 and 26. We focus on respondents who had completed their education at age 26. Pupils who are still in some form of higher education are dropped (341 observations). Students share a similar family background, as measured by father's social class, compared to other respondents. However, they have higher test scores, therefore their exclusion from the sample might slightly bias our results. The family background variables are similar to those defined for the NCDS but they were collected when the child was 10. The main difference in the definition of the variables concerns the measure of ability and paternal income. For children observed in the BCS, family income and ability are measured at age 10, rather than, respectively, 16 and 7 in the NCDS. We rely on ability tests taken at age 10 as they are more similar to the NCDS tests, however, they may be correlated with early schooling achievement.

The data are summarized in Table 1, by cohort and gender. As Scotland has a different educational system than England and Wales, children living in Scotland are dropped from this analysis. The number of years in education has increased by nearly one year for the younger cohort with the average school leaving age being nearly 18 years[8]. As the NCDS cohort was the first to experience a compulsory school leaving age of 16, this difference in educational achievement comes from a change in schooling decisions. As Figure 1 shows, among the younger cohort a smaller proportion left school at the earliest opportunity, 47% against 60% for the older cohort, and a larger proportion completed some form of higher education,

**Table 1**. Summary statistics

| | NCDS: Cohort 1958 | | S: Cohort 1970 | |
|---|---|---|---|---|
| Variable | Women | Men | Women | Men |
| Left school at 16 | 0.5780 (0.4940) | 0.6477 (0.4778) | 0.4436 (0.4969) | 0.5224 (0.4991) |
| Left school at 17 | 0.1262 (0.3321) | 0.0948 (0.2931) | 0.1376 (0.3446) | 0.0946 (0.2927) |
| Left school at 18 | 0.1456 (0.3527) | 0.1058 (0.3076) | 0.1600 (0.3666) | 0.1183 (0.3230) |
| Left school at 19/20 | 0.0438 (0.2048) | 0.0462 (0.2099) | 0.0573 (0.2324) | 0.0535 (0.2216) |
| Left school at 21 | 0.1064 (0.3084) | 0.1054 (0.3072) | 0.2016 (0.4012) | 0.2111 (0.4082) |
| Mother: compulsory ed. +1 | 0.1154 (0.3195) | 0.1121 (0.3156) | 0.1526 (0.3597) | 0.1438 (0.3509) |
| Mother: compulsory ed. +2 | 0.0834 (0.2765) | 0.0793 (0.2703) | 0.0727 (0.2596) | 0.0812 (0.2732) |
| Mother: compulsory ed. 3/4 | 0.0482 (0.2142) | 0.0441 (0.2053) | 0.0506 (0.2193) | 0.0617 (0.2407) |
| Mother: compulsory ed. +5+ | 0.0367 (0.1880) | 0.0328 (0.1781) | 0.1519 (0.3590) | 0.1507 (0.3578) |
| Father compulsory ed. +1 | 0.0949 (0.2931) | 0.0822 (0.2747) | 0.1209 (0.3260) | 0.1079 (0.3104) |
| Father compulsory ed. +2 | 0.0787 (0.2694) | 0.0765 (0.2659) | 0.0538 (0.2256) | 0.0548 (0.2277) |
| Father compulsory ed. +3/4 | 0.0633 (0.2435) | 0.0561 (0.2301) | 0.0541 (0.2263) | 0.0613 (0.2400) |
| Father compulsory ed. +5+ | 0.0453 (0.2080) | 0.0465 (0.2107) | 0.1705 (0.3761) | 0.1658 (0.3720) |
| Father pay: 0–50£ | 0.0327 (0.1779) | 0.0398 (0.1956) | 0.0492 (0.2164) | 0.0531 (0.2243) |
| Father pay: 50–100£ | 0.3059 (0.4609) | 0.2965 (0.4568) | 0.2557 (0.4363) | 0.2444 (0.4298) |
| Father pay: 100–150£ | 0.3598 (0.4800) | 0.3688 (0.4826) | 0.3371 (0.4728) | 0.3519 (0.4777) |
| Father pay: 200–250£ | 0.0579 (0.2335) | 0.0564 (0.2308) | 0.0674 (0.2508) | 0.0561 (0.2302) |
| Father pay: 250 +£ | 0.0737 (0.2613) | 0.0716 (0.2578) | 0.0604 (0.2383) | 0.0592 (0.2360) |
| Nbr sibling | 3.0288 (1.4924) | 3.0176 (1.4711) | 2.4537 (0.9827) | 2.4845 (0.9772) |
| Council estate | 0.4202 (0.4937) | 0.4108 (0.4921) | 0.2155 (0.4112) | 0.2185 (0.4133) |
| Father present | 0.9533 (0.2111) | 0.9577 (0.2013) | 0.8753 (0.3304) | 0.8877 (0.3158) |
| Mother present | 0.9759 (0.1533) | 0.9707 (0.1686) | 0.9762 (0.1523) | 0.9801 (0.1396) |
| White | 0.9687 (0.1741) | 0.9556 (0.2061) | 0.9728 (0.1628) | 0.9741 (0.1589) |
| Father soc 1 | 0.0500 (0.2179) | 0.0596 (0.2368) | 0.0625 (0.2421) | 0.0643 (0.2454) |
| Father soc 2 | 0.1822 (0.3861) | 0.1675 (0.3735) | 0.2218 (0.4155) | 0.2116 (0.4085) |
| Father soc 3n | 0.1006 (0.3009) | 0.0949 (0.2931) | 0.0765 (0.2658) | 0.0911 (0.2878) |
| Father soc 3m | 0.4299 (0.4952) | 0.4439 (0.4969) | 0.3626 (0.4808) | 0.3800 (0.4855) |
| Father soc 4 | 0.1707 (0.3763) | 0.1675 (0.3735) | 0.0978 (0.2971) | 0.0902 (0.2866) |
| Father soc missing | 0.0155 (0.1234) | 0.0190 (0.1367) | 0.1554 (0.3624) | 0.1382 (0.3452) |
| Math test: 25/50 | 0.2671 (0.4425) | 0.2620 (0.4398) | 0.2564 (0.4367) | 0.2198 (0.4142) |
| Math test: 50/75 | 0.2304 (0.4212) | 0.2384 (0.4262) | 0.2585 (0.4379) | 0.2275 (0.4193) |
| Math test: 75 + | 0.1636 (0.3699) | 0.1996 (0.3998) | 0.1952 (0.3965) | 0.2949 (0.4561) |
| Read test: 25/50 | 0.2746 (0.4464) | 0.2884 (0.4531) | 0.2469 (0.4313) | 0.2327 (0.4227) |
| Read test: 50/75 | 0.2480 (0.4319) | 0.2109 (0.4080) | 0.2812 (0.4497) | 0.2522 (0.4343) |
| Read test: 75 + | 0.2243 (0.4172) | 0.1569 (0.3638) | 0.2295 (0.4206) | 0.2310 (0.4216) |
| Observations | 2782 | 2836 | 2863 | 2316 |

Omitted categories are parents no compulsory education, father pay 150–200£ per week, father social class 5, bottom quartile of mathematics and English tests.

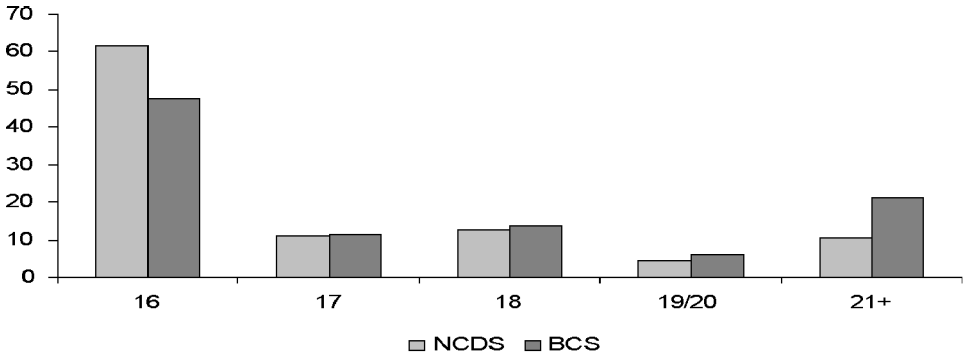


**Fig. 1**. Distribution of school leaving age by cohort

21% against 10% respectively[9]. The exit rates of education in between, largely due to dropping out are left virtually unchanged.

In both cohorts, women receive more schooling than men, but the difference is never significant. Parental schooling also differs between the two cohorts, the difference being the largest for parents with more than 4 years of post-compulsory schooling (Tertiary education)[10]. In the NCDS, 4% of fathers have achieved this level; the corresponding figure in the BCS is 17%, a similar observation can be made about the mothers' education. The paternal income is reported in 1980 prices using the retail price index. The average family size dropped from an average of 3 children per family for the older cohort to 2.5 in the BCS.

## 4. Empirical Results

We wish to measure the economic determinants described in the model presented in Section 2 for the five education/leaving age groups we observe: left school at minimum age, left school at 17, 18, 19 or 20, and older than 20. For the ordered probit estimation, the categories are numbered from 1 for pupils who left education after their 20th birthday to 5 for those who left school at 16. The reasons for the reverse ordering are purely technical and are explained below. Part of our analysis is based on the evolution of the ratio of marginal revenue to marginal cost (see equation 6) which is computed from the cut-off values obtained form the ordered probit. Since we define five education groups, we generate four thresholds hence the ratio cannot be identified for one education group. As most pupils leave school at 16, we decided to compute the ratio of revenue over cost for school leavers rather than for typically university graduates, which explains our reverse ordering of the school leaving groups. Furthermore, as this ratio of marginal revenue over marginal cost is decreasing in years of education and converges towards zero, the ratio of marginal revenue over marginal cost is likely to be small for graduates.

Estimates of the determinants of school-leaving age are presented in Tables 2a and 2b for women and for men. The parameters and mean marginal effects are reported in columns 1 and 3 for NCDS and BCS, respectively, with a specification that does not include ability measures. Due to the ordering of the dependent variable, a negative coefficient indicates a greater likelihood of transition. The results are consistent with the previous literature (among others Dearden, 1998, using the NCDS data). Parental education, father's social class and belonging to a

**Table 2a**. Determinants of age left education[1]: Women

| | NCDS: cohort 1958 – Women | | BCS: cohort 1970 – Women | |
|---|---|---|---|---|
| Mother: compulsory + 1 | **−0.2583 (0.0711)** **[−0.0510]** | −0.2575 (0.0708) [−0.0509] | −0.3423 (0.0624) [−0.0312] | −0.2547 (0.0632) [−0.0200] |
| Mother: compulsory + 2 | **−0.5348 (0.0831)** **[−0.1025]** | −0.4595 (0.0866) [−0.0924] | −0.3842 (0.0819) [−0.0344] | −0.3089 (0.0827) [−0.0239] |
| Mother: compulsory + 3/4 | **−0.7236 (0.1017)** **[−0.1457]** | −0.6514 (0.1007) [−0.1319] | −0.5818 (0.1039) [−0.0494] | −0.4597 (0.1049) [−0.0344] |
| Mother: compulsory + 5 + | **−0.9161 (0.1316)** **[−0.1826]** | −0.8444 (0.1342) [−0.1702] | −0.3018 (0.0862) [−0.0277] | −0.2505 (0.0855) [−0.0198] |
| Father: compulsory + 1 | **−0.2912 (0.0804)** **[−0.0577]** | −0.2884 (0.0803) [−0.0572] | −0.1096 (0.0716) [−0.0103] | −0.0326 (0.0745) [−0.0027] |
| Father: compulsory + 2 | **−0.3494 (0.0864)** **[−0.0697]** | −0.3261 (0.0887) [−0.0650] | −0.2576 (0.0999) [−0.0236] | −0.2403 (0.1020) [−0.0189] |
| Father: compulsory + 3/4 | **−0.4272 (0.0997)** **[−0.0856]** | −0.4277 (0.0988) [−0.0860] | −0.2308 (0.1007) [−0.0213] | −0.2148 (0.1011) [−0.0169] |
| Father: compulsory + 5 + | **−0.3881 (0.1263)** **[−0.0777]** | −0.3513 (0.1294) [−0.0703] | −0.3102 (0.0872) [−0.0285] | −0.2623 (0.0874) [−0.0207] |
| Father pay:0–50£ | 0.0800 (0.1489) [0.0151] | 0.0381 (0.1456) [0.0072] | 0.1378 (0.1174) [0.0133] | 0.0232 (0.1216) [0.0019] |
| Father pay:50–100 £ | **0.1990 (0.0734)** **[0.0375]** | 0.1527 (0.0741) [0.0289] | 0.1667 (0.0670) [0.0160] | 0.1452 (0.0679) [0.0119] |
| Father pay:100–150 £ | **0.1597 (0.0683)** **[0.0303]** | 0.1413 (0.0691) [0.0269] | 0.0778 (0.0595) [0.0074] | 0.0636 (0.0602) [0.0052] |
| Father pay:200–250 £ | 0.0877 (0.1023) [0.0166] | 0.0792 (0.1033) [0.00150] | **−0.2588 (0.0904)** **[−0.0224]** | −0.2121 (0.0907) [−0.0167] |
| Father pay:250 + £ | −0.0532 (0.0941) [−0.0120] | −0.0874 (0.0947) [−0.0170] | **−0.4176 (−0.0370)** **(0.0986)** | −0.3681 (0.1003) [−0.0281] |
| Math test: 25/50 | | −0.0376 (0.0653) [−0.0072] | | −0.1155 (0.0631) [−0.0093] |
| Math test: 50/75 | | **−0.1442 (0.0689)** **[−0.0280]** | | −0.2406 (0.0701) [−0.0191] |
| Math test: 75 + | | **−0.3164 (0.0782)** **[−0.0625]** | | −0.4863 (0.0834) [−0.0396] |
| Read test: 25/50 | | **−0.0109 (0.0721)** **[−0.0021]** | | −0.1881 (0.0667) [−0.0150] |
| Read test: 50/75 | | **−0.3918 (0.0748)** **[−0.0774]** | | −0.4191 (0.0720) [−0.0328] |
| Read test: 75 + | | **−0.5413 (0.0790)** **[−0.1079]** | | −0.8421 (0.0859) [−0.0613] |
| Cut off 1 | −1.8366 (0.2555) | −2.0637 (0.2642) | −0.7424 (0.2908) | −0.9839 (0.3051) |
| Cut off 2 | −1.5745 (0.2567) | −1.7901 (0.2634) | −0.5123 (0.2917) | −0.7275 (0.3062) |
| Cut off 3 | −0.9416 (0.2560) | −1.1301 (0.2624) | 0.0128 (0.2924) | −0.1567 (0.3067) |
| Cut off 4 | −0.5142 (0.2564) | −0.6839 (0.2630) | 0.4171 (0.2928) | 0.2741 (0.3071) |
| Observation | 2782 | 2782 | 2863 | 2863 |
| Pseudo $R^2$ | 0.1244 | 0.1471 | 0.0918 | 0.1330 |

**Table 2b**. Determinants of age left education: Men

| | NCDS: cohort 1958 – Men | | BCS: Cohort 1970 – Men | |
|---|---|---|---|---|
| Mother: compulsory + 1 | **−0.3585 (0.0728)** **[−0.0806]** | −0.3057 (0.0742) [−0.0685] | −0.2717 (0.0743) [−0.0283] | −0.2136 (0.0754) [−0.0223] |
| Mother: compulsory + 2 | **−0.5457 (0.0873)** **[−0.1253]** | −0.4989 (0.0884) [−0.1146] | −0.5233 (0.0959) [−0.0504] | −0.4081 (0.0971) [−0.0406] |

| | NCDS: cohort 1958 – Men | | BCS: cohort 1970 – Men | |
|---|---|---|---|---|
| Mother: compulsory + 3/4 | **−0.4754 (0.1031)** **[−0.1091]** | −0.4460 (0.1054) [−0.1024] | −0.7227 (0.1048) [−0.0647] | −0.5974 (0.1092) [−0.0561] |
| Mother: compulsory + 5 + | **−0.5064 (0.1374)** **[−0.1166]** | −0.4111 (0.1452) [−0.0941] | −0.3184 (0.0973) [−0.0329] | −0.2044 (0.0968) [−0.0214] |
| Father: compulsory + 1 | **−0.3499 (0.0851)** **[−0.0789]** | −0.3100 (0.0863) [−0.0697] | −0.1225 (0.0846) [−0.0131] | −0.0058 (0.0863) [−0.0006] |
| Father: compulsory + 2 | **−0.3492 (0.0916)** **[−0.0788]** | −0.2972 (0.0918) [−0.0668] | −0.1158 (0.1177) [−0.0123] | −0.0212 (0.1198) [−0.0023] |
| Father: compulsory + 3/4 | −0.1836 (0.0945) [−0.0405] | −0.1759 (0.0960) [−0.0388] | **−0.2940 (0.1015)** **[−0.0301]** | −0.2987 (0.1034) [−0.0304] |
| Father: compulsory + 5 + | **−0.4876 (0.1323)** **[−0.1119]** | −0.4318 (0.1354) [−0.0990] | −0.1952 (0.0970) [−0.0206] | −0.1398 (0.0972) [−0.0148] |
| Father pay:0−50£ | **0.4359 (0.1505)** **[0.0821]** | 0.3804 (0.1481) [0.0728] | 0.3126 (0.1350) [0.0349] | 0.3180 (0.1446) [0.0350] |
| Father pay:50−100£ | **0.2403 (0.0766)** **[0.0498]** | 0.2297 (0.0785) [0.0476] | 0.1852 (0.0764) [0.0204] | 0.1602 (0.0786) [0.0174] |
| Father pay:100−150 £ | 0.0943 (0.0698) [0.0199] | 0.0810 (0.0718) [0.0171] | 0.1184 (0.0668) [0.0129] | 0.1097 (0.0680) [0.0118] |
| Father pay:200−250 £ | −0.1359 (0.1123) [−0.0298] | −0.1302 (0.1130) [−0.0285] | −0.1901 (0.1139) [−0.0200] | −0.1844 (0.1168) [−0.0193] |
| Father pay: 250 + £ | −0.0619 (0.0947) [−0.0133] | −0.0368 (0.0962) [−0.0079] | **−0.3048 (0.1073)** **[−0.0312]** | −0.3228 (0.1115) [−0.0327] |
| Math test: 25/50 | | −0.1254 (0.0716) [−0.0271] | | −0.0623 (0.0828) [−0.0067] |
| Math test: 50/75 | | **−0.2643 (0.0763)** **[−0.0581]** | | −0.2231 (0.0891) [−0.0235] |
| Math test: 75 + | | **−0.4090 (0.0809)** **[−0.0918]** | | −0.6881 (0.0949) [−0.0676] |
| Read test: 25/50 | | **−0.1993 (0.0699)** **[−0.0433]** | | −0.1866 (0.0817) [−0.0197] |
| Read test: 50/75 | | **−0.4937 (0.0757)** **[−0.1107]** | | −0.3131 (0.0854) [−0.0327] |
| Read test: 75 + | | **−0.6798 (0.0829)** **[−0.1550]** | | −0.7703 (0.0967) [−0.0742] |
| Cut off 1 | −1.8884 (0.2698) | −2.3120 (0.2712) | 0.0528 (0.3532) | −0.2770 (0.3634) |
| Cut off 2 | −1.6280 (0.2695) | −2.0395 (0.2708) | 0.2599 (0.3539) | −0.0404 (0.3644) |
| Cut off 3 | −1.1660 (0.2692) | −1.5502 (0.2705) | 0.6534 (0.3533) | 0.4018 (0.3637) |
| Cut off 4 | −0.8277 (0.2695) | −1.1832 (0.2703) | 0.9428 (0.3534) | 0.7224 (0.3637) |
| Observation | 2836 | 2836 | 2316 | 2316 |
| Pseudo $R^2$ | 0.1182 | 0.1503 | 0.1001 | 0.1571 |

Note: Coefficient (se) [marginal effect]
The regression also includes a set of dummies for paternal class, family structure, region of residence, the number of siblings in the household, race of child and neighbourhood effect (Council estates). Robust standard errors in parentheses. Bold characters indicate significance at 5% level.
[1]A negative coefficient indicates a greater probability of transition to a higher grade.

racial minority[11] are positively correlated with more education whereas lower family income, number of siblings, and living in a council estate (not reproduced in the Tables) reduce the likelihood of transition to a higher grade. These results are similar for both genders and cohorts. For the older cohort, the paternal income effect is significant for pupils whose fathers' earnings are in the bottom of the distribution. For example, those whose fathers earn between £50 and £100 net per month (1980 prices) are 5% (4% for women) less likely to invest in post-compulsory education than pupils whose father earns between £150 and £200. For the younger cohort, pupils whose fathers are in the top earnings category are significantly more likely to stay longer in education than those whose father earns between £150 and £200. However, the differential in schooling achievement between children whose father earns more than £250 a week and children whose father earns between £50 and £100 has stayed rather similar over time[12].

In a first attempt to differentiate between direct and indirect effect of paternal income, a measure of ability is included in the model since ability is a positive function of the unobserved characteristics of the family background. Columns 2 and 4 of Tables 2a and 2b report the estimated schooling determinants when accounting for the child ability. For each test, pupils in the lowest test quartile define the omitted category. As in Gregg and Machin (1999) using the NCDS, we find that early ability tests have a large positive effect on schooling achievement. The reading test appears to have a slightly stronger effect than the maths test on the probability of investing more in education. No substantial differences between boys and girls are observed.

Family characteristics, by affecting the development of the child (as measured by ability), have a significant effect on schooling attainment. However, family income remains over and above its effect on ability, a significant determinant of schooling. The inclusion of the test scores variables does not affect our previous conclusions concerning the remaining explanatory variables. Large differences in schooling attainment appear to be explained by the financial situation of the family.

We may now compare the educational determinants for the two cohorts. First, we calculate the marginal revenue-marginal cost ratio by gender and by cohort. The ratios are defined at each cut off value as the exponential of the cut-off value divided by the exponential of $-\overline{X}\beta$, see (6). For comparison purposes, the marginal revenue-marginal cost ratio for pupils who left school at 17 is used as a base, and is set equal to unity. Figure 2 illustrates the evolution of the marginal revenue-marginal cost ratio between the two cohorts for females. The normalized ratios are similar between cohorts, indicating that schooling determinants have remained similar over time and are almost linearly decreasing in years of education. Pupils who quit school at the first opportunity have the highest relative return, whereas those who invest more in their own education see the marginal return of their investment reduced.

The stability of the determinants of education between the two cohorts may be tested more formally. To make the parameters (see Table 2a and Table 2b) comparable between equations, we divide all of them by an estimate who has remained similar between the two cohorts (number of siblings) so that the ratios are independent of the scale parameter. The null hypothesis is that the coefficients are similar between cohorts. It can be shown that this test statistic is distributed as a $\chi^2((j-1)\,(k-1))$, where $j$ is the number of cohorts and $k$ is the number of parameters. For both genders and specifications, we cannot reject the null hypothesis that the coefficients are identical between the two cohorts (see Table 3). Despite the

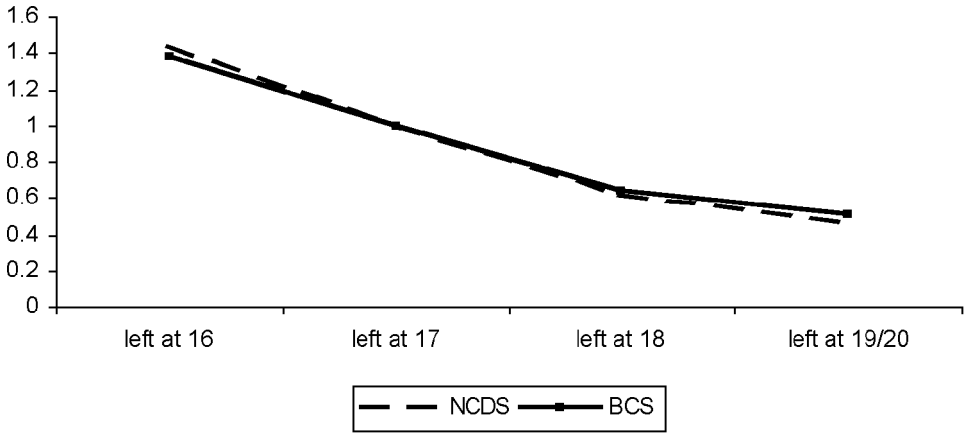


**Fig. 2.** Ratio marginal revenue-marginal cost

Note: The marginal revenue-marginal cost ratios presented are derived from the estimates based on the equations including ability for females. The exclusion of the ability measures does not change the general trend. The ratios for men follow a similar pattern.

observed changes in the educational attainment between the two cohorts previously observed, the determinants of the school choice have remained stable over time. This result holds for both genders and with the inclusion of the ability measures.

Family income during childhood appears to have a major effect on the schooling decision. However, this income effect may be spurious as some characteristics of the family may jointly explain the poor financial situation and the educational decision of the children[15]. The scope for income support policies depends clearly on the relative magnitudes of financial constraints versus other family characteristics. Therefore, we model the effect of a financial transfer on schooling decision. This technique allows us to relax the financial constraints but maintain the family characteristics, thus capturing the pure effect of family income on the schooling decision.

## 5. The Effect of a Financial Transfer

An Educational Maintenance Allowance (EMA) is currently being piloted in 15 areas, it provides 16–19 year-olds from poorer families (annual income lower than £13,000) with a financial allowance of £30 or £40 per week depending on the piloting area, if they remain in full-time education after year 11[16]. The scheme is means-tested and the amount received declines linearly down to £0 for children

**Table 3.** Test of stability of the educational determinants between cohorts

| | Female | Male | $X^2$ critical value, p = 0.0025 |
|---|---|---|---|
| Without ability measure | 12.32 | 13.9 | $X^2(30) = 16.80$ |
| With ability measure | 7.87 | 4.08 | $X^2(36) = 20.91$ |

from a family with an annual taxable income higher than £30,000. Children from families with taxable income greater than this threshold are not eligible for EMA. The amount of the EMA is not deducted from any other benefits that the family may receive and is therefore a real increase to the family income. The piloting areas are divided between area where the EMA is allocated to the pupil and others where the money goes to the mother. Additionally, bonuses are paid on performance to encourage educational effort and not only attendance.

To assess the potential effect of financial transfer on schooling decisions, we would like to estimate the effect of a simpler and more generous scheme (+ £30 a week for all pupils, given to the father) on the pupils from the 1970 cohort. The difficulty encountered is that the available data on paternal income for the cohort of interest (BCS, pay measured in 1980) are grouped. Hence we are faced with two solutions. Either we match this information with another survey, derive a continuous earnings variable and then implement the policy of income support. Or, as the mapping technique is rather cumbersome, we check the validity of our results with a much cruder method where we move each father into the above income category. This corresponds to an increase in earnings of £50 in 1980 (£122 in 1999 price) that is three times more than the piloted EMA.

We first present the mapping using data from the Family Expenditure Survey (FES). The FES is an annual survey of 10,000 households in the UK, which provides extensive information on earnings but none on children's education. We use the 1980 survey of the FES to map on to the earnings variable from the BCS[17].

The model we present and estimate in the previous sections can be understood as a model where the endogenous latent variable, $y^\star$, is the part of the cost which is individual specific *i.e.* $y^\star = \ln(\varphi(x)) + \ln(\varepsilon)$, see equation (2). We will assume that the correct specification is:

$$y^\star = x\beta + f(z) + \ln(\varepsilon) \tag{7}$$

where $x$ is a vector of observable family characteristics, $z$ is a continuous variable in earnings, $f(z)$ is some non-linear function of $z$ that can be represented exactly by a polynomial of order q, and we have $f(z) = \mathbf{z}\alpha$. The row vector $\mathbf{z}$ is such that the $p$th column in $\mathbf{z}$ is $z^{p-1}$, for $p \in \{1,2,\ldots q\}$, and where $\alpha$ is the vector of parameters that defines the polynomial. The assumptions made above ensure that $\ln(\varepsilon)$ is distributed independently of $x$ and $\mathbf{z}$. Clearly, the observed schooling levels $s$ are transformations of latent dependent variable $y^\star$; (see equation (4)).

We consider two samples. The FES sample (sample **A**) contains information about $x$ and $z$, but no information about $s$. The BCS sample (sample **B**) is such that we observe $s$ but we do not observe $z$. Instead, we observe for all observations whether $z$ belongs to a given interval among a set of $m$ disjoint intervals which cover the range of $z$, that is the information about $z$ is summarized by a vector of $m$ dummy variables. The vector $x$ has to be identical between the two samples, thus we simplify our previous specification and keep only variables on mother's and father's education, number of siblings, parents' marital status as well as the regional dummies. The major disadvantage of this basic specification is that ability measures can no longer be included, as they are included only in one sample, thus we are likely to overestimate the income effect. Since we introduce five dummies to describe the father's pay distribution in the BCS, we fit a quartic polynomial in earnings[18]. The estimated polynomial function, (f (z)), is slightly decreasing in earnings (see Figure 3). The difference between males and females is not statistically significant.

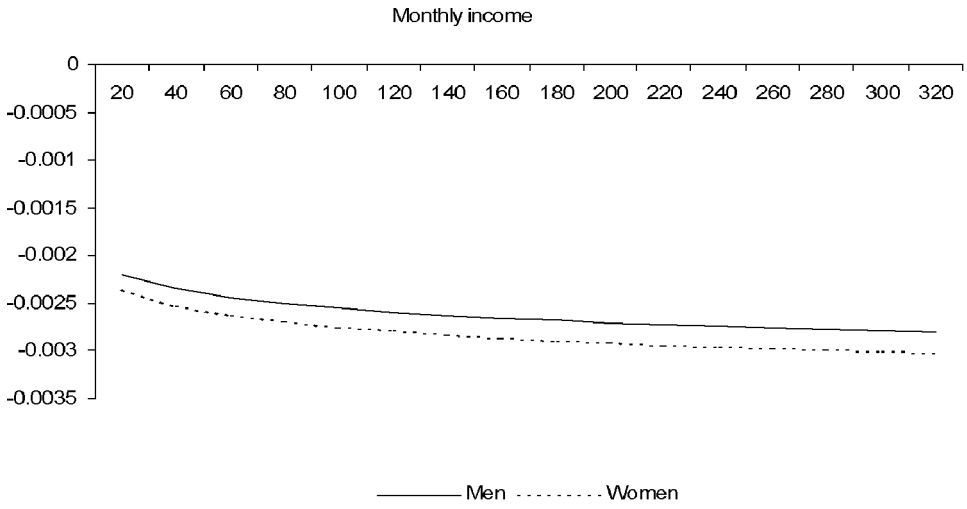


**Fig. 3.** Corrected estimates of the earnings effect on educational choice
Note: The height is defined up to an additive constant, thus only relative analysis can be conducted.

The estimated values of the schooling determinants are used to calculate the contribution to the cost function, which depends on family background: ($\varphi(x) = \exp(-X\beta)$). Using the distribution of the costs, the cut-off values of the ordered probit are corrected so that at the mean cost, the probabilities defined are identical to the probabilities observed in the original sample (BCS). Using the corrected threshold values and the earnings polynomial, we define corrected values of the ratio of marginal revenue-marginal cost for men and women. The ratio for pupils leaving school at 17 is fixed at unity for comparison purposes. These ratios of the marginal revenue-marginal cost are represented in Figure 4. The pattern is similar to the one observed without correction (see Figure 2).

Now that we have a continuous paternal earning variable, we can estimate the effect of our hypothetical policy. Adding £30 to the father weekly earnings and keeping the other family characteristics constant also gives us an idea of the relative effect of family characteristics and financial constraints on the decision to invest in

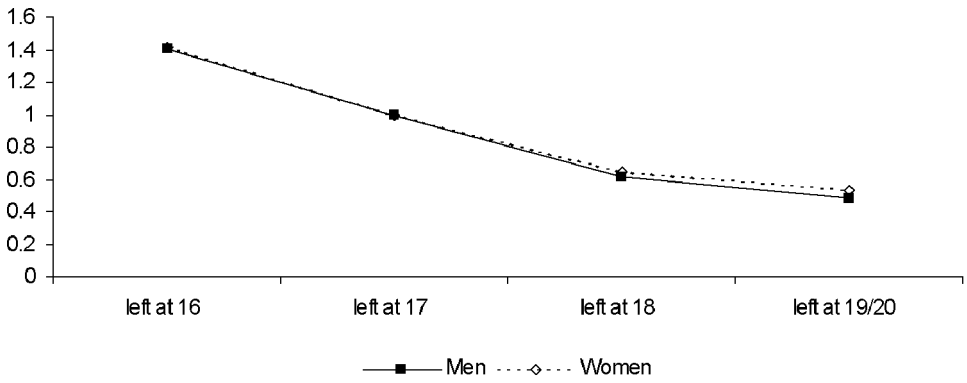


**Fig. 4.** Corrected marginal revenue–marginal cost ratio

post-compulsory schooling. Table 4 reports the probabilities of leaving school at a given age for men[19]. In the upper panel of Table 4, the school leaving age probabilities are reported by cost decile (*i.e.* deciles of $\varphi(x)$ corrected for the introduction of the polynomial in earnings). Pupils with the highest educational cost (decile 1) have a probability of exiting education at the first opportunity of 87%. This probability of quitting school at compulsory schooling age is only 4% for children with the most favourable background.

We calculate the effect on the school leaving age distribution of an educational allowance that would affect all children irrespectively of their paternal pay and add £12.30 per week (equivalent of £30 in 1999) to all fathers' earnings. The results of this transfer on schooling decisions are reported in the right hand side of the upper panel of Table 4. At all levels of the cost distribution, the effect of such an educational allowance on the school leaving age probability is marginal.

We replicate the calculations when grouping the population by paternal earnings decile, as a positive effect is more likely for poorer pupils. These results are presented in the lower panel of Table 4. The variation in the school leaving age distribution by income decile is not as severe as with the cost-deciles; pupils from the poorest decile have a probability of 62% of leaving school at 16, whereas for the richest the probability is 30%. However, the effects of a financial transfer are again

**Table 4**. Probability of leaving school: men

| Cost decile | Before the reform | | | | | Post-school maintenance reform | | | | |
|---|---|---|---|---|---|---|---|---|---|---|
| | Age 21 + | Age 19/20 | Age 18 | Age 17 | Age 16 | Age 21 + | Age 19/20 | Age 18 | Age 17 | Age 16 |
| 1 | 0.0163 | 0.0124 | 0.0429 | 0.0518 | 0.8765 | 0.0165 | 0.0126 | 0.0434 | 0.0523 | 0.8753 |
| 2 | 0.0385 | 0.0247 | 0.0751 | 0.0797 | 0.7820 | 0.0390 | 0.0250 | 0.0757 | 0.0801 | 0.7802 |
| 3 | 0.0664 | 0.0369 | 0.1024 | 0.0986 | 0.6957 | 0.0672 | 0.0372 | 0.1030 | 0.0990 | 0.6935 |
| 4 | 0.0916 | 0.0462 | 0.1206 | 0.1091 | 0.6324 | 0.0927 | 0.0465 | 0.1212 | 0.1095 | 0.6301 |
| 5 | 0.1123 | 0.0528 | 0.1324 | 0.1148 | 0.5878 | 0.1134 | 0.0531 | 0.1330 | 0.1150 | 0.5854 |
| 6 | 0.1399 | 0.0605 | 0.1449 | 0.1196 | 0.5351 | 0.1413 | 0.0609 | 0.1454 | 0.1198 | 0.5327 |
| 7 | 0.1682 | 0.0673 | 0.1546 | 0.1221 | 0.4877 | 0.1698 | 0.0677 | 0.1551 | 0.1222 | 0.4853 |
| 8 | 0.2438 | 0.0808 | 0.1692 | 0.1213 | 0.3850 | 0.2457 | 0.0811 | 0.1694 | 0.1212 | 0.3826 |
| 9 | 0.4422 | 0.0937 | 0.1636 | 0.0962 | 0.2042 | 0.4446 | 0.0937 | 0.1633 | 0.0959 | 0.2026 |
| 10 | 0.7871 | 0.0599 | 0.0784 | 0.0334 | 0.0412 | 0.7887 | 0.0596 | 0.0779 | 0.0331 | 0.0407 |
| Income decile | Before the reform | | | | | Post-school maintenance reform | | | | |
| | Age 21 + | Age 19/20 | Age 18 | Age 17 | Age 16 | Age 21 + | Age 19/20 | Age 18 | Age 17 | Age 16 |
| 1 | 0.1336 | 0.0446 | 0.1080 | 0.0938 | 0.6200 | 0.1346 | 0.0449 | 0.1085 | 0.0941 | 0.6180 |
| 2 | 0.1198 | 0.0427 | 0.1043 | 0.0913 | 0.6418 | 0.1208 | 0.0430 | 0.1047 | 0.0916 | 0.6398 |
| 3 | 0.1266 | 0.0441 | 0.1071 | 0.0933 | 0.6290 | 0.1276 | 0.0443 | 0.1075 | 0.0936 | 0.6270 |
| 4 | 0.1819 | 0.0506 | 0.1149 | 0.0941 | 0.5585 | 0.1830 | 0.0508 | 0.1153 | 0.0942 | 0.5566 |
| 5 | 0.1674 | 0.0508 | 0.1189 | 0.0993 | 0.5637 | 0.1686 | 0.0510 | 0.1193 | 0.0995 | 0.5616 |
| 6 | 0.1977 | 0.0570 | 0.1286 | 0.1029 | 0.5137 | 0.1990 | 0.0573 | 0.1290 | 0.1030 | 0.5117 |
| 7 | 0.2251 | 0.0640 | 0.1382 | 0.1058 | 0.4669 | 0.2266 | 0.0642 | 0.1385 | 0.1058 | 0.4649 |
| 8 | 0.2553 | 0.0598 | 0.1266 | 0.0967 | 0.4616 | 0.2568 | 0.0600 | 0.1268 | 0.0967 | 0.4597 |
| 9 | 0.2479 | 0.0591 | 0.1252 | 0.0957 | 0.4721 | 0.2493 | 0.0593 | 0.1254 | 0.0958 | 0.4702 |
| 10 | 0.4556 | 0.0621 | 0.1111 | 0.0730 | 0.2982 | 0.4571 | 0.0621 | 0.1110 | 0.0729 | 0.2968 |

small, changing the probability of exiting after compulsory schooling by a few tenths of a percent. This indicates that children's schooling achievement is dominated by the effect of family characteristics. Paternal earnings have a statistically significant effect on the educational choice of children, but a financial transfer policy does not generate any economically significant changes.

We finally test whether these results are robust. Going back to the BCS where the paternal earnings variable is discrete, we shift each individual to the above category (with the exception of the top group), this is equivalent to a £50 (in 1980 prices) transfer. This fictitious benefit is three times more important than the piloted one. For an annual cost of £3,600 per pupil, it will decrease the average probability of leaving school at 16 by 6% (from 52% to 46%) for males and 11% (from 44% to 33%) for females. Also, as stated previously, this could be an over-estimate of the effect of the education benefit, as it does not account for possible substitution effects. This latest projection confirms the limited impact that financial incentives have on the probability of staying on in post compulsory education. Unless large transfers of money are provided, family characteristics are likely to dominate financial constraints as a determinant of children educational choices.

## 6. Conclusion

It is commonly advocated that financial constraints prevent pupils from the poorer backgrounds investing in their own education. Previous research has shown the negative impact on educational attainment of being brought up in a poorer household. However, the effect of family income on the child's educational attainment is unclear, as it is related to other family characteristics that also affect the schooling decision. We propose a methodology that separates these effects, by holding constant the family characteristics while allowing for changes in income. Similarly to Harmon and Walker (2000), we find that the effect of family income on a child's schooling attainment is rather limited and is dominated by the effect of other family characteristics, mostly the parental education[20]. It is worth pointing out that the current piloting may lead to different results as our study is based on an older cohort, when attendance to post-compulsory education was much lower, and our financial transfer goes to the father.[21]

A policy of financial transfers appears to be economically insignificant at increasing schooling achievement, and educational choices appear to mostly reflect family characteristics. It is therefore arguable that a policy of financial transfer would be the most effective at increasing post-compulsory education decision. Its effects may be too late for an adolescent to revise his/her human capital investment strategies. It would be of interest to compare the relative effectiveness of a financial transfer and other policies aiming to increase children's ability at an earlier age (e.g. STAR experiment, Head Start) or reducing disparities due to differences in family characteristics (*e.g.* enriching the information set of adolescents).

## Notes

1. Alternatively, the improvement of childhood conditions for children at risk is generally viewed as a promising policy reducing inequality (see Heckman, 1999 for a survey).
2. Also, a reduction of the inequality in educational attainment is usually seen as a way to reduce intergenerational transmission of poverty (see Dearden *et al.* (1997) for evidence).

3. Note that the model is not observationally distinct from a model where the revenue function and the cost function have the following functional forms:

$$\text{either } R(s|x,\eta,\varepsilon) = R(s)\psi(x,\eta,\varepsilon)$$

$$\text{and } C(s|x,\eta,\varepsilon) = C(s)\psi(x,\eta,\varepsilon)\varphi(x)\varepsilon, \text{ or } R(s|x,\eta,\varepsilon) = R(s)\frac{\psi(x,\eta,\varepsilon)}{\varphi(x)\varepsilon}$$

$$\text{and } C(s|x,\eta,\varepsilon) = C(s)\psi(x,\eta,\varepsilon),$$

   where $\eta$ is some unobservable and $\psi(x,\eta,\varepsilon)$ is any positive function of $x$, $\eta$ and $\varepsilon$. Indeed the expression for the probability of observing a given level of schooling does not change. The identification of the cost and return functions is thus impossible.
4. A large part of the Cameron and Heckman (1998) contribution studies the condition under which the model is non-parametrically identified. The data we use does not allow us to identify non-parametrically the distribution of the unobserved heterogeneity.
5. Another model of schooling decision for the UK could take the form of an ordered probit with only three categories: exit education at 16, 18 or 21 (see Figure 1). It can be argued that these three exit points are the ones where a schooling decision is made, the other exit points are mostly drop-outs. Such a model assumes that drop-outs are being failed by the system. However, we assume that a drop-out reflects the reconsideration of one's educational choice.
6. Hanusheck (1992) and Feinstein and Symons (1999) stress the importance of parental interest in the child's education (time spent with child) as a significant factor explaining schooling attainment. Parents from higher socio-economic groups tend to spend more time with their children either because they have fewer children or because they value education more than other parents.
7. In 1974, miners strikes led to power failures; a number of industries reacted by cutting their working week to three days.
8. We use years of education rather than qualifications (as in Blundell *et al.*, 2000) since we are not interested in returns to education but in the effect of an EMA on educational investment, whatever this investment consists of.
9. These figures based on age left full-time education understate the education attained by the respondents, as a large proportion of pupils would have gone on to apprenticeship and other forms of part-time education. However, as the current EMA scheme is targeted at 16 year olds in full-time education, we believe that this is the variable of interest.
10. We use post-compulsory schooling as opposed to years of schooling since the minimum school leaving age was increased in 1948 from 14 to 15. The observed increase in education between the two generations of parents is therefore not picking up the effect of the change in minimum school leaving age but instead a real increase in the decision to invest in post-compulsory education.
11. The data used does not allow us to differentiate between the different ethnic minorities. On average, ethnic minorities have a greater likelihood to stay in education after compulsory schooling but variations between ethnic groups are important (see Leslie and Drinkwater, 1999).
12. Marginal effects are estimated at the mean characteristics of each sample, therefore comparisons between samples are possible, only if the means are similar.
13. We also included interactions between paternal income and parental education; however, the interaction terms were not found to be significant.
14. We also included interactions between paternal income and parental education, however, the interaction terms were not found to be significant.
15. We also included interactions between paternal income and parental education, however, the interaction terms were not found to be significant.
16. To benefit from EMA, the pupil, one parent and an educational institution have to complete a learning agreement. Payment would be suspended as soon as the pupil breaks the agreement (truancy, exclusion).
17. The details concerning the mapping process can be obtained from the authors.
18. Exact details of the procedure is available from the authors upon request.
19. Results for females are similar and are not reproduced here.

20. The lack of significance of the paternal income for some regressions could be due to the colinearity of father's pay with father's education. This may have led us to underestimate the effect of father's income and therefore undermine the positive effect of a financial transfer.
21. Duflo (1999) provides evidence that males may be bad agent for their family.

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