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The Relative Effect of Family and Financial Characteristics on Educational Achievement

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Non-Technical Summary

Britain is characterised by a low rate of post compulsory schooling compared to other European countries. To reduce this disparity, the British government has been testing an Education Maintenance Allowance (EMA) where 16 to 19-year olds are given financial support to attend schooling when the family income falls below a threshold. This paper attempts at first separating family and income effects and second estimating the impact of a financial transfer on educational attainment.

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 which prevents them from investing in the human capital of their offspring, thus, policies of financial support could be efficient at reducing schooling inequality. Or, some parents may be endowed with characteristics that make them less successful on the labour market and worse at parenting. Then, direct support to the children would be more efficient than financial support at reducing inequality in schooling achievements.

We propose a methodology that separates financial and familial effects. By simulating a financial support policy, we are able to maintain the observed and unobserved characteristics of the family constant, and estimate the direct effect of an education benefit on post compulsory schooling decisions.

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. These results may nevertheless be dependent on our methodology. Due to data constraints, we have allocated the financial transfer to the father income, whereas in the currently tested scheme, the EMA goes either to the child or to the mother. These changes in the recipient of the allocation may have a large effect on our conclusion.

The Relative Effect of Family Characteristics and Financial Situation: the Effect of Financial Transfers on Educational Achievement

Arnaud Chevalier and Gauthier Lanot

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Introduction

Schooling attainment and other choices made during adolescence reflect the conditions in which children are growing up (see Haveman and Wolfe, 1995 for the US and Gregg *et al.*, 1999, for the UK, for recent reviews). Children from poorer backgrounds are generally observed to have lower outcomes (less schooling, more crime, higher teenage pregnancy rate), 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 theories 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 offspring. The effect of family income on child's attainment is direct, thus, policies of financial support could be efficient at reducing the differences between children from different backgrounds.

Second, poorer parents may be endowed with characteristics that make them less successful on the labour market and worse at parenting (Mayer, 1997). Additionally, the family background characteristics might 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 colinearity of the 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. 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. The argument of these authors is affected by the method used (see Heckman, 1995 or Golderberg and Manski, 1995 for reviews). 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[1]. Hence, the efficiency of income support policies in helping pupils from less favourable backgrounds to invest in their schooling is questionable (see also Cameron and Taber, 2000 or Shea, 2000 for the US)[2]. Harmon and Walker (2000) for the UK rely on schooling contingent income to identify income effect, but find no significant effect on the staying on probability.

On the other hand, Rice (1987) for the UK and more recently Acemoglu and Pischke (2000) and Dynarski (1999, 2000) for the US provide some evidence that financial support can be efficient and cost effective. For example, Dynarsky (1999) uses 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. 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 elasticity of 0.14.

Britain is characterised by a low rate of post compulsory schooling compared to 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[3] 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. The preliminary evaluation of the first year of EMA indicates that participation in post compulsory education increases by 3 to 11 percentage points in pilot areas compare to areas where EMA was not available (DfEE, 2001).

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 estimate the direct effect of financial transfers on

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[1] Feinstein (2000) stresses that parental education and income as well as measures of the child's psychological development at age 10, have a major effect on school attainment, even when controlling for ability. The psychological attributes can be seen as being the outcome of long-term family characteristics.

[2] 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).

[3] Also, a reduction of the inequality in educational attainment is usually seen as a way to reduce intergenerational poverty.

schooling decisions. Our estimate would be biased upwards if parents reallocate transfers to their children after the implementation of the benefit.

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 child development effect *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. The differences with the pilot scheme results could be due to methodological differences but also to dissimilarity in the definition of education benefit. In the pilot scheme, the financial transfer goes directly to the child, or to the mother depending on area, whether in our estimation, the benefit goes to the father. The disparity in the effect of an education benefit may therefore indicate that fathers are not good agents for their children[4][5].

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 returns to schooling, $R(s)$, are defined as a concave function of years of schooling independent of the individual characteristics. The intersection between the marginal cost and returns functions defines the optimal level of schooling. To insure 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:

[4] The decision to register in post-compulsory education was not significantly different between areas where EMA was received by the child and those where the mother was the recipient (DfEE, 2001).

[5] See also Duflo (1999) for gender effect in within family financial transfer.

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

The optimal amount of schooling $s*$ is then the unique solution to the maximisation:

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

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

$$C(s|x) = C(s) j\,(x) e\ , \quad (2)$$

where $j\,(x)$ is a function of family characteristics and the observed ability and e is a random variable accounting for the 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:

$$\mathrm{E}[e] = 1\,,\ e > 0 \text{ and } j\,(x) > 0\,.$$

The following system of inequalities guarantees that $s*$ is the optimal level of schooling.

$$\begin{cases} R(s*) - C(s*) j\,(x) e \geq 0 \\ R(s*) - C(s*) j\,(x) e \geq R(s*-1) - C(s*-1) j\,(x) e \\ R(s*) - C(s*) j\,(x) e \geq R(s*+1) - C(s*+1) j\,(x) e \end{cases} \quad (3)$$

Thus, for each individual, at the optimal educational level, s^*, the unobserved component of the cost function, e, is bounded[6]:

[6] Note that the model is not observationally distinct from a model where the revenue function and the cost function have the following functional forms: either $R(s|x,h,e) = R(s) y\,(x,h,e)$ and

$$\frac{R(s^*)-R(s^*-1)}{[C(s^*)-C(s^*-1)]j(x)} \geq e \geq \frac{R(s^*+1)-R(s^*)}{[C(s^*+1)-C(s^*)]j(x)} \tag{4}$$

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

$$\text{Prob}(s\,|\,x)=\Pr\left[\frac{R(s^*+1)-R(s^*)}{[C(s^*+1)-C(s^*)]*j(x)} \leq e \leq \frac{R(s^*)-R(s^*-1)}{[C(s^*)-C(s^*-1)]*j(x)}\right] \tag{5}$$

This model will take the familiar form of an ordered probit model[7] where $j(x)$=exp(-Xb) and $l(s)=\ln\left(\frac{R(s^*+1)-R(s^*-1)}{[C(s^*+1)-C(s^*-1)]}\right)$, and assuming that $\ln(e)$ 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.

$$\frac{mR(s=j)}{mC(s=j)}=\frac{\exp(m_j)}{\exp(-\overline{X}\hat{b})} \tag{6}$$

where m_j is the cut-off point defining the j^{th} educational group *i.e.* Pr(s=j)=$F(m_j - \overline{X}\hat{b})$-$F(m_{j-1} - \overline{X}\hat{b})$,and $\overline{X}\hat{b}$ is measured at the average characteristics of the cohort.

Pupils who invest least in their education have a cost function with the largest first derivative. Also, returns to education are concave; hence the ratio of marginal return over marginal cost is decreasing in years of education and tends towards zero[8].

$C(s\,|\,x,h,e)=C(s)y(x,h,e)j(x)e$, or $R(s\,|\,x,h,e)=R(s)\frac{y(x,h,e)}{j(x)e}$ and $C(s\,|\,x,h,e)=C(s)y(x,h,e)$, where h is some unobservable and $y(x,h,e)$ is any positive function of x, h and e. 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.

[7] 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.

[8] 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 and are truly external choice. However, we assume that drop-outs reflect

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.

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[9], 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 example). However, this is the constraint that the adolescent is facing while making his choice of investing in more education. Additionally, many interviews in 1974 were conducted during the "three-day week"[10]. 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

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students reconsidering their educational choice.

[9] 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.

[10] In 1974, miners strikes led to power failures, a number of industries reacted by cutting their working week to three days.

child was 5, 10, 16 and 26. We focus on respondents who had completed their education at age 26. Pupils who are still in 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. Earnings at age 16 were poorly reported.

The data are summarised 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[11]. 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, 21% against 10% respectively[12]. 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)[13]. 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

[11] 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.

[12] 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.

[13] 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.

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. The cut-off values obtained from the ordered probit measure the critical ratio of marginal revenue to marginal cost and define the threshold for being in a given category. Since we define five education groups, we generate four thresholds; these values are used to calculate the ratio of marginal revenue over marginal cost for the four educational groups. The ordering of the leaving age group is therefore important, and allows us to calculate either the ratio for 16 year olds or for graduates. 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 revenue over cost can be approximated as being null for graduates.

Estimates of the determinants of school-leaving age are presented in Table 2a for women and Table 2b 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). Parental education, father's social class and belonging to a racial minority[14] are positively correlated with more education whereas lower family income, number of siblings, and living in a council estate (not reproduced in Tables) reduce the likelihood of transition to a higher grade. These results are similar for both

[14] 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).

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 6% (4% for women) less likely to invest in post-compulsory education than pupils whose father earns more than £250. For the younger cohort, pupils whose fathers are in the top earnings category are significantly more likely to stay longer in education. 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[15].

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 quartile define the omitted category. On average pupils with better scores in reading and mathematics at an early age achieve substantially more schooling than other children. As in Gregg and Machin (1999) using the NCDS, we find that early ability tests have one of the largest effects 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. As ability measured at an early schooling stage is an important determinant of educational success, policies aiming at providing support during childhood, *e.g.* child care or access to library facilities, might be promising ways to reduce inequalities between children[16].

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 adolescent. However, it can be argued that this relation is spurious as parents with poor parenting ability may also be less successful on the labour

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[15] Marginal effects are estimated at the mean characteristics of each sample, therefore comparisons between sample are possible, only if the means are similar.

[16] The effect of this type of measure is usually difficult to measure. For example, access to a library increases the likelihood of investing in education, however, if libraries are more often frequented by pupils from a middle class background, more libraries might increase the dispersion of educational achievement between poor and other children.

market. The scope for income support policies depends clearly on the relative magnitudes of financial constraints versus other family characteristics[17].

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 $-\overline{X}\vec{b}$, see (6) for the definition used. 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 normalised 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 a chosen estimate (here the coefficient of the 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 the test statistic is distributed as a $c^2((j\text{-}1)\ (k\text{-}1))$, 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 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 not be independent of other characteristics of the family that may jointly explain the poor financial situation and the educational decision. 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.

[17] We also included interactions between paternal income and parental education, however, the interaction terms were not found to be significant.

5. The Effect of a Financial Transfer

So far we have found that being brought up in a poorer family has a significant negative effect on educational attainment. As education is a well-proven device to reduce adult poverty, traditional policies attempting to reduce intergenerational transmission of poverty have been concerned with helping pupils from poorer backgrounds invest in post-compulsory education. Additionally, a more educated workforce is associated with higher economic growth (see among others Gemmel, 1996). The British government has been testing an Educational Maintenance Allowance (EMA). This scheme, initially piloted in 15 areas (now extended to 56), provide 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[18]. The scheme is means-tested and the amount received declines linearly down to £5 for children from a family with an annual taxable income nearing £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[19].

To assess the potential effect of financial transfer on schooling decisions, we estimate the effects of a simpler and more generous scheme (+ £30 a week for all pupils, given to the father) on 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 in 6 categories only. Hence we are faced with two solutions. Either we match this information with another survey providing a continuous earnings variable, which is a cumbersome technique. Or we replace each income dummy by the following one, which 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

[18] 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).

[19] Piloting started in September 1999, see DfEE (2001) for further details on the scheme, its testing and the first results.

information on earnings but none on children and their education. We use the 1980 survey of the FES to map on to the earnings variable from the BCS (see Appendix for details).

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

$$y^* = xb + f(z) + \ln(e) \qquad (7)$$

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 $f(z) = \mathbf{z}a$. The row vector $\mathbf{z}$ is such that the p th column in $\mathbf{z}$ is z^{p-1}, for $p \in \{1,2,...q\}$, and where a is the vector of parameters which define the polynomial. The assumptions made above ensure that $\ln(e)$ is distributed independently of x and $\mathbf{z}$. Clearly, the observed schooling levels s are transformations of latent dependent variable y^*; 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 summarised by a vector of m dummy variables. The vector x of continuous variables 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. The estimated polynomial function (f (z)) is slightly decreasing in earnings (see Figure 3). The difference between males and females is not statistically significant.

The estimated values of the schooling determinants are used to calculate the contribution to the cost function, which depends on family background: ($j\,(x) = \exp(-Xb)$). 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.

Table 4 reports the probabilities of leaving school at a given age for men[20]. In the upper panel of Table 4, the school leaving age probabilities are reported by cost decile (*i.e.* deciles of $j(x)$ 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 but accounting for their family characteristics. We add £12.30 per week (equivalent of £30 in 1999) to all fathers' earnings. The results of this transfer are reported in the right hand side of the top 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; the poorest 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 insignificant, changing the probability of exiting after compulsory schooling by a few tenths of a percent. The results indicate that children's schooling achievement is dominated by the effect of parental education and family structure (marital status, number of siblings). The change in income generated by the transfer shifts the cost function only marginally.

We finally test that these results are not a result of the procedure that we use to define the polynomial function in earnings. We use the BCS data where the paternal earnings variable is discrete. Each earnings dummy represents a range of £50 in 1980, equivalent to £122 in 1999. We shift each individual to the above category (with the exception of the top

[20] Results for females are similar and are not reproduced here.

group). This fictitious benefit is three times more important than the piloted one. For a cost of £5000 per pupil, it will decrease the probability of leaving school at 16 for the average individual 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 a possible reduction in the transfers from parents to children that may take place with the introduction of the transfer. This latest projection confirms the limited impact that financial incentives have on the probability of staying on past compulsory education.

6. Conclusion

Governments look at incentives to increase the educational attainment of youths for two main reasons: to reduce intergenerational transmission of inequality and to increase future economic growth. 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[21].

Thus, as families do not appear to be financially constrained a policy of financial transfers does not appear to be successful at increasing schooling achievement. However, this lack of significance, which is in contradiction with the first evidence from the piloting, may come from our methodology. We have assumed that the financial transfer was made to the father, however its effects may be drastically different if the child receives the money directly. Parents may be bad agent for their children; the current experience of EMA tests this hypothesis: some children receive directly the EMA whereas for another group, the transfer is through the mother. Preliminary results suggest no differences between areas

[21] 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.

where the EMA is allocated to the child and those where the mother is the recipient (DfEE, 2001). However, Duflo (1999) shown some evidence, that females are better agents than males to redistribute within the family. Additionally, we do not account for bonuses that will be paid depending on attendance and results. Finally, children affected now, were born 15 years later than those we observed and our results may reflect some cohort effect. All those may explain the differences between our results and the piloting results.

As family characteristics appear to be more important than the financial situation on the family when the adolescent is making his education choice, it is arguable that a policy of financial transfer is the most effective at increasing post-compulsory education decision. Its effects may be too belated as an adolescent' characteristics may make a revision of his investment in human capital strategies impossible. It would be thus 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 (STAR experiment, Head Start) or reducing disparities due to differences in family characteristics (*e.g.* enriching the information set of adolescents).

Table 1: Summary statistics

Variable	NCDS: Cohort 1958		BCS: Cohort 1970	
	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.

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	-.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

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 type of accommodation (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.

Table 2b: Determinants of age left education[1]: 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]**
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 type of accommodation (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.

Table 3: Test of stability of the educational determinants between cohorts

	Female	Male	χ^2 Critical value , p=0.025
Without ability measure	12.32	13.9	$\chi^2(30)=16.80$
With ability measure	7.87	4.08	$\chi^2(36)=20.91$

Table 4: Probability of leaving school: Men

	Before the reform					Post school maintenance reform				
Cost decile	**Age 21+**	**Age 19/20**	**Age 18**	**Age 17**	**Age16**	**Age 21+**	**Age 19/20**	**Age 18**	**Age 17**	**Age16**
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

	Before the reform					Post school maintenance reform				
Income decile	**Age 21+**	**Age 19/20**	**Age 18**	**Age 17**	**Age16**	**Age 21+**	**Age 19/20**	**Age 18**	**Age 17**	**Age16**
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

Figure 1: Distribution of school leaving age by cohort

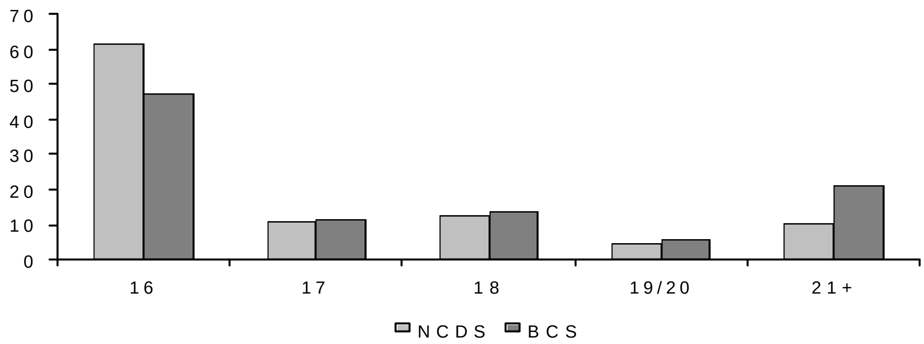
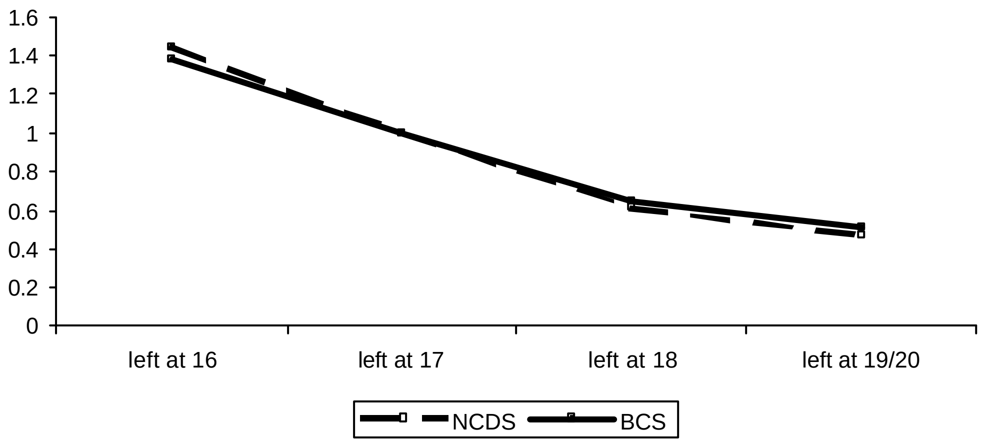
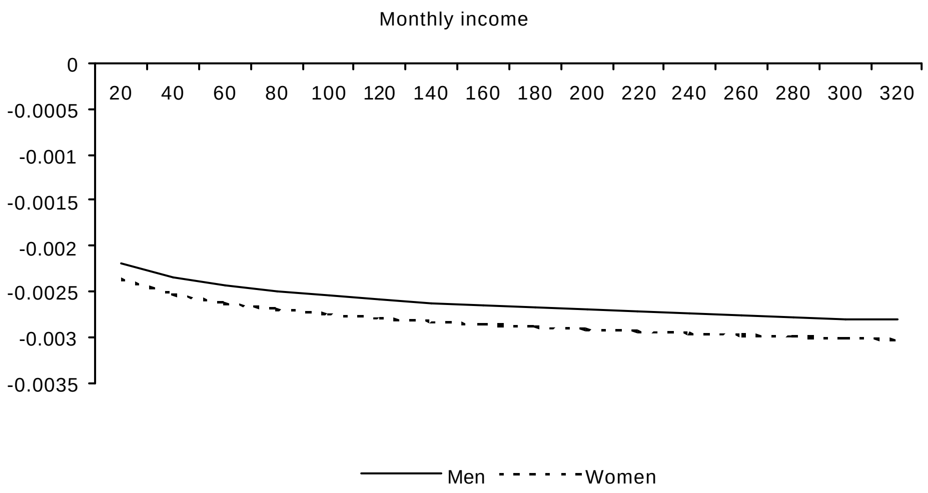


Figure 2: Ratio marginal revenue-marginal cost



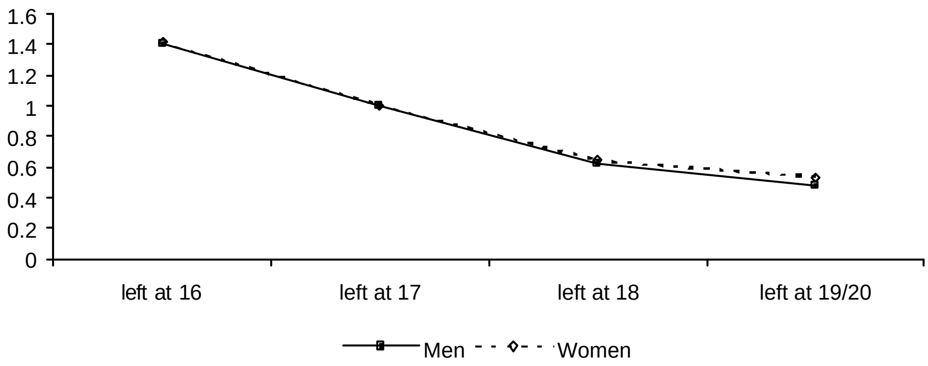
Note: The marginal revenue-marginal cost ratios presented are derived from the estimates based on the equations including ability for females. The exclusion of the ability measures does not change the general trend. The ratios for men follow a similar pattern.

Figure 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.

Figure 4: Corrected marginal revenue-marginal cost ratio



Appendix

Sample **A**'s information (FES), $N_{\mathbf{A}}$ observations, is collected in the matrices $\mathbf{X_A}$, $\mathbf{Z_A}$ and $\mathbf{?}_{\mathbf{A}}$. $\mathbf{Z_A}$ is such that column p, for $p \in \{1,2,\ldots q\}$, contains the values of z^{p-1} for each individual observation. $\mathbf{?}_{\mathbf{A}}$ collects the individual vectors of earnings dummy variables. Sample **B**'s information (BCS), $N_{\mathbf{B}}$ observations, is collected in the matrices $\mathbf{y_B}$, $\mathbf{X_B}$, and $\mathbf{?}_{\mathbf{B}}$. $\mathbf{?}_{\mathbf{B}}$ collects the individual vectors of earnings dummy variables for this sample.

Equation (7) can be estimated by the following misspecified model, using the information available in sample **B**:

$$y^* = x\hat{\tilde{b}} + \Delta\hat{q} + \ln(\hat{e}) \tag{A1}$$

where Δ is a vector of dummies for the earnings variable and $\ln(\tilde{e}) = \ln(e) + f(z) - \Delta\hat{q}$, and $\hat{\tilde{b}}$ and $\hat{q}$ are consistent estimates of the pseudo-true value of $\tilde{b}$ and q, obtained for example from the maximisation of the ordered probit likelihood. Asymptotically these estimate of the pseudo-true value are such that:

$$\mathrm{E}\left[\left(y^* - x\hat{\tilde{b}}_\infty - \Delta\hat{q}_{\infty+1}\right)\left(x \quad \Delta\right)\right] = 0,$$

where the subscript δ refers to asymptotic limit. That is the ordered probit likelihood applied to the misspecified model (because of the imperfect observation of paternal income) imposes orthogonality between the pseudo-errors and the explanatory variables (this is an assumption of the misspecified model). Hence, asymptotically and provided all the relevant quantities below exist, the estimates $\hat{\tilde{b}}_\infty$ and $\hat{q}_\infty$ verify the following relationships:

$$\hat{\tilde{b}}_\infty = \left\{\mathrm{E}\left[x'x\right] - \mathrm{E}\left[x'\Delta\right]\mathrm{E}\left[\Delta'\Delta\right]^{-1}\mathrm{E}\left[\Delta'x\right]\right\}^{-1}\left\{\mathrm{E}\left[x'y^*\right] - \mathrm{E}\left[x'\Delta\right]\mathrm{E}\left[\Delta'\Delta\right]^{-1}\mathrm{E}\left[\Delta'y^*\right]\right\}$$

$$\hat{q}_\infty = \left\{\mathrm{E}\left[\Delta'\Delta\right] - \mathrm{E}\left[\Delta'x\right]\mathrm{E}\left[x'x\right]^{-1}\mathrm{E}\left[x'\Delta\right]\right\}^{-1}\left\{\mathrm{E}\left[\Delta'y^*\right] - \mathrm{E}\left[\Delta'x\right]\mathrm{E}\left[x'x\right]^{-1}\mathrm{E}\left[x'y^*\right]\right\}$$

From these expressions, some tedious calculus gives us the relationships between the pseudo-true values and the true parameters of the correctly specified model as follows:

$$\hat{\tilde{b}}_\infty = b + \left\{ \mathrm{E}\left[x'x\right] - \mathrm{E}\left[x'\Delta\right]\mathrm{E}\left[\Delta'\Delta\right]^{-1}\mathrm{E}\left[\Delta'x\right] \right\}^{-1} \left\{ \mathrm{E}\left[x'\mathbf{z}\right] - \mathrm{E}\left[x'\Delta\right]\mathrm{E}\left[\Delta'\Delta\right]^{-1}\mathrm{E}\left[\Delta'\mathbf{z}\right] \right\} a,$$

$$\hat{q}_\infty = \left\{ \mathrm{E}\left[\Delta'\Delta\right] - \mathrm{E}\left[\Delta'x\right]\mathrm{E}\left[x'x\right]^{-1}\mathrm{E}\left[x'\Delta\right] \right\}^{-1} \left\{ \mathrm{E}\left[\Delta'\mathbf{z}\right] - \mathrm{E}\left[\Delta'x\right]\mathrm{E}\left[x'x\right]^{-1}\mathrm{E}\left[x'\mathbf{z}\right] \right\} a,$$

where we have eliminated the terms with zero expectation. The previous expressions can be rearranged as follows:

$$a = \left\{ \mathrm{E}\left[\Delta'\mathbf{z}\right] - \mathrm{E}\left[\Delta'x\right]\mathrm{E}\left[x'x\right]^{-1}\mathrm{E}\left[x'\mathbf{z}\right] \right\}^{-1} \left\{ \mathrm{E}\left[\Delta'\Delta\right] - \mathrm{E}\left[\Delta'x\right]\mathrm{E}\left[x'x\right]^{-1}\mathrm{E}\left[x'\Delta\right] \right\} \hat{q}_\infty,$$

$$\begin{aligned} b = {} & \hat{\tilde{b}}_\infty - \\ & \left\{ \mathrm{E}\left[x'x\right] - \mathrm{E}\left[x'\Delta\right]\mathrm{E}\left[\Delta'\Delta\right]^{-1}\mathrm{E}\left[\Delta'x\right] \right\}^{-1} \left\{ \mathrm{E}\left[x'\mathbf{z}\right] - \mathrm{E}\left[x'\Delta\right]\mathrm{E}\left[\Delta'\Delta\right]^{-1}\mathrm{E}\left[\Delta'\mathbf{z}\right] \right\} \\ & \left\{ \mathrm{E}\left[\Delta'\mathbf{z}\right] - \mathrm{E}\left[\Delta'x\right]\mathrm{E}\left[x'x\right]^{-1}\mathrm{E}\left[x'\mathbf{z}\right] \right\}^{-1} \left\{ \mathrm{E}\left[\Delta'\Delta\right] - \mathrm{E}\left[\Delta'x\right]\mathrm{E}\left[x'x\right]^{-1}\mathrm{E}\left[x'\Delta\right] \right\} \hat{q}_\infty, \end{aligned}$$

These expressions suggest some feasible asymptotic bias corrections using appropriate empirical moments from the two samples. We have [A1]:

$$\begin{aligned} \hat{a} &= \left(\frac{\Delta'_A \mathbf{M}_{\mathbf{x}_A} \mathbf{Z}_A}{N_A} \right)^{-1} \frac{\Delta'_B \mathbf{M}_{\mathbf{x}_B} \Delta_B}{N_B} \hat{q}, \\ \hat{b} &= \tilde{b} - \left(\frac{\Delta'_B \mathbf{M}_{\Delta_B} \mathbf{Z}_B}{N_B} \right)^{-1} \left(\frac{\Delta'_A \mathbf{M}_{\Delta_A} \mathbf{Z}_A}{N_A} \right) \left(\frac{\Delta'_A \mathbf{M}_{\mathbf{x}_A} \mathbf{Z}_A}{N_A} \right)^{-1} \left(\frac{\Delta'_B \mathbf{M}_{\mathbf{x}_B} \Delta_B}{N_B} \right) \hat{q}, \end{aligned} \qquad \text{(A2)}$$

[A1] Given the relationship between the asymptotic pseudo-true value, the true value and the error term it is straightforward to obtain an expression for the asymptotic variance-covariance of $\left(\hat{b}, \hat{a}\right)$. The feasible estimator depends clearly on the estimated variance-covariance of $\left(\hat{\tilde{b}}, \hat{q}\right)$.

where $\mathbf{M}_{\mathbf{X}_{\mathbf{B}}} = \mathbf{I}_{N_{\mathbf{B}}} - \mathbf{X}_{\mathbf{B}} \left(\mathbf{X}'_{\mathbf{B}} \mathbf{X}_{\mathbf{B}} \right)^{-1} \mathbf{X}'_{\mathbf{B}}$ and similarly for $\mathbf{M}_{\mathbf{X}_{A}}$, and

$\mathbf{M}_{\mathbf{?}_{\mathbf{B}}} = \mathbf{I}_{N_{\mathbf{B}}} - \mathbf{?}_{\mathbf{B}} \left(\mathbf{?}'_{\mathbf{B}} \mathbf{?}_{\mathbf{B}} \right)^{-1} \mathbf{?}'_{\mathbf{B}}$, and similarly for $\mathbf{M}_{\mathbf{?}_{A}}$.

References

Acemoglu, D. and Pischke, S. (2000), 'Changes in the Wage Structure, Family Income and Children's Education', Centre for Economic Performance, mimeo.

Becker, G. and Tomes, N. (1986), 'Human Capital and the Rise and Fall of Families', Journal of Labor Economics, 4, pp.S1-39.

Blundell, R., Dearden, L., Goodman, A. and Reed, H. (2000), 'The Returns to Higher Education in Britain: Evidence from a British Cohort', Economic Journal, Vol. 110, pp.F82-F99.

Cameron, S. and Heckman, J. (1998), 'Life Cycle Schooling and Dynamic Selection Bias: Models and Evidence for Five Cohorts of American Males', Journal of Political Economy, No 106, pp.262-332.

Cameron, S. and Heckman, J. (1999), 'The Dynamics of Educational Attainment for Blacks, Hispanics and Whites', Working Paper, No. W7249, NBER.

Cameron, S. and Taber, C. (2000), 'Borrowing Constraints and the Returns to Schooling', Working Paper No. W7761, NBER.

Card, D. (1999), 'The Causal Effect of Education on Earnings', in O. Ashenfelter and D. Card, (eds.), *Handbook of Labor Economics*, Elsevier: Amsterdam.

Dearden, L. (1998), 'Ability, Families, Education and Earnings in Britain', Working Paper No. W98/14, Institute for Fiscal Studies.

Department for Education and Employment (2001), 'Education Maintenance Allowance: the first year', Research Brief, 257

Duflo, E. (1999), 'Child health and household resources in South Africa: Evidence from old age pension program', MIT, mimeo

Dynarski, S. (1999), 'Does Aid Matter? Measuring the Effect of Student Aid on College Attendance and Completion', Working Paper No. W7422, NBER.

Dynarski, S. (2000), 'Hope for Whom? Financial Aid for the Middle Class and Its Impact on College Attendance', Working Paper No. W7756, NBER.

Feinstein, L. (2000), 'The Relative Economic Importance of Academic, Psychological and Behavioural Attributes Developed in Childhood', Discussion Paper No. 443, Centre for Economic Performance, London School of Economics.

Feinstein, L. and Symons, J. (1999), 'Attainment in Secondary School', Oxford Economic Papers, 51, pp.300-321.

Gemmel, N. (1996), 'Evaluating the Impacts of Human Capital Stocks and Accumulation on Economic Growth: Some New Evidence', Oxford Bulletin of Economics and Statistics, 58, pp.9-28.

Golderberg, A. and Manski, C. (1995), 'Review Article: "The Bell Curve" by Herrnstein and Murray', Journal of Economic Literature, 33, pp.762-776.

Gregg, P., Harkness, S. and Machin, S. (1999), 'Child Development and Family Income', Joseph Rowntree Foundation.

Gregg, P. and Machin, S. (1999), 'Childhood Disadvantage and Success or Failure in the Labour Market', in D. Blanchflower and R. Freeman (eds.), *Youth Employment and Joblessness in Advanced Countries*, NBER: Cambridge, MA.

Hanushek, E. (1992), 'The Trade-Off Between Child Quality and Quantity', Journal of Political Economy, 100, pp.84-117.

Harmon, C. and Walker, I. (2000), 'Child Outcomes and Child Poverty: Provisional Results on Early School Leaving', University of Warwick, mimeo.

Haveman, R. and Wolfe, B. (1995), 'The Determinants of Children's Attainments: A Review of Methods and Findings', Journal of Economic Literature, Vol. 33, pp.1829-1878.

Heckman, J. (1995), 'Cracked Bell', Reason Magazine, March.

Heckman, J. (1999), 'Policies to Foster Human Capital', Working Paper No. W7288, NBER.

Herrnstein, R. and Murray, C. (1994), *The Bell Curve*, Free Press: NY.

Leslie, D. and Drinkwater, S. (1999), 'Staying on in Full-Time Education: Reasons for Higher Participation Rates Among Ethnic Minority Males and Females', Economica, 67, pp.63-77.

Mayer, S. (1997), '*What Money Can't Buy*', Harvard University Press: Cambridge.

Mickewright, J. (1986), 'A Note on Household Income Data in $NCDS_3$', City University, NCDS user support group, Working Paper No 18.

Rice, P. (1987), 'The Demand for Post-Compulsory Education', Economica, Vol. 54, pp.465-474.

Shea, J. (2000), 'Does Parent's Money Matter?', Journal of Public Economics, Vol. 77, pp.155-184.

Wolfe, B., Haveman, R., Ginther, D. and Bum An, C. (1996), 'The "Window Problem" in Studies of Children's Attainments: A Methodological Exploration', Journal of the American Statistical Association, 91, pp.970-982.