Participation in higher education in Britain: The effect of ability and parental income*

Tania Oliveira†
University of Leicester

Luisa Zanchi‡
University of Leeds

January 2004

Abstract

We test the existence of an income bias in higher education decisions: children from better-off households are over-represented among university students. A theoretical model shows that, even with perfect access to capital markets, the market equilibrium exhibits an income bias. We estimate a binary choice model of higher education participation as a function of a set of individual, family and secondary school characteristics, ability and parental income. Both ability and parental income have a positive effect on the participation decision. Lines of equal probability of participation in the income-ability space are, as the theory predicts, downward sloping.

Keywords: Education, Ability, Income
JEL: I21, C25

* We thank Gianni De Fraja, Karen Mumford, Indrajit Ray and Joanna Swaffield for useful comments. NCDS data have been provided by the UK Data Archive; the authors alone are responsible for the analysis and interpretation in this paper.
† I am grateful for the financial support provided by FCT – Fundação para a Ciência e Tecnologia, Ministério da Ciência e do Ensino Superior, Portugal.
‡ Correspondence: Economics Division, Leeds University Business School, University of Leeds, Leeds LS2 9JT, UK. Email: lz@lubs.leeds.ac.uk

1 Introduction

To what extent is the personal choice of investing in higher education explained by factors such as familiar background instead of ability differences among individuals? This question has troubled policy makers over the years, as it raises the problem of equality of treatment and opportunity[1]. This paper intends to test the presence of an income bias in the individual education decision: children from better-off households are over-represented among university students. We follow De Fraja (2002b) developing a theoretical model that shows how, even if households have access to perfect capital markets, the free market equilibrium exhibits an income bias. The bias occurs due to the shape of the individual utility curves that reflects the risk aversion of the agents. To test this result we use a methodology similar to the one in Epple *et al.* (2000). We estimate a model of binary choice, using the participation in higher education as the dependent variable. Considering the variables ability and parental income, it is possible to draw a set of lines of equal probability of participating in higher education. The theory predicts downward sloping lines reflecting the absence of equal opportunities in the access to higher education.

The paper will review the relevant literature in section 2. The theoretical model is described in section 3, while the empirical methodology is presented in section 4. In section 5 we describe the dataset used and some characteristics of the variables chosen. The empirical evidence is discussed in section 6, with the different effects grouped in three subsections: income and ability, background variables, and school type and peer effects. Finally, section 7 concludes and raises some general points.

[1] In this context, equality of opportunity is interpreted in the sense that university attendance is independent of parental income. For a discussion of this definition see De Fraja (2002b).

2 Review of the literature

Having a theoretical background that explains participation decisions in higher education is a recent feature. The traditional approach, the Rosen schooling model, is a partial equilibrium analyses that focuses on the cost associated with foregone income during education years. Epple and Romano (1998) and De Fraja (2002a) illustrate the recent interest in explaining the choice of education in a general equilibrium framework. Both works consider that students differ in their ability and their household income, and try to find the equilibrium properties of the higher education market.

Most work on schooling and higher education has been made in the context of the returns to education using earnings equations[2]. The attempt to solve self-selection bias resulted in the estimation of participation equations as suggested firstly in Willis and Rosen (1979). They develop an instrumental variables (IV) procedure that solves the endogeneity of the higher education participation, recognising that earnings might play a part in the schooling decision. A structural model of demand for college education is constructed based on the theory of comparative advantage, the existence of self-selection and unobserved components Using probit analysis they study the college participation decision in the NBER-Thorndike dataset, including background variables such as father's education and occupation, religion, mother's participation in the labour market, and existence of siblings. Four ability measures were used, two related with IQ type of ability, and two others more associated with manual skills. The estimation of structural probit models, that show how the anticipated gains in earnings affect the decision to attend college, is seen as another original contribution to empirical analysis. These estimates support the economic hypothesis that expected gains in earnings influence the decision to attend college, and the selection patterns

[2] Card (1999) provides an extensive survey of the literature that uses instrumental variables or related techniques to address the problems caused by omitted ability measures.

support the comparative advantage hypothesis. Harmon and Walker (1995) used IV methods to correct biases in OLS estimates of the returns to schooling. They use the exogenous changes in the educational distribution caused by the raising of the minimum school-leaving age in the UK as instruments for schooling. A selectivity model is also estimated using age, region of origin, the minimum school-leaving age, and a set of year dummies as explanatory variables in the ordered-probit analysis of the years of schooling. This analysis shows the presence of a large and negative bias in the OLS estimate of the returns to schooling, but does not address the question of participation in any level of schooling. More recently Dearden (1999) uses proxy or matching methods to solve possible ability and composition biases and IV methods to deal with measurement errors in schooling variables, finding that conventional OLS estimates of the returns to education which control for age, gender, ethnicity, and region of residence and ignore endogeneity of education are a fair approximation to the causal effect of education on earnings.

The endogeneity of schooling can also arise due to the correlation between its quantity and its quality. Harmon and Walker (2000), Dearden *et al.* (2002), Dustmann *et al.* (2003), and Strayer (2002) consider how school quality affects wages. The results confirm other evidence that returns to education quantity are larger than returns to quality. Dearden *et al.* (2002) find that the pupil-teacher ratio has no impact on educational qualifications, once controlling for ability, family background, local authority and neighbourhood characteristics. Their results show that the strongest impact of educational inputs on attainment occurs via the type of school attended. Dustmann *et al.* (2003) use micro data for England and Wales (the National Child Development Study (NCDS)) to examine the effects of class size on the decision to stay on the decision to stay on in full time education at 16 and wages later in life. They find that class size has a significant effect on the decision to stay on. Wage equations show that the effect of staying on is

significantly positive. Combining these two findings, the authors conclude that class size significantly affects future wages. The use of reduced form models leads to less accurate and insignificant estimates. They conclude that caution should be used in accepting outcomes of non-experimental studies that state no effects of class size on wages. Strayer (2002) also focus on the structural effects of high school quality on earnings. The results suggest that school quality influences earnings by affecting the university choice behaviour, while the direct effect of school quality on earnings is less evident. The author uses a unique dataset that combines the National Longitudinal Survey of Youth (NLSY) with the Integrated Postsecondary Data System (IPEDS), that provides information on individuals' background, the secondary schools and universities they attend, and post-school wages and jobs. The proportion of teachers with graduate degrees in the school and a smaller class size have a significant positive effect on the probability of choosing to attend university. The type of university selected, in turn, affects positively post-school earnings.

For the UK, in the empirical microeconomic literature, the debate about the determinants of participation in full-time education has focused on the decision of staying on after compulsory education.(e.g. Leslie and Drinkwater, 1999; Rice, 1987; Micklewright, 1989; Armstrong, 1999; and Johnson, 2002). Rice (1987) develops a model of post-compulsory schooling. A probit model is estimated using data from the 1976 Family Expenditure Survey. The author includes current employment status, occupational class and industrial sector of the head of the household as background variables. The measures of household income are its normal gross weekly income as a proxy for permanent income, and accounts for its composition making the distinction between income derived from market sources and income received from state benefits and pensions. A measure of the relative unemployment rate in the region as well as a set of regional dummies are included to test the hypothesis that a higher rate of

unemployment decreases the opportunity cost of further education. The results reveal significant differences between sexes. For males, household income is not a significant factor in the decision to continue in post-compulsory education, while for females it is. The socio-economic background of the household has an influence on the probability of undertaking post-compulsory education. The regional unemployment rate shows the expected effect, although regional differences remain significant. Micklewright (1989) looks at the probability of leaving school at the first legal opportunity using the National Child Development Study (NCDS) dataset. Three groups of variables are used. The first refers to the child's family background, including parental presence in the household, father's and mother's social class when employed, the number of younger and older siblings, household income, a dummy variable that serves as a proxy for low-income households, and regional dummies for London and the Southeast and Wales. The second relates to the characteristics of the school attended during the final year of compulsory education, including dummies for the type of school (comprehensive, grammar, and independent or direct grant). The third group reflects academic ability at age 16, using the sum of scores in maths and English comprehension tests. Separate estimates for boys and girls show that family background variables have the expected signs, and remain significant when the author controls for ability and school type, providing grounds to conclude that "there does not exist equality of educational outcomes for children with equal ability following compulsory schooling". Leslie and Drinkwater (1999) study the ethnic minorities' greater tendency to stay in full-time education beyond compulsory age. The method used is to fit a joint leaving and employment equation for males and females aged 18-24 using data from the British Labour Force Survey and the Sample of Anonymized Records from the 1991 Census. The leaving equation includes predicted earnings, a set of socio-economic background variables, a set of age dummies, and a set of dummies that reflects ethnicity. The bivariate

probit results confirm the hypothesis that individuals with a higher probability of unemployment will have higher probability of staying on in full-time education, after controlling for other factors. Attempts to assess how school characteristics influence the decision to stay on in education and labour market success can also be found in the literature. Armstrong (1999) uses microeconomic survey data (Status 0 Survey) to investigate the impact of school performance on the educational and labour market choices in Northern Ireland. Multinomial logit models are estimated at age 16 for four main destinations of young people: further education, employment, vocational training, and unemployment. The models include measures of school performance (overall examination performance and attendance rates), gender, local unemployment rates, indicators of religion, number of older and younger siblings, parental employment status. The results show that young people from schools which perform relatively well are more likely to remain in full-time education, ceteris paribus. Higher participation rates are also found amongst women and Catholics. Johnson (2002) presents a structural model of demand for educational investment. Ability controlled earnings profiles are estimated by matching individuals from the General Household Survey to individuals in similar occupations from the NCDS. The results show that expected earnings profiles vary according to observed ability and educational choice and expected lifetime earnings have a significant impact on educational choice.

Besides this type of microeconomic approach, some work as also been done using time-series evidence to identify the factors influencing the propensity of 16-year-olds to stay in full-time education. Pissarides (1981) analyses the period between 1955 and 1978 and finds that changes in growth of participation rates were a result of movements in the relative earnings of qualified workers and differences in household income. Whitfield and Wilson (1991) find that the main factors are the rate of return to education, changing social class structure, unemployment rates, and training measures

as the Youth Training Scheme. More recently, McVicar and Rice (2001) analyse the behaviour of the rate of participation in further education since the mid-50s using full information methods. The results show that the rapid increase of the last decade was driven by improvements in GCSE attainment together with the expansion of the higher education sector. The authors also conclude that fluctuations in labour demand play a significant role, and that the rise in youth unemployment of the early 1990s contributed to the growth in participation at this time. The results suggest that for females the improvement in attainment levels is a response to the greater returns in the market to workers with further educational qualifications.

The study of how household income affects a child's educational attainment has produced conflicting results, the effectiveness of policies that aim to reduce inequalities remaining unclear. Chevalier and Lanot (2002) identify a direct effect of parental income related to financial constraints preventing further investment in their children's education, and/or an indirect one reflecting the fact that socio-economic conditions under which a child was brought up may influence schooling achievements. They propose a methodology that separates those effects using an ordered probit for estimating the leaving education decision for five age groups, from 16 to 20 years of age, relying on the NCDS and the British Cohort Study (BCS). They use a set of background characteristics, a variable that tries to capture neighbourhood effects, father's earnings as a proxy for family income, and results of reading and mathematics tests at age 7, in the NCDS, and at age 10, in the BCS. The results show that family income appears to have a significant effect on the schooling decision. However as this income effect might be correlated with other family characteristics that determine the educational choice, the authors model the effect of an educational maintenance allowance on the choice. The estimates suggest that the schooling achievement is dominated by the effect of parental education and family structure. Shea (2000) finds that changes in parents' income due to

luck have a negligible effect on children's ages, earnings, years of schooling, and total family income. Only for children in families whose father has less than 12 years of schooling does the parents' income have a beneficial impact. The author uses a sample from the Panel Study of Income Dynamics (PSID) to perform two-stage least squares regressions of children's income on demographic characteristics, fathers' observable skills, and measures of parents' income, using father's union membership, industry and job loss variables as instruments for parents' income. Cameron and Taber (2000) test the presence of borrowing constraints using four different methods: instrumental variables; regression models allowing for interaction between observables and presence of a college in the county of residence; structural estimation allowing borrowing rates to depend on observables; and structural estimation allowing for unobservable heterogeneity in borrowing rates. They find no evidence of borrowing constraints using any of the methods, although that does not mean that credit market constraints would not exist in the presence of the programs currently available.

3 The Theoretical Model

We use the general equilibrium theoretical model presented by De Fraja (2002b). In the economy there exists an infinite number of households, each constituted by a mother, representing the older generation, and a daughter, the younger generation counterpart. Households differ in the mother's after tax income and daughter's innate ability to earn income in the future.

The mother's income is denoted by Y. The daughter's ability is denoted by θ. Income and ability are distributed in $\left[\underline{Y}, \overline{Y}\right] \times \left[\underline{\theta}, \overline{\theta}\right]$. The joint distribution is denoted by $H(Y, \theta)$, with joint density $h(Y, \theta)$.

Within the household, the mother makes the decisions maximising a utility function given by the sum of her own utility $(u_m(c_m))$ and the

daughter's utility of consumption $(u(c-\psi(\eta)))$, where $\psi(\eta)$ is a function measuring the cost of effort in the labour market (η). The mother has to choose the amount she allocates to private consumption, and the amount she invests in her daughter's education through a monetary transfer (t). To simplify the model, education only takes two values: either going to university or not. Besides the mother's transfer, the daughter receives also labour income. The individual income is a random variable whose realisation depends on education, innate ability, and labour market effort. If the daughter goes to university, she either receives labour market income y_H with probability $P(\theta,\eta)$, or $y_L < y_H$ with probability $1-P(\theta,\eta)$. This probability is strictly increasing in both arguments, $P_{\eta\eta} \leq 0$, and it is assumed that $P_{\theta\eta} \geq 0$. If she doesn't go to university, she exerts effort η_N and receives y_N.

The hypothesis of perfect capital markets implies that the mother can borrow against the daughter's future income at the market interest rate. Although very unrealistic, this hypothesis is introduced to show that even in that extreme case there is scope for government intervention. In the literature it is possible to find some empirical support for this hypothesis in Shea (2000) and Cameron and Taber (2000)[3]. Perfectly competitive insurance markets are also assumed, implying that the mother can transfer to her daughter, according to the realisation of the labour market income, t_H if income is y_H and t_L if income is y_L. Perfect competition among insurers ensures that the cost of acquiring insurance (t) equals the expected payout $(P(\theta,\eta)t_H + (1-P(\theta,\eta))t_L)$. Participation in higher education has a direct cost $k > 0$, the tuition fee, taken as given by the households and determined

[3] For the UK, it can be argued that the grant system in existence at the time of reference of our data and the absence of tuition fees in public universities, until the recent introdution, could justify such an assumption.

endogenously at the market equilibrium.

The decision to invest in higher education depends on the comparison between expected utilities under the two choices. Effort is not contractible and is decided by the daughter after the insurance contract is known. If she does attend university, she exerts effort $\hat{\eta}$, a function of her ability that maximises her expected utility; otherwise she exerts effort η_N. If the mother decides to invest in higher education for her daughter, the household utility is:

$$\begin{aligned} U(\theta, Y, k) = \max_{t,\, t_H,\, t_L} \{ & u_m(Y - t - k) + P(\theta, \hat{\eta}) u(y_H + t_H - \psi(\hat{\eta})) + \\ & + [1 - P(\theta, \hat{\eta})] u(y_L + t_L - \psi(\hat{\eta}))\} \\ \text{s.t.: } & t = P(\theta, \hat{\eta}) t_H + [1 - P(\theta, \hat{\eta})] t_L \end{aligned}$$

The utility achieved if the daughter does not go to university is:

$$U_N(Y) = \max_{t_N} \{ u_m(Y - t_N) + u(y_N + t_N - \psi(\eta_N))\}$$

The mother decides to invest in higher education for her daughter as long as $U(\theta, Y, k) > U_N(Y)$. It is also assumed that, whatever the parental income, the ablest children $(\theta = \overline{\theta})$ go to university, and the least able $(\theta = \underline{\theta})$ do not:

Assumption *For every* $Y \in [\underline{Y}, \overline{Y}]$, $U(\underline{\theta}, Y, k) < U_N(Y)$ *and* $U(\overline{\theta}, Y, k) \geq U_N(Y)$.

The main result of the model predicts the household decision regarding university participation. With effort non-contractible, under strict concavity of the mother's utility function $(u''_m(c_m) < 0)$, and the daughter's decreasing

absolute risk aversion over consumption $d\left(\frac{-u''(c)}{u'(c)}\right)/dc < 0$:

Proposition *There exists a strictly decreasing function defined in* $\left[\underline{Y}, \overline{Y}\right]$, *denoted by* $\theta^m(Y)$, *and defined the* market indifference curve, *such that a household with income* Y *sends the daughter to university if and only if she has ability at least* $\theta^m(Y)$.

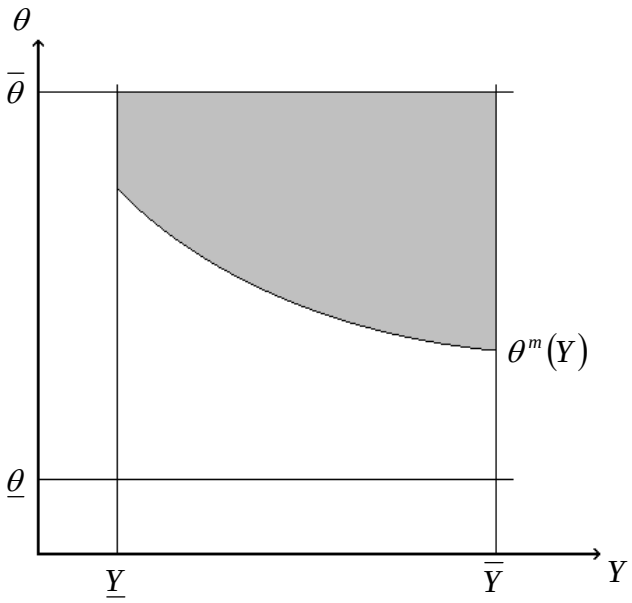

Figure 1: The Market Indifference Curve

As shown in Figure 1, the market indifference curve $\theta^m(Y)$, in the income-ability space, is decreasing. This implies that individuals whose parents have higher income are more likely to attend university. The free market bias occurs as a result of effort being not contractible and of individuals showing decreasing absolute risk aversion, which translates into considering a specific shape for the utility curves. Investment in education represents a sacrifice today against uncertain future returns. The presence of

moral hazard prevents this risk to be fully insured. The uninsured risk has a cost in terms of utility that is lower for higher income households. These households are therefore more willing to take the risk and invest in education. The main conclusion is then that the free market does not provide equality of opportunity even when borrowing and full insurance against labour market risk is assumed. The testable prediction of the model is that the market indifference curve is downward sloping.

4 The Empirical Model

To test the prediction of the theoretical model we begin by estimating a model of binary choice, using participation in higher education as the dependent variable. From these estimates, it is then possible to derive a set of lines of equal probability of participating in higher education in the parental income-ability space. The theoretical model predicts that such lines should be negatively sloped, reflecting the absence of equality of opportunity in accessing higher education.

The probability of participating in higher education is modelled as a function of a set of individual, family, schooling, and regional characteristics, the individual's ability, and parental income. Ideally, the required measure of an individual's ability would be an indicator of his or her "innate" ability. However, our dataset (like others) provides ability measures observed at some stage of the educational development of individuals – in our case, at the age of 7 and 11. Such measures of ability, therefore, may be themselves affected by socio-economic and schooling characteristics also affecting the probability of participating in higher education directly, in particular by parental income. The full empirical model can be specified as follows:

$$A_i = \boldsymbol{\beta}_1' \mathbf{X_i} + \beta_2 Y_i + u_i \quad (1)$$

$$HE_i = \boldsymbol{\delta}_1' \mathbf{X_i} + \delta_2 Y_i + \delta_3 A_i + v_i \quad (2)$$

where A_i is a measure of the observed individual *i*'s ability, HE_i is a dummy variable which takes the value 1 if individual *i* participated in higher education and 0 otherwise, $\mathbf{X_i}$ is a set of individual and family socio-economic factors and school characteristics, and Y_i is parental income.

We assume that the disturbances u_i and v_i are uncorrelated, i.e. that the system is a fully recursive model. This entails $\text{cov}(A_i, v_i) = \text{cov}(\boldsymbol{\beta}_1' \mathbf{X_i} + \beta_2 Y_i + u_i, v_i) = 0$. Therefore in equation (2) the full set of right-hand variables is uncorrelated with the disturbance v_i and the parameters in equation (2) may be consistently estimated. Also, since we use a set of different ability measures for various dimensions of ability observed at different ages, the actual esimated equation is:

$$HE_i = \boldsymbol{\delta}_1' \mathbf{X_i} + \delta_2 Y_i + \boldsymbol{\delta}_3 \mathbf{A_i} + v_i \quad (2')$$

Given the binary nature of the dependent variable HE_i, we employ a probit estimator. We use both a comprehensive sample of males and females together, including a gender dummy among the explanatory variables, and separate samples for males and females, to explore the possible differences in the behaviour of the two genders. To find if the participation decision is significantly different across income ranges, we divide the sample in two classes using £60[4] weekly net income as a cut off point. Again, for the lower income group, separate equations are estimated for males and females.

[4] The proportion of households with weekly net income over £60 in 1974 in the sample is 21%.

5 The Data

We use the data from the National Child Development Study (NCDS), a panel survey of the cohort of individuals born between the 3rd and the 9th of March in 1958 in Britain, that have been followed from birth until the age of 42. The NCDS is a longitudinal study that started with a survey carried out in the week of birth and continued with six follow-up surveys, when the subjects were 7, 11, 16, 23, 33, and 42. The data set has its origins in the Perinatal Mortality Survey (PMS), designed to examine the social and obstetric factors associated with stillbirth and early infancy death. The scope of the survey was broadened in subsequent waves to include information on the physical, educational and social development of the subjects. It includes very detailed information on family background, ability test scores, educational qualifications, characteristics of the schools attended, and parental income. As most longitudinal studies, the survey suffers from incomplete information and attrition[5].

The dependent variable is binary, assuming the value 1 if the individual participated in higher education straight after secondary education and 0 otherwise, although we allow for a short break and consider the individuals that were still engaged in university degrees by age 23. This variable was constructed based on the information about the highest qualification obtained until that age using waves 4, 5, and 6[6]. We refer only to participation in higher education and not to higher education qualifications, because for some cases we cannot infer if the individuals actually finished the course they were attending. As determinants of the participation decision we consider a set of individual and family background socio-economic characteristics, the individual's ability, secondary school characteristics,

[5] For a list of the drawbacks of the data, see Micklewright (1989). Connelly *et al.* (1992) provide the most extensive study of attrition in this panel.

[6] When conflicting information arose we took the information conveyed by the wave closest to the moment of acquisition of the education in question.

including a measure of peer group effects, and parental income at age 16.

Among the family background variables we include the parents' education, the number of siblings, the household's type of accommodation, the mother's participation in the labour market, and the type of place of residence, either urban or rural, at age 16. The parents' education is summarised by four dummy variables indicating if the father and the mother have attended secondary or higher education. We take has reference the case where both parents only have compulsory education. The variables were constructed based on the age the parents left school. The father's social class is included as a dummy variable that assumes the value of one if the father's occupation is non-manual. Regarding the type of accommodation, we consider a binary variable that assumes unity value if the household accommodation is owner occupied. A binary variable for Scotland is kept in the model for the reason illustrated in section 6.2. Preliminary estimates using a full set of region of residence dummies resulted in insignificant marginal effects for all the other regions.

We use a dummy variable for the sex, that assumes the value 1 if the individual is female. To clarify potential differences in the behaviour of the two genders we also estimate separate models for men and women.

The NCDS contains data on arithmetic and reading ability tests performed by the children at the aged of 7. At the age of 11, the dataset provides information on three types of test: mathematics, reading and general ability. The results for the test scores were all re-scaled to vary between 0 and 1, so that the estimates can be directly comparable.

The three school types considered are grammar, secondary modern and 'public' (private) schools. The reference individual is assumed to have attended a comprehensive school. We also control for the fact that the school attended is single sex only. To measure possible peer-group effects, we include a set of dummy variables indicating the percentage of children in the school whose fathers have non-manual occupations.

The parental income is measured at age 16. That was the only information available regarding the household's economic situation. The information is either in weekly or in monthly terms, but the relevant questions grouped the several sources of income in intervals that correspond to the same income per unit of time. We summed across mid-points of each range creating a single 'quasi-continuous' variable for the household's income. There are some doubts about the accuracy of these values, but comparison with the Family Expenditure Survey (FES) data for the same year (Micklewright, 1986) provide ground to consider it valid as a representation of the income distribution. The household income has a mean of £44.766 per week, and a standard deviation of 20.579, ranging from £2 to £167 per week. The household income not earned by the father figure (percentage of non-father income) is also considered to try and assess the influence of other sources of income. An initial attempt using only the father's income showed that this had no influence on the participation decision, indicating that the existence of other sources of income may be relevant to the choice of participating in higher education.

The final sample has 7729 observations, 3835 of which are women. We have excluded individuals for whom the information on education, parental income and ability at age 7 and 11 is missing. In the final sample 9.7% of individuals participated in higher education before reaching the age of 23[7]. Table 1 presents summary statistics for the variables used in specification presented.

[7] The proportion of participation is higher (15.9%) if we consider university participation before reaching the age of 42. The difference between the two percentages reflects a feature specific to this cohort, in which a rather high proportion of individuals engaged in some form of higher education at later stages in life.

6 Empirical Evidence

The results for the whole sample are presented in Table 2. Dividing the sample in two income classes resulted in the outcomes presented in Table 3.

6.1 Income and ability effects

The results for a probit model of the decision of participating in higher education for the comprehensive, female and male samples are presented in Table 2. Both the ability variables and total household income have a positive effect on the participation decision. This means that there is a trade-off between the two variables in their influence on the dependent variable. Analysing the ability variables, we can conclude that the age 11 maths and reading test scores are significant; also the age 7 arithmetic results have a significant positive effect, although this evidence is not so strong. It is possible to say that an increase of 10% in the score of the maths test increases the probability of participating in higher education by 1.043%, while the reading test has a slightly stronger effect of 1.165%. Regarding parental income, the effect of an increase of £10 in weekly household income is an increase of about 0.2% in the probability of attending university. This evidence is not as statistically strong as the one found for the two ability variables. Nevertheless, since we include some variables that can be used as proxies for parental income in the set of explanatory variables – such as the father's social class and house ownership, used by some authors as indicators of the household economic conditions instead of income – we find that it is relevant that such an income effect is found at all. Considering each of the significant ability variables in turn, we can draw sets of isoprobability lines in the income-ability space (Figures 2 and 3). As predicted by the theoretical model presented in Section 3, the graphs depict negatively sloped lines.

Women have a lower probability of participating in higher education than men with the same characteristics, although such evidence is not significant. This made us consider a split in the sample as differences between the two genders may not be well captured by the gender dummy. We find significant differences in the behaviour of individuals of each sex. For males, the total household income variable is not significant and ability has a stronger positive marginal effect on participation, of about 1.239% and 1.174% respectively for increases of ten percentage points in the maths and reading tests completed at age 11. For females, the same effects are only 0.752% and 1.045% and total household income increases the probability of participation by about 0.2%. This result can be attributed to the higher risk of investment in women's education. The female labour market participation rate in the UK in that period was about 50%. This means that the expected return from investment is lower for women. Even if women do work, the expected wages are lower than for men with the same characteristics, which supports the argument of a higher risk. The percentage of non-father income has significant effects for both sub-samples, but in opposite directions. While for females an increase of ten percentage points in non-father income increases the probability of participation by 0.157%, for males it reduces the probability of participation by 0.278%.

To distinguish these effects across different economic backgrounds, we divide the sample into two sub-groups, using total household income to sort the individuals. The group with weekly total income under £60 is further split by gender. The participation rate of the individuals in the first group (with weekly parental income over £60) is 18.14%, while for the second group (with weekly parental income under £60) it is 6.5% for females and 8.26% for males. The ability variables effects are much stronger for the richer household group. Again, these effects are stronger for males than females. In neither of these groups the income variable is significant. This is probably due to the smaller variability of this variable within each group.

6.2 Background variables

For the comprehensive sample, the father's higher and secondary education increases the probability of participation, as well as the mother's higher education. The strongest absolute effect and evidence is found when the father himself participated in higher education. In the male sample, the positive effect of parents with higher education is stronger, while for the female case these variables are no longer significant. Ownership of the house lived in also seems to increase the probability of participation in higher education. The father's social class has also a significant positive effect, although this effect no longer subsists in the two gender equations. There is some evidence that the effect of the regional dummy for Scotland is also positive, which may reflect historically higher participation rates in this country. However, this finding holds only for the male specification. The marginal effects for all the remaining family background variables are not significantly different from zero, showing no statistical effect on the educational decision.

In the group of families with higher income, the significant background variables – specifically father with non-manual occupation and house ownership – have a greater influence on the participation in higher education than in the comprehensive specification. For the lower income group, the results for females and males confirm the findings of the previous equations.

6.3 School type and peer effects

The attendance of a grammar school increases the probability of higher education participation by about 0.1%, while attending a secondary modern school decreases such probability by about 1.44%. The effect of attending a secondary modern school prevails in the female sample, but the marginal effects for all the other variables are not significantly different from zero in

both gender specifications. The analysis of the peer-group effects shows that no significant differences in participation in higher education can be attributed to attending a school with a certain percentage of fathers with non-manual occupations. For the lower income group, the school type has a significant negative effect in the case of secondary modern schools. For males with total household income over £60 per week in 1974, the attendance of a private secondary school has a positive effect on higher education participation. For the higher income group, all the variables related to school do not have significant effects in the decision under study.

7 Conclusion

The evidence of a trade-off between ability and parental income in their influence on the participation in higher education is a new addition to the understanding of the workings of investment in education. It shows that equality of opportunity is not a characteristic of the higher education market in the UK, even in the absence of borrowing constraints. The comparison of different risk groups and the different results obtained for women and men seems to support the risk aversion explanation for this feature of higher education. Besides this main point, the analysis of the determinants of participation in tertiary education brought some features into light: a relevant role of fathers' education, differences in the effects of secondary school characteristics and the potential different behaviour of women and men.

References

Armstrong, D. M. (1999). 'School performance and staying on: A micro analysis for Northern Ireland', *Manchester School*, vol. 67, pp. 203-230.

Cameron, S. and Taber, C. (2000). 'Borrowing constraints and the returns to schooling', NBER, Working Paper No. 7761.

Card, D. (1999). 'The causal effect of education on earnings', in: Ashenfelter, O. and Card, D. (eds.), *Handbook of Labor Economics*, Amsterdam: North-Holland, vol. 3A, pp. 1801-1863.

Chevalier, A. and Lanot, G. (2002). 'The relative effect of family characteristics and financial situation on educational achievement', *Education Economics*, vol. 10, pp. 165-181.

City University Social Statistics Research Unit (2000). *National Child Development Study composite file including selected perinatal data and sweeps one to five, 1958-1991 (computer file)*, 2nd ed., National Birthday Trust Fund, National Children's Bureau, City University Social Statistics Research Unit (original data producers); Colchester, Essex: UK Data Archive (distributor), SN: 3148.

Connolly, S., Micklewright, J. and Nickell, S. (1992). 'The occupational success of young men who left school at sixteen', *Oxford Economic Papers*, vol. 44, pp. 460-479.

De Fraja, G. (2002a). 'The design of optimal education policies', *Review of Economic Studies*, vol. 69, pp. 437-466.

De Fraja, G. (2002b). 'Equal opportunities in education: Market equilibrium and public policy', mimeo, University of York.

Dearden, L. (1999). 'Qualifications and earnings in Briatin: How reliable are conventional OLS estimates of the returns to education?', Institute for Fiscal Studies, Working Paper No. W99/7.

Dearden, L., Ferri, J. and Meghir, C. (2002). 'The effect of school quality on educational attainment and wages', *Review of Economics and Statistics*, vol. 84, pp. 1-20.

Dustmann, C., Najma, R. and van Soest, A. (2003). 'Class size, education, and wages', *Economic Journal*, vol. 113, pp. F99-F120.

Epple, D., Figlio, D. and Romano, R. (2000). 'Competition between private and public schools: Testing stratification and pricing predictions', NBER, Working Paper No. 7956.

Epple, D. and Romano, R. E. (1998). 'Competition between private and public schools, vouchers, and peer-group effects', *American Economic Review*, vol. 88, pp.33-62.

Harmon, C. and Walker, I. (1995). 'Estimates of the economic return to schooling for the United Kingdom', *American Economic Review*, vol. 85, pp. 1278-1286.

Harmon, C. and Walker, I. (2000). 'The returns to the quantity and quality of education: Evidence for men in England and Wales', *Economica*, vol. 67, pp. 19-35.

Johnson, D. (2002). 'Lifetime earnings, discount rate, ability and the demand for post-compulsory education in men in England and Wales', *Bulletin of Economic Research*, vol. 54, pp. 233-247.

Joint Centre for Longitudinal Research (2001). *National Child Development Study and 1970 British Cohort Study (BCS70) follow-ups, 1999-2000 (computer file)*, Colchester, Essex: UK Data Archive (distributor), SN: 4396.

Leslie, D. and Drinkwater, S. (1999). 'Staying on in full-time education: Reasons for higher participation rates among ethnic minority males and females', *Economica*, vol. 66, pp. 63-77.

McVicar, D. and Rice, P. (2001). 'Participation in further education in England and Wales: An analysis of post-war trends', *Oxford Economic Papers*, vol. 53, pp. 47-66.

Micklewright, J. (1986). 'A note on household income data in NCDS3', NCDS User Support Group, Working Paper No. 18, Social Statistics Research Unit, City University, London.

Micklewright, J. (1989). 'Choice at sixteen', *Economica*, vol. 56, pp. 25-39.

Pissarides, C. A. (1981). 'Staying-on at school in England and Wales', *Economica*, vol. 48, pp. 345-363.

Rice, P. (1987). 'The demand for post-compulsory education in the UK and the effects of educational maintenance allowances', *Economica*, vol. 54, pp. 465-475.

Shea, J. (2000). 'Does parents' money matter?', *Journal of Public Economics*, vol. 77, pp. 155-184.

Strayer, W. (2002). 'The returns to school quality: College choice and earnings', *Journal of Labor Economics*, vol. 20, pp. 475-503.

Whitfield, K. and Wilson, R. A. (1991). 'Staying on in full-time education: The education participation rate of 16-year-olds', *Economica*, vol. 58, pp.391-404.

Willis, R. J. and Rosen, S. (1979). 'Education and self-selection', *Journal of Political Economy*, vol. 87, pp. S7-S36.

Table 1: Descriptive Statistics

	Whole Sample		Females		Males		Income > £60 pw		Income ≤ £60 pw Females		Income ≤ £60 pw Males	
Variables	Mean	St. Dev.	Mean	St. Dev.	Mean	St. Dev.	Mean	St. Dev.	Mean	St. Dev.	Mean	St. Dev.
Higher education participation	0.097	0.296	0.089	0.285	0.105	0.306	0.181	0.385	0.065	0.247	0.083	0.275
Female	0.496	0.500	1.000	0.000	0.000	0.000	0.505	0.500	1.000	0.000	0.000	0.000
Ability test scores												
Arithmetic age 7	0.521	0.245	0.508	0.244	0.534	0.246	0.573	0.240	0.496	0.244	0.518	0.244
Reading age 7	0.792	0.230	0.819	0.217	0.765	0.238	0.843	0.199	0.807	0.224	0.751	0.243
Maths age 11	0.433	0.257	0.425	0.251	0.440	0.263	0.519	0.261	0.402	0.243	0.416	0.258
Reading age 11	0.467	0.177	0.466	0.170	0.467	0.184	0.519	0.179	0.451	0.167	0.454	0.181
General ability age 11	0.551	0.196	0.564	0.193	0.538	0.198	0.611	0.190	0.548	0.192	0.521	0.195
Parental income at 16												
% non-father income	0.405	0.328	0.408	0.331	0.403	0.324	0.343	0.199	0.425	0.357	0.419	0.349
Total household income	44.766	20.579	44.769	20.761	44.763	20.401	73.788	13.391	36.639	14.002	37.085	14.152
Family background												
Father with higher education	0.080	0.271	0.077	0.266	0.082	0.275	0.177	0.381	0.053	0.223	0.054	0.225
Father with secondary education	0.313	0.464	0.324	0.468	0.303	0.459	0.386	0.487	0.303	0.460	0.284	0.451
Mother with higher education	0.054	0.226	0.058	0.234	0.051	0.219	0.132	0.338	0.036	0.187	0.030	0.170
Mother with secondary education	0.453	0.498	0.456	0.498	0.449	0.497	0.521	0.500	0.436	0.496	0.433	0.496
House ownership	0.492	0.500	0.481	0.500	0.502	0.500	0.698	0.459	0.424	0.494	0.447	0.497
Father non-manual	0.307	0.461	0.302	0.459	0.311	0.463	0.539	0.499	0.242	0.428	0.245	0.430
Scotland	0.110	0.313	0.114	0.318	0.106	0.308	0.086	0.280	0.121	0.326	0.113	0.316
Number of siblings	2.383	1.790	2.388	1.794	2.378	1.786	2.177	1.569	2.451	1.844	2.427	1.839
Mother works	0.681	0.466	0.678	0.467	0.683	0.465	0.809	0.393	0.643	0.479	0.649	0.477
Urban	0.451	0.498	0.453	0.498	0.449	0.497	0.467	0.499	0.449	0.498	0.443	0.497
School type												
Grammar	0.101	0.301	0.111	0.314	0.092	0.289	0.152	0.359	0.094	0.292	0.080	0.272
Secondary modern	0.191	0.393	0.191	0.393	0.192	0.394	0.152	0.359	0.202	0.401	0.202	0.402
Private	0.019	0.136	0.018	0.133	0.020	0.139	0.045	0.207	0.012	0.109	0.012	0.108
Single sex school	0.202	0.401	0.214	0.410	0.191	0.393	0.288	0.453	0.188	0.391	0.170	0.375

Continued

Table 1: Descriptive Statistics

	Whole Sample		Females		Males		Income > £60 pw		Income ≤ £60 pw Females		Income ≤ £60 pw Males	
Variables	Mean	St. Dev.	Mean	St. Dev.	Mean	St. Dev.	Mean	St. Dev.	Mean	St. Dev.	Mean	St. Dev.
Peer effects (% of non-manual fathers in school)												
10-19	0.146	0.353	0.153	0.360	0.139	0.346	0.095	0.293	0.168	0.374	0.152	0.359
20-29	0.161	0.368	0.164	0.370	0.158	0.365	0.128	0.334	0.173	0.378	0.167	0.373
30-39	0.108	0.311	0.106	0.308	0.110	0.313	0.084	0.278	0.113	0.316	0.117	0.322
40-49	0.068	0.251	0.060	0.238	0.074	0.263	0.063	0.244	0.059	0.236	0.078	0.268
50-59	0.066	0.249	0.066	0.249	0.067	0.249	0.070	0.254	0.066	0.248	0.065	0.247
60-69	0.054	0.227	0.057	0.232	0.052	0.222	0.082	0.274	0.049	0.217	0.045	0.207
70-79	0.025	0.156	0.021	0.142	0.029	0.168	0.043	0.203	0.017	0.131	0.022	0.148
80-89	0.050	0.217	0.055	0.229	0.044	0.205	0.121	0.326	0.036	0.186	0.024	0.154
Number of observations	7729		3835		3894		1654		3000		3075	

Table 2: Probit Estimates for Participation in Higher Education: Whole, Female and Male Samples

	Whole Sample			Females			Males		
Variables	Marginal Effects[a]	S.E.[b]	P-value	Marginal Effects[a]	S.E.[b]	P-value	Marginal Effects[a]	S.E.[b]	P-value
Female	-0.0034	0.0028	0.224						
Ability test scores									
Arithmetic age 7	0.0157	0.0072	0.026	0.0118	0.0088	0.170	0.0236	0.0105	0.023
Reading age 7	0.0011	0.0122	0.929	-0.0022	0.0178	0.899	0.0046	0.0153	0.766
Maths age 11	0.1043	0.0116	0.000	0.0752	0.0139	0.000	0.1239	0.0173	0.000
Reading age 11	0.1165	0.0156	0.000	0.1045	0.0205	0.000	0.1174	0.0215	0.000
General ability age 11	0.0083	0.0130	0.524	0.0202	0.0158	0.206	-0.0071	0.0193	0.711
Parental income at 16									
% non-father income	-0.0029	0.0052	0.574	0.0157	0.0063	0.010	-0.0278	0.0085	0.001
Total household income	0.0002	0.0001	0.014	0.0002	0.0001	0.014	0.0001	0.0001	0.301
Family background									
Father with higher education	0.0192	0.0075	0.001	0.0138	0.0089	0.055	0.0250	0.0117	0.006
Father with secondary education	0.0073	0.0036	0.034	0.0056	0.0043	0.175	0.0087	0.0054	0.087
Mother with higher education	0.0160	0.0080	0.014	0.0109	0.0089	0.144	0.0213	0.0133	0.040
Mother with secondary education	0.0023	0.0032	0.467	-0.0012	0.0039	0.764	0.0068	0.0048	0.144
House ownership	0.0161	0.0036	0.000	0.0163	0.0046	0.000	0.0144	0.0052	0.004
Father non-manual	0.0083	0.0035	0.011	0.0075	0.0043	0.061	0.0091	0.0051	0.061
Scotland	0.0125	0.0061	0.020	0.0080	0.0073	0.209	0.0162	0.0095	0.046
Number of siblings	-0.0008	0.0010	0.425	-0.0021	0.0013	0.088	0.0008	0.0015	0.594
Mother works	-0.0007	0.0034	0.844	-0.0029	0.0043	0.490	0.0033	0.0047	0.495
Urban	-0.0027	0.0029	0.346	-0.0007	0.0035	0.837	-0.0046	0.0043	0.278
School type									
Grammar	0.0095	0.0051	0.035	0.0048	0.0056	0.347	0.0134	0.0083	0.058
Secondary modern	-0.0144	0.0037	0.001	-0.0164	0.0043	0.001	-0.0110	0.0056	0.077
Private	0.0110	0.0113	0.258	-0.0003	0.0099	0.976	0.0287	0.0229	0.095
Single sex school	0.0044	0.0041	0.258	0.0085	0.0056	0.081	-0.0013	0.0054	0.805

Continued

Table 2: Probit Estimates for Participation in Higher Education: Whole, Female and Male Samples

	Whole Sample			Females			Males		
Variables	Marginal Effects[a]	S.E.[b]	P-value	Marginal Effects[a]	S.E.[b]	P-value	Marginal Effects[a]	S.E.[b]	P-value
Peer effects (% of non-manual fathers in school)									
10-19	0.0039	0.0056	0.468	0.0016	0.0066	0.801	0.0036	0.0081	0.638
20-29	0.0063	0.0054	0.208	0.0056	0.0069	0.377	0.0063	0.0077	0.383
30-39	0.0041	0.0056	0.442	0.0112	0.0083	0.117	-0.0023	0.0071	0.751
40-49	0.0045	0.0064	0.458	0.0041	0.0087	0.616	0.0043	0.0088	0.605
50-59	0.0098	0.0067	0.099	0.0129	0.0091	0.088	0.0037	0.0087	0.654
60-69	0.0048	0.0067	0.434	0.0019	0.0072	0.783	0.0063	0.0106	0.513
70-79	0.0181	0.0116	0.049	0.0239	0.0172	0.058	0.0117	0.0143	0.336
80-89	0.0058	0.0067	0.340	0.0092	0.0088	0.225	0.0002	0.0089	0.979
Number of observations	7729			3835			3894		
Wald chi^2	1035.70		0.0000	546.36		0.0000	541.49		0.0000
Pseudo R^2	0.3548			0.3601			0.3636		
Link test for model specification: pred.^{2}	-0.0020	0.0013	0.174	-0.0022	0.0013	0.189	-0.0017	0.0019	0.421

Notes: [a] Estimated marginal effects at the sample means and their standard errors. For dummy variables, they are estimates for a discrete change from 0 to 1.
[b] Hubert-White robust standard errors.

Table 3: Probit Estimates for Participation in Higher Education: Income > £60, Income ≤ £60/Female and Income ≤ £60/Male Samples

	Income > £60 pw			Income ≤ £60 pw Females			Income ≤ £60 pw Males		
Variables	Marginal Effects[a]	S.E.[b]	P-value	Marginal Effects[a]	S.E.[b]	P-value	Marginal Effects[a]	S.E.[b]	P-value
Female	0.0080	0.0136	0.553						
Ability test scores									
Arithmetic age 7	0.0351	0.0338	0.301	0.0023	0.0062	0.713	0.0253	0.0090	0.003
Reading age 7	-0.0120	0.0632	0.849	0.0017	0.0135	0.902	0.0018	0.0122	0.886
Maths age 11	0.2772	0.0482	0.000	0.0497	0.0121	0.000	0.0799	0.0150	0.000
Reading age 11	0.3955	0.0663	0.000	0.0486	0.0158	0.000	0.0775	0.0183	0.000
General ability age 11	-0.0248	0.0644	0.699	0.0240	0.0114	0.037	-0.0045	0.0156	0.773
Parental income at 16									
% non-father income	0.0194	0.0414	0.639	0.0079	0.0049	0.088	-0.0184	0.0072	0.005
Total household income	0.0007	0.0005	0.111	0.0000	0.0001	0.732	0.0003	0.0001	0.082
Family background									
Father with higher education	0.0253	0.0244	0.265	0.0133	0.0091	0.037	0.0256	0.0132	0.005
Father with secondary education	0.0159	0.0172	0.348	0.0041	0.0032	0.170	0.0046	0.0043	0.257
Mother with higher education	0.0192	0.0247	0.410	0.0123	0.0098	0.080	0.0134	0.0130	0.183
Mother with secondary education	-0.0082	0.0165	0.617	-0.0005	0.0028	0.870	0.0065	0.0040	0.082
House ownership	0.0378	0.0170	0.035	0.0089	0.0033	0.002	0.0114	0.0045	0.004
Father non-manual	0.0485	0.0159	0.003	0.0024	0.0029	0.390	0.0047	0.0041	0.216
Scotland	0.0150	0.0297	0.594	0.0053	0.0054	0.243	0.0138	0.0080	0.033
Number of siblings	0.0054	0.0051	0.286	-0.0018	0.0009	0.030	-0.0004	0.0012	0.748
Mother works	0.0040	0.0178	0.825	-0.0032	0.0033	0.292	0.0022	0.0038	0.564
Urban	-0.0182	0.0140	0.198	0.0006	0.0025	0.811	-0.0020	0.0034	0.573
School type									
Grammar	0.0095	0.0198	0.622	0.0065	0.0052	0.115	0.0063	0.0065	0.262
Secondary modern	-0.0303	0.0201	0.173	-0.0087	0.0034	0.019	-0.0108	0.0041	0.025
Private	-0.0093	0.0298	0.765	0.0076	0.0128	0.456	0.0561	0.0413	0.021
Single sex school	0.0139	0.0175	0.416	0.0045	0.0044	0.222	0.0000	0.0046	0.998

Continued

Table 3: Probit Estimates for Participation in Higher Education: Income > £60, Income ≤ £60/Female and Income ≤ £60/Male Samples

	Income > £60 pw			Income ≤ £60 pw Females			Income ≤ £60 pw Males		
Variables	Marginal Effects[a]	S.E.[b]	P-value	Marginal Effects[a]	S.E.[b]	P-value	Marginal Effects[a]	S.E.[b]	P-value
Peer effects (% of non-manual fathers in school)									
10-19	-0.0175	0.0243	0.504	-0.0007	0.0044	0.868	0.0118	0.0082	0.082
20-29	0.0009	0.0243	0.970	0.0057	0.0053	0.226	0.0066	0.0067	0.275
30-39	0.0145	0.0324	0.638	0.0063	0.0057	0.194	-0.0005	0.0058	0.927
40-49	-0.0022	0.0290	0.939	-0.0004	0.0056	0.948	0.0101	0.0088	0.160
50-59	0.0238	0.0308	0.399	0.0023	0.0054	0.650	0.0092	0.0087	0.208
60-69	-0.0147	0.0232	0.558	0.0026	0.0057	0.622	0.0163	0.0130	0.095
70-79	0.0198	0.0353	0.544	0.0226	0.0165	0.033	0.0163	0.0176	0.216
80-89	0.0170	0.0252	0.472	0.0043	0.0070	0.479	-0.0038	0.0071	0.635
Number of observations	1654			3000			3075		
Wald chi^2	330.03		0.0000	332.16		0.0000	370.61		0.0000
Pseudo R^2	0.3251			0.3580			0.3608		
Link test for model specification: pred.^{2}	0.0017	0.0085	0.840	-0.0010	0.0009	0.408	-0.0008	0.0016	0.649

Notes: [a] Estimated marginal effects at the sample means and their standard errors. For dummy variables, they are estimates for a discrete change from 0 to 1.
[b] Hubert-White robust standard errors.

Figure 2: Isoprobability Lines – Ability Measure: Reading Score at Age 11

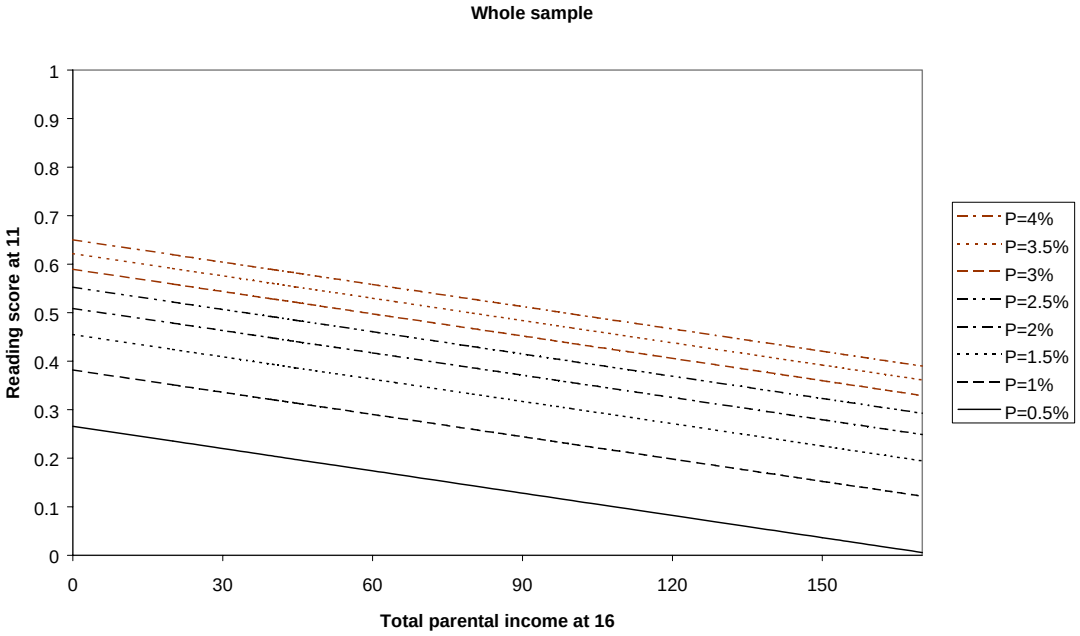

Figure 3: Isoprobability Lines – Ability Measure: Maths Score at Age 11