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Centre for Economic Policy Research

DISCUSSION PAPERS

CHOICE AT SIXTEEN
John Micklewright
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CHOICE AT SIXTEEN

John Micklewright
Queen Mary College
University of London

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CHOICE AT SIXTEEN

John Micklewright

Queen Mary College
University of London

and

Department of Economics
Research School of Social Sciences
Australian National University

February 1988

Abstract

The proportion of sixteen-year olds in Britain who stay on at school is low by OECD standards. The paper examines the probability of completing education at the minimum legal age using micro-data on individuals. Parameter estimates of a reduced form logit model of the leaving probability are obtained for both boys and girls. The rich data set used allows the separate effects of family, school and ability to be assessed. Family background in the form of class and parental education is shown to have a large effect even when ability and school type are controlled for.

Acknowledgements

This research was supported by ESRC grant FOO23 2294 ("Young People's Early Careers: Analysis of Employment History Data in NCDS4"). The part played by a Visiting Fellowship at the Australian National University during Summer 1986 in furthering the work is thankfully acknowledged. I am grateful to the NCDS User Support Group at The City University for their help with the data, to Stephen Nickell and Manuel Arellano for useful discussions, to Joan Payne for advice and for comments on an earlier version of the paper, and to seminar participants in Canberra, Hull, London and Oxford. I also thank two referees for their suggestions.
INTRODUCTION

The role of human capital theory in explaining earnings differentials stimulates one's interest in discovering what distinguishes the group of people which chooses to acquire no further schooling beyond the compulsory level. The proportion of children in Britain who complete their full-time education at the minimum school-leaving age of sixteen – more than 50 per cent – is far higher than that in several other OECD countries.¹ Public concern is often expressed about this situation with blame laid at the doors of a variety of causes. The 1987 Conservative government's Secretary of State for Education said he wished it to become "an almost natural thing for young people to stay-on at school" and claimed that there was a need to change the attitudes of both parents and children.² Others have seen the problem of being one of family finances with, for example, the Labour Party manifesto at the 1987 General Election reviving an old proposal for maintenance allowances for children who stay-on at school.³

This paper looks at the probability of leaving school at the first legal opportunity using micro data on individuals.⁴ The emphasis is thus on elucidating the roles of such factors as family background, income and individual ability at one point in time rather than of changing macroeconomic conditions. (The demand for post-compulsory education in Britain has been studied using aggregate time-series data by Pissarides, 1981.) There is currently little micro evidence in Britain within the labour economics literature on the determinants of school-leaving choice. This contrasts with the situation in the U.S. where there has been quite extensive modelling of schooling decisions using data on individuals.⁵

The data used here are drawn from the National Child Development Study (NCDS), a large and nationally representative cohort study of children born in 1958. The next section describes the NCDS data, bringing out on the one hand their remarkably rich nature and on the other, the problems that they present. Section II defines the particular sample of data and explanatory variables that are used in estimation of a reduced form logit of the probability of leaving at the minimum age. Separate results for boys and girls in England and Wales are presented in Section III. Section IV concludes.
1. THE NATIONAL CHILD DEVELOPMENT STUDY

The NCDS is a longitudinal study of those children born in Great Britain in the week of March 3rd–9th 1958—some 17,500 births. The data collected at birth formed the Perinatal Mortality Survey (PMS) with the four follow-ups at age 7, 11, 16 and 23 being known as NCDS1–4. Information has been obtained from a variety of sources—parents, schools, family doctors, and (increasingly) the children themselves (the most recent sweep at age 23 in 1981 was based on interviews with the cohort members only). The structure of the NCDS is pictured in Figure 1, which also shows the number of children with data at each sweep. It is important to note at the outset that the data obtained at age 16 in 1974 were collected in the spring of that year, several months before the children became eligible to leave school for the first time. This means that it is not until age 23 that we observe the leaving date, hence the inability to conduct the present research before the recent release of NCDS4 data. But it also implies that we have a very detailed picture of each child and his or her family and educational circumstances in the months immediately preceding the choice to leave or stay on.

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**THE NATIONAL CHILD DEVELOPMENT STUDY**

<table>
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<tr>
<th></th>
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<td>Questionnaire</td>
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<td>Questionnaire</td>
<td>Tests</td>
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<td>Subjects</td>
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<td></td>
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<td>Interview</td>
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<td>Nos. with data</td>
<td>17414</td>
<td>15468</td>
<td>15503</td>
<td>14761</td>
<td>14376</td>
<td>12637</td>
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</table>
The subjects covered in the data are wide-ranging. With the exception of income, standard household survey information was collected at each sweep up to age 16, e.g. parental occupations and education, class, tenure, household demographic structure, etc. Parental income data were sought only once but happily for the present research this was at NCDS3 when the children were 16. The information obtained from the schools includes the results of tests of maths and English comprehension administered at age 7, 11 and 16. As well as the measure of ability afforded by these tests, the NCDS also records the results of public examinations taken by the children. Collected in 1978, distinction can be made between total exam success by age 20 and that in the summer of 1974 when the children reached minimum school leaving age.

The breadth of information in the NCDS is of considerable import. Take, for example, the role of parental socio-economic group. While this is likely to be an influence on the school leaving decision one would want to be able to control for other family characteristics which it may proxy, such as income and parental education. Similarly, the level of academic ability at 16 is needed if we wish to isolate the influence of family background at that time as opposed to its accumulated effect over the previous 16 years via the attainment of ability. The wide coverage of the NCDS makes it an attractive source in this respect.

In the present paper, the data obtained at age 23 are used only to identify the leaving date (as described above) and the information collected on wages on entry to the labour market and at 23 are not exploited in the reduced form model that is estimated. These latter data would be relevant in a structural model of schooling choice where the leaving decision would be modelled as a function of expected income streams in the different options. The approach adopted in their paper avoids the modelling problems that would be involved and does not impose any restrictive structure on the data, although it is acknowledged that it cannot indicate the full effects of policies such as educational maintenance allowances that alter the relative values of income streams resulting from different schooling choices.7

What of the drawbacks of the data? Firstly, the NCDS has suffered a substantial element of attrition and the numbers with data at each sweep shown in Figure 1 give an erroneous impression of the success of the panel. By no means all of the 12,537 cases responding at age
23 have data at each of the earlier sweeps. Nor does presence at one sweep necessarily imply that data were successfully collected at that time from all the different sources concerned. For example, the parents might respond and the child might not or vice versa. Even where an interview was carried out, the answers to certain key questions may be missing, e.g. income data are missing from some 20 per cent of cases where a parental interview at NCDS3 was conducted.8

Secondly, we need to acknowledge that the NCDS children reached minimum school leaving age in 1974 at a time when the youth labour market was radically different from now. Only 4 per cent of members of the NCDS cohort who left school in 1974 were unemployed in January 1975. In January 1985, however, 22 per cent of 16 year olds who joined the labour market the previous year were unemployed and a further 49 per cent were without a job and on the government Youth Training Scheme (Social Trends, 1987, Table 3.10). Nevertheless, it is of considerable interest to see how decisions were made in a period more or less free of demand-side constraints and specific government intervention in the youth labour market. Moreover, although the fate of those who do leave has altered radically, the change during the period in the proportion of successive birth cohorts that has left at the minimum age has not been that great.9

Thirdly, there is one specific feature of the cohort which relates particularly to the subject under investigation here. The minimum school leaving age, currently 16, was raised from 15 in September 1972, a level it had been at for some 25 years. Children born in 1958, i.e. the NCDS children, were the first full birth year to be affected. The NCDS families thus faced a new institutional framework which removed their previous free choice at 15. It is also possible that the youth labour market during 1973 was supply constrained with employers starved of their previous stream of 15 year-olds. The conditions facing the NCDS children in 1974, therefore, may have been particularly good, making the comparison with 1987 even more marked.10

II. DATA

The data used in estimation relate to a sub-sample of 5,304 cases out of the 12,537 responding at age 23 to NCDS4. The young people selected were all living in England and Wales at age 16, those in Scotland being discarded on account of the differences in educational
system from the rest of Britain. However, the main reason for the reduction in sample size is missing data at one or more NCDS sweeps – a feature of the panel discussed in Section 1. Full details on the sample may be found in the Appendix.

Table 1 shows that 62.9 per cent of boys in the sub-sample left full-time education at the minimum age (LEAVER = 1) and 55.0 per cent of girls, figures that reflect the differences between the sexes shown by the full NCDS4 sample as well as by administrative data. Girls have traditionally had a lower minimum-age leaver rate due to the lack of employer-based training schemes (apprenticeships) available to them at this age. Table 1 also gives means of explanatory variables used in estimation (further details of their construction are given in the Appendix). All these variables refer to the position when the child was 16. The rich nature of the NCDS data means that it is not difficult to find a large number of variables relating to child, family or school that appear to have significant partial correlations with the leaving probability. However, the economic or sociological explanations which would allow one to interpret these associations as causal relationships are often less obvious. Such problems caution in favour of a reasonably parsimonious specification.

The first group of variables refers to the child's family background. The means of the dummy variables FASTAY and MASTAY when compared to those for LEAVER reflect post-war (and earlier) rises in the proportion of children staying on after the minimum age (itself lower for the parents' generation). The family class dummies PAPROF, PAINT, FASKNMIN AND MAPPINT were selected after experimenting with a full set of parental status variable (the base category includes all manual parents, parents not working and parents not present). Two variables measure the number of older and younger siblings – OLD SiBS and YNG SiBS. There is a considerable literature in the US on the effect of siblings on schooling and earnings – see Behrman and Taubman (1986) for a review. In general the weight of argument is that siblings reduce the amount of schooling received, with this being particularly true for the effect of older siblings. One of the more plausible arguments is that parental time spent with each child furthers the child's academic ability and that this time decreases with family size, but with older children obviously getting more attention prior to the birth of their younger siblings.
### TABLE 1: VARIABLE DEFINITIONS AND MEANS

<table>
<thead>
<tr>
<th>Variable (X)</th>
<th>X(Boys)</th>
<th>X(Girls)</th>
<th>Definition</th>
</tr>
</thead>
<tbody>
<tr>
<td>LEAVER</td>
<td>0.629</td>
<td>0.550</td>
<td>Completed full-time education by Sept. 1974</td>
</tr>
<tr>
<td>PASTAY</td>
<td>0.243</td>
<td>0.245</td>
<td>Father stayed on at school post minimum age</td>
</tr>
<tr>
<td>MASTAY</td>
<td>0.263</td>
<td>0.273</td>
<td>Mother stayed on at school post minimum age</td>
</tr>
<tr>
<td>PAPROF</td>
<td>0.055</td>
<td>0.054</td>
<td>Father Social Class I (Professional)</td>
</tr>
<tr>
<td>PAINT</td>
<td>0.191</td>
<td>0.205</td>
<td>Father Social Class II (Intermediate)</td>
</tr>
<tr>
<td>PASKINN</td>
<td>0.095</td>
<td>0.078</td>
<td>Father Social Class II (Skilled Non-Man)</td>
</tr>
<tr>
<td>MAPPINT</td>
<td>0.110</td>
<td>0.111</td>
<td>Mother Social Class I or II (Prof/Inter)</td>
</tr>
<tr>
<td>OLDSIBS</td>
<td>1.072</td>
<td>1.119</td>
<td>Nos. older siblings</td>
</tr>
<tr>
<td>YNGSIBS</td>
<td>1.147</td>
<td>1.157</td>
<td>Nos. younger siblings</td>
</tr>
<tr>
<td>INCOME</td>
<td>49.99</td>
<td>50.30</td>
<td>Net Household Income (Weekly £s 1974)</td>
</tr>
<tr>
<td>BADFIN</td>
<td>0.123</td>
<td>0.123</td>
<td>Child in family receives free school meals or parents seriously troubled financially in past 12 months.</td>
</tr>
<tr>
<td>LOWSE</td>
<td>0.154</td>
<td>0.144</td>
<td>Living in London or South-East</td>
</tr>
<tr>
<td>WALES</td>
<td>0.075</td>
<td>0.071</td>
<td>Living in Wales</td>
</tr>
<tr>
<td>COMPREN</td>
<td>0.568</td>
<td>0.526</td>
<td>Comprehensive School (non-selective state run)</td>
</tr>
<tr>
<td>GRAMMAR</td>
<td>0.117</td>
<td>0.165</td>
<td>Grammar School (higher ability state run)</td>
</tr>
<tr>
<td>PRISCH</td>
<td>0.052</td>
<td>0.046</td>
<td>Private School</td>
</tr>
<tr>
<td>ABILITY</td>
<td>40.4</td>
<td>38.9</td>
<td>Sum of Age 16 Maths (0-31) and Comprehension (0-35) test scores</td>
</tr>
<tr>
<td>n</td>
<td>2,639</td>
<td>2,665</td>
<td>Sample size</td>
</tr>
</tbody>
</table>

**Notes**

1. All variables refer to the position when the child was 16.
2. "Father" and "Mother" refer to the persons acting as the father and mother figures at NCDS3 and not necessarily the natural parents.
3. The means of INCOME refer only to the 2,027 boys and 2,021 girls whose parents provided complete income data at NCDS3 - see main text for discussion.
4. Full definitions given in the Appendix.

INCOME should have a negative effect on the probability of leaving school at the minimum permissible age, working in particular through discount rates with lower income households being constrained in their choices (e.g. Pissarides 1981, Atkinson et. al., 1983, Rice 1987). The data used to construct this variable are present only for about three-quarters of the sample and in the majority of other cases is missing because the parents refused to provide or did not know one or more of the three bits of information required: father's pay, mother's pay and other household income. Even when present, the data are not in an ideal form since the information from each source was requested merely in the form of discrete ranges. In order to obtain a single income variable, the sum was taken across the relevant mid-points of the 11
bounded ranges and an extraneous estimate of the median of the top unbounded range—see Micklewright (1986) for further details. The resulting measure must be treated with considerable caution. As an alternative to INCOME, the dummy variable BADFIN is used as a proxy for low income households. This is available for the full sample and takes the value one if the family receives means-tested free school meals for any of its children or answers in the affirmative to a question about financial trouble. The final variables in the family background group are two regional dummies, LONSE and WALES, chosen after preliminary estimation with a full set of such variables. Region may reflect a variety of influences, including preferences (e.g. strong tradition of valuing education), school quality and labour demand.

The second group of variables relates to the child’s school during the final year of compulsory schooling, and comprises three dummies indicating the type of school attended—COMPRHN, GRAMMAR and PRISCH—with the base being secondary modern schools (lower ability state run). The importance of type of school in determining length of schooling has been emphasised by Halsey, Heath and Ridge (1980). School type in part is a reflection of ability—as in the case of grammar schools which are selective on ability at age 11— but given that we can control for ability in the NCDS, school type is included here for other reasons. The type of school may reflect peer pressure from other pupils and career guidance from teachers and in the case of PRISCH, parental preferences, income, and capital as well.13

The last variable is a measure of the child’s academic ability and is the sum of scores in Maths and English Comprehension tests at age 16. This should be seen principally as affecting further earnings streams and with this interpretation the information contained in the NCDS on success in public examinations (O-levels and CSEs) in the summer of 1974 might at first seem more appealing. After all, these qualifications give a public demonstration of academic performance to employers making wage offers and indeed many jobs would not be open to those who do not possess a certain number of O-levels. However, the exam results are almost certainly endogenous to the school leaving decision. If it has already been decided that a child will leave school as soon as possible then he or she might not try very hard to pass these exams if secure in the knowledge that a job had already been obtained (the exams may not even be entered). The age 16 tests have the advantage of being taken on an equal basis by all
the children in the sample without regard to their decisions about future education.\textsuperscript{14}

\section*{II RESULTS}

Results were obtained from maximum likelihood estimation of the probability, $p$, of leaving school at the minimum permissible age where the functional form adopted for $p$ is the logit given by

$$p = 1/[1 + \exp(-\beta^T\mathbf{x})]$$

The likelihood function for the sample is therefore

$$L = \prod_{j=1}^{n} \frac{1 + \exp(-\beta^T\mathbf{x}_j)}{1 + \exp(-\beta^T\mathbf{x}_j)} \prod_{k} \frac{\exp(-\beta^T\mathbf{x}_j)}{1 + \exp(-\beta^T\mathbf{x}_j)}$$

where $j = 1$ to $m$ is the set of leavers and $k = 1$ to $n$ is the set of stayers.

As an aid to judging the importance of each estimated parameter, note that

$$\frac{dp}{dX_i} = p(1 - p)B_i \text{ where } X_i \text{ is the } i\text{th element of } \mathbf{x}.$$  \hspace{1cm} (3)

Thus at $\beta = 0.5$ the estimated effect on the predicted probability of a unit change in a continuous variable, or the turning on of a dummy variable, is approximately equal to $\beta/4$.

**Family Background, School Type and Ability**

We begin with the role of family background, school and ability, leaving for the moment the effect of income. Separate estimates of $\beta$ for boys and girls are contained in Table 2. The specification in columns 1 and 3 excludes the school type variables and ABILITY; the results therefore show the "full" effect of family background, making no distinction between the direct effect on the school-leaving decision and its indirect effect coming through ability and school type. Thus when we see that a child from a certain class background has a lower probability of leaving this may be due to one or both of two effects. Firstly, the class in question may be associated with a high rate of investment in human capital from infancy onwards, resulting in a more able child and a school type that is selective or private, both of which effects may result in an increased change of staying on at school. Secondly there may be a direct effect whereby parents from that class encourage or coerce a child to stay on
irrespective of his or her ability or type of school. Sociologists have referred to these indirect and direct impacts as the "primary" and "secondary" effects respectively of social class stratification with British educational policy traditionally concerning itself with the latter, aiming to eliminate class differences among those with equal ability (Halley, Heath and Ridge, 1980). Comparison of results with and without ability and school type will allow the two effects to be separated out.

The family background variables in columns 1 and 3 of Table 2 all have the expected sign and several are well determined. Parental education (PASTAY and MASTAY) not surprisingly has a large and significant effect and this is true of both parents. Using equation (3) and setting \( \hat{\beta} = 0.5 \) we can see that the *ceteris paribus* effect of either parent staying-on after the minimum leaving age is to reduce the child's probability of leaving by some 20 per cent points. Note that the apparent differences between the coefficients of PASTAY and MASTAY are not significant within or across equations. Father's class is important for both sexes; the non-manual social class II and III dummies (PAINT and PASKNMN) have roughly speaking the same impact for the boys as the parental education variables. The effect of PAPROF is much more substantial. By contrast, the coefficient of MAPFINT is quite large and significant for the girls but small and barely significant for the boys, suggesting that professional mothers have more influence on their daughters than their sons. As expected, siblings significantly increase the probability of early leaving. The results for boys provide some evidence that the effect of elder brothers and sisters is greater than that of younger siblings (t of difference = 2.7). Note that the coefficients are such that it takes at least three siblings to outweigh the impact of a parent having stayed on. Finally, the regional coefficients provide some illustration of the importance of geographical location with a boy in London and a girl in Wales having substantially lower probabilities of leaving.

The specification in columns 1 and 3 is designed to show the "full" effect of family background on the leaving decision, including that coming through the child's academic ability and type of school attended. We should not be surprised therefore by the considerable reductions in many of the family background coefficients in columns 2 and 4 when school type and ability are introduced into the equations. In terms of goodness of fit these latter equations
<table>
<thead>
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<th>3</th>
<th>4</th>
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<tr>
<td>Constant</td>
<td>0.969</td>
<td>6.402</td>
<td>0.435</td>
<td>4.479</td>
</tr>
<tr>
<td></td>
<td>(0.094)</td>
<td>(0.317)</td>
<td>(0.089)</td>
<td>(0.261)</td>
</tr>
<tr>
<td>PASTAY</td>
<td>-0.791</td>
<td>-0.448</td>
<td>-0.689</td>
<td>-0.326</td>
</tr>
<tr>
<td></td>
<td>(0.115)</td>
<td>(0.137)</td>
<td>(0.114)</td>
<td>(0.129)</td>
</tr>
<tr>
<td>MASTAY</td>
<td>-0.895</td>
<td>-0.510</td>
<td>-0.785</td>
<td>-0.440</td>
</tr>
<tr>
<td></td>
<td>(0.108)</td>
<td>(0.128)</td>
<td>(0.107)</td>
<td>(0.121)</td>
</tr>
<tr>
<td>PAPROF</td>
<td>-2.080</td>
<td>-1.339</td>
<td>-1.828</td>
<td>-1.369</td>
</tr>
<tr>
<td></td>
<td>(0.270)</td>
<td>(0.304)</td>
<td>(0.275)</td>
<td>(0.298)</td>
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<tr>
<td>PAINT</td>
<td>-0.880</td>
<td>-0.512</td>
<td>-0.765</td>
<td>-0.469</td>
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<td></td>
<td>(0.120)</td>
<td>(0.143)</td>
<td>(0.116)</td>
<td>(0.130)</td>
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<td>PASONMY</td>
<td>-0.700</td>
<td>-0.467</td>
<td>-0.513</td>
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<td>(0.175)</td>
<td>(0.160)</td>
<td>(0.178)</td>
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<td>-0.302</td>
<td>-0.445</td>
<td>-0.944</td>
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<td>(0.152)</td>
<td>(0.179)</td>
<td>(0.158)</td>
<td>(0.177)</td>
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<td>0.299</td>
<td>0.229</td>
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<td></td>
<td>(0.044)</td>
<td>(0.050)</td>
<td>(0.038)</td>
<td>(0.044)</td>
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<td>YNG SIBS</td>
<td>0.192</td>
<td>0.077</td>
<td>0.243</td>
<td>0.106</td>
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<td>(0.042)</td>
<td>(0.050)</td>
<td>(0.039)</td>
<td>(0.044)</td>
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<td>LONSE</td>
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<td>(0.124)</td>
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<td>(0.127)</td>
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<td>-0.746</td>
<td>-0.608</td>
<td>-0.908</td>
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<td>(0.174)</td>
<td>(0.202)</td>
<td>(0.171)</td>
<td>(0.190)</td>
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<td>COMP RHN</td>
<td>-0.192</td>
<td>-0.105</td>
<td>-0.563</td>
<td>-0.647</td>
</tr>
<tr>
<td></td>
<td>(0.132)</td>
<td>(0.120)</td>
<td>(0.175)</td>
<td>(0.194)</td>
</tr>
<tr>
<td>GRAMMAR</td>
<td>-1.500</td>
<td>-1.189</td>
<td>-0.988</td>
<td>-1.021</td>
</tr>
<tr>
<td></td>
<td>(0.342)</td>
<td>(0.319)</td>
<td>(0.007)</td>
<td>(0.006)</td>
</tr>
<tr>
<td>PRISCH</td>
<td>1.405.31</td>
<td>1.070.33</td>
<td>1.517.17</td>
<td>1.260.13</td>
</tr>
<tr>
<td>-log-likbd</td>
<td>0.193</td>
<td>0.385</td>
<td>0.173</td>
<td>0.313</td>
</tr>
<tr>
<td>Pseudo R²</td>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
</tbody>
</table>

**NOTES**

1. Estimated standard errors in brackets.
2. Pseudo $R^2$ calculated as $1 - \frac{\text{model log-likelihood/} \log \text{ likelihood with optimal constant only}}{\log \text{ likelihood with optimal constant only}}$. See Maddala (1983) p.40.
3. Log-likelihood with optimal constant only = -1,740.35 (boys) and -1,833.84 (girls).

represent a considerable improvement. With only four extra parameters log likelihood improves by over 300 points for the boys and 250 points for the girls. These changes are reflected in the significance of the coefficients of several of the new variables, in particular ABILITY, the sum of age 16 test scores. Starting from a predicted probability of leaving of one half, a *ceteris paribus* increase in ABILITY of 12 points (about one standard deviation for both sexes)
would roughly reduce the probability of leaving for a girl to one quarter and a boy to one fifth. Despite family background and now ability being in the model, school type has some importance. There is no significant difference associated with being at a comprehensive school compared with a secondary modern school (the base category) but the coefficients on GRAMMAR and PRISCH are better determined with t-ratios in the order 3 or 4. Being a grammar school pupil might be characterised as being worth about one half-standard deviation of ABILITY score while the PRISCH coefficients are really quite large, being greater than that of any other dummy variable for the boys. As discussed in Section II school type may be proxying a number of effects.  

Ability and school type are obviously important, but what now of family background? The results show clearly that family background does have a direct or secondary effect over and above its influence through ability and schooling. Several of the parental education and class dummies retain t ratios in excess of 3.5 with the coefficients being typically half to three-quarters of their previous values. The same applies to the effect of older siblings but the coefficients of YNGSIBS suggest that most of the effect of younger brothers and sisters comes through a deleterious impact on home investment in the child prior to the leaving decision rather than on the decision itself. The most important differences between the results for girls and boys relate to the constant terms and the coefficients of ABILITY. However, the pattern of the latter is the opposite of that needed to help explain the higher staying-on rate for girls, in that their coefficient is smaller in absolute size. The difference in the coefficients has a t of 2.5 and at the average level of ability represents a difference in effect equivalent to some of the larger dummy variable coefficients in either equation. The lower average leaving rate for girls remains unexplained by the model and this is reflected in the very large difference in the estimated constants (given the similarity in the coefficients and means of other variables).

Table 3 summarises the results from columns 2 and 4 of Table 2 with some calculations of predicted probabilities for particular combinations of characteristics.
Table 3: Predicted Probabilities of Leaving for Given Characteristics

(Based on Columns 2 and 4 of Table 2)

<table>
<thead>
<tr>
<th></th>
<th>Boys</th>
<th>Girls</th>
</tr>
</thead>
<tbody>
<tr>
<td>1)</td>
<td>Base chars. except COMPRHN=1 plus median ABILITY</td>
<td>0.777</td>
</tr>
<tr>
<td>2)</td>
<td>as (1) except top decile ABILITY</td>
<td>0.335</td>
</tr>
<tr>
<td>3)</td>
<td>as (2) except GRAMMAR=1</td>
<td>0.242</td>
</tr>
<tr>
<td>4)</td>
<td>as (2) expect OLDSII=1 YNGSII=1</td>
<td>0.392</td>
</tr>
<tr>
<td>5)</td>
<td>as (1) except PASTAY, MASTAY, PAINT, MAPFINT=1</td>
<td>0.339</td>
</tr>
<tr>
<td>6)</td>
<td>as 95) except top decile ABILITY</td>
<td>0.069</td>
</tr>
</tbody>
</table>

Note: Median ABILITY is 41 for boys and 39 for girls. Top decile are 57 and 55 respectively.

The top line gives the probabilities with base characteristics, except the school is taken as a comprehensive, the type attended by the majority of the sample. The parents did not stay on at school, they are manual workers or do not work and the child has no siblings. With these characteristics and the median level of ability, the predicted leaving probability is about three-quarters for a boy and two-thirds for a girl. The second line shows that with the same family background, being in the top 10 per cent of ability at 16 is associated with a greatly increased probability of remaining at school, but there is still a one-third chance of leaving if a boy and one-quarter if a girl. Keeping the same ability and parents, lines 3 and 4 contrast being an only child and at a grammar school with having two siblings and attending a comprehensive school. Lines 5 and 6 show the probabilities if the parents themselves stayed-on and are from social class II (I or II for the mother). At the median ability scores (line 5), the leaving probabilities are 1 in 3 if a boy and just 1 in 6 if a girl. (If the father was Professional and the child was at a private school – not shown – the probabilities would drop still further to around 1 in 20 and 1 in 36 respectively, holding other characteristics constant.) With top decile scores the chances of leaving are just 7 per cent and 4 per cent respectively.

**Income**

The equations in Table 2 contain no measure of income, a possible influence on leaving decisions of considerable policy interest. There is certainly a gross correlation between higher
income and staying on at school. For those cases in the sample where data on net household income are available, mean INCOME for boys who left at 16 is £47.3 per week compared with £54.6 for households where the boy stayed on (the position for girls is very similar). However, this situation may merely reflect differences in social class composition of the two groups; if we accept that class and other family influences such as parental education do play an underlying role then we must look for the partial effects of income when controlling for these factors. The equations in Table 2 were re-estimated using the smaller samples where income data were present and the coefficients of INCOME are shown in Table 4.

The addition of income clearly makes no difference for the boys; INCOME is quite insignificant irrespective of whether ability and school type are included or not. The picture for the girls is slightly different. With no ability measure or school dummies present (column 3) the coefficient has the expected sign although is not particularly well determined. The results indicate that an increase in INCOME of £20 (one standard deviation) would reduce a \( \beta \) of 50 per cent by only 4.5 per cent points.19 But as with the boys, income adds very little to

<table>
<thead>
<tr>
<th>Table 4: The Effect of INCOME</th>
</tr>
</thead>
<tbody>
<tr>
<td><strong>BOYS</strong></td>
</tr>
<tr>
<td>( n = 2,027 )</td>
</tr>
<tr>
<td>Column in Table 2 with same specification</td>
</tr>
<tr>
<td>1</td>
</tr>
<tr>
<td>(no ability or schools)</td>
</tr>
<tr>
<td>INCOME/10</td>
</tr>
<tr>
<td>(0.033)</td>
</tr>
<tr>
<td>- log likelihood</td>
</tr>
<tr>
<td>- log likelihood in restricted model (no INCOME)</td>
</tr>
</tbody>
</table>

log-likelihood when ABILITY and school dummies are included and although the coefficient retains the right sign the \( t \)-ratio is insignificant at conventional levels.

In interpreting these results, it should be noted that INCOME contains considerable
measurement error given the method of its construction and the collection of the data concerned. The failure to detect income effects should not be taken therefore as conclusive evidence of their absence. As an alternative to INCOME, the dummy variable BADFIN described in Section II gives an indication of a low income family, defined using responses to a different set of questions. The introduction of this variable into the equations in Table 2 produces a qualitatively similar pattern of results to the use of INCOME. Insignificant for boys whether included in the presence or absence of ability and school type, the addition of BADFIN results in a small increase in log-likelihood of 5.6 points for girls when family background (and regional dummies) only are included. The estimated coefficient of 0.513 is not small (standard error 0.157) suggesting a ceteris paribus increase in leaving probability for children from low income families of around 10 per cent points (depending on the β adopted in [3]). However, with ability and school type in the model, BADFIN drops away for the girls also.

A final handle on this subject is the self-reported information contained in the NCDS on the influence of income in the decision over school-leaving. For obvious reasons this information cannot be regarded as particularly reliable but nevertheless it is of some interest. On the one hand there is the view from the parents who were asked at NCDS3 whether they would have liked their children to leave at 15 and their reasons for this if they answered in the affirmative. Multiple answers were permitted but the parents of only 53 boys and 54 girls would admit to "financial reasons" as a contributory factor. The children were (quite separately) asked when they thought they would leave school and the reasons for this. Some 8 per cent of boys and 9 per cent of girls thought they were both likely to leave school at 16 and gave their family needing the money as one reason. If introduced into the family background equations in Table 2, a dummy variable indicating this combination of responses was large and significant for both sexes, with the effect surviving, albeit diminished in strength, when ability and school type were also included. The interpretation of such a variable within the model is difficult but it is perhaps worth noting that yet again the increase in log-likelihood and size of coefficient was notably larger for the girls.

Drawing together the different strands of evidence on the effect of income, the results suggest a stronger effect for girls but that with neither sex does the impact leap out from the
data. The finding that household income is insignificant for boys confirms that of Rice (1987) in her study of labour force participation of 16 and 17 year olds using the UK Family Expenditure Survey. This source has vastly superior income data to the NCDS but on the other hand Rice was unable to control for parental education, or for the child's ability and school type which have been shown here to reduce the income effect to insignificance. Finally we should, like Rice, note that a reduced form model of the type estimated here cannot be used to deduce the full impact of the introduction of a system of educational maintenance allowances providing payments conditional on further schooling. This would require the estimation of a structural model of the leaving decision using estimates of streams of expected future income in both options.

III CONCLUSIONS

This paper has presented results from the estimation of reduced form logit models of the probability of leaving school in England and Wales at the minimum school leaving age, using microdata from the National Child Development Study. Separate estimates were produced for boys and girls. Not surprisingly, family background as measured by parental education, class and numbers of siblings was found to have a substantial impact. Of more interest and immediate policy relevance, is that between half and two-thirds of this effect remains when controls for the children's academic ability and type of school are introduced. This is illustrated by considering the situation of a child in the top 10 per cent of academic ability as measured in this paper and attending a non-selective state school. A manual background with parents who did not themselves stay on, leads to a predicted probability of leaving of 33 per cent if a boy and 27 per cent if a girl, even if he or she had no siblings. With professional parents who themselves had post-compulsory schooling the probabilities would be negligible. It is clear that there does not exist equality of educational outcomes for children with equal ability following compulsory schooling.
Appendix: Definition of Variables (including NCDS codes)

Data used in estimation were cases satisfying the following criteria:

(i) Resident in England or Wales at 16
(ii) Educational exam history at 20
(iii) Ability tests and parental, school and individual data at 16
(iv) Ability tests and parental data at 7
(v) Parental data at birth.

LEAVER = 1 if CTAES < 198 i.e. full-time education completed by September 1974, the beginning of the first possible voluntary academic year (see also footnote 11)

PASTAY = 1 if natural father present at NCDS3; (N2375=1) and natural father at NCDS1 reported as staying on (N194=2) or natural father not present and father-figure reported at NCDS3 as leaving school at age 16+ (N2396 > 5)

MASTAY = 1 if natural mother present at NCDS3 (N2371=1) and natural mother at PMS reported staying-on (N537=3, 4 or 6) or natural mother not present at NCDS3 but mother figure reported leaving at age 16+ (N2397 > 5).

PAPROF = 1 if father employed (N2383=6) and social class I (N2384=1).

PAINT = 1 if father employed and social class II (N2384=2).

PASKNMM = 1 if father employed and social class III and skilled non-manual (N2384=3).

MAPINT = 1 if mother works (N2392=1) and social class 1 or II (N2393=1 or 2).

OLDSIBS = Number of older brothers and sisters with same natural mother (N2367 + N2369 minus values set to zero). NB Coding notes indicate that adopted and step-children may have been included.

YNGSIBS = as OLDSIBS, but N2368 + N2370.

INCOME = sum of midpoints of ranges of weekly income from father’s pay, mother’s pay and other household income (N2462–N2467). £60+ ranges set at £72.5, see Micklewright (1986).

BADFIN = 1 if N2440 = 1 or N2441 = 1.

LONSE = 1 if N2350=6.

WALES = 1 if N2350=10.

COMPRHN = 1 if N2101=1 (Comprehensive).

GRAMMAR = 1 if N2101=2 (Grammar)

PRISCH = 1 if N2103=1 or 2 (Independent or Direct Grant).

ABILITY = sum of Mathematics Comprehension score aged 16 (N2930) and Reading Comprehension, aged 16 (N2928).
Footnotes

1. The proportion of UK 16-18 year olds in full-time education in 1981 was only 32% compared for example with 58% in France, 69% in Japan and 79% in the USA (Social Trends, 1987, 3.11).


3. See Rice (1987) for a summary of the history of such proposals for maintenance allowances. The suggestion that a single system of allowances could be set up to support a young person in either education or employment-based training can be found in The Economist (2nd May 1987, p.20) and in the 1987 SDP/Liberal Alliance manifesto.

4. Throughout this paper "leaving school" should be taken as meaning leaving full-time education. A substantial number of 16 year olds in Britain do in fact leave school but stay in full-time education in Colleges of Further Education or Tertiary Colleges where courses are taken which may be the same as those at school. Such children are included amongst those who stay on.

5. One British study using micro data is that of Rice (1987) who looks at the participation of 16 and 17 year olds using the 1976 Family Expenditure Survey. See also Atkinson et. al. (1983). The work of sociologists Halsey, Heath and Ridge (1980) should also be noted. US studies include Kohn et. al. (1976), Willis and Rosen (1979), Mare (1980) and Behrman and Taubman (1986).

6. Earlier sweeps of the NCDS have seen extensive use e.g. see Fogelman (1983) for a collection of papers based on NCDS3. The NCDS data are available from the ESRC Data Archive at the University of Essex.

7. The possible use of NCDS4 earnings data in a structural model, following Willis and Rosen (1979), is discussed in Micklewright (1987). Problems include the length of the NCDS panel being very short for the estimation of expected lifetime earnings (the latest wage being recorded as a maximum of only 7 years after entry to the labour market compared with 20 years in the Willis and Rosen data) and the variation in further schooling amongst those staying on at the minimum age (this problem is discussed in detail by Willis and Rosen who admit their failure to produce a satisfactory resolution). The reduced form approach also implies that no attempt is made here to discover whether the effect of any variable comes through expected earnings, discount rates, preferences or a combination of the three.

8. NCB comparisons of NCDS4 data with information drawn from the Census conducted in the same year are reasonably encouraging (National Children's Bureau, 1984, Chapter 6).

9. The proportion of 16 year olds in Great Britain who were in full-time education in January 1985 and eligible to leave in the summer of 1984 is given in Social Trends as 45 per cent compared with 41 per cent in 1975/6 (1986, Table 3.7 and 1987, Table 3.19).

10. The total number of school leavers of all ages in G.B. who joined the labour force in Summer 1973 was about half that of earlier years due to the raising of the minimum age (DE Gazette, Dec. 1975, p.1269).
11. Had children living in Scotland at 16 been included, the figures would have been 63.1 and 53.1 per cent and this compares with proportions in the total NCDS4 sample (including those with missing information) of 64.7 per cent for boys and 57.1 per cent for girls. The variable used to define LEAVER is one derived by the NChi from the raw data. This ignores gaps in education of up to 6 months but the importance of this for 16 year old leaving is minimal. A handful of cases were excluded with CTAE2 missing. The 8 per cent point gap between the leaver rates for boys and girls in the NCDS data does appear somewhat high compared to that indicated by administrative sources. Social Trends (1986, Table 3.7) gives a difference of only 4 points in GB figures for 1975/6 (61 per cent for boys and 57 per cent for girls) although by 1983/4 this had widened to 10 points (60 per cent and 50 per cent respectively. 

12. The arguments Behrman and Taubman review range from biological (children with older siblings ceteris paribus have older mothers, thus resulting in higher incidence of births defects) through to the possibility of a declining marginal utility of parenthood with family size. Note that the specification used in this paper of numbers of older and younger siblings is equivalent to total number of siblings and birth order used by Behrman and Taubman.

13. It should be noted that 1974 came towards the end of a period of much re-organisation of state schooling and school type at 16 does not necessarily reflect school type throughout 11 to 16. About one half of the comprehensive schools attended by the NCDS children had previously been secondary modern schools.

14. The reader who nevertheless wishes to see results from equations using success in exams as a measure of ability is referred to Micklewright (1987). Note that the discrete nature of the O-levels and CSEs precludes their satisfactory instrumentation by standard regression methods. Some 50 per cent of the boys and 44 per cent of the girls have no O-levels (or equivalent CSE grade 1s) while 38 per cent of each sex has no CSEs grades 2-5.

15. As estimate of the standard error of the difference between two estimated coefficients, \( \hat{\beta}_1 \) and \( \hat{\beta}_2 \), within an equation is given by

\[
\text{SE}(\hat{\beta}_1 - \hat{\beta}_2) = \sqrt{\text{Cov}(\hat{\beta}_1, \hat{\beta}_2) + \text{Cov}(\hat{\beta}_1, \hat{\beta}_2)^2}
\]

16. A more general equation including a full set of family class and employment status variables showed no significant differences between different manual classes of parent or between working and non-working parents, and no significant effects from single parenthood.

17. Experimenting with a set of seven dummies for different ABILITY levels showed a linear specification for this variable to be quite appropriate. The restriction of equality of coefficients of separate scores for maths and comprehension tests embodied in the specification in columns 2 and 4 of Table 2 was accepted by the data with ease, twice the difference in log-likelihood from the unrestricted form being only 2.0 for the boys and 0.9
19

for the girls. No significant interactions of ability with parental class or school type could be found.

18. An additional variable relating to the school which proved completely unsuccessful in the presence of school type and ability was the school's own report of the proportion of its pupils in the relevant age group that stayed on in the year prior to the raising of the school leaving age. It is rather surprising that this variable did not pick up variation in school quality, locality or peer pressure not controlled for by the simple school type dummies.

19. Results from logarithmic specification for income in an equation with no ability or school dummies indicates an elasticity of the predicted leaving probability for girls at $\beta = 0.5$ of 0.189, again suggesting a fairly small response.
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