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"The Marginal and Average Returns to Schooling"

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The Marginal and Average Returns to Schooling

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Abstract:

The existing literature now features many examples where log wages are linear in years of schooling and which effectively attempt to correct for least squares bias using instruments based essentially on a single variable. Two recent developments, taken together, cast some doubt on the downward bias in least squares estimates of the return to schooling that have been an important feature of the recent literature. The first is the realization that instrumental variables (IV) only estimate the effect of some treatment if the effect is the same for everyone. The second is that IV may only estimate the effect of the treatment on the individuals whose choices are affected by the instrument in question.

The main features of this paper are that: because of its importance in the IV context, the linearity assumption is tested; and the possibility that different instruments may affect different margins is explored. The paper finds: a large and significant downward bias in the least squares estimate of the returns to schooling; the returns to schooling are highest at low and high levels of schooling suggesting significant sheepskin effects; and the estimates are stable with respect to the choice of instruments. In particular, instruments which might be expected to affect children likely to be constrained by the minimum school leaving age seem to imply the same estimated returns as do instruments likely to be the most important in determining participation in higher levels of schooling.

Our findings suggest that individual heterogeneity in returns to schooling may not be the most important explanation of the negative least squares bias reported in this and earlier studies. Rather, our results suggest that the presence of downward bias in least squares could be attributed to measurement error in the schooling variable.

* Material from the General Household Survey used in this study have been made available to us by the Economic and Social Research Council's (ESRC) Data Archive with the permission of the Controller of Her Majesty's Stationary Office. Helpful comments from Michael Devereux, John Fitzgerald, Paddy Geary, Gauthier Lanot, Costas Meghir, Anthony Murphy, Donal O'Neill and participants in the Dublin Economics Workshop, and seminars at Maynooth and University College Dublin. Much of this research was completed while the first author was a member of staff at the Economics Department, Maynooth College. The authors gratefully acknowledge ESRC support for the Centre for Microeconomic Analysis of Fiscal Policy at IFS, and the European Union Human Capital Mobility Programme grant number ECR 930225. The usual disclaimer applies.

1. Introduction

Simultaneity between schooling on earnings can arise for a number of reasons and recent empirical research suggests that this has caused least squares estimates of the rate of return to schooling to be biased in a downwards direction - usually by quite sizable amounts. Typically this bias is attributed either to the presence of individual heterogeneity not directly accounted for, or to the existence of measurement error in the schooling variable (Zvi Griliches, 1977). It is important to try to distinguish between these two explanations since the former implies that the estimated returns cannot be used to make general policy prescriptions - such as inferring the effect of some policy change on a different marginal group to that identified by the instrumenting procedure used. Two recent developments, taken together, cast some doubt on this downward bias in least squares estimates of the return to schooling that has been an important feature of the recent literature. The first is the suggestion that the instrumental variables estimation method (IV) only picks up the effect of some treatment if the effect is the same for everyone (James J. Heckman, 1995). The second is that IV picks up the effect of the treatment of the individuals whose choices are affected by the instrument in question (David Card, 1994).

The existing literature now features many examples which correct for least squares bias using instruments based on essentially a single variable. For example Kristin F. Butcher and Anne Case (1995) use sibling composition and obtain an estimated return of 18%. Card (1993) uses the proximity of the individual to college, and Orley Ashenfelter and Alan B. Krueger (1994) based on cross reporting by one twin of the others schooling and obtain estimates close to 10%. A recent study on the education system in the Netherlands by Adriaan Kalwij (1996) estimated the return to schooling by IV procedures at 15%, over twice the corresponding OLS estimate. However Card (1994) and Kevin Lang (1993) suggest that the instruments used have typically been relevant for the educational decisions of individuals with low levels of education by relying on 'interventions' that have affected children who may have had higher than typical discount rates. For example Colm Harmon and Ian Walker (1995) instrumented education by years of compulsory education and hence rely on interventions that **only** affected those who wished to leave school below the prevailing minimum age. That study found that the rate of return was 17%, precise, and approximately double the least squares estimate, confirming several US studies. However in subsequent work Angrist and Krueger (1992) using

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the Vietnam draft lottery, which you might expect to have most impact upon the top of the schooling distribution, also obtained a return of 10%.

Moreover the specification of the earnings function in the literature has typically assumed log wages are linear in years of schooling, which suggests that each additional year of schooling yields the same return. Recent contributions, such as that of Jin Heum Park (1994), have been critical of the assumption of linearity with respect to schooling, suggesting a non-linear specification may be more realistic, because of the likely impact of credentials and the arguments of the signalling literature (Andrew Weiss, 1995).

The main contributions of this paper are that: the possibility that different instruments may affect different margins is explored; and, because of its potential importance, the linearity in schooling assumption is tested. The paper finds: a large and significant downward bias in the least squares estimate of the returns to schooling; the returns to schooling are significantly different from being linear; and that the estimates are stable with respect to the choice of instruments. In particular, instruments which might be expected to affect youths likely to be constrained by the minimum school leaving age seem to imply the same returns as instruments likely to be more important in determining participation in higher levels of schooling. These results are significant because they suggest that recent findings in the literature are more robust than might have been imagined. Thus, while we can have greater confidence in existing estimates in the literature which use a single instrument, our results, being based upon such a large sample, are so precise that we cannot condone the standard practice of imposing linearity on the schooling/earnings relationship. Moreover, the results suggest that the least squares bias arises because of measurement error rather than discount rate heterogeneity suggesting that our findings could be used to analyse specific policy questions that affect any margin.

The plan of the paper is as follows. Section 2 motivates the choice of explanatory variables and instruments by considering the arguments regarding the specification of models of schooling returns and the use of instrumental variable techniques. Section 3 describes the data used and in particular describes time series changes in the educational opportunities open to individuals in the data. Estimates are presented in Section 4. Section 5 concludes.

2. Motivation

In a recent survey of the area, Card (1994) reviewed the findings of a number of studies into the schooling/earnings relationship which purport to use exogenous sources of variation in the education distribution to form an IV estimate of the return to schooling. The majority of the studies examined clearly supported the finding that OLS estimates were biased downward, although the point is made that these results are relatively imprecise. In another recent contribution Harmon and Walker (1995) support the finding of downward bias and with considerably improved precision.

The conventional approach estimates the following two-equation system describing log earnings (y_i) and years of schooling (S_i).

$$(1) \quad y_i = \mathbf{X}_i' \delta + \beta S_i + u_i$$

$$(2) \quad S_i = \mathbf{Z}_i' \alpha + v_i$$

where $\mathbf{X}$ and $\mathbf{Z}$ are vectors of observed attributes, $E(\mathbf{X}_i, u_i) = E(\mathbf{Z}_i, v_i) = 0$, and β is interpreted as the rate of return to schooling. As is well documented, u_i and v_i must be uncorrelated for OLS to yield an unbiased estimate of β . The potential for such bias to occur has led to numerous investigations, often based on imaginative uses of variables as instruments.

One criticism levelled at the recent literature is that the instruments used have related to interventions or 'experiments' which affect just one sub-sample of the population. For example, if compulsory schooling laws, as in Angrist and Krueger (1991) and Harmon and Walker (1995), affect just those who would have left school earlier then it is not possible to generalize from the resulting estimates. Card (1994) argues that when an intervention which only affects one subgroup is used, the resulting IV estimate of β will equal the marginal rate of return in that subgroup. That is, individuals invest in schooling until the marginal return to schooling equates to their specific marginal discount rate. If the variation in ability is relatively small then low participation in schooling could be attributed to high discount rates and those with low schooling will tend to have a higher than average marginal returns. In Card's framework, OLS yields the average return for the population as a whole while the IV procedure, based on interventions that impact on the individuals with low schooling and high discount rates, will generate a higher marginal return than does OLS.

A second, related, issue is the assumption of linearity with respect to years of schooling. This has been an important assumption in the literature, but one which has not been examined extensively in recent research (see Card and Krueger, 1992). The assumption of linearity in schooling is that each additional year of schooling should bring an equal percentage increase in earnings. This may not be realistic for a number of reasons, such as signaling arguments and the so-called 'sheepskin' effect of Thomas Hungerford and Gary Solon (1987) which suggests that the return to years where credentials are earned could attract a premium. Recent research by Heckman *et al.* (1995) has been critical of this approach in the context of the school quality/earnings issue. However, relatively little attention has been given to relaxing the linearity assumption in the context of the schooling returns literature¹.

In response to these recent developments we adopt a similar approach in Section 4 to the studies cited by Card (1994) and we contrast the effects of using instruments that should affect the top end of the educational distribution with instruments that may be expected to have their impact at the bottom of the distribution. In addition we allow for non-linearity in the schooling/earnings relationship by using dummy variables to capture the effects of successive increments in schooling, where we control for the endogeneity of these dummies with an ordered probit model of schooling to generate selectivity-corrected estimates of the earnings equation.

3. Data and Instruments

The UK is a particularly appropriate laboratory for research on schooling since there have been important changes in the education system over a relatively short period of time. There has been a considerable increase in the participation rate in schooling in the UK, as in most developed countries over the course of this century, a point well documented by economic historians in the context of models of growth and convergence (Kevin O'Rourke and Jeffrey Williamson, 1995) and in one recent paper specifically on the UK (Patricia G Rice and Duncan McVicar, 1996). However despite this general increase in participation a number of specific interventions seem likely to have directly influenced the distribution of schooling, particularly elective post-compulsory and, in particular, 'higher' education. Specifically in the

context of England and Wales, recommendations from the Education Acts in 1902 and 1944, in conjunction with the recommendations of the Robbins' Report (Robbins Committee, 1963) have led to increases in the minimum school-leaving age from 14 to 15 in 1947 and from 16 to 17 in 1973 (Albert H. Halsey *et al.*, 1980)². However, of particular interest in this study, given the arguments presented in Section 2, are policy changes specifically designed to influence participation at higher education levels. For the period 1948-1987 we have collected data from UK Department for Employment and Education and other statistical sources³ on a number of changes to the higher education system and the economic environment of youths to provide a wider range of potential instruments. Figure 1 plots these variables.

Figure 1 Time Series of Instrumental Variables

² These increases (discussed in John Micklewright *et al.*, 1989) were influential in raising not only the level of education of those that would have left earlier but also the participation rate for post-compulsory education. Moreover, the introduction of school meals and school milk schemes would have influenced the decision to participate for children from poorer backgrounds. The gradual development of the "comprehensive" school system, aimed at leveling the inequality in education opportunities, and the introduction of new opportunities through unifying the examination structure, led to a marked increase in the average level of attainment (Rice and McVicar, 1996).

³ For full details and sources see Appendix.

¹ Park (1994) is an exception, exploring a 'dip' in the schooling/earnings profile at 15 years of schooling.

The availability of places in "higher" education and a shift towards greater participation in post-compulsory schooling raised the distribution of schooling at the highest level. The development and expansion of the university sector as a result of the Robbins' Report led to a marked increase in the supply of university places, as well as the development of the "polytechnic" sector (which also offered degree courses, but largely in vocational subjects) and a "further" education sector providing sub-degree level courses. Figure 1 charts the participation in the university sector by plotting the number of entrants to university as a percentage of the 18-20 year old population for the period 1948-1987⁴. We also plot the numbers in university as a percentage of the same population grouping. As can be clearly seen the impact of the Robbins reforms are felt in both the large increase in the flow of new entrants as a proportion of the population in the mid-60's onwards, and in the stock measure of the student population to total population ratio which almost doubles in the space of a decade. Entry into higher education in the UK has been rationed by admissions criteria based on performance at examinations taken around the age of 18. The higher the ratio of university places to the population of the appropriate age (here measured as those aged 18 to 20) the lower the admission criterion needs to be to fill the available capacity.

A second element clearly important to the decision to participate at higher levels of education is the level of finance and its ease of availability. The recommendations of the Robbin's committee also led to the establishment of mandatory maintenance awards (known as student grants) for university education and these awards have since been extended to other, non-degree courses of study. In Figure 1 we show the variability in the real grant measured in real terms (1974 prices) since its introduction in 1962⁵. Overall the grant has declined in real terms over the 1962-87 period and, certainly since the high point in 1978, the decline has become more steady over the eighties⁶. Figure 1 also shows clearly that the proportion of university entrants being awarded a grant increased substantially over the course of this data, from an average of approximately 43% of entrants holding an award at the start of the 1950's to an average of over 85% over the 1980's.

⁴ A new expansion in the numbers of university level institutions since the late eighties has also impacted on the supply of places.

⁵ This relates to data on maximum allowable grants for students resident outside London. A cost of living adjustment is made for students within London. A larger grant was payable to Oxford and Cambridge students. The grants are means-tested against parental income.

⁶ Since 1992 the grant has been frozen in nominal terms and students have been eligible for top-up loans at zero real interest rates with favorable repayment terms.

Finally, although not directly controlled by policy intervention, important developments have taken place in the youth labour market (Stephen Nickell, 1993). One measure we also consider is the change in the earnings of youths over the period. We illustrate the ratio of the earnings of 16/20 year old males to the earnings of their adult (20+) peers. This acts as a measure of the opportunity cost of schooling for the 16 year olds facing the option of further participation.

The microeconomic data used in this paper is drawn from the UK General Household Surveys (GHS). The GHS is an annual survey of approximately 8,000 households⁷. For this analysis we pool 20 consecutive surveys, 1974 to 1994, providing a sample of 68676 prime age males born, and currently resident, in England and Wales⁸. Summary statistics are provided in the Appendix. The earnings measure used is the average real hourly earnings in January 1974 prices⁹. Figure 2 reports the median and upper/lower quartile earnings for different exit points from schooling. Figure 3 shows the schooling distribution.

Figure 2 *Quartiles of Earnings by Age of Leaving School*

⁷ For further information on the GHS see David Blanchflower and Andrew Oswald (1995).

⁸ Legislation introduced in the 1944 Education Act and in the Robbins Report did not include Scotland and Northern Ireland. In addition, by excluding those born outside the England and Wales we are eliminating the possible presence of migrants educated outside the system and clearly not affected by education policy in their schooling choices.

⁹ After 1983 the GHS does not report overtime hours, so the individuals reported earnings are replaced by their usual weekly earnings where a difference is reported, and deflated by the usual hours of work per week. See Blanchflower and Oswald (1995) and David Blackaby *et al.* (1994) on this issue.

Figure 3 Schooling Distribution

Our instruments are defined as follows. A dummy variable (MSLA16) is generated for those individuals entering their 14th year after 1971 and who were subsequently affected by the raising of the minimum school leaving age from 15 to 16. The variables AWARDENT and GRANTLEV are defined as the log of the ratio of awards to entrants and the log of the real grant level in 1974 prices respectively. These are meant to capture variations in the opportunity cost of higher education. FULLPOP and ENTPOP represent the ratio of university students to 18 to 20 year olds in the population, and the ratio of entrants to university to the 18 to 20 age cohort. These variables pick up rationing in higher education, which was essentially achieved through variations in the exam performance admission criterion. Finally YOUTWAGE defines the ratio of youth earnings to adult earnings which also captures opportunity cost. We code these variables into the GHS file on the basis of the year in which the individual in our data reaches 16. For example if the person was 16 in 1960 the values for the instruments in 1960 would be applicable to that individual¹⁰.

¹⁰ As our time series runs from 1948 we exclude individuals born prior to 1933 to ensure that the persons in the dataset all reached 16 years of age after 1948. Similarly, due to the unavailability of data for some of our instruments at present we must exclude those who reached age 16 after 1987.

4. Estimates

The criticisms levelled by Heckman (1995) and Paul Bingley and Niels Westergaard-Nielsen (1994) amongst others, discussed in Section 2, suggest a non-linear specification of the schooling relationship is necessary. In addition schooling is not a continuous variable, and in any data set there will be a considerable portion of the data leaving at the minimum level or at discrete points, typically defined by school examinations. We therefore pursue an extension of the Heckman two-step approach, adopted by John Garen (1984), where S_i^* is a latent variable corresponding to S_i and we treat this as an ordered probit. (See Heckman 1979, 1990). Thus, in this specification,

$$(3) \quad y_i = \mathbf{X}_i' \delta + S_i \beta + u_i$$

$$(4) \quad S_i^* = \mathbf{Z}_i' \alpha + v_i$$

Preference over education level is captured by the index S_i^* , where $S_i = \{S_i^j\}$ and $S_i^j = 1, j = 14, 15, \dots$ if $\mu_j < S_i^* \leq \mu_{j+1}$ and zero otherwise. The μ 's are unknown parameters, estimated together with α , which indicate the threshold values for moving through the participation decision. We therefore derive a vector of variables for education based on information on the age the respondent left school. The dummy variables for each school leaving age are added to the wage equation. Those who left school at 15 are the omitted category. As is well documented (see, for example, Charles F. Manski, 1989) identification is provided by including variables in the $\mathbf{Z}_i$ that can be legitimately excluded from $\mathbf{X}_i$. In this study identification is provided from a combination of our instruments defined in Section 3. The log earnings equation is selectivity corrected by including the relevant hazard from the schooling ordered probit.

Table 1 provides a summary of the estimates for a number of different specifications, utilizing various combinations of instruments.¹¹ Controls for yearly and regional variation are included in the specification of the model, although not reported in the main tables summarizing the findings with respect to schooling.¹² Age and age squared are included rather than the more conventional experience, as measurement error in education will induce error in

¹¹ See Appendix for details of the alternative reduced form schooling equations.

¹² Full estimates can be provided on request.

the derived experience measure. Column A reports the OLS estimates and the corresponding standard errors for the earnings equation (3).¹³ Column B uses the full set of instruments, column C uses only those instruments that we regard as having their effect through raising participation in schooling at low levels of schooling (MSLA16 and YOUTWAGE), column D uses only those instruments that we expect to affect the participation decision at high levels of schooling - particularly participation in higher or further education beyond age 18 (ENTPOP, FULLPOP, GRANTLEV, AWARDENT).

While the differences compared to the OLS estimates are significant throughout, the results are remarkably consistent across the different instrumented specifications with differences in the coefficients never showing statistical significance.¹⁴ One interesting feature of the estimates in column D is that they are based on omitting the instruments YOUTWAGE and MSLA16, the variables most likely to impact upon the high discount rate individuals.¹⁵ Contrary to the expectations discussed by Card (1994) the instrumented results were higher than the OLS equivalents. The flexibility of the non-linear approach appeared to be supported by the estimates and the use of different instrument sets did not seem to be affected by this. The strong conclusion of endogeneity was supported by the Mill's ratio (λ) which was significant and negative throughout. The validity of the instruments was verified by the tests proposed by John Bound et al (1995). The F-statistic on the excluded instruments in the reduced form schooling equation was computed. In addition, we calculated the partial $\bar{R}^2$, by regressing schooling against our potential instruments, having partialled out common exogenous regressors. Across the specifications in columns (B) to (D) the results compared favourably to those of Bound et al. (1995) and suggested that the instruments were legitimate.

¹³ To obtain a single estimated return consistent with the previous literature we also estimated models corresponding to those estimated in Table 1 but restricting the parameters on the dummy variables to be proportional to years of schooling. Thus the return corresponding to column A was 6%.

¹⁴ We note that while Table 1 reports the results from a number of instrument combinations we have tested this specification using all combinations of four, five and six instruments. In each case the estimated returns match closely those reported in the table. This is also true for the diagnostic statistics for instrument relevance. The estimates are also stable with regard to time variation, a feature of the estimates for the UK reported in Harmon and Walker (1995).

¹⁵ Estimated linear probability models of the school leaving age variables showed joint significance of the instruments with high F-statistics. However individual t-statistics on the instruments suggest significance of the YOUTWAGE and, more particularly, MSLA16 is concentrated on low S. Conversely the individual significance of the instruments concerned with education grants and participation rates is most apparent in the schooling regressions. This would confirm our selection of the instruments on the grounds of different effects for individuals with different discount rates for schooling.

Table 1 Estimates from the General Household Surveys 1974-1994

	(A)		(B)		(C)		(D)	
	β	se(β)	β	se(β)	β	se(β)	β	se(β)
Constant	-2.147	0.020	-2.199	0.022	-2.207	0.022	-2.214	0.022
S=16	0.155	0.004	0.230	0.014	0.243	0.015	0.254	0.015
S=17	0.215	0.005	0.372	0.030	0.398	0.032	0.420	0.031
S=18	0.255	0.006	0.495	0.046	0.535	0.048	0.567	0.047
S=19	0.255	0.010	0.576	0.062	0.629	0.064	0.672	0.063
S=20	0.281	0.011	0.683	0.078	0.749	0.080	0.803	0.079
S=21+	0.395	0.004	0.877	0.092	0.956	0.095	1.020	0.093
AGE x 10	0.998	0.011	1.003	0.011	1.004	0.011	1.004	0.011
AGE ² x 100	-0.114	0.002	-0.112	0.002	-0.114	0.002	-0.114	0.002
λ	--	--	-0.049	0.009	-0.057	0.009	-0.064	0.009
$\bar{R}^2$	0.339		0.338		0.339		0.339	
Instruments Used			MSLA16 YOUTWAGE AWARDENT GRANTLEV ENTPOP FULLPOP		MSLA16 YOUTWAGE		ENTPOP FULLPOP AWARDENT GRANTLEV	
F-test on excluded instruments	N/A		65.62		116.02		91.17	
Partial $\bar{R}^2$ x 100	N/A		0.44		0.28		0.22	
F-test for Linear functional form.	138.3		116.8		114.9		116.7	

Note: Dependent variable is the log of the real wage using the Retail Prices Index (all items) as the deflator. N=68676. Also included in specification are dummies controlling for Year and Region. Omitted categories are Year=1974, Region=North, and educated to 15 years of age.

The generalization of the 'Bound' test proposed by John Shea (1996) is relevant when collinearity of the instrument set may pose a problem: for example, in a time series model. However Shea, in discussing Roger J. Bowden and Darrell A. Turkington (1984), proposes the computation of canonical correlations between the instrument set and the endogenous schooling variables as a statistical measure of instrument validity. Moreover this approach has been adopted in a study by Alastair Hall *et al* (1996) which shows that β_{IV} is identical to β_{OLS} .

all the canonical correlations converge to nonzero limits. They develop a likelihood ratio statistic for a null that the smallest canonical correlation is zero. This test, asymptotically distributed as χ^2 , is conclusively passed by our instrument combinations.

Figure 7 plots the estimated schooling dummy variable parameters for each specification and for the corresponding linear restricted versions. This clearly illustrates the strong downward bias in least squares and the non-linearity in the effect of schooling (Park [1994], Heckman et al. [1995]). The F-test of the linearity restriction is reported in the table and conclusively rejects the linear specification in each case.

Figure 4 Estimated Schooling Returns

5. Conclusion

This paper estimates the returns to schooling for men from a standard human capital framework, exploiting the exogenous change in the education distribution generated from a number of policy interventions. The paper also tests the stability of the estimates to differences in the choice of instruments and, in particular, to instruments which might be expected to be

valid for children who seem likely to have been constrained by the minimum school-leaving age compared to instruments that seem likely to be most important at determining participation at higher levels of schooling.

Unlike the conventional approach we adopt a non-linear specification in schooling by way of an ordered probit specification. The corrected estimates reinforce recent research which has suggested the presence of a large and negative bias in least squares estimates. As in Harmon and Walker (1995) we support the finding of returns that are approximately double the corresponding least squares estimate. The results are stable and consistently well determined throughout. Our finding that the estimated return is stable across different instrument sets suggests that individual heterogeneity in returns may not be the cause of the downward bias in OLS estimates, rather that this bias should be attributed to measurement error. This implies that our findings can be used for addressing the impact of policies which affect any margin of the population.

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Appendix

Table A1 General Household Survey 1974-1994

	All		Left School aged 15		Left School aged 16 or 17		Left School aged 18 to 20		Left School aged 21 or 22		Left School after 22	
	Mean	s.d.	Mean	s.d.	Mean	s.d.	Mean	s.d.	Mean	s.d.	Mean	s.d.
Age	34	10	38	9	30	9	30	9	33	8	36	8
Ln Wage	0.083	0.471	0.005	0.378	0.027	0.483	0.158	0.531	0.387	0.495	0.430	0.484
Schooling	11.90	2.83	10.00	0.00	11.22	0.41	13.53	0.75	16.45	0.50	20.08	1.95
Year=75	0.06	0.23	0.08	0.28	0.04	0.19	0.04	0.19	0.04	0.19	0.04	0.20
Year=76	0.06	0.23	0.08	0.27	0.04	0.20	0.04	0.20	0.04	0.19	0.04	0.20
Year=77	0.05	0.22	0.07	0.25	0.04	0.19	0.04	0.21	0.04	0.20	0.04	0.20
Year=78	0.05	0.23	0.06	0.25	0.05	0.21	0.05	0.22	0.04	0.21	0.05	0.22
Year=79	0.05	0.23	0.06	0.24	0.05	0.22	0.05	0.21	0.05	0.22	0.05	0.21
Year=80	0.06	0.23	0.06	0.24	0.05	0.23	0.05	0.22	0.05	0.21	0.05	0.23
Year=81	0.06	0.23	0.06	0.23	0.06	0.23	0.05	0.22	0.05	0.21	0.05	0.21
Year=82	0.05	0.21	0.05	0.21	0.05	0.22	0.05	0.21	0.04	0.20	0.04	0.21
Year=83	0.05	0.21	0.05	0.21	0.05	0.21	0.04	0.20	0.04	0.21	0.04	0.20
Year=84	0.04	0.21	0.04	0.19	0.05	0.21	0.05	0.21	0.05	0.21	0.05	0.21
Year=85	0.05	0.21	0.04	0.20	0.06	0.23	0.05	0.22	0.05	0.22	0.05	0.21
Year=86	0.05	0.22	0.04	0.20	0.06	0.23	0.05	0.22	0.05	0.23	0.05	0.22
Year=87	0.05	0.22	0.04	0.20	0.06	0.24	0.05	0.23	0.05	0.24	0.06	0.24
Year=88	0.04	0.19	0.03	0.17	0.04	0.20	0.04	0.20	0.04	0.20	0.04	0.19
Year=89	0.05	0.22	0.04	0.19	0.06	0.24	0.07	0.25	0.06	0.24	0.06	0.23
Year=90	0.05	0.21	0.03	0.18	0.06	0.23	0.06	0.23	0.06	0.23	0.06	0.23
Year=91	0.05	0.21	0.03	0.18	0.06	0.23	0.05	0.23	0.06	0.24	0.05	0.23
Year=92	0.05	0.21	0.03	0.18	0.05	0.23	0.06	0.23	0.06	0.24	0.07	0.25
Year=93	0.04	0.20	0.03	0.16	0.05	0.21	0.06	0.23	0.06	0.24	0.06	0.25
Year=94	0.01	0.10	0.01	0.08	0.01	0.10	0.02	0.12	0.02	0.13	0.01	0.12
S=15	0.39	0.48	--	--	--	--	--	--	--	--	--	--
S=16	0.29	0.45	--	--	--	--	--	--	--	--	--	--
S=17	0.08	0.27	--	--	--	--	--	--	--	--	--	--
S=18	0.06	0.54	--	--	--	--	--	--	--	--	--	--
S=19	0.02	0.14	--	--	--	--	--	--	--	--	--	--
S=20	0.02	0.12	--	--	--	--	--	--	--	--	--	--
S=21+	0.13	0.32	--	--	--	--	--	--	--	--	--	--
N	68676		26922		25731		6823		4777		4423	

Note: Males, aged 16-65 in year of interview. Excludes those born outside England or Wales, and individuals reporting education outside of England and Wales.

Table A2 Sources of Data for Constructing Instruments

Variable Name	Series Used	Sources
Grantlev	Real grant level (1974 Price)	Supplied by Research Division, National Union of Students, London.
Awardent	Number of New Awards Made by Local Authorities, Number of New Entrants	Annual Abstract (Her Majesty's Stationary Office, London)
Entpop	Number of New Entrants to University, 18/20 Year Old Population	Annual Abstract (Her Majesty's Stationary Office, London)
Fullpop	Numbers enrolled in University, 18/20 Year Old Population	Annual Abstract (Her Majesty's Stationary Office, London)
Youtwage	Average Earnings 16/20 year old Males, Average Earnings 20+ year old Males	1947-1979 : Department of Employment 'October Survey'. 1980 - : New Earnings Survey. Nickell (1993), Wells (1987)

Table A3 Reduced Form Schooling Equations

	B		C		D	
	γ	se (γ)	γ	se (γ)	γ	se (γ)
Constant	0.3570	0.822	-2.9238	0.557	-4.0440	0.449
MSLA16	0.1431	0.030	-0.0323	0.025	--	--
YOUTWAGE	0.4863	0.258	1.4069	0.212	--	--
ENTPOP	0.3274	0.116	--	--	0.4350	0.115
FULLPOP	0.0209	0.113	--	--	-0.1615	0.112
GRANTLEV	0.0025	0.003	--	--	-0.0005	0.002
AWARDENT	0.3440	0.044	--	--	0.0806	0.037
Log L	-99455.01		-99537.36		-99517.12	

Note: N=68676. Also included in the specification are age, year and region controls. Omitted categories are Year=1974, Region=North.