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No 783

WARWICK ECONOMIC RESEARCH PAPERS

DEPARTMENT OF ECONOMICS

THE UNIVERSITY OF WARWICK

Different returns to different degrees? Evidence from the British Cohort Study 1970*

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December 2006

Abstract

As in many other countries, government policy in the UK has the objective of raising the participation rate of young people in higher education, while also increasing the share of the costs of higher education borne by students themselves. A rationale for the latter element comes from evidence of a high private return to university undergraduate degrees. However, much of this evidence pre-dates the rapid expansion in the graduate population. In the current paper, we use evidence from a cohort of people born in 1970 to estimate hourly wage returns to a university degree. Among other results, we find (i) that compared to an earlier 1958 birth cohort the average returns to a first degree for men changed very little, while the return for women declined substantially and (ii) substantial evidence of differences in returns to a first degree according to subject area of study and class of degree awarded.

JEL: J3, J4, I2

Keywords: degree, return, subject, UK, university

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*We would like to thank Martyn Andrews, Arnaud Chevalier, Luca Flabbi, Susan Harkness, Derek Leslie, Pedro Martins, Peter Sloane, Mark Stewart, Jan van Ours, Anna Vignoles, Ian Walker and participants at the ESPE 2005 Conference (Paris) and at seminars at CentER (Tilburg University) and the University of Pavia, for helpful comments. The usual disclaimer applies.

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1 Introduction

Higher education policy in Europe is in flux, not least in the UK which has witnessed considerable and ongoing policy change over the last half-century. One aspect of the UK experience has been a steady shift in the burden of funding higher education (HE) away from the taxpayer and towards students and their families. Maintenance grant provision has been removed substantially and has been replaced by a system of repayable loans. Furthermore, since 1998, uniform university tuition fees have been paid by all full-time UK university students from within the European Union. Following recent legislation, universities are able to charge top-up fees up to a regulated maximum, differentiated by university and by course, from Autumn 2006. Contemporaneously, there has been a significant expansion in the HE participation rate since the late 1980s, associated both with a reduction in the prior academic performance required for university admission and in the unit of resource in the teaching of university undergraduates.

In this context of ongoing policy change, it is important to examine the magnitude of private returns to HE and the extent to which they have changed over time. Using data on the 1958 birth cohort from the National Child Development Study (NCDS), Blundell *et al.* (2000) report an estimated hourly wage return to a degree of around 17% for men and 37% for women, relative to a control group who obtained one or more A-levels (the highest secondary school qualification) but who did not proceed into HE. In part, estimates of sizeable private returns to university degrees have been cited as evidence in support of policies shifting the burden of costs on to students. Graduates in the cohort analysed by Blundell *et al.* (2000) would most typically have graduated *circa* 1979, at just about the time that public sector financial support to university students began to decline significantly. Also at this time, UK government policy changes sought to raise substantially the HE participation rate. Rapid expansion of student numbers since the early 1980s is likely to have exerted downward pressure on average returns to a degree. Against this, skill-biased technical change (SBTC) during the last two decades of the twentieth century is likely to have increased the demand for graduate labour. The direction of the net effect of these changes on graduate returns is ambiguous. It seems timely, therefore, to update our understanding with estimates based on more recent cohorts. We also note that, while research has concentrated on average returns to qualifications, the issue of variations according to level of performance, given qualifications, is surprisingly under-explored. In the current paper, we examine both the average returns to a degree and also variations by specific factors. In particular, we address the argument that over a period in which the graduate population has expanded, better-performing graduates might have experienced a wage premium to a 'good' degree performance (see Naylor and Smith, 2006).[1]

Section 2 of the paper provides a brief review of evidence on trends in the returns to a degree in the UK. The subsequent analysis conducted in this paper is based on data from the 1970 British Cohort Study (BCS70); members of this cohort attaining HE qualifications would typically have graduated *circa* 1991. In view of the various supply and demand-side changes occurring between the late 1970s and the early 1990s, this 1970 cohort is interesting to contrast with the 1958 cohort. It is also of particular interest to address the question of how wage returns to higher education vary by both (i) the class of degree awarded and (ii) the subject of the degree studied.

Variation in returns by class of degree has received relatively little attention in the literature. This is largely a consequence of the fact that few datasets contain adequate information on class of degree awarded. The issue is of interest, however, for two reasons. First, if there is significant variation by degree class around the average return to a degree, then the investment in HE could yield a low return to poor-performing students. Shifting the burden of university fees further towards students then risks generating a greater disincentive to HE participation than would be the case with relatively little variation around the average: a narrow focus on the average return may be inadequate for policy purposes.[2] Second, it is of general interest to examine the extent to which the labour market rewards the graduate's class of degree. Estimates of returns to education

[1] In the UK degrees are classified in descending order as first, upper second, lower second, third class, non-honour degrees, fail. First and upper second class degrees are often referred to as 'good' degrees.

[2] It is interesting to note that in 2006, following the introduction of top-up fees, there has been a 5% fall in UK-based applications.

have tended to focus on years of schooling or on levels of qualifications. Yet, as there is substantial clustering of labour market entrants on both these criteria, one would expect employers to discriminate between candidates on factors such as grades achieved: that is, on degree class awarded in the context of higher education in the UK. This itself is likely to vary with the proportion of a cohort investing in a university degree.

Variation in returns by degree subject has received more attention, as we discuss in more detail below. Since the introduction of flat-rate fees, a number of authors have argued that there is a theoretical case for differentiating fees by subject (see, for example, Greenaway and Haynes, 2003). The strength of the case for differentiating fees depends in part on the strength of evidence that the return to a degree varies by subject studied and/or by institution attended. Our data do not enable us to estimate *ceteris paribus* variations in returns by institution of study. On this issue, see Chevalier and Conlon (2003).

The rest of the paper is organised as follows. Following a brief survey in section 2 of recent evidence on returns to HE in the UK, section 3 provides a description of the dataset and the sample selection procedure used in our analysis. In section 4, we discuss the issue of the endogeneity of educational qualifications and describe a way of addressing it, the so-called *proxying and matching method*. Section 5 reports estimates of the wage return to HE qualifications and to degree class and degree subjects. Section 6 explores the possibility of heterogeneity in the wage returns to a first degree, degree class and degree subjects using propensity score matching. Finally, section 7 summarises the main findings and concludes.

2 Evidence on the returns to a degree in the UK

An important paper on the estimation of the returns to a degree in Britain is that of Blundell *et al.* (2000). This study used data from the National Child Development Study (NCDS), an ongoing survey of all individuals born in Britain in a particular week in March 1958, to estimate the impact of different levels of HE on gross hourly wages at age 33. The study compares individuals with HE qualifications with those individuals who did not go on to HE but whose secondary school qualifications (A-levels) would have permitted them admission to HE, and estimate the raw wage returns to a first degree to be 21% for men and 39% for women.[3] When the full set of controls is included in the estimation, the estimated wage returns to a first degree fall substantially in the case of men - to only 12% - and only slightly in the case of women - to 34%. Without controls for ability at age 16 or A-level score, the estimated wage returns are 17% for men and 37% for women.

There have been a number of other studies using a variety of data sources in order to estimate the private return to a university first degree in the UK. Dearden (1999), also using NCDS, reports an estimated wage return to a degree of 17% for men and of 32% for women, based on OLS, and also finds that the conventional OLS estimates are reasonable approximations of the true causal impact of higher education on wages. Harkness and Machin (1999) examine changes in wage returns to education in the UK between 1974 and 1995 using data from the General Household Survey (GHS). They report time-varying estimates of the wage premium associated with various educational qualifications. For the period 1979-81, the estimated wage premia to a first degree, relative to A-level qualifications, are 14% for men and 21% for women. By the period 1993-95, these estimated premia have risen to 20% and 26%, respectively. Harkness and Machin (1999) conclude that despite a rise in the relative supply of workers who have a degree in the UK, the fact that the return to a degree was rising in the 1980s and 1990s suggests that relative demand - for example induced by SBTC - rose faster than relative supply. Walker and Zhu (2001), using Labour Force Survey (LFS) data from 1993-2000, estimate the average return to a degree over A-level to be approximately 25% for men and 30% for women. The return to a first degree was 20% for men in 1993 and about 26% in 2000, while for women it was 33% in 1993 and about 25% in 2000. These figures suggest, therefore, an increase over time in the return to HE for men and a decrease for

[3] Heckman et al. (2003) stress that in estimating rates of return it is necessary to take account of, among other factors, the direct and indirect costs of schooling, taxes, and the length of working life. In what follows, we often use the term wage 'return' although it should be interpreted in the narrow sense of a log-wage premium.

women.

The differences in the estimates from different studies referring to the same period often stem from the specification adopted which in turn depends on the nature of the data used. Longitudinal studies, such as those based on the NCDS or BCS70, are rich in information on family background, ability-related and past educational variables, which are important to address the issue of ability bias and whose inclusion often results in a reduction in the estimated return to education (see Card, 1999, and Blundell *et al.*, 2003, among others). For the same reason, the studies using other data sources where these variables are not available (such as the LFS) estimate higher returns. Moreover, Heckman *et al.* (2003) discussing the differences between cross-sectional and cohort-based estimates of the return to education, suggest that the latter should be used when the purpose is to estimate historical returns and make comparisons over time, since cohort changes are likely to affect the cross-section estimates slowly as more and more individuals from the new cohorts enter the labour market.

A number of studies have investigated the extent to which returns to a university degree vary by subject studied. Because of problems of small cell size, most studies consider broad subject groups. Blundell *et al.* (2000) find that returns for men tend be relatively low in Biology, Chemistry, Environmental Sciences, and Geography and for women tend to be relatively high in Education, Economics, Accountancy and Law and in 'Other social sciences'. Lissenburgh and Bryson (1996) using the Youth Cohort Study estimate returns of 9% for Science relative to Arts and Social Sciences for both males and females combined. Harkness and Machin (1999) find that for men Social Sciences always give the highest wage premium with respect to A-level (25% in 1995) while Science ensures the highest premium for women (45%).[4] It should be observed that while male graduates generally do not have statistically significant wage premia from undergraduate degrees in Arts, female Arts graduates do earn significantly higher wages in all years considered, especially in 1995 when the wage of Arts graduates is higher than that of Social Science graduates (with premia of 31% and 23%, respectively).

[4] Including controls for age, age squared, dummies for degree subject, teacher status, region and industry.

Walker and Zhu (2001) use a quite disaggregated definition of subjects (13 in total), but based on their disaggregated estimates, for males (females) in 1999 the returns with respect to A-level are 19% (42%), 24% (46%) and 4% (21%) for Science, Social Science and Arts and Humanities, respectively.[5] Therefore, both males and females appear to obtain higher returns for Social Science degrees. Moreover, women have higher returns than men in all degree subjects, and in particular in Arts and Humanities. Neither Harkness and Machin (1999) nor Walker and Zhu (2001) control for family background variables, and this may have inflated their estimates of the return to undergraduate degrees.

Using follow-up surveys of samples of graduates, Dolton *et al.* (1990) analyse earnings data from the 1986 survey of one in six of the 1980 UK university graduates (5,002 graduates). Dolton *et al.* (1990) find significant earnings premia for Science and Social Science students compared to Humanities or Education students. A positive earnings premium for Mathematics-related degree courses is a common finding in studies using the graduate sample follow-ups: see Chevalier *et al.* (2002), Belfield *et al.* (1997), and Battu *et al.* (1999) for results pertaining to the 1996 follow-ups of the 1985 and 1990 graduate cohorts. Chevalier *et al.* (2002) analyse 1998 earnings data for a sample of 8,264 graduates from the 1995 graduate cohort. They report that relative returns are highest for Mathematics (at 29% for men and 19% for women), compared to Education studies. They make the important point that differences in relative returns across cohorts are to be interpreted with care given differences across cohorts in the method of classifying degree subjects. Chevalier *et al.* (2002) provide a comprehensive survey of estimates of returns to HE.

With respect to differences in returns to a degree according to the degree class obtained, Battu *et al.* (1999), using graduate cohort data, estimate a significant log-pay premium associated with a first class over lower and upper second class degrees for

[5] Science includes Health, Nursing, Science, Maths, Engineering, Arthitecture. Social Science includes Economics, Law and Social Studies. Arts and Humanities includes Language, Education, Art and Combined degrees. Their specification includes controls for age, age squared and dummies for marital status, race, union status and region.

graduate earnings 6 years after graduation. Naylor *et al.* (2003) match administrative data on the entire population of UK university students - as collected formerly by Universities' Statistical Record (USR) and now by the Higher Education Statistics Agency (HESA) - to the information contained in the responses to the first destination survey of all graduates for all 1993 graduates, and estimate an occupational earnings premium of 4% for a first class degree relative to an upper second class degree for both men and women. The premium for a first over a third class degree is estimated to be 14% for men and 9% for women; there is also strong evidence that the premium for a first class degree has been growing over time. One hypothesis for this is that as the population of graduates has grown, greater importance is attached by employers to the signal emitted by a graduate who has performed well at university. For a more formal treatment of this hypothesis, see Naylor and Smith (2006). One focus of the current paper is to test for corroborating evidence on the extent of any degree class premium from a different data source. Using BCS70, our attention focuses on a cohort of young people who, typically, would have been graduating in the very early 1990s - the period of time for which Naylor *et al.* (2003) estimate significant relative premia for a good degree performance.

3 Data and sample selection

In this paper we use data drawn from BCS70, a dataset based on the cohort of 16,135 babies born in England, Wales, Scotland and Northern Ireland between the 5^{th} and the 11^{th} of April 1970. There are currently five complete follow-up surveys available: at periods 5, 10, 16, 26 and 30 years after the original survey. We use data collected in the 30-year follow-up survey on gross hourly wages and highest educational qualification achieved, while family background and individual characteristics come from the 10-year follow-up survey. Based on the sample of respondents to the 30-year follow-up survey (11,261 individuals), and in analogy with Blundell *et al.* (2000, p. F84), we select only individuals who have obtained at least A-level qualifications,[6] which is our population of interest, and analyse the wage return to HE qualifications with respect to those

[6] Or an equivalent level of education such as a Scottish higher or sixth form college.

individuals who did not complete any form of HE. In order to check the sensitivity of the estimated wage returns to the selection of the comparison group, we report both estimates using individuals with at least one A-level and, separately, those with at least two A-levels as the comparison group.

From our sample of individuals who have at least one A-level or equivalent (a total of 4,296 cases), we focus on those who also replied to the 10-year follow-up survey (3,978 individuals). This is done since in the estimation of the log-wage regressions we include individual and family background variables which are provided by the 10-year follow-up. We exclude those who did not report data on wages (942) and obtain a sample with 3,036 individuals working as full-time or part-time employees. In order to maintain the sample size, individuals with missing values in the covariates are kept in the dataset and missing value dummy variables included in the regressions.

From Table 1 we see that the mean hourly wage of male students with an undergraduate degree is £12.65; this is 21% higher than the average for those with just one or more A-levels. For females, the mean wage rate is £10.81 for those with an undergraduate degree; 31% higher than for those with just one or more A-levels. This suggests that, on average, gender wage differences are less pronounced at the higher education level. There is little if any difference in the average wage rate between those with one or more and those with two or more A-levels. Of those with an undergraduate degree, the raw data indicate a premium associated with having obtained a good, rather than a lower class of degree (that is, lower second, third or below): for males, the premium is 14% while for females it is just 4%. There are also some differences according to degree subject area; for males the premium for a Social Science degree over an Arts and Humanities degree is 11%; for females it is 21%. The wage differences between Science and Arts and Humanities are quite modest.

The BCS70 follow-up surveys were affected by panel attrition. From an original sample size of 16,135 individuals, the sample reduced to 14,875 individuals in the 10-year follow-up, to 11,622 individuals in the 16-year follow-up and to 11,261 individuals in the 30-year follow-up.[7] The rate of non-response was particularly high in the 16-year

[7] The first three figures are taken from Office for National Statistics (1999, p.11) while the fourth

follow-up. Together with a higher incidence of item non-response in the 16-year wave, this is the main reason for our use of family background variables at age 10, along with the availability of an indicator of 'innate' or 'early' ability (the British Ability Scale score, see Elliot *et al.*, 1979) at this age. As to the representativness of the different waves, the Office for National Statistics (1999, p. 11) states: 'Analysis of differential response comparing achieved samples and target samples for any follow-up, using data gathered during the birth and earlier follow-ups, shows that the achieved samples are broadly representative of the target sample'.

4 OLS, endogeneity and the proxying and matching method

When we estimate the wage returns to different educational qualifications, we consider the effect of a multiple treatment, namely educational qualifications, denoted as $j = 1, ..J$, on individual wages, w_i . We consider four different educational qualification: A-level only ($j = 1$, the reference group), non-degree Higher Education ($j = 2$), undergraduate (UG) degrees ($j = 3$) and postgraduate (PG) degrees ($j = 4$). If we indicate with w_i the gross hourly wage of individual i, our model can then be written as follows:

$$\ln w_i = mX_i + \sum_{j=2}^{J} b_j Q_{ij} + u_i. \tag{1}$$

where mX_i is a linear function of the observed variables X_i, which we will refer to as the no-treatment outcome, Q_{ij} are dichotomous variables assuming value 1 if individual i has as her/his highest educational qualification a qualification of level j and 0 otherwise, and the b_j's are the effects of these educational qualifications on log-wages; i.e., they are our parameters of interest. We abstract for the moment from problems concerning the correct specification of the no-treatment outcome and assume that a linear function is an appropriate representation of the log-wage data generating process, as this is the usual assumption in most of the existing empirical literature on the returns to education. In the case $E(u_i|X_i, Q_{ij}) = 0$, the b_j parameters can be estimated without bias using ordinary least squares (OLS, hereafter). Assuming no heterogeneity in the returns to

refers to the number of observations in the microdata file released by the UK Data Archive.

education, the Average Treatment on the Treated (ATT), the Average Treatment on the Non-Treated (ATNT), and the Average Treatment Effect (ATE) all coincide and are recovered by the b_j's.

However, there are several reasons why we may expect a non-zero correlation between educational qualifications and the error term in the log-wage equation. These include:

1. *Ability bias.* We might assume that the error term u_i in equation (1) consists of two components, i.e. $u_i = \alpha_i + \epsilon_i$, one reflecting unobserved earnings capacity (α_i), with $E(\alpha_i|X_i, Q_{ij}) \neq 0$ and the other some unobserved factors uncorrelated with all covariates included in the wage regression $E(\epsilon_i|X_i, Q_{ij}) = 0$. It is the non-zero correlation between unobserved earnings capacity (also referred to in the literature as ability) and education which causes the so-called 'ability bias'. In particular, we may expect high ability individuals both to acquire more education and to earn higher wages. Earnings capacity is potentially observed by the individual but not by the analyst;

2. *Return bias.* The returns to the different educational qualifications may not be homogeneous across individuals. Let the individual's return to qualification j be specified as $b_j + b_{ij}$, where b_{ij} is an educational qualification-specific idiosyncratic component pertaining to the individual i. In this case, we will have a distribution of b_{ij}'s.[8] There is a return bias when $E(b_{ij}|X_i, Q_{ij} = 1) \neq 0$, i.e. individuals self-select into the different educational qualifications according to their idiosyncratic returns, which depend in turn on characteristics that are observable to the individual but not to the researcher;

3. *Measurement error bias.* The educational variables may be measured with error. In our case, where education is a categorical variable, measurement error is non-classical and in general it is not possible to say anything on the direction and magnitude of the bias (see Kane, Rouse and Staiger 1999).

In our analysis in the current paper, we focus only on the first source of bias, i.e.

[8] In this specification b_{ij} is a random coefficient.

ability bias, and assume that the return bias is small or absent for the following reasons. Although heterogeneous returns according to unobserved characteristics may exist, there is a return bias only if individuals are able to predict correctly their idyosioncratic gains in the return distribution, that is they know b_{ij}, and use this information to choose their level or type of educational qualification, which is a strong assumption. In this regard, there is an interesting stream of literature on students' income expectations. Betts (1996) using US data finds that students can predict their starting salaries quite well and better than life-time earnings profiles and tend to underestimate wages in fields outside their own. He also finds that the most widely used source of information for wages are newspapers and magazines, which would suggest a substantial homogeneity in income expectations. Dominitz and Manski (1996) using US data find that students are very uncertain about their own future earnings, both at ages 30 and 40 and tend to be more uncertain about their earnings with a university degree than about earnings with only secondary school. The authors also find substantial heterogeneity in students' beliefs about the actual earnings distribution. Wolter and Zbinden (2002) use Swiss data and find that students' expectations are much closer to actual wages at the time of graduation while their prediction errors are higher when considering the pattern of wage increase during the first 10 years of their careers.

Therefore, most studies show that individuals are able to predict more accurately their starting wages, while their predictions are much less precise for earnings later on in the life-cycle, which we consider in this paper since individuals from the BCS70 with a first degree typically have in 2000 about 9 years of labour market experience. Blundell *et al.* (2005) using NCDS data find the absence of both an ability and a return bias when interactions between educational qualifications and individuals' observed characteristics are included in the log-wage equation estimated through OLS. Finally, we think that the third source of bias should be less severe when including educational qualifications, as we do, rather than the number of years of schooling, for the simple fact that recall errors on the highest educational qualification should be only minor for 30 year old individuals.

A possible approach to tackle endogeneity issues when the dataset is particularly rich, as in our case, is the so-called *proxy and matching method*. This approach is followed in Blundell *et al.* (2000) and consists of including among the individual characteristics X_i factors which might affect both the educational qualification achieved and wages, and by proxying the unobserved component α_i with observed factors highly correlated with it, so that $u_i = \epsilon_i$. Equation (1) can be viewed as a form of regression-based linear matching. It follows that the estimates which we now present in section 5 can be argued to have been obtained using a method which addresses the issue of endogeneity of education. We also discuss the results of a control function approach and, in Section 6, results based on propensity score matching.

5 Results from the proxying and matching method

5.1 Returns to HE qualifications

The application of the *proxying and matching method* requires the availability and inclusion among the X_i's of a wide set of individual characteristics affecting education and wages.

In particular, we include among the X_i's in our wage equation:

1. Personal characteristics: region of residence at age 10, ethnicity. We conduct separate analyses by gender.

2. Family background variables: father's education, mother's education, family social class (as the highest between father's and mother's social class), presence of the father, family income, number of younger siblings, number of elder siblings, parental interest in child's education; all at age 10.

3. Ability at age 10: score in the verbal and non-verbal sections of the British Ability Scales questionnaire, as proxies for verbal and quantitative innate (or early) ability.[9]

[9] The BAS score is missing for many observations (about 21%) since not all individuals in the BCS70 were administered the BAS tests, and we include a dummy variable for missing BAS score in order not to reduce the sample size.

We follow a specification similar to those used by Blundell *et al.* (2000) and Blundell *et al.* (2005) for the NCDS data. Like the latter, but differently from the former, we do not include the employer's characteristics for two main reasons. First, they may be endogenous in the sense of being choice variables for the individual and jointly determined with wages. Second, employers' characteristics may be affected by educational qualifications, and by excluding them we estimate the 'overall' effect of education, both on wages and on the likelihood of working for certain types of employers (see for instance Blundell *et al.*, 2005, and Pereira and Martins, 2004).[10]

Table 2 shows the estimates obtained using the proxying and matching method both when the comparison group is set to individuals with at least one A-level and when it is set to individuals with at least two A-levels. In the first case, the estimated coefficient on an UG degree is 0.17 for men and 0.20 for women: these convert into wage premia of 19% and 22%, respectively, using the $e^{\beta} - 1$ calculation. Hereafter we will continue to report the unconverted coefficients referring to them as log-wage returns, or 'wage returns' for short. Male workers with non-degree HE and PG degrees do not earn significantly more than those with A-levels only. It must be noted that, unlike men, women with a non-degree HE or PG degrees earn more than those with A-levels only (the wage returns being 0.08 and 0.12, respectively). The estimated returns to HE qualifications are very similar when one considers as the comparison group individuals with two or more A-levels, although they tend to decrease by between 0.01 and 0.02 points. The wage return to an UG degree is now 0.15 for men and 0.19 for women, in both cases statistically significant at the 1%.

In order to check the appropriateness of the assumption of exogeneity of education, we implement a Control Function Approach (CFA, hereafter).[11] In our specific context, this

[10] We have also estimated wage regressions including educational information (number and grades) collected in the 30-year follow-up on S (Supplementary), A (Advanced) and AS (Advanced Supplementary) levels, that is education at age 18, and O (Ordinary) levels, CSE (Certificate of Secondary Education) and GCSE (Certificate of Secondary Education), that is education at age 16. The effect was to reduce the return to HE qualifications. However, in the current version of the paper we present only the results of the regressions excluding these variables since they may subject to a considerable measurement error.

[11] See Vella and Verbeek (1999).

method consists of estimating an ordered probit for the highest educational qualification and then estimating a wage equation which includes an additional regressor, called the generalised residual (or inverse Mill's ratio), obtained from the ordered probit equation. The CFA offers a direct test for endogeneity of educational qualifications, which can also be interpreted as a specification test in the spirit of Heckman (1979). In particular, the absence of endogeneity can be tested by testing whether the coefficient on the generalised residual equals zero. Implicitly, this tests whether or not the omitted variables in the wage equation and in the education equation are correlated, and therefore whether or not the educational qualifications dummies are correlated with u_i. For the effect of the educational qualification to be identified, other than purely on functional form, it is necessary that at least one variable that enters the ordered probit model is excluded from the wage regression. We use as identifying variables parents' education, including them only in the child's education equation. All the other explanatory variables listed above are included in both the education and the wage equations. Previous research has shown that parents' educational qualifications are highly correlated with children's education (see Ermisch and Francesconi, 2001, and Chevalier and Lanot, 2002, among others). The Wald tests reported in Table 3 show that parents' educational qualifications are highly significant in the child's education equation while they are not significant in the wage regression. Despite this not being a formal test, it nonetheless provides a raw indication of the potential validity of our 'instruments' (or identifying variables) in the spirit of Bound *et al.* (1995). Table 3 reports the estimates obtained from the CFA and these are very similar to those of Table 2. The null hypothesis that the educational qualifications are exogenous in the wage regressions cannot be rejected in our data.

5.2 Differences by degree class

In the previous section, we considered an undergraduate education to be a homogeneous commodity. However, students may be more or less successful in completing their UG studies. In particular, previous work has shown the positive effect of a 'good' degree performance on graduates' earnings, see Battu *et al.* (1999) and Naylor *et al.* (2003). However, neither of these papers is able to address the issue of returns to degrees relative to non-graduate outcomes as they are based on graduate data only, with no control group of non-graduates.

BCS70 provides degree class for UG degrees, and so we are able to investigate differences in the wage return to an UG degree according to the class of degree awarded. In order to avoid small cell size problems, we consider only two broad degree classes: 'good' degree and 'lower' degree classes. This distinction is also suggested by the common practice of some employers of conditioning job offers on the attainment of a 'good' degree result.

The estimation results are shown in Table 4 and are based on excluding parents' education from the covariates. The average wage return to an undergraduate degree is 0.16 (0.20) for men (women), when one or more A-levels is set as the default.[12] The premium for a good degree is estimated to be 0.21 (0.23) for men (women), while that for a lower degree class is 0.12 (0.15), relative to the default case. A Wald tests for the equality of returns between good and lower degree classes reject the hypothesis of equality at the 5% statistical level for both genders. The difference in the wage return between good and lower degree classes is remarkably similar by gender, at 0.09 points for males and 0.08 points for females. These results are robust to the choice of reference group.

5.3 Differences by degree subject

In this section, we consider another possible source of heterogeneity in the wage return to UG degrees: by degree subject studied. We focus on the following aggregation of subjects: Science (Medicine and Dentistry, Subjects Allied to Medicine, Biological Sciences, Agriculture, Physical Sciences, Mathematical Sciences, Computing, Engineering, Technology and Architecture), Social Science (Social Studies, Economics, Law and Politics, Business and Mass Communications) and Arts and Humanities (Classics and Literature, Modern European Languages, Other Languages, Creative Arts, Education and Other).

Table 5 shows the results based on a specification which excludes parental education,

[12] Averages returns to a UG degree reported in Table 4 differ slightly from those reported in Table 2 as we are now excluding parental education, following the analysis of Section 5.1

following the analysis of Section 5.1. Starting with the estimates using the individuals with at least one A-level as the comparison group, for men our estimated wage returns for the different subjects are not very different from those of Walker and Zhu (2001). Compared to an average wage return to a first degree of 0.16, Social Science graduate have the highest wage return (0.26), and Arts and Humanities the lowest wage return (0.10), which is not statistically different from zero at the 5% level. The wage return for Science is intermediate at 0.19. Wald tests for the equality of wage returns across all degree subjects cannot be rejected for men at the 10% statistical level. When we consider Social Science versus Arts and Humanities, the difference is statistically significant at the 5% level, while the differences betwen Arts and Humanities and Science and that between Science and Social Science are not statistically significant.

For women, we observe the same ordering of subjects as for men, although the spread of the estimates around the average wage return of 0.20 is much tighter, with Social Science having the highest wage return (0.24) and Arts and Humanities the lowest wage return (0.15). Again, only the null hypothesis of equality between the wage return to Social Science and Arts and Humanities degrees is rejected at the 5% level.

For both men and women, the estimated effects from using individuals with two or more A-levels as the comparison group are very similar to those already reported, though slightly lower.

6 The case of heterogeneous returns: propensity score matching analysis

Our previous analysis using the CFA suggests the absence of an ability bias. However, as in the case of selection exclusively on observables, OLS estimates will recover the unbiased ATT only if the no-treatment outcome has been correctly specified. This requires that the model is correctly specified in terms of the (linear) functional form chosen and that the treatment effect is homogeneous across individuals with different observed characteristics (i.e., treatment has only an intercept and not a slope effect). A semiparametric method that allows us to relax these assumptions and to highlight the problem of the so-called common support is the estimation of ATT based on propensity score matching (PSM): see Caliendo and Kopeinig (2005). In this section, we estimate the wage return to (i) HE qualifications, (ii) a good as opposed to lower degree class, and (iii) different subjects studied, using PSM.

Let us define: X_i as a vector of variables affecting both educational qualifications and wages; Q_i as the treatment variable, that equals one for the treated and zero for the non-treated (in our case it will be the dummies for having a having a first degree, or for degree class awarded or degree subject studied), and w_{1i} and w_{0i} the log-wage for individual i in the case of treatment and no-treatment, respectively. Following Rosenbaum and Rubin (1983) the propensity score is defined as:

$$p(X_i) \equiv Pr\{Q_i = 1|X_i\} = E\{Q_i|X_i\}, \tag{2}$$

i.e., the conditional probability of receiving a treatment given pre-treatment characteristics. Rosenbaum and Rubin (1983) show that if the following two hypotheses hold:

1. *Balancing hypothesis*: If $p(X_i)$ is the propensity score, then $Q_i \perp X_i|p(X_i)$;
2. *Unconfoundedness hypothesis*: Suppose that assignment to treatment is unconfounded,[13] i.e. $w_{1i}, w_{0i} \perp Q_i|X_i$. Then assignment to treatment is unconfounded given the propensity score, i.e. $w_{1i}, w_{0i} \perp Q_i|p(X_i)$;

then the ATT can be estimated as follows:

$$\begin{aligned} ATT &= E\{w_{1i} - w_{0i}|Q_i = 1\} \\ &= E\{E\{w_{1i} - w_{0i}|Q_i = 1, p(X_i)\}\} \\ &= E\{E\{w_{1i}|Q_i = 1, p(X_i)\} - E\{w_{0i}|Q_i = 0, p(X_i)\}|Q_i = 1\}. \end{aligned} \tag{3}$$

In our case, PSM and ATT are implemented using kernel matching. We prefer kernel matching to other methods since it appears to be more suitable to the characteristics of

[13] This hypothesis is also called the Conditional Independence Assumption, i.e. selection only on observables, and cannot be tested within the propensity score-ATT framework.

our samples of treated and control individuals, which are not very large. Kernel matching uses all information available (since the counterfactual is built by using all individuals in the control group) and therefore there is a higher likelihood of obtaining significant ATT estimates even with small samples compared to methods using few control individuals to build the counterfactual. When using kernel matching the choice of the bandwidth implies a trade-off between efficiency and bias. In our case the bandwidth was selected optimally using cross-validation (see Härdle 1991).

The ATT estimates, using as the control group individuals with at least one A-level, are reported in Table 6. PSM is successful in balancing the covariates in the samples of treated and control individuals, as the small pseudo R^2 in the matched samples shows.[14] The percentage of observations out of the common support is generally low, showing that lack of common support is not an issue in our samples.

The estimates obtained using PSM are generally close to those obtained with the proxying and matching method; compare the reported estimates in Table 6 with the respective estimates reported in Tables 2 and 5. The only notable difference compared to earlier results is that the wage return to Social Science degrees tends to increase when using PSM. However, the precision of the estimates is lower, compared to the OLS estimates, probably due to the smaller sample sizes and the fact that standard errors are bootstrapped to take into account the fact that propensity scores are estimated. Similar results are obtained when using individuals with two or more A-levels as the contol group.

7 Concluding remarks

In this paper we have estimated the wage return to a first degree using birth cohort data from the 1970 British Cohort Survey. We estimate that there is a log wage return to an undergraduate degree of 0.16 (0.20) for men (women) relative to a control group of individuals with one or more A-level qualifications, but without higher education.

[14] Moreover, although is not reported in Table 6 the null hypothesis of joint exclusion of all covariates from the probit model can be never rejected at conventional statistical levels.

Our estimate for men is very similar to estimates obtained previously for the 1958 birth cohort. However, our findings suggest that, in contrast, the wage return to a first degree for women has fallen substantially across the two cohorts: by between one-third and one-half. The HE wage return to women is now only a little greater than that for male graduates.

We have also analysed differences in wage returns according to both degree class and degree subjects. Our estimates show the existence of a positive wage return for a good degree class compared to a lower degree class. For both men and women, the premium for a good over a lower degree class is about 8 percentage points. Our results qualitatively confirm previous findings by Battu *et al.* (1999) and Naylor *et al.* (2003), who also found earnings premia for a 'good' degree performance. Our analysis of log-wage differences by degree subjects also confirms findings from related work. As far as the ranking of subjects is concerned, for instance, we have in decreasing order: Social Science, Science and Arts and Humanities, for both men and women. Moreover, Arts and Humanities degrees are associated with a positive return (relative to workers with A-levels) only in the case of women. Although our estimates suggest the presence of differences by degree subjects, the effects are not always precisely estimated and only the difference between Social Science and Arts and Humanities degrees appears statistically significant.

Our analysis has clear policy relevance. Students in the UK - and beyond - are faced with an increasing burden of financing their higher education. In this paper, we find that the average wage return to an undergraduate degree is substantial, making the investment decision of participating in higher education seem an attractive proposition. However, we also find that there is significant evidence of quite marked variation around this average wage return, according both to the class of degree the student is awarded and to the degree subject studied; rendering the investment decision of whether to participate in higher education potentially much more marginal.

8 References

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Table 1: Hourly wage rate by educational qualification (BCS70)

Highest Educational Qualification	Males				Females			
		% sample	Wage (£)			% sample	Wage (£)	
	Obs.	(1545 obs.)	Mean	S.D.	Obs.	(1491 obs.)	Mean	S.D.
Qualification Level								
1+ A-level	223	14.43	10.47	6.25	232	15.56	8.28	4.36
2+ A-level	175	11.33	10.57	6.30	163	10.93	8.28	3.23
Non-degree HE	560	36.25	10.37	10.06	550	36.89	9.09	10.59
UG degree	576	37.28	12.65	8.82	506	33.94	10.81	10.75
PG degree	186	12.04	11.25	5.30	203	13.62	9.46	2.95
UG degree class								
Good degree	274	17.73	13.53	10.74	277	18.58	11.01	4.89
Lower degree	298	19.29	11.87	6.56	228	15.29	10.60	15.09
UG degree subject								
Sciences	212	13.72	13.14	8.24	151	10.13	10.55	4.00
Social Sciences	103	6.67	14.38	9.99	113	7.58	12.78	21.12
Arts and Humanities	105	6.80	12.99	10.66	150	10.06	10.56	4.92

Notes: Wage refers to gross hourly wage rate at age 30. % of sample refers to the size of the sample including 1+ A-level control group.

Table 2: Estimates of the log-wage premia (wage 'returns') to HE qualifications (BCS70) - OLS, including parents' education

HE qualification	Control group					
	1+ A-level			2+ A-level		
	Coef.		s.e.	Coef.		s.e.
Men						
Non-degree HE	0.025		0.035	0.006		0.038
UG degree	0.171	***	0.038	0.152	***	0.041
PG degree	0.070		0.055	0.053		0.056
N.obs.		1,545			1,497	
R^2		0.090			0.085	
Women						
Non-degree HE	0.084	**	0.031	0.075	*	0.035
UG degree	0.200	***	0.032	0.187	***	0.036
PG degree	0.120	**	0.037	0.107	**	0.040
N.obs.		1,491			1,422	
R^2		0.114			0.114	

Notes. The dependent variable is the natural logarithm of gross hourly wages. The wage equation also includes all the variables listed in section 5. Standard errors are robust to the presence of heteroskedasticity. ***Significant at the 1% level; **significant at the 5% level; *significant at the 10% level.

Table 3: Estimates of the log-wage premia (wage 'returns') to HE qualifications (BCS70) - Control Function Approach

HE qualification	Control group					
	1+ A-level			2+ A-level		
	Coef.		s.e.	Coef.		s.e.
Men						
Non-degree HE	0.023		0.035	0.007		0.040
UG degree	0.164	***	0.036	0.146	***	0.041
PG degree	0.064		0.056	0.052		0.054
$\rho\sigma$	0.001		0.001	0.000		0.001
N.obs.		1,545			1,497	
Wald test on parents' education (p-value)						
Education equation		0.000			0.000	
Wage equation		0.147			0.187	
Women						
Non-degree HE	0.082	**	0.031	0.070	*	0.035
UG degree	0.194	***	0.033	0.178	***	0.035
PG degree	0.116	**	0.040	0.099	**	0.037
$\rho\sigma$	0.000		0.000	0.000		0.000
N.obs.		1,491			1,422	
Wald test on parents' education (p-value)						
Education equation		0.001			0.001	
Wage equation		0.900			0.886	

Notes. The dependent variable is the natural logarithm of gross hourly wages. The wage equation also includes all the variables listed in section 5. The model is identified by parents' education that is included only in the education equation. Standard errors are bootstrapped with 500 replications since the model is estimated in two stages. [a]Wald test for the exclusion of parents' education in the education equation and the wage equation.

***Significant at the 1% level; **significant at the 5% level; * significant at the 10% level.

Table 4: Estimates of the log-wage premia (wage 'returns') by degree class (BCS70) - OLS, without parents' education

HE qualification	Control group					
	1+ A-level			2+ A-level		
	Coef.		s.e.	Coef.		s.e.
Men						
Good degree class	0.213	***	0.044	0.195	***	0.047
Lower degree class	0.124	**	0.043	0.106	*	0.045
UG degree (average)	0.164	***	0.037	0.146	***	0.040
Wald test Good=Lower (p-value)	0.028			0.030		
N.obs	1,541			1,493		
R^2	0.085			0.080		
Women						
Good degree class	0.233	***	0.036	0.216	***	0.039
Lower degree class	0.154	***	0.037	0.139	***	0.040
UG degree (average)	0.195	***	0.032	0.178	***	0.035
Wald test Good=Lower (p-value)	0.046			0.029		
N.obs	1,490			1,421		
R^2	0.114			0.113		

Notes. The dependent variable is the natural logarithm of gross hourly wages. The wage equation also includes all the variables listed in section 5, except parents' education. Standard errors are robust to the presence of heteroskedasticity. ***Significant at the 1% level; **significant at the 5% level; *significant at the 10% level.

Table 5: Estimates of the log-wage premia (wage 'returns') by degree subject (BCS70)-OLS, without parents' education

HE qualification	Control group					
	1+ A-level			2+ A-level		
	Coef.		s.e.	Coef.		s.e.
Men						
Science (S)	0.187	***	0.044	0.177	***	0.046
Social Science (SS)	0.257	***	0.057	0.247	***	0.058
Arts and Humanities (AH)	0.096		0.064	0.085		0.065
UG degree (average)	0.164	***	0.037	0.146	***	0.040
Wald test S=SS (p-value)		0.253			0.239	
Wald test S = AH (p-value)		0.189			0.198	
Wald test SS = AH (p-value)		0.038			0.037	
Wald test all subjects = (p-value)		0.115			0.113	
N.obs		1,545			1,497	
R^2		0.089			0.085	
Women						
Science (S)	0.183	***	0.036	0.162	***	0.038
Social Science (SS)	0.236	***	0.049	0.214	***	0.050
Arts and Humanities (AH)	0.154	***	0.041	0.133	**	0.042
UG degree (average)	0.195	***	0.032	0.178	***	0.035
Wald test S=SS (p-value)		0.258			0.255	
Wald test S = AH (p-value)		0.172			0.174	
Wald test SS = AH (p-value)		0.036			0.035	
Wald test all subjects = (p-value)		0.110			0.109	
N.obs		1,491			1,422	
R^2		0.112			0.113	

Notes. The dependent variable is the natural logarithm of gross hourly wages. The wage equation also includes all the variables listed in section 5, except parents' education. Standard errors are robust to the presence of heteroskedasticity. ***Significant at the 1% level; **significant at the 5% level; *significant at the 10% level.

Table 6: Estimates of the log-wage premia (wage 'returns') using PSM-ATT, control group are 1+ A-level individuals. (BCS70)

Treatment	ATT[a]		s.e.[b]	No. treated	No. controls	Probit pseudo-R^2 before[c]	Probit pseudo-R^2 after[d]	% obs. out of support	% bias reduction[e]
Men									
Non-degree HE	0.051		0.048	548	223	0.09	0.01	1.53	64.96
UG degree	0.156	***	0.043	552	223	0.08	0.01	3.00	68.66
PG degree	0.070		0.066	177	223	0.10	0.08	2.20	73.64
Good degree class	0.205	***	0.059	259	223	0.10	0.01	3.02	77.35
Lower degree class	0.135	***	0.050	284	223	0.10	0.01	2.69	66.23
Science	0.200	***	0.058	205	222	0.15	0.01	1.61	82.26
Social Science	0.290	***	0.080	100	222	0.21	0.03	0.92	75.10
Arts and Humanities	0.114		0.093	98	222	0.19	0.02	2.14	54.63
Women									
Non-degree HE	0.079	*	0.045	516	232	0.07	0.01	4.35	70.71
UG degree	0.191	***	0.035	466	232	0.10	0.01	5.42	82.98
PG degree	0.099	**	0.046	195	232	0.17	0.01	1.84	63.07
Good degree class	0.217	***	0.040	250	232	0.15	0.01	5.30	77.00
Lower degree class	0.154	***	0.042	218	232	0.10	0.01	2.17	70.36
Science	0.199	***	0.045	145	229	0.15	0.01	1.58	63.59
Social Science	0.260	***	0.067	112	232	0.22	0.02	0.29	65.91
Arts and Humanities	0.147	**	0.059	143	232	0.22	0.03	1.83	63.75

Notes. The outcome variable is the natural logarithm of gross hourly wages, and the control group individuals with at least one A-level. [a] Average Treatment on the Treated (ATT) computed using PSM, namely kernel matching. Optimal bandwidth selected using cross-validation (Härdle 1991). [b] Bootstrapped standard errors, 500 replications; [c] Pseudo R^2 of the probit model used to compute the propensity scores before matching; [d] Pseudo R^2 of the probit model in the matched samples; [e] % reduction in median bias (see Rosenbaum and Rubin, 1983) ***Significant at the 1% level; **significant at the 5% level; *significant at the 10% level.