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Unpacking Unequal Pay Between Men and Women Across Cohort and Lifecycle

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Abstract

This paper analyses the pay gap between men and women in the two British birth cohort studies using the new data collected in 2000 when their subjects had reached the ages of 30 and 42 respectively. The paper also includes new analysis of improved data on the 1958 cohort at 33 in 1991, and a comparison with our earlier analyses of the 1946 cohort at 32 in 1978 and the 1958 cohort at 33 in 1991. The analysis is of hourly earnings in full-time jobs, where the impact of the Equal Pay Act might be expected to be more complete, given the lack of male comparators in the extensive but low paid part-time employment sector for women. We decompose the wage gap at each age, and the change in the components of the average wage gap over time. We also examine the distribution of estimated gender premia across our samples and relate them to the wage level. For people in their early thirties, the crude wage gap closed between 1978 and 2000, but this was mainly due to improved human capital characteristics of the women in full-time employment at that stage of their lives. Unequal treatment also fell, but not by much. When following the 1958 cohort from age 33 to age 42 in 2000, men's real wages rose more than women's. The increased gap was roughly equally due to widening differentials in characteristics and deteriorating rates of remuneration for women entering middle age. Although the 42 year-old employees included women with less exceptional qualifications, who had returned to the labour force with interrupted employment histories, women who had been relatively continuously in employment also experienced the rising gender penalty over time.

Acknowlegements

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Introduction

Wage differentials between men and women have been a concern of public policy since the Equal Pay Act of 1970, and remain under review (Kingsmill 2001, Equal Pay Task Force 2001). The conventional economic analysis of the pay gap attributes it to components representing differences in human capital characteristics and differences in the rate at which these characteristics are remunerated. Any unequal treatment may be due to direct, now illegal, discrimination, or to a number of other factors. These include features of the employment context, such as occupational segregation, monopsony, differential union coverage, preferences, risks or domestic commitments which may result in women and men of equivalent education and employment experience being differently paid (see Joshi and Paci 1998 Chapter 2). Our concern in this paper is not what these features of the employment context may be, but how the components of the gender pay gap- human capital ('explained') and unequal treatment (or 'discrimination' or 'unexplained') - have been changing over time. The pioneering work of Juhn, Murphy and Pierce (1993) has developed the tools for accounting for the evolution of differentials and their components over time. In a previous paper (Makepeace et al 1999), the present authors analysed differentials between the earnings of male and female workers in full-time employment from the British cohort studies of individuals born in 1946 and 1958 (NSHD and NCDS), compared in their early thirties. Here, we exploit data newly collected in 2000 for the second and third British cohort studies (NCDS and BCS70) at ages 42 and 30 respectively. We ask how gender earnings differentials have evolved across cohort, for individuals employed full-time in their early thirties, comparing the individuals born in 1970 and 1958, and we also ask how the full-time differential has evolved within a cohort as the 1958 cohort advances through ages 33 to 42. Given our earlier finding that Equal Pay legislation had not equalized remuneration of men's and women's human capital by 1991, we were interested to see how much further, if at all, equal pay might have progressed during the 1990s, both for the new generation of thirty-year–olds and for the cohort advancing into middle age.

These data are particularly interesting since there have been numerous labour market changes during the time when these people progressed through the education system and entered the labour market. Most notably the opportunities for women have changed remarkably. The participation rates of women in the labour market continued to rise and young women have continued to improve their basic human capital. The participation in Higher education has risen from 16% to 33% and girls now systematically out-perform boys in terms of examination results. This study aims to estimate the extent to which these structural effects have impacted on the male-female earnings differential. It will also quantify and analyse the extent that the life course transition from early thirties to early forties has led to an opening up of gender earnings differentials.

The cohort data sets provide a remarkably good data source for this kind of analysis. Both of the two cohorts have been followed since birth and have excellent information on family and social background as well as educational achievement. The common CAPI style interview of both cohorts with identical questions in 2000 provides a unique opportunity to study people in detail across age. The data offer many advantages for an economist. They contain some of the best work history data available for Britain and give a detailed record of educational performance and training, the former back to primary school assessments. The sample sizes are large facilitating robust estimation of parameters.

Our previous results analyzed the large fall in the raw earnings differential between men and women employed full-time in their early thirties from 1978 to 1991(Makepeace et al 1999). . There was also a substantial fall in the Oaxaca-Blinder measure of unequal treatment but the average female full-time women would still have earned an estimated 17% more if she had been paid the same as a man. The explained proportion of the differential fell considerably reflecting the improvement in the qualifications and work experience of women. BCS70 data for age 26 suggests that the labour market position of young women has continued to improve (Joshi and Paci 1997), but what will the newly released data for age 30 in 2000 show? There is an important policy dimension to this issue because many of the social reforms that seek to improve the status of women will have their greatest impact on new entrants. The policies to facilitate career continuity would also be expected to have an impact at preserving women's human capital and maintaining their wage level. It will also be apparent how far such developments are generating a polarisation among women, already apparent in the 33 year olds of 1991.

The present results, in a nutshell, show that by 2000 the raw wage gap for full-timers in their early thirties had continued to fall, but this was mostly because of the improvement of the human capital of those women who were in full-time employment at that age. Differential treatment of the sexes, though reduced remained substantial.

Following the 1958 cohort from age 33 to age 42 in 2000 showed a widening gender gap on average, due partly to the fact that by that time the composition of those employed had changed to include relatively more women with lower qualifications and interrupted work experience, but the estimate of unequal treatment had also increased over these ages.

The estimates of differential treatment were not neutral with respect to experience and qualifications. For the 30 year olds in 1970, men seemed to benefit more from accumulating experience. For the 42 year olds in 2000 tertiary qualifications seemed to protect women's wages partially from the opening gap.

Comparison of the distributions of individual estimates of differential treatment under different condition shows that the change between 1991 depends on weighting and position in the distribution. The gains for women aged around 30 between 1991 and 2000 are not unambiguous.

Decompositions of the differential

Background

We adopt the familiar Oaxaca-Blinder methodology incorporating where appropriate extensions to deal with changes over time (Oaxaca 1973, Blinder, 1973). We begin by briefly reviewing these procedures.

The earnings equations for cohort *h* are:

$$\ln W_h^f = X_h^f \beta_h^f + u_h^f \qquad h\text{=N33, N42, BCS} \qquad 1.$$

$$\ln W_h^m = X_h^m \beta_h^m + u_h^m \qquad h\text{= N33, N42, BCS} \qquad 2.$$

where the superscript *f* represents women and *m* men. W is real hourly earnings, *X* is a vector of regressors and β the corresponding vector of coefficients. *N33 and N42 refer respectively to the 1991 and 2000 sweeps of NCDS,*

The predicted logarithms of earnings for someone with characteristics *X** are

$$E(\ln W_h^s \mid X^*) = X^* \beta_h^s \qquad s\text{=}f,m; \qquad h\text{= N33, N42, BCS.}$$

Intra-cohort comparisons

The difference between the predicted logarithms of earnings for a man and a woman in the same cohort at a given age are:

$$\ln W_h^m - \ln W_h^f = (X_h^m - X_h^f)\beta_h^m + X_h^f(\beta_h^m \text{-} \beta_h^f) \qquad \text{(decomposition } a\text{)} \qquad 3.$$

$$\ln W_h^m - \ln W_h^f = (X_h^m - X_h^f)\beta_h^f + X_h^m(\beta_h^m \text{-} \beta_h^f) \qquad \text{(decomposition b)} \qquad 4.$$

where h= N33, N42, BCS in equations 3 and 4.

Define E_h^a and U_h^a as the explained and unexplained differentials for cohort *h* using decomposition *a*.

$$E_h^a = (X_t^m - X_h^f)\beta_h^m \qquad U_h^a = X_h^f(\beta_h^m \text{-} \beta_h^f) \qquad h\text{= N33, N42, BCS.}$$

U_h^a compares the logarithms of wages when a woman is paid according to the men's and women's earnings schedules. Since it is not easy to interpret U_h^a, other measures are used. A more intuitive concept is the 'discrimination' coefficient, D_h^a. This is

$$D_h^a = e^{U_h^a} - 1 = \frac{\exp(X_h^f \beta_h^m)}{\exp(X_h^f \beta_h^f)} - 1 \qquad h\text{= N33, N42, BCS.} \qquad 5.$$

D_h^a shows the proportional increase in predicted earnings when a woman is paid according to the men's schedule rather than the women's. Since there is some disagreement over the appropriateness of the word 'discrimination', we shall refer to this coefficient the 'index of unequal treatment'.

Extending the notation

$$E_h^b = (X_t^m - X_h^f)\beta_h^f \qquad U_h^b = X_h^m(\beta_h^m \text{-} \beta_h^f) \qquad h\text{= N33, N42, BCS.}$$

$$D_h^b = e^{U_h^b} - 1 = \frac{\exp(X_h^m \beta_h^m)}{\exp(X_h^m \beta_h^f)} - 1 \qquad h\text{= N33, N42, BCS.} \qquad 6.$$

It is traditional to evaluate these measures for an individual defined as someone with the value of each regressor set equal to the mean value in sample. Since a regression goes through its sample mean, this means that the effect of the error terms can be ignored. We shall follow this practice.

Inter-cohort comparisons

Writing decomposition (*a*) explicitly for the two different cohorts at the same approximate age.

$$\ln W_{BCS}^m - \ln W_{BCS}^f = \left(X_{BCS}^m - X_{BCS}^f\right)\beta_{BCS}^m + X_{BCS}^f\left(\beta_{BCS}^m - \beta_{BCS}^f\right) \qquad 7.$$

$$\ln W_{N33}^m \text{ - } \ln W_{N33}^f = \left(X_{N33}^m \text{ - } X_{N33}^f\right)\beta_{N33}^m + X_{N33}^f\left(\beta_{N33}^m \text{ - } \beta_{N33}^f\right) \qquad 8.$$

We can clearly apply this logic to intra-cohort comparisons by substituting NCDS 2000 for BCS.

explained differential

We can decompose the inter-cohort change in the explained differential using decomposition (a).

$$E_{BCS}^{a} - E_{N33}^{a} = \left(X_{BCS}^{m} - X_{BCS}^{f}\right)\beta_{BCS}^{m} - \left(X_{N33}^{m} - X_{N33}^{f}\right)\beta_{N33}^{m}$$

$$E_{BCS}^{a} - E_{N33}^{a} = \left(\Delta X^{m} - \Delta X^{f}\right)\beta_{\mathrm{BCS}}^{m} + \left(X_{BCS}^{m} - X_{BCS}^{f}\right)\left(\beta_{\mathrm{BCS}}^{m} - \beta_{\mathrm{N33}}^{m}\right) \qquad 9.$$

where $\Delta X^{s} = X_{BCS}^{s} - X_{N33}^{s}$ is the change in the values of the characteristics for gender *s*.

The change in the explained differential is due to changes in characteristics (the first term) and changes in coefficients (the last term). This arises because the weighting of the difference in the characteristics (β_{h}^{m}) changes over time even when we agree to weight each period's change in characteristics by the coefficients for the men's equations.

For this reason, it is customary to identify the effect of the change in the characteristics over time as the first term of this expression. (See Blau and Kahn (1994)) or Juhn, Murphy and Pierce (1993).

The second term shows the change in the coefficients for men over time. Although it does contain a "gender" characteristics effect, the effect is specific to one cohort and does not change over time. It is called the observed price effect and measures the changing returns to characteristics over time.

unexplained differential

The unexplained differential can be written as:

$$U_{BCS}^{a} - U_{N33}^{a} = X_{BCS}^{f}\left(\beta_{BCS}^{m} - \beta_{BCS}^{f}\right) - X_{N33}^{f}\left(\beta_{N33}^{m} - \beta_{N33}^{f}\right)$$

$$U_{BCS}^{a} - U_{N33}^{a} = X_{BCS}^{f}(\Delta\beta^{h} - \Delta\beta^{f}) + \left(X_{BCS}^{f} - X_{N33}^{f}\right)\left(\beta_{N33}^{m} - \beta_{N33}^{f}\right) \qquad 10.$$

where $\Delta\beta^{h} = \beta_{BCS}^{h} - \beta_{N33}^{h}$ is the change in the coefficients for gender *h* (*h=f,m*).

The unexplained differential shows the change in earnings that is due to changes in coefficients. We would like the change in the unexplained differential to also reflect changes in coefficients. However the decomposition above reflects changes in coefficients ($\Delta\beta^{h} - \Delta\beta^{f}$) and changes in characteristics ($X_{BCS}^{f} - X_{N33}^{f}$).

The first term is the 'pure' coefficients effect.

Comparisons of distributions

An unexplained differential has the form:

$$U_{h} = X\beta_{h}^{m} - X\beta_{h}^{f} \qquad h= \text{N33, N42, BCS.}$$

The two components of the difference are evaluated at the same value for *X* and can, of course, be evaluated for a set of *X* values. Jenkins (1994) argues that the comparisons should be based on the characteristics for females. We follow his example and evaluate the unexplained differentials for different sets of women. We also evaluate the differentials for NCDS and BCS using the same set of women. This corresponds to the 'pure' coefficient effect in the inter-cohort decomposition of the unexplained differential.

We also apply the same techniques to the associated measures of the 'discrimination' coefficient and the wage gap.

Data

Data source

This paper examines the changes in the earnings differentials over time using the British Birth Cohort Studies (Bynner, Ferri and Wadsworth, forthcoming, Ferri 1993). The main characteristics of the three birth cohorts used here are summarized in Table 1. The studies began with the MRC National Survey of Health and Development, which initially surveyed all births during the week of March 3-9 1946. Subsequently one third of original births were followed up into adulthood. Among these contacts was postal survey in December 1977 that was returned in early 1978 when the individuals were nearly 32. 3340 individuals were contacted and the resulting data were analysed in our previous work (Makepeace 1999, Joshi and Paci 1998). The 1958 and 1970 birth cohorts, National Child Development Study, (NCDS) and The British Cohort Study (BCS70) respectively, also followed one week's births, at variable intervals into adulthood, but without any sub-sampling, so that by the time we observe adult wages (age 33 and 42 for the 1958 cohort and age 30 there were around 12,000 individuals in each cohort still in touch with the study out of around 17,000 births.

The advantages of cohort data are well known. Individuals have experienced the same macroeconomic conditions over time so these influences are automatically controlled for. Cohort studies often give access to very high quality data. In the present context we have data on actual work experience and ability variables proxied by performance on maths and reading tests when young. These tests were specific to the cohort studies and outside of the formal examination system. We would anticipate less preparation for them and that they may be a valuable independent indicator of competence attained if not 'innate' ability.

The earnings equation is based on a simple human capital specification in which earnings depend on the highest qualification obtained, work experience and region. This specification enables us to compare our results with the earlier study so that the changes in earnings differentials can be tracked over 3 cohorts spanning 22 years. The specification is constrained by the data available in the MRC survey although we have also taken the opportunity to upgrade the NCDS 1991 data. This makes the 1991 data consistent with the 2000 data and also reflects slightly changed definitions of the qualifications variables[1]. We have also recoded the qualifications data and undertaken a massive re-working of the work history data. Finally the region in 1991 was previously incompletely coded and referred to the region of the interviewer. The present data is based on region of residence.

Although we can clearly improve the fit of the model by including other variables,[2] there is merit in considering a simple human capital specification since many of the additional variables may incorporate discriminatory practices. For example, dummies for industry and occupation and public-sector job may explain part of the differentials by occupational segregation. Region is included, as a simple dichotomy to allow crudely for different levels of prices and wages in different parts of the country, although London and the South-East also show different gender-specific patterns of commuting.

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[1] They are now ultimately defined with reference to NVQ level.

[2] We have shown this elsewhere by including job and firm characteristics in our analysis of the 1991 NCDS data (Paci, Joshi and Makepeace, 1995, Joshi and Paci 1998). Anderson et al (2001) report analyses of British wages for employees of all ages in 1998 and 1999/2000 which all include a number of workplace characteristics, as well as commuting time in the case of the 1999/2000 Labour Force Survey). The presence of these terms in their model means that the pay gap remaining unexplained cannot be treated as an estimate of unequal treatment and is not directly comparable with ours.

We use reading and mathematics scores as our measures of competence or 'ability'. Although both the NCDS and BCS cohorts undertook tests in early childhood, the form of the test was slightly different for each cohort. We have standardised the scores to make them comparable. The standardised scores are derived from the values of the reading and mathematics scores in largest sample of observations for each cohort. They are computed for pooled samples of men and women. We obtain the standardised score by subtracting the mean and dividing by the standard deviation. Observations with missing values are given a standardised score of zero[3].

The present study distinguishes full-time work from part-time work and also includes a variable for tenure with the current employer[4]. The sample is: full time employees with recorded wages. The present sample for NCDS 1991 contains 19% more men and 21% more women than the previous sample. This is due to better treatment of the qualifications, region and work history variables. The present sample sizes are 3679 for men and 1718 for women. Work experience was previously only available for 3177 men and 1446 women before the age of 26. (A large number of the observations with missing values also have missing tenure.) The qualifications data was missing for a further 67 men and 20 women in the earlier sample. Missing values for tenure account for the remaining differences in the sample. No attempt is made to control for selection bias at this stage, although in the work reported by Makepeace et al (1999) it was not found to be a significant factor. This is possibly because of the narrow age span included and the unusually rich information on early capabilities which is usually unobserved.

Descriptives

Table 2 shows the means of the variables entering the regressions as dependent and explanatory variables for each sex at the contact around age 30 and, for the NCDS, 1958 cohort at age 42 in 2000. The dependent variable is the log of hourly earnings, corrected for changes in price levels between survey dates. Real wages rose between 1991 and 2000 if one considers respondents in the 30-33 age band: but more for women (an increase of .07 points on the log) than men (+.01). The raw log pay gap declined from 0.165 to 0.082. One purpose of this paper is to see how far this catching up at age 30 is due to differential changes in the composition of those employed full-time, or to less unequal treatment. Following the earlier cohort from age 33 to age 42, when the earning of continuous workers would still be expected to be on a rising curve, men's log wages went up by 0.18 points, whereas those of women employed full-time at the two cross-sections rose by only 0.04. The raw gap in log pay at 42 increased to 0.303. The other purpose of this paper is to see how far greater employment continuity of the men accounts for their greater wage earnings growth. Of the 3679 men in the sample at age 33, 2719 were also in the sample at age 42, but the overlap in sample membership was much less for women. We have 1718 women in our sample here as full-timers at age 33, and a larger sample at 42, 2270, but only 991, just over a half of the 33 year sample and just under half of the age 42 sample appear in both samples. At age 42 there are substantial numbers of 'returners' with considerable earnings interruptions behind them, and substantial numbers of those who were in full-time employment at 33 had delayed labour forces interruption or switched to part-time employment over this period.

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[3] Our earlier study used general ability.

[4] The NSHD data divided work experience into work done before and after the age of 26 and was the sum of part-time and full-time work. Here there is no distinction by age and part-time is separated from full-time work experience.

The women employed full-time around 30 have high earning characteristics, esp. NCDS - their 'ability' scores are 0.08 to 0.10 standardised units above those of men[5] who reflect the nearly the whole cohort (89%) rather than around one third of women participating in full-time employment. There is also a lead for test scores in BCS70 full-timers for women though less dramatic. In 1991 37% of women employed full-time had either a degree or diploma compared to 34% of the men. For BCS70 30 year olds in full-time work these proportions are each 47% and 38%. In each cohort the men have more work experience than the women, but for BCS70 at 30 the gap is only about 1 year (and not much more for NCDS if part-time experience counts).

By age 42, more NCDS women were in full-time employment and their average human capital had gone down - women of lower earning power had rejoined the labour market and some of the highly qualified who postponed childbearing into their 30's would have dropped out of full-time work. Average test scores of male and female workers are no longer very different, nor are educational qualifications. There is however a big gap in employment experience (6 years on full-time experience) as a result of the recruitment of women with employment interruptions.

Thus the measured characteristics would lead one to expect the smallest wage gap for the 30 year olds in BCS70, followed by the 33 year olds in 1991, with the women in their forties, including the less qualified returners experiencing the lowest rates of pay relative to their male contemporaries. This is the pattern we observe in the new data, this is the pattern we find in the "explained" component of the pay gap, but might we not expect that the residual unexplained pay gap for persons of given human capital would then have been reduced to a minimal level consistent with the spirit of Equal Pay law?

Results

BCS 70

Table 3 shows separate regression analyses of the log hourly earnings by gender[6]. All the estimates significantly different from zero with t-values much above 2 except part-time work and math missing for men. Math missing is marginally significant for women. We retain math missing for consistency with NCDS.

The coefficients in the men's equation are significantly different from those in the women's equation but surprisingly the differences between the equations can be summarised by a simple dummy for gender[7]. These results suggest that there are no differences in the rates of return to the different characteristics. The rates of return, for instance, to a degree and to full-time work experience are the same for men and for women. Any differences between the sexes are accounted for by the intercept shift. Our estimate suggests that the pay of women is about 11% lower than that of men.

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[5] This is not an artefact due the treatment of missing values. The differences in 'ability' scores remain when the observations with missing values are omitted.

[6] We began with a specification that included missing value indicators for all variables. They are omitted from the specification in Table 3 because they are insignificant at the 5% level when they were included (including reading ability). There is a perfect correlation between the missing values for mathematics and reading ability in NCDS. Although there are differences in BCS, they affect only 30 observations out 6849. It seems reasonable to omit one missing value indicator for consistency with NCDS. We omit the one for reading.

[7] The hypothesis that the coefficients are identical for men and women except for a constant is accepted at the 5% level.

The intercept shows the log earnings of someone who has none of the attributes affecting earnings in our equation. It is sometimes argued that such an individual can only make a limited contribution in the labour market and that the different intercepts measure the different valuations of raw physical prowess. In this context, the difference in the constants may not be discriminatory. While this is true, it is hard to see why this argument should apply to individuals with degrees and high levels of work experience, yet that is what the specification implies.

We also attempted to identify where any differences between the two sets of estimates might lie by including a set of interactions between 'being female' and the regressors. Examining the interactions, there are no obviously robust differences although full-time work experience and London & South East had t-values of 1.7 and 1.9. Both estimates are negative indicating that these two attributes count against women in the labour market. If we apply stepwise regression to the specification with all the interactions we are left with an equation that has significant interactions for O-level, FT work experience and living in London. This that the rates of return on full-time work experience are lower for women than for men but that the opposite is true for O-levels. Women also fare relatively badly in London & South East.

The differential return to full-time work experience is consistent with the view that women are promoted less often (Jones and Makepeace (1996)), or at any level in an occupational hierarchy, experience less wage growth. Booth et al (2002), suggest that either of these may occur because men receive more outside offers which raise wages whether or not they actually move jobs. Another variant of this interpretation is that expectations (of employers and perhaps women themselves) that continuous employment may not be sustained into the 30s, keeps women out of the 'fast track' (or fast wage growth occupations) in their twenties. The result for London and the South-East is unexpected[8]. We might expect that women would do relatively well in a large and active labour market. However the London labour market also involves a lot of commuting. Traditional family roles may mean that men travel longer distances while women are more restricted to local labour markets. Anderson et al, 2001 provide evidence for workers of all ages (including part-timers) that men's and women's different commuting times contribute modestly to explaining pay differentials and that women's return on commuting time is less than men's.

NCDS 1991

Table 4 shows the estimates of separate earnings equations for men and women. With the exception of part-time work experience, the coefficients are statistically robust and have the expected signs. The coefficients are similar to those in Makepeace et al (1999).

Unlike the BCS 70, we reject the null hypothesis that the differences between the results for men and women can be summarised by a simple dummy for gender. When a set of interactions for all the regressors is included, the interaction dummies for female and London & South East have large t-values and the interaction dummies for diploma, degree and tenure have t values bordering on significance[9]. The shift dummy for is large and negative and the advantage that women obtain from working in London is smaller than that of men.

We applied stepwise regression techniques to locate any potential differences in parameter estimates. The resulting estimates show that the earnings of a woman are lower through an intercept shift, if she has an O-level and if she works in London but her return on tenure is

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[8] It applies almost equally to those living in Greater London and outside it

[9] The t-values are 1.9, 1.7 and 1.7 respectively.

higher. We did not examine the differences between the specifications for men and women in our earlier paper. Nonetheless the differences in the coefficients in that paper appear consistent with the effects observed here. Women face an immediate fall in earnings of about 13% due to the shift in the intercept and a further falls of 7% if they have an O-level and of 5% if they work in or near London[10]. In contrast to the wages of 30 year olds in 2000, the return to experience favours women, if anything. The return to tenure is under ½% per year. The median value for tenure is 1 year so the effect of tenure is very small for most individuals.

NCDS 2000

Table 5 displays the regression estimates for the NCDS 2000 sample. As with the 1991 sample, the coefficients are jointly significantly different from one another and the differences cannot be summarised by an intercept change. We argue below that the main differences lie in the constants[11], and the return to early school attainment, and living in London and the South East, which favour men, whose effects are offset for some women by tertiary qualifications and full time work experience.

We identified where the significant differences in the estimates might lie by estimating a regression that included interaction terms for gender on all the regressors. The results for this unrestricted specification show, (with t-values in parenthesis), that the intercept is lower for women by 0.34 (-5.48), and the estimates of the coefficients for reading ability are smaller by 0.04 (-1.87) and region 0.05 (-1.87) while estimates of the coefficients for diploma 0.07 (1.81), degree 0.10 (2.2) and year of full time work experience 0.005 (1.73) are higher for women. These six interactions are the only ones retained when stepwise regression is applied to this specification. The t-values of the estimates for reading ability (-3.0) and work experience (2.1) increase in absolute value while the remaining t-values for the included interactions remain more or less the same.

A major difference between the four sets of coefficients for NCDS is that the men have a substantial increase in the constant term (and some further returns on their childhood-rated abilities). Otherwise the earnings functions are quite similar within gender between age 33 and age 42.

Decompositions

Table 6 shows the standard Oaxaca-Blinder decomposition. The crude log pay gap is shown in the top row, starting at 0.305 for 32 year olds in 1978, falling through approx 0.166 for each sample of NCDS 33 year olds in 1991 to 0.303 for the 30 year olds in 2000. Notice that gap for 42 year olds in 2000 returns to the original level observed for 32 year olds in 1978. The two standard decompositions divide the gap into components explained by human capital regressors and those attributable to different coefficients, i.e. differential treatment, unequal treatment or 'discrimination. The first method weights parameter differences by the mean value of women's regressors and the differences in mean characteristics by men's

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[10] Further investigation showed that for the NCDS at both 33 and 42, this difference applied particularly in the Southeast outside London. Anderson et al (2001) show that although commuting times are highest for those living in central London, gender differences in commuting time are greatest for those living in the rest of the South east. The actual commuting times of the cohort members remain to be investigated.

[11] At least given the particular set of of reference categories chosen for the dummy variables, sub O-level qualifications and residence outside the South East

coefficients. Decomposition (b) reverses these weights. The index D_h^f, expresses the coefficient gap, weighted by women's characteristics as the percentage by which the average woman's wage would increase if her characteristics were remunerated on the men's rates of remuneration. Similarly D_h^m, derived for decomposition (b) describes the percentage by which the average man's wage would fall if remunerated on the women's rates.

The first thing to notice is that the two decompositions give similar results. We proceed to concentrate on the first version, yielding D_h^f, along with much of the literature (Jenkins 1994). Secondly the 'explained ' differences in all samples are smaller than the unexplained. In samples of women working full-time many variables show higher human capital than the less select samples of men. Indeed, the negative term for BCS 70 shows that the 30-year women full-timers in 2000 had on average better human characteristics (worth some 3% on their average pay) than men. The different composition of women employed at 42 increases the explained component of the increased pay gap to an advantage for men of around 12%. The unexplained component of these gaps is the largest at each point, though across all the samples in their early thirties it falls over time, as one would expect give the implementation of Equal Pay Policy. Expressed as D^f, the gender premium fell from 24% in 1978 through 15-17% in 1991 to 12% in 2000. However in 2000 there was still more unequal treatment than was apparent in the (often quoted) crude full-time pay differential. Furthermore, the 1958 cohort experienced growing levels of 'discrimination' across the 1990s, although the D^f of 21% for 42 year olds in 2000 does not quite revert to the 24% for the 1978 sample, despite the similarity in the crude ratios noted above. There is more measured difference between men and women employed at 42 than there was at age 32 twenty-two years earlier.

The comparison of pay compositions over time is not straightforward as the indices are weighted averages, and they may change because of changes in coefficients, characteristics or changing weights. The Juhn, Murphy and Pierce (1993 method isolates the 'pure effect of changing characteristics in the explained component from changed 'prices' (or weights) and for isolating the 'pure' effect of changing coefficients from effects of changing weights in the 'unexplained component. The results of applying their formula to the two inter-cohort comparisons and the intra-cohort comparison are shown in Table 7. Note that the inter-cohort changes are toward more equal pay, more equal characteristics, and more equal treatment, while for the 33 to 42 comparison within the 1958 cohort all these terms are moving the opposite direction. The effects of changing weights are relative minor, though usually offsetting the general trend. In the explained component, the underlying change in characteristics is somewhat stronger than suggested by the change in the 'explained' term from Table 6. The 'unexplained' term reasonably reflects the underling improvement of coefficients for the two inter cohort comparisons, but within NCDS the effect of deteriorating coefficients (0.09) is twice the apparent change in the unexplained component.

So far we have looked at the estimates of unequal treatment for an average person. We would like to know how the index of unequal treatment varies across individuals. It is sometimes suggested that it is associated with pay. We divided the wage data into quintiles. Figure 1 shows how the mean values of the index of unequal treatment for BCS70 and NCDS change across the quintiles. It shows a clear tendency for the value of index of unequal treatment to increase with pay for BCS and to fall for NCDS at age 33. By age 42 the indices of unequal treatment for NCDS are the highest, around 20% across all quintiles but the association with pay is not clear cut. However, higher wages are associated with less unequal treatment for 33 year olds in 1991 and more for 30 year olds in 2000[12]

The differentials on particular coefficients help to account for this. The interaction of gender and education in NCDS at 33 would appear to lower the D^f for those with high wages, but that at 30 the interaction of experience and gender does not produce a strong relationship with wages. The high gender premium on the intercept at 42 leads to a high D^f across all wage levels, with the pro-male interaction on 'ability' scores countering the pro-female interaction on qualifications.

Table 8 summarises the distribution of the indices of unequal treatment implied by the 5 sets of parameters[13] for a common set of characteristics, those of women working full-time in NCDS at 33 in the present sample[14]. Since the set of women, and therefore the set of characteristics, is being held constant, any changes are due to changes in coefficients providing an alternative to the standard decomposition in Table 7. The two versions of NCDS at age 33 give similar estimates for the mean and standard deviation but with the present estimate of the average some 2% lower at 15%. The MRC estimates from 1978 are much higher (28% at the mean) and much more dispersed. The BCS 70 estimates from 2000 are marginally lower than 1991 and less dispersed. This interpretation is not substantially altered by considering the median instead of the mean. However, this ranking is reversed at the lowest levels of the distribution. The 10th percentile for the MRC (3%) is smaller than the corresponding estimate for NCDS 1991 while the 10th percentile for BCS 70 (8½%) is higher than the corresponding estimate for NCDS 1991.

Table 9 presents summary statistics for the distributions using the sample of women in BCS 70 for the inter-cohort comparison and the NCDS 2000 sample for the intra-cohort comparison[15]. The mean of the index of unequal treatment falls from 15% in NCDS 1991 to 12% in BCS 70 supporting the interpretation given for Tables 6 and 7 that there has been a noticeable fall in the index of unequal treatment for people in their early thirties over time which is due to changing coefficients. Figure 2 compares the distributions of the differential. We can see that the distribution has shifted to the left for much of its range over time leading to a fall in the value of the index of unequal treatment at most points in the distribution. However the fall has not been uniform and the distribution has shifted to the right for its smallest values. The smallest values of D^f, say values under 6%, have become relatively less common. Residual discrimination, as measured by the average value of D^f, and its dispersion have fallen over time so informally we would say that women have improved their position relative to men. However the crossing of the empirical distribution functions suggests that an unambiguous welfare ranking is not possible. Nonetheless it would require an extra-

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[12] This has been confirmed by regressing the value of D^f on the dummies for the wage quintiles. D^f varies systematically with pay for 30/33 year olds but not for 42 year olds. Adding a dummy for public sector job shows that the value of D^f is much lower in public sector jobs.

[13] These are the MRC and NCDS 1991 estimates from the JHR and the NCDS 1991, BCS70 2000, NCDS 2000 estimates from the present paper.

[14] The summary statistics are derived from the distributions of predicted values for the women in the NCDS 1991 sample. The mean for NCDS 1991 (15.4%) differs from that in Table 6 (15.1%) because the mean of the values for D^f is not the same as the value of D^f evaluated at the means of the regressors. The figure for BCS 70 of 14½% differs from the figure for BCS 70 in Table 6 (11.9%) primarily because the sets of women on which the figures are based are, respectively, NCDS 1991 and BCS 70.

[15] This is essentially equivalent to changing the base group for the comparison as can be done with all these decompositions.

ordinarily large weight to be placed on the lower end of the distribution to make the NCDS distribution better than the BCS one.

The Oaxaca-Blinder estimates in Table 6 suggest that there has been a substantial increase in the unexplained differential as the NCDS cohort aged from 33 to 42. The within-cohort decomposition in Table 7 suggests that the change in the unexplained differential due to coefficient changes may be substantially under-estimated by a simple comparison of the unexplained differentials in Table 6. This is confirmed by the intra-cohort comparisons in Tables 8 and 9.

The comparison of the two sets of parameters estimated for NCDS, evaluated using 1991 characteristics, produces a sizable upward shift in the distribution, from 15.4% to 27.6% at the mean. However this comparison freezes experience at that gained by age 33, and then confronts the age 33 characteristics with parameters estimated on a sample who had had the chance to accumulate 9 more years experience and tenure. If these terms were allowed to grow over time, this would reduce estimated gender penalty experienced at 42. The estimates in Table 9 are lower but the change in the means from 10.4% to 20.7% is still huge. This conclusion is not sensitive to the choice of average since the medians are much the same as the means. Figure 3 illustrates what has happened empirically, namely that the whole distribution has moved to the right.

Another approach to understanding the source of the change in differentials from age 33 to age 42, is to consider the subset of 991 women who were observed in full-time employment in both surveys. These women can mostly be thought of as having been employed continuously. 89% had been employed for at least 8 of the 9 years involved, and 93% for at least 7years[16]. The women observed twice in full-time jobs had higher early academic test scores than the wider cross-sectional samples of women and men, and more tertiary education. They also had more full-time experience since age 16 than other women, but less than men, and less part-time experience than other women, but more than men. Table 10 shows that they initially had higher mean wages than the age 33 cross-sectional sample and that their lead against other women in employment had widened by age 42. They experienced a growth in the mean of real wages of 27% over the 9 years[17]. The corresponding growth in real mean wages for all men in full time employment at both dates was 31%.

If we apply the parameters estimated across all 33 and 42 year olds to their characteristics at each survey, their experience of wage discrimination at age 33 is estimated at 14.9%, and at 42 as 20.0%, an increase of unequal treatment by 5 percentage points over the period. Despite their rather distinctive educational profile and experience of full-time employment, these estimated values for the mean of D^f are virtually identical to those estimated for the two cross-sectional samples in Table 6. Their experience is not uniform. When we compute the difference in the values of D^f at ages 33 and 42 for this group, slightly more than a fifth (21%) have decreases in the value of D^f while another fifth (21.2%) have increases in the value of D^f greater than 10%. (see Figure 4). Nevertheless, the greater continuity of

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[16] The corresponding figures for men who were in full-time employment at both 1991 and 2000 interviews were 96% (for 8 or more years) and 97% (for 7 or more years).

[17] The mean real hourly earnings for women in employment in 1991 and 2000 grew by 27% from £8.48 to £10.74. The mean real hourly earnings for all women in employment grew by 6% from £8.62 in 1991 to £9.10 in 2000. The mean real hourly earnings for all men in employment grew by 28% from £9.99 in 1991 to £12.81 in 2000. The mean real hourly earnings for men in employment in 1991 and 2000 grew by 31% from £10.14 to £13.32.

employment of this sub-sample did not protect them from the deteriorating relative terms on which women in this cohort were treated as they moved into middle age.

A regression of the unequal treatment terms for the 991 women who were full time employees in 1991 and in 2000[18] suggests a general upward shift and that those facing the greatest gender penalties at 33 also tended to face them at 42. It is likely that pay discrimination varies by workplace characteristics (occupational segregation private sector, unionisation, firm size) and that women will tend to be the same types of job at both dates. The general increase in the discrimination estimate might then either be due to movement into the relatively worse paying types of job, or a general tendency for men to experience more wage growth over these years than even continuously employed women staying in originally rather gender-neutral types of job.

Conclusions

This paper has investigated the differential pay of men and women currently in full-time work in Britain. It does not attempt, as in the literature on sample selection, to infer the wages facing those who are not in paid work. We focus where policy is focussed, on the observed differentials in the full-time labour market, which usually provides the 'headline' indicator of the gender pay gap.

The Cohort Studies provide comparisons of the pay at around age 30 in three cohorts: in 1978 in 1991 and in the year 2000. They also provide evidence of the change through time of the 1958 Cohort between the age of 33 in 1991 and 42 in the year 2000. The general expectation is that male and female wages are converging given the convergence of men and women's educational attainment and of their labour force experience, and also given the increasing spread of equal opportunities policies and practices.

Do these data suggest that equal pay is progressing in a uniform way across these cohorts? If we look at the crude differential between men and women aged just over 30 in 1978, and then in 1991, we see a narrowing of the pay gap from around 36% to around 18% in 1991. When we look to the year 2000 and to the workers born in 1970 who were then aged 30, the crude pay gap had diminished even further, it had halved again to just over 8%, but if we look within the 1958 cohort, to the samples who were in full-time jobs at age 33 and at age 42, the pay gap does not close, it opens up. It is back to 35% again in the year 2000 for 42 year olds, just about where it was for the 32 year olds, 22 years earlier in the 1946 birth cohort.

Can these changes in pay differentials be explained by the changing characteristics of the workers in full-time employment or do they reflect changes in the relative treatment of men and women? Our results suggest that the explained component of all of these gaps remains small and is indeed negative for the 1970 cohort at age 30. Women who are in full-time work are better qualified and have almost as much work experience as the men in full-time work on average. There is more pro-masculine difference in characteristics between male and female workers at age 42 than there is at around age 30, but nevertheless in all cases the unexplained component, the component of the wage gap which is not explained by human

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[18] $D^f_{2000} = 0.117 + 0.555 D^f_{1991}$ t =(21.6) R^2=0.320

capital variables, is larger than that which is. This unequal treatment component falls over time with respect to workers around age 30, but it increases over time if one considers workers in the same cohort as time goes by and the cohort gets older.

The estimates of unequal treatment vary by the characteristics of the workers. In the 1970 cohort to age 30 it appears that men gain more from accumulating work experience than women. In all of the 1991 and 2000 observations they gain more from living in the South-east or London. However, for the 30 year olds in 1970 there is not much difference on the most other characteristics (apart from employment experience) in the treatment of men and women. For the 33 year olds in 1991 the interaction with work experience was in the other direction. Women gained more from having more job tenure and there was a slight tendency for women to benefit more from being qualified than men. By the year 2000, those 42 year olds of the 1958 cohort, who were then in full-time employment, showed differential treatment in favour of women in respect of some characteristics, but on the whole the unexplained pay gap had opened up across the board.

These structural influences on this degree of unequal treatment are also apparent when individual estimated indices of unequal treatment are plotted against wages. For the 1970 cohort aged 30, discrimination rises with wages and for the 1958 cohort at age 33, it falls. For the 42 year olds the level of wages is not strongly associated with wage discrimination. Women who were in full-time employment at both points for the 1958 cohort were hardly penalised less for their gender than were those who had not been employed at age 33 and who had re-entered the labour market full-time at age 42. This suggests that there are features of the lifecycle that are not included in our human capital model, which intensify the obstacles to high earning by women as they get older, or conversely that increase men's chances relative to women's of wage growth and promotion as the lifecycle proceeds. Men are seldom penalised in the wage market for having spouses and children (Greenhalgh, 1980, Davies and Peronaci, 1977). Women may be penalized, or at least not rewarded, and this is subject to further investigation, particularly with respect to part-time jobs, not considered here. It will also be worth investigating whether it makes a difference at what ages the work experience recorded at 42 was accumulated.

A widening of gender wage differentials over the lifecycle for continuous full-time workers was also inferred from a cross-sectional analysis of earnings from the BHPS to 1994 (Davies et al 1997, elaborated in Rake (ed) 2000), although this suggested that tertiary qualifications were a more effective protection against a growing gender penalty than appears here. The present findings confirm that the advance of equal treatment is by no means uniform across workers or across age groups. However we have not confirmed a simple picture of polarization among women whereby those with better human capital also receive better treatment. There was some sign of such a process among 33 year olds at 1991, but by 2000 it had disappeared among 42 year olds, and perhaps reversed among 30 year olds.

These findings are not the whole story about equal treatment of men and women in the British labour market, even in these two birth cohorts, because they have paid no attention whatsoever to the wages available in part-time work which are generally lower, especially when done by women. This remains to be investigated, as do the issues of whether the estimated wage equation is affected by a selection into full-time work, and of the workplace and family circumstances that may be associated with different levels of pay.

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Table 1: British cohort studies

	Year of birth	Observations of wages used here	
		Date	Age
National Survey of Health and Development	1946	1978	32
National Child Development Study	1958	1991	33
British Cohort Study 1970	1970	2000	30
National Child Development Study	1958	2000	42

Table 2: Means

	Men BCS 30	Women BCS 30	Men NCDS 33	Women NCDS 33	Men NCDS 42	Women NCDS 42
Log hourly wage, 2001 prices	2.1853	2.1031	2.1988	2.0337	2.3816	2.0782
Maths ability: z score	0.09127	.1182	.1279	.2030	.1230	0.07372
Reading ability: z score	-0.00595	.1973	.1077	.2086	0.09572	0.08445
O level	.3002	.2842	.2903	.3108	.2715	.2978
A level	.1784	.1491	.1884	.1473	.1810	.1339
Diploma	.1680	.2040	.1628	.1909	.1831	.2044
Degree	.2092	.2740	.1544	.1781	.1779	.1705
FT work experience	10.6888	9.8227	13.9212	11.8149	21.9457	16.2145
PT work experience	.1746	.4332	0.07504	.7150	.1223	2.7941
Tenure	5.2317	4.9664	6.5191	5.7915	10.8211	8.0123
London or S. East	.3058	.3201	.3047	.3254	.2876	.2749
Math missing	.2454	.2495	.1370	.1403	.1325	.1436
Sample size	4120	2730	3679	1718	3856	2270

The z-scores are derived from the values of the reading and mathematics scores in largest sample of observations for each cohort. The z-scores are computed for pooled samples of men and women. We obtain the z-score by subtracting the mean and dividing by the standard deviation. Observations with missing values are given a z-score of zero.

Note the z-scores only have a mean of 0 and a standard deviation of 1 in the sample from which they are derived.

Table 3: BCS 70 JHR Specification Men and women separately

	Men				Women		
	Coeff	S. Error	t-value		Coeff	S. Error	t-value
Constant	1.594	.039	40.852		1.560	.044	35.587
Maths ability	0.0677	.011	6.269		0.0520	.012	4.202
Reading ability	0.0240	.011	2.242		0.0419	.013	3.293
O level	0.0794	.020	3.928		0.117	.027	4.368
A level	0.216	.023	9.506		0.208	.030	6.995
Diploma	0.344	.023	14.825		0.361	.029	12.649
Degree	0.595	.027	22.157		0.573	.032	18.114
FT work experience	0.0212	.003	6.961		0.0126	.003	3.806
PT work experience	-0.0095	.008	-1.147		-0.0156	.006	-2.545
Tenure	0.0093	.002	6.040		0.0093	.002	5.100
London or S.East	.236	.014	17.238		0.199	.015	13.360
Math missing	-0.0179	.015	-1.223		.0275	.016	1.702
Sample							
R^2	0.291				0.322		

Dependent Variable: Natural logarithm of hourly wage at age 30

Table 4: NCDS 1991 Men and women separately

		Men				Women	
	Coeff	S. Error	t-value		Coeff	S. Error	t-value
Constant	1.602	0.039	40.6		1.366	0.044	30.905
Mathematics ability	0.0604	0.010	6.214		0.0589	0.014	4.135
Reading ability	0.0250	0.009	2.650		0.01748	0.015	1.169
O-level	0.147	0.018	8.153		0.112	0.028	4.070
A-level	0.215	0.020	10.560		0.249	0.034	7.436
Diploma	0.358	0.022	16.502		0.432	0.031	13.835
Degree	0.538	0.027	20.197		0.613	0.036	17.151
FT work experience	0.0194	0.003	7.451		0.0240	0.003	7.441
PT work experience	-0.0119	0.010	-1.188		-0.0060	0.006	-.965
Tenure	0.0047	.000	4.051		0.0085	.000	4.426
London or S. East	0.210	.013	16.029		0.162	.019	8.591
Math missing	0.00943	.017	0.540		0.0554	.026	2.158
Sample	3679				1718		
R^2	.306				.376		

Table 5: NCDS 2000 Men and women separately

	Men				Women		
	Coeff	S. Error	t-value		Coeff	S. Error	t-value
Constant	1.746	.049	35.600		1.405	.036	38.622
Maths ability	0.0766	.012	6.559		0.0621	.013	4.846
Reading ability	0.0472	.011	4.137		0.0122	.013	.908
O level	0.117	.023	5.070		0.141	.024	5.867
A level	0.248	.026	9.703		0.264	.030	8.892
Diploma	0.361	.026	13.824		0.432	.027	16.018
Degree	0.592	.030	19.930		0.691	.031	22.624
FT work experience	0.0118	.002	5.523		0.0168	.002	9.290
PT work experience	-0.0083	.009	-.952		0.0185	.003	.696
Tenure	0.0519	.001	5.625		0.0070	.001	5.361
London or S.East	0.217	.016	13.247		0.170	.018	9.609
Math missing	-0.0171	.022	-.783		0.0314	.022	1.396
Sample	3856				2270		
R^2	0.298				0.428		

Table 6: Oaxaca-Blinder decomposition of differences in the logarithm of wages for full time male and female workers

	MRC 1978 JHR	NCDS 1991 JHR		NCDS 1991	BCS 70 2000		NCDS 1991	NCDS 2000
Crude gap	0.305	0.167		0.165	0.082		0.165	0.303
Decomposition (a)								
Explained	0.091	0.011		0.023	-0.030		0.023	0.117
Unexplained	0.214	0.156		0.142	0.112		0.142	0.186
Decomposition (b)								
Explained	0.074	0.003		0.032	-0.038		0.032	0.120
Unexplained	0.231	0.164		0.133	0.120		0.133	0.183
D_h^f	23.9	16.9		15.2	11.9		15.2	20.5
D_h^m	26.0	17.8		14.4	12.7		14.4	20.1

Notes:

1. The crude gap is $\overline{\ln W_h^m} - \overline{\ln W_h^f}$.
2. Decomposition (a) weights the unexplained differential by the characteristics of females explained=$\left(\overline{X_h^m} - \overline{X_h^f}\right)\beta_h^m$, unexplained=$\overline{X_h^f}\left(\beta_h^m - \beta_h^f\right)$, $D_h^f = \exp\left(\overline{X_h^f}\left(\beta_h^m - \beta_h^f\right)\right)-1$
3. Decomposition (b) weights the unexplained differential by the characteristics of males. explained=$\left(\overline{X_h^m} - \overline{X_h^f}\right)\beta_h^f$, unexplained=$\overline{X_h^m}\left(\beta_h^m - \beta_h^f\right)$, $D_h^m = \exp\left(\overline{X_h^m}\left(\beta_h^m - \beta_h^f\right)\right)-1$

Table 7: Juhn, Murphy and Pierce decompositions of the changes in explained and unexplained components of the gender gap

	NCDS-NSHD 33-32		**BCS-NCDS** 33-30		**NCDS** 42-33
	Δ Explained component:		*Δ Explained component:*		*Δ Explained component:*
Total change (explained) (0.011-0.091)	-0.079[a]	Total change (explained) (-0.030-0.023)	-0.053	Total change (explained) (0.117-0.023)	0.093[a]
Changes in characteristics $\left(\overline{\Delta X^m} - \overline{\Delta X^f}\right)\beta^m_{N33}$	-0.108	Change in characteristics $\left(\overline{\Delta X^m} - \overline{\Delta X^f}\right)\beta^m_{BCS}$	-0.058	Changes in characteristics $\left(\overline{\Delta X^m} - \overline{\Delta X^f}\right)\beta^m_{N42}$	0.114
Price effect $\left(\overline{X^m_{MRC}} - \overline{X^f_{N32}}\right)\left(\beta^m_{N33} - \beta^m_{N32}\right)$	0.029	Price effect $\left(\overline{X^m_{BCS}} - \overline{X^f_{BCS}}\right)\left(\beta^m_{BCS} - \beta^m_{N33}\right)$	0.005	Price effect $\left(\overline{X^m_{N33}} - \overline{X^f_{N33}}\right)\left(\beta^m_{N42} - \beta^m_{N33}\right)$	-0.021
	Δ Unexplained component:		*Δ Unexplained component:*		*Δ Unexplained component:*
Total change (unexplained) (0.156-0.214)	-0.059	Total change (unexplained) (0.112-0.142)	-0.030	Total change (unexplained) (0.186-0.142)	0.045
Changes in coefficients $\overline{X^f_{N33}}\,(\Delta\beta^m - \Delta\beta^f)$	-0.077	Changes in coefficients $\overline{X^f_{BCS}}(\Delta\beta^h - \Delta\beta^f)$	-0.027	Changes in coefficients $\overline{X^f_{N42}}\,(\Delta\beta^m - \Delta\beta^f)$	0.090
Changes in characteristics $\left(\overline{X^f_{N33}} - \overline{X^f_{N32}}\right)\left(\beta^m_{N32} - \beta^f_{N32}\right)$	0.019	Changes in characteristics $\left(\overline{X^f_{BCS}} - \overline{X^f_{N33}}\right)\left(\beta^m_{N33} - \beta^f_{N33}\right)$	-0.003	Changes in characteristics $\left(\overline{X^f_{N42}} - \overline{X^f_{N33}}\right)\left(\beta^m_{N33} - \beta^f_{N33}\right)$	-0.045

[a] There is some rounding error in the figures quoted.

*N*32 refers to the NSHD sample, *N*33 to the NCDS 1991 sample, and *N*42 to the NCDS 2000 sample

Table 8: Distribution of the 'discrimination coefficient' (D_h^f) based on characteristics NCDS women employed full-time in 1991

Percentile	MRC JHR	NCDS JHR	NCDS 1991	BCS 2000	NCDS 1991	NCDS 2000
10	3.0	9.5	7.2	8.5	7.2	19.1
20	10.6	11.8	9.5	10.6	9.5	21.7
25	13.1	12.5	10.4	11.6	10.4	22.9
30	15.9	13.4	11.5	12.4	11.5	23.8
40	21.5	14.8	13.4	13.6	13.4	25.7
50	27.0	16.4	15.2	14.6	15.2	27.4
60	31.6	18.1	17.1	15.7	17.1	29.1
70	36.4	20.0	19.2	16.9	19.2	30.8
75	39.6	21.2	20.4	17.6	20.4	32.2
80	43.3	22.6	21.4	18.3	21.4	33.5
90	54.1	25.2	24.4	20.2	24.4	36.5
Mean	27.8	17.0	15.4	14.5	15.4	27.6
Standard deviation	20.5	6.2	6.7	4.6	6.7	6.8
coefficient of variation	0.74	0.36	0.43	0.32	0.43	0.25
inter-quartile ratio	0.98	0.53	0.65	0.40	0.65	0.34

Table 9: Summary statistics for the distribution of the 'discrimination coefficient' (D_h^f) for full time women workers in BCS 70 and in NCDS 2000.

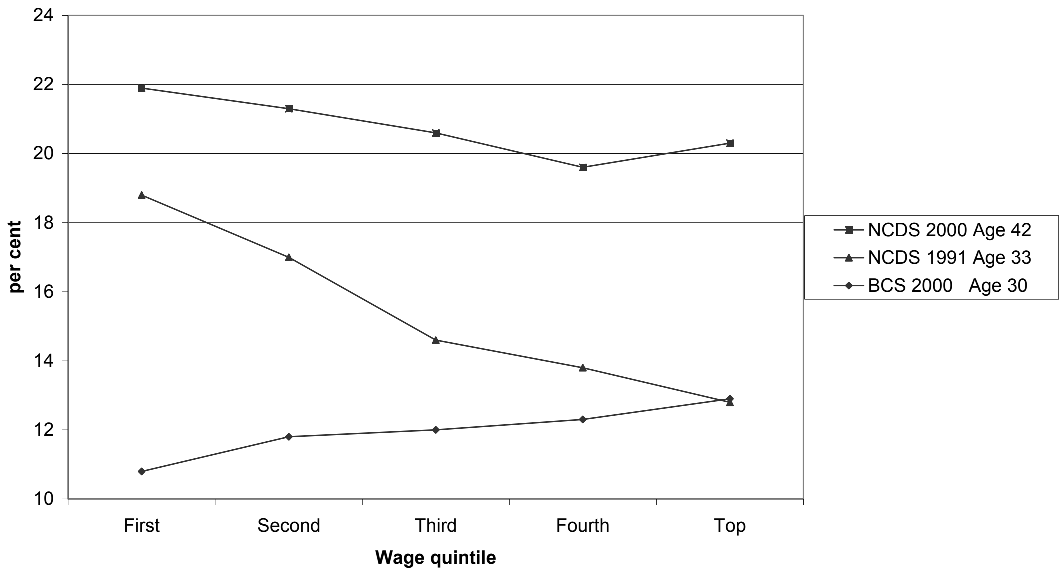
Reference group	BCS	BCS		NCDS 2000	NCDS 2000
Coefficients	BCS	NCDS 1991		NCDS 1991	NCDS 2000
Mean	12.0	15.1		10.4	20.7
Std. Deviation	4.0	6.6		7.6	7.1
Coefficient of variation	33.9	43.7		73.1	34.3
Median	11.9	14.7		10.3	20.6
Interquartile range	5.1	9.0		10.9	9.6
Interquartile ratio	42.8	61.2		105.8	46.6

Table 10

Mean of log earnings: women full-timers in NCDS

	Actual at 33	Predicted at 33
All women in 1991	2.0337	2.0337
Women in both samples	2.0600	2.0450
	Actual at 42	Predicted at 42
All women in 2000	2.0782	2.0782
Women in both samples	2.2588	2.2283

Figure 1. Mean Discrimination Coefficient, D^f, by Wage Quintile



Data for Figure 1

The mean of the discrimination coefficient, D^f, by wage quintiles

Wage quintile	BCS 2000	NCDS 1991	NCDS 2000
First	10.8	18.8	21.9
Second	11.8	17.0	21.3
Third	12.0	14.6	20.6
Fourth	12.3	13.8	19.6
Top	12.9	12.8	20.3

Figure 2: Cumulative relative frequency distributions of the discrimination coefficient using BCS women

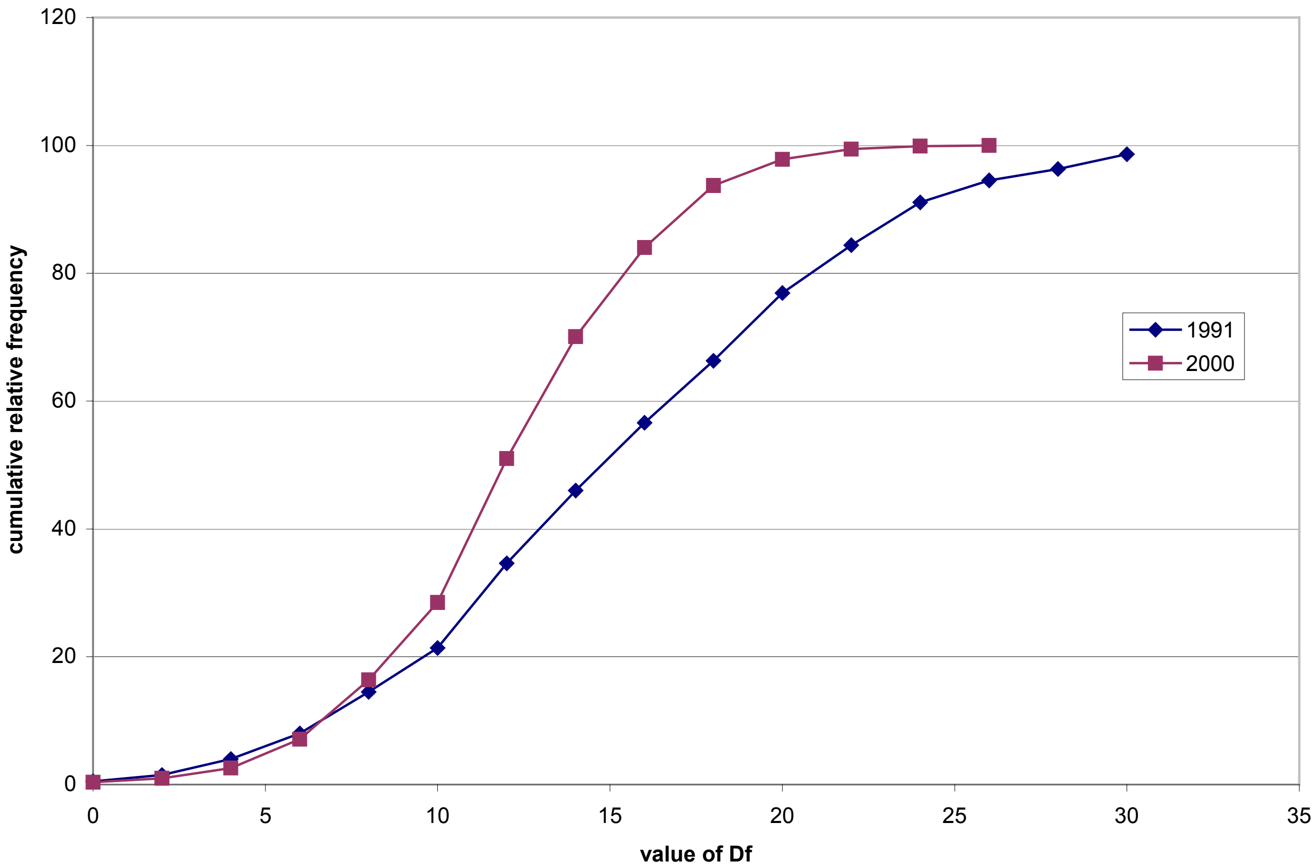


Figure 3: Cumulative relative frequency distributions for the discrimination coefficient using NCDS 2000 women

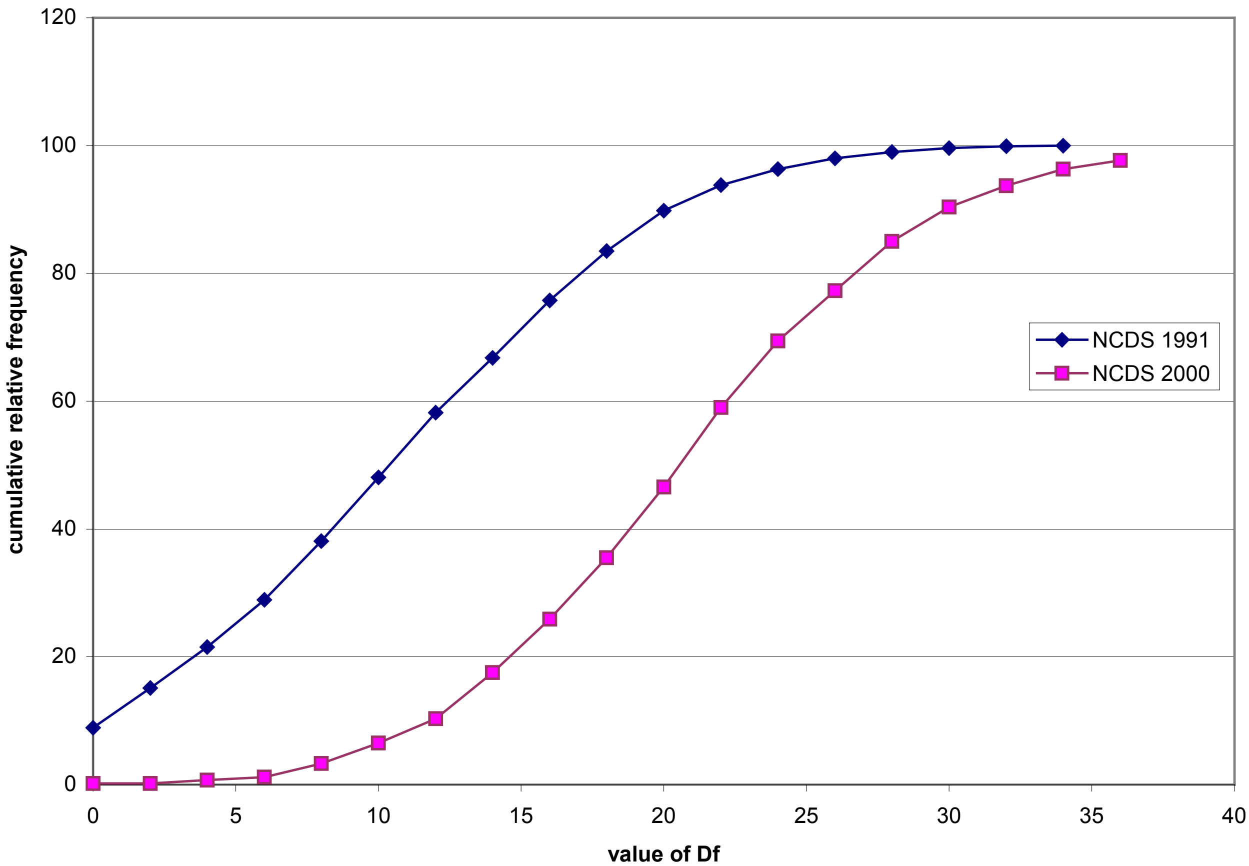
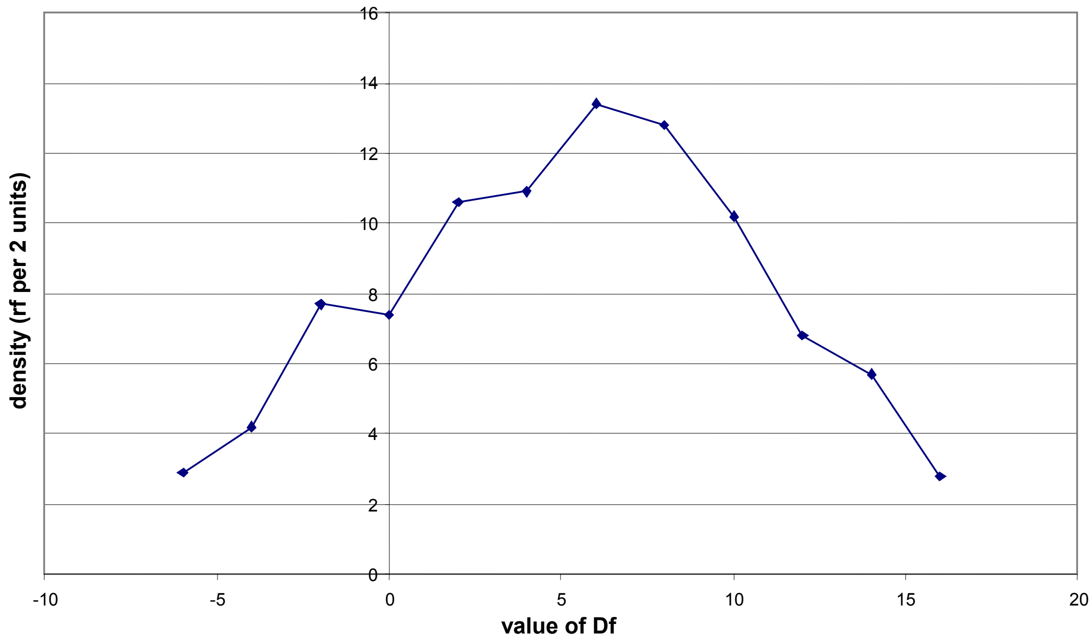


Figure 4: Relative frequency polygon for the difference in Df for NCDS samples



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