MANCHESTER
1824

The University
of Manchester

Cathie Marsh Centre for Census and Survey Research

# A quantile regression analysis of gender differences in wage mobility

CCSR Working Paper 2007-13
Nikos Tzavidis and Shirley Dex
nikos.tzavidis@manchester.ac.uk

Wage growth between ages 33 and 42 was modeled as a function of age 33 wages, gender and other labour market relevant characteristics using a quantile regression approach. The data are from a cohort of British men and women, born in 1958, who were employed full time at both ages. The age 33 starting wage and wage squared explain much of the gender differences in age 42 wages and the fact that men are more likely than women to work in high wage and male dominated occupations. Quantile regression revealed differences in predictors of wage growth across the wage distribution. .

# A quantile regression analysis of gender differences in wage mobility


**Shirley Dex**

**Centre for Longitudinal Studies**

**Institute of Education, University of London**

**and**

**Nikos Tzavidis**

**Centre for Census and Survey Research,**

**University of Manchester**



**November 2007**

Words(excluding Abstract ) = 7793

# ABSTRACT

Wage growth between ages 33 and 42 was modeled as a function of age 33 wages, gender and other labour market relevant characteristics using a quantile regression approach. The data are from a cohort of British men and women, born in 1958, who were employed full time at both ages. The age 33 starting wage and wage squared explain much of the gender differences in age 42 wages and the fact that men are more likely than women to work in high wage and male dominated occupations. Quantile regression revealed differences in predictors of wage growth across the wage distribution.

Words=100.


## ACKNOWLEDGEMENTS


We acknowledge funding for the research contained in this paper from the ESRC grant reference RES-225-25-2001, as part of the Priority Gender Network, GeNet, coordinated from Cambridge University. We also wish to thank Dr. Jane Elliott for providing us with ready made indicators of the occupational segregation of 1990 occupation groups at the 3- digit level.

## 1. Introduction

Although the gender wage gap in the UK has been declining over time, it has not gone away (Kingsmill, 2001, Women and Work Commission, 2006) and it is also persistent, to varying degrees in other European countries. The difference between rates of pay remains between women in part-time jobs and men in full-time jobs, and within the full-time labour force, which features in headline measurements of the pay gap. Equality appears closer among workers in their twenties, so it is tempting to conclude that unequal pay is being phased out, and will disappear as the younger generation grows older, the 'cohort replacement' theory. However, this conclusions may be premature. Purcell et al (2006) found that gender wage gaps between graduates are evident by seven years after graduation. It is also possible that the apparently equal pay of younger workers may yet hide unequal remuneration of women's increasingly superior qualifications. In this paper we focus on the wage growth of a selected group of more mature employees, born in 1958, who are most likely to be equal in their labour market experiences, as they moved across their thirties over the 1990s. It is important to understand how far the pay gap among mature workers in full-time employment is to be explained by women at these ages having less earning power than men, for example, if they have less continuous employment records or fewer qualifications, and how far the labour market is remunerating equivalent characteristics unequally.

Wage growth can occur in a number of ways – through changing job within the same occupation, or changing job and occupation, or by remaining in one occupation and moving up set salary scales. It is not clear a priori whether individuals who move jobs have more potential to improve their wages and experience wage growth than those who stay in the same job. Occupational change can be thought of as a route through which the accumulation of human capital is rewarded, or failure to maintain it is penalized.

However, some studies suggest there are gender differences in wage outcomes from staying within the same job, versus moving jobs (Booth et al, 2003). Particularly in the face of gendered occupational segregation, it is pertinent to enquire if these strategies and their rewards or penalties affect men and women equally. Studies of this particular period in the 1990s have found that despite being a time of increased insecurity for men, the likelihood of a substantial hourly pay rise (over 30%) was much higher for men who changed jobs (27% in 1996) than for those who stayed in the same job (12%) year on year and these rates rose between 1991 and 1996 (Nickell, et al, 2002) [1].

The structure of the paper is as follows. In Section 2 we review the existing literature on the gender wage gap. Section 3 describes the available data. In Section 4 we outline our modeling approach, which employs quantile regression. Section 5 discusses the results and our conclusions are presented finally in Section 6.

## 2. Earlier studies – explaining the gap.

Concerns with gender differences in pay have led to a series of studies attempting to understand the size and causes of the gap in wages between employed men and women. The gender gap has been attributed to women who interrupted their employment histories for 'home time', imputed on the basis of children born (Zabalza and Arrufat, 1986), to unequal remuneration for education and experience and low pay in part-time jobs (Ermisch and Wright, 1993; Manning and Petrongolo, 2004; Joshi and Paci 1998). Other authors introduced further information to account for gender wage differences, for example, training, travel to work distances, firm size and unionization and these have all contributed to explaining some of the gap (Anderson, 2001). It is also supposed that occupational segregation keeps women's wages low even for full-timers, in that it is more

[1] If men were employed continuously in the same job they were more likely, than in earlier periods, to face a 10% decline in their real wage over the 1990s.

difficult to implement the 1986 Equal Pay for Equal Value legislation where there are few male comparators. Also, women may be crowded into these fields, thereby keeping pay low. Sex segregation of an occupation might lower pay for both women and men. When segregation is recorded at the level of the workplace, there is some evidence of a pay penalty to working in a feminized workplace (Millward and Woodland 1995). Hakim (1998) concluded that vertical sex segregation is much more important than horizontal segregation across occupations, accounting for around three quarters of the gender earnings differential. This implies that women's earnings were held back more by their failure to make occupational advances than by their choice of, or selection into, stereotypically female types of job on entry to the labour market.

Greater family responsibilities have tended to be associated with higher wages for men (Akerloff, 1998) and lower wages for women, mirroring and possibly reinforcing, gender differences in family roles. Greenhalgh (1980) and Davies and Peronaci (1997) attempted to disentangle productivity and selection in the marital status differentials in men’s wages. Joshi and Paci (1998) failed to find much direct impact of motherhood on women’s pay after allowing for its indirect effects through work experience and part-time employment. When models of the pay gap which do not control for family responsibilities find an unexplained gap between men’s and women’s pay, one reason may be the asymmetrical effects on wages of family responsibilities.

The gender pay gap has also been found to vary across ages and across the wage distribution (Arulampalan et al, 2006), being higher at the top and bottom ends of the distribution. The gap can in principle be created in a number of ways; by men having more favourable market characteristics than women; or by men being rewarded more favourably for their market characteristics than women.

Booth and Francesconi (1999) reported that among full-timers remaining in the labour force over a 5-year period, men and women had remarkably, and surprisingly, similar chances of promotion. This would appear to suggest that women full-timers were indeed breaking through the glass ceiling, though there is some doubt about how well promotions are identified in the data. Booth et al (2003) then pointed out that women were nevertheless at some disadvantage, as they had higher chances of quitting and gained less wage growth between the rungs of the promotion ladder than the men, a 'sticky floor'. They also conjectured, though they could not directly observe, that women's slower wage growth may be due to being less aggressive or less risk taking in the job market,[2] conjectures also supported by Blackaby et al (2005) and Manning's (1996) view of the importance of monopsony in the setting of lower wages for women. Manning and Swaffield (2006) and Joshi et al (2007) examining wage rate levels over the first two decades of labour force years found that they cannot be explained solely by men accumulating more employment experience. The rates of pay offered to identical, full-time, continuous workers show a greater advantage for men over women after 10 years or so in the labour market than on entry. These earlier analyses form the starting point of the current investigation.

## 3. Data

Our wage mobility data for analysis come from the employment histories of the 1958 British Birth Cohort members, the National Child Development Study (NCDS). NCDS is a nationally representative longitudinal survey of over 17,000 births in 1958. There has been some attrition to around 11,000 in 1991 and 2000, although Hawkes and Plewis's (2006) examination of non-response in the NCDS data found few significant predictors of

[2] Booth et al (2003) argued that women were being treated as 'loyal servants' whose assumed aversion to moving, possibly due to domestic responsibilities or partnership ties, kept their wages down.

attrition. This supports the view that the data are still reasonably representative of this population.

NCDS contains detailed employment, occupational and education histories for its members up to age 42, at the time of writing this paper. Information about the wages of the employed and their job tenure were only available about jobs held at the survey sweeps, ages 42, 33 and 23. For the age 42 contact/interview in 2000, NCDS data achieved 11,419 interviews, 5624 with cohort men and 5795 with the women. The cohort members who gave complete information about their weekly working hours and earnings is less than those who were interviewed at both sweeps of NCDS partly because these data were only available for those employed at the interview, but also because earnings data, not uncommonly, had a higher item non-response than most other questions in the interview. In addition, we decided to focus on those employed full time both at age 33 and at age 42, a sample of 3558 full-time employees, 2606 men and 952 women, with full information on all other variables entered. All wages have been adjusted to reflect inflation changes, all being deflated or inflated to year 2000 values. A full list of the variables used, their means and standard deviations are presented in Appendix, Table A1.

These men and women, employed full time at both 33 and 42 varied from the rest of the interviewed sample at age 42 in terms of their qualifications, partnership status and family at age 42. Both men and women in our sample were more highly educated than those who were not working full time at this point, and women were more highly educated than the men; 34.0 per cent of the men in our sample and 39.4 per cent of the women had a qualification at NVQ level 4 (degree level) or above by age 42. Among those interviewed at age 42, but not in our sample, 20.4 per cent of the men and 21.8 per cent of the women had such degree level qualifications at this age. Men in our sample tended to be more likely to be partnered at age 42 (83.2%) compared with men who were not in our sample (66.1%), and than women in our sample (67.7%) who were also less likely to be partnered than women who were not in our sample (83.6%). Men in our sample (63.9%) were more likely to have a dependent child than men out of our sample (52.4%) and women in our sample were far less likely to have such a child (43.5%) than women out of our sample (82.7%) at age 42. Qualifications, partnership status and having a dependent child appear to be associated with the selection of men and women into this sample, with opposite effects for men and women in the case of family status. Dependent children are particularly uncommon among women employed full-time at both 33 and 42.

The majority of our sample, working full time at 33 and 42, are men (73%). The women are a select group who should be best placed to take advantage of equal pay policy. In focusing on women employed full time at two points in time across the 1990s, we are removing some of the normal variation that occurs in women's wages. Many British women work part time when they have children, and move in and out of the labour market generating an intermittent employment career. Those who were working full time in their thirties are likely to be at the top end of the employment attachment spectrum, working continuously. They will not all be committed high-flyer career women depicted as having higher qualifications and having delayed childbearing, if any, into their thirties. However, the high flyers will be in this group, along with others who are working full time out of financial necessity, and possibly some mothers who had their children early and had returned to full-time employment.

The early 1990s was a period of recession until 1993 which gave way to more buoyant labour market conditions to the end of the decade. For the sample who were working full time both at age 33 and age 42, women's raw hourly wage rate at age 33 was £8.47 per hour compared with a men's wage rate of £10.10 at the same age and time after RPI adjustment. By age 42, on an adjusted basis, women's hourly wage rate had increased to £9.84, a 16.2 per cent rise, and men's rates to £12.36, a 22.4 per cent rise. Thus controlled for working full time, the female to male ratio for this NCDS cohort started out at 83.9 per cent but declined by age 42 to 79.6 per cent, a fall of 4.3 percentage points over this 9 year period in mid-life over the 1990s. National UK figures in 2000 for full-time employees put the gender pay gap at between 18.9% and 20.2%, depending on the data set (WEU, 2004), which is within the same range as our NCDS figure even though NCDS data consist of only one British age group. However, WEU (2004) chose to focus on the fact that this was a reduction in the gap since 1997/98, whereas according to NCDS data for people in their 30s, it was an increase in gender wage gap for full timers since 1991.

The means of our dependent log hourly wage growth variables (Appendix Table A1) are geometric means; these suggest, unlike the arithmetic means cited above, that there is little difference between men's and women's wage growth, at the mean, over the period, 12 compared with 13 per cent.[3] However, plotting our wage growth measure over the percentiles of the distribution, rather than solely at the mean, shows there is considerable variation. Figure 1 plots the male to female ratio of wage growth across the percentiles of wage at 33 and shows that the growth gap between men and women is much higher at the high paid end over the 9 year period. Figure 2 plots the ratio of hourly wages at age 42 over age 33 and shows that people with high salaries are commandeering disproportionate percentage pay rises over the period, up to four times bigger than those in lower earnings. The differences in wage growth across the distribution is consistent

[3] The geometric mean minimizes the effect of extremes (high earners in this case) in its calculation, unlike the arithmetic mean that emphasizes high earners.

with figures from UK national statistics; from 1992 to 1997 gross hourly earnings (excluding over time for adults over 21 years) increased by 13.9 per cent at the bottom decile compared with 17 per cent at the median and 20 per cent at the highest decile (Butcher, 2005). This constituted hourly earnings growing roughly in line with retail prices at the bottom end of the distribution but having much higher real increases at the top end of the distribution. Our figures also overlap with earlier findings on age distributions by Arulampalam et al (2005). At the end of this period in April 1999, the National Minimum Wage was introduced and this benefited low paid women more than low paid men (Butcher, 2005), but this was too late to have had a major effect on the earnings data collected for the NCDS cohort in 2000.

**Figure 1: Log ratio of male/female hourly wage by percentiles of log wage at 33 and interview age.**

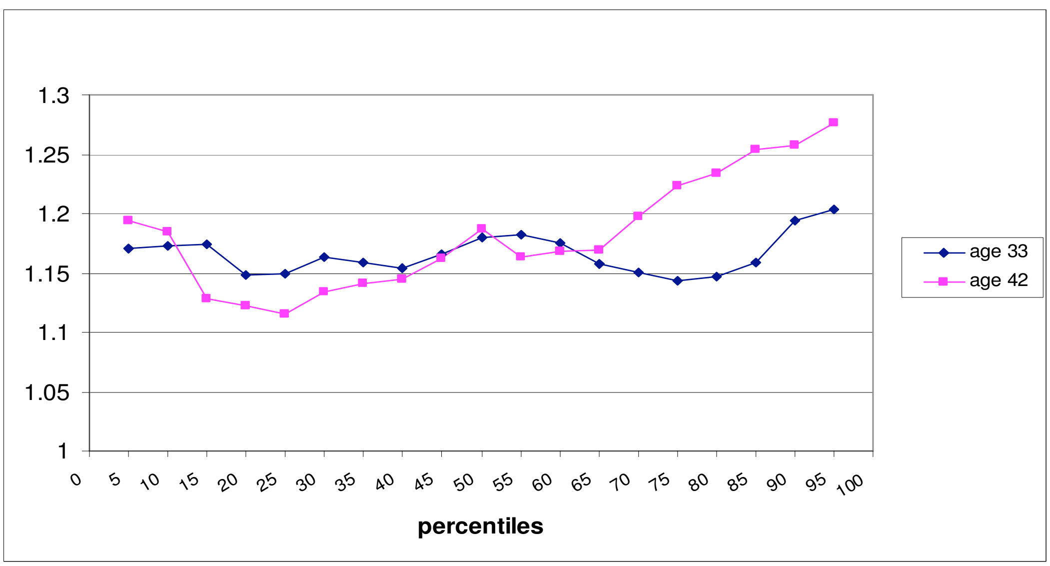
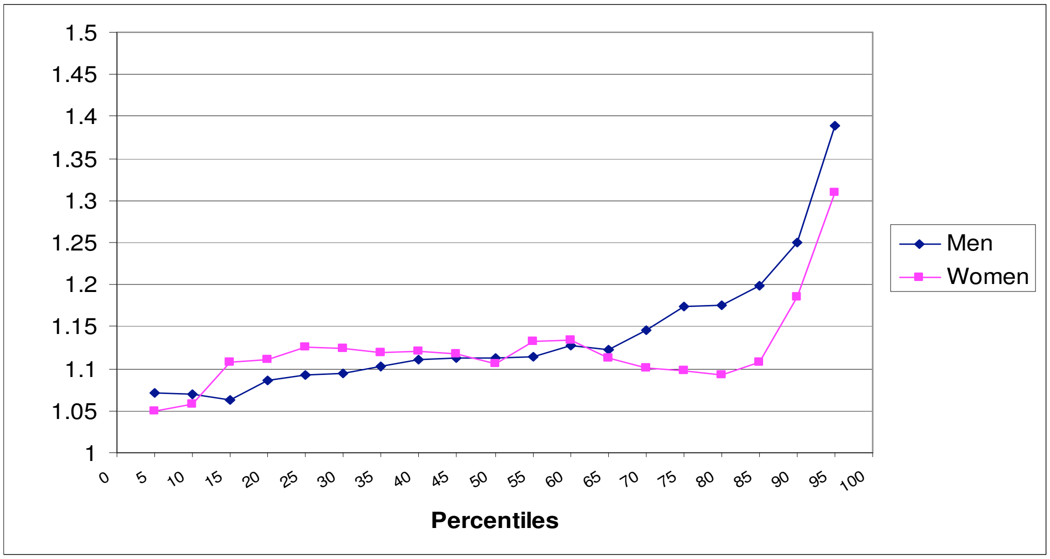


**Figure 2. Log ratio of age 42/age 33 hourly wage by percentiles of log wage at 33 and gender**




## 4. The Model

Our main approach to understanding wage growth is through the human capital theory (Becker, 1975; Mincer, 1974). Current wages at time t are expected to reflect earlier investments in an individual's human capital. The most obvious sources of earlier investments are educational qualifications achieved through schooling and work experience and training accumulated since leaving full-time education. Periods out of work are expected to have caused depreciation of the stock of human capital up to that point. Such periods also halt the accumulation of human capital, leading to subsequent declines in the potential wages of those who have been out of the labour market.

Women, as reviewed above, tend to have lower wage rates than men and in part at least, this is due to different earlier investments in human capital. Wage growth became part of the early debates over whether women expected to have time out of the labour market to have children and consequently invested less in human capital as result. Another alternative, suggested by Mincer and Polachek (1974), was that women would select an occupation bearing in mind their expectations to leave the labour market for a time, and choose occupations that incurred lower penalties (depreciation rates) for time out of work. The issue this raised was that it took for granted a slower rate of wage growth on returning to work after a labour market gap. What if there was a faster rate of wage growth, big enough for women to recoup the losses due to labour market absence? Some evidence emerged to support the idea that wage growth was faster after returns to work (Dex, 1985). However, over time, behaviour has changed quite substantially. Educational qualifications have equalized among young men and women, occupational choices have moved to be more similar, although far from being identical, work experience has not equalized although some women's gaps out of work for childbirth have drastically reduced, and wages have grown closer (particularly for younger workers). In this paper we focus on the period from age 33 to 42 and examine wage growth of men and women over this 9 year period of the 1990s. Individuals' wage rates at age 33 embody their past human capital investments up to that point.

A number of possible measures of hourly wage growth were considered. Hourly wages were tightly cut off at the lower end of the distribution and had a very long upper tail. We ruled out using the ratio of raw values of men's and women's hourly wages since they were very unlikely to be Gaussian errors since these were not symmetrically distributed and hence they may potentially be affected by extreme values or outliers. We also rejected using the measure of the absolute £ per hour difference in wages from age 33 to 42. Since pay rises are mostly in percentage rather than absolute terms, there were likely to be enormous differences between high and low earners. This would imply a multiplicative dependence on the starting wage that would not be adequately modeled using the raw wage values. We selected log hourly wage values in order to reduce the influence of outliers and a model of log hourly wage at time t (age 42) as a function of log hourly wage at time t-1 (age 33), log hourly wage at time t-1 squared, to allow for a non-linear relationship, gender and an interaction terms between gender and log hourly wage at time t-1.

Since wage growth is likely to vary across the values of wages at time t-1, we adopt a quantile regression approach. An ordinary least squares (OLS) regression is based on the mean of the conditional distribution of the regression's dependent variable. OLS is used because one implicitly assumes that possible differences in terms of the impact of the covariates along the conditional distribution are unimportant. However, this may prove inadequate in some situations. Quantile regression (Koenker and Basset 1978; Breckling, and Chambers 1988; Chambers and Tzavidis 2006), may be viewed as an extension of classical least squares estimation of conditional mean models to the estimation of an ensemble of models for several conditional quantile functions. Unlike OLS, quantile regression models allow for a full characterisation of the conditional distribution of the dependent variable. The central case is the median regression estimator which minimizes a sum of absolute errors. Other conditional quantile functions are estimated by minimizing an asymmetrically weighted sum of absolute errors. In a wage equation setting, the quantile regression model can be written as:

$$Q\left(\mathrm{Ln}\ (\mathrm{W_i}) \mid X_i\right) = X_i\ \beta(\mathrm{q}),$$

where $X_i$ is the vector of covariates, $\beta(\mathrm{q})$ is the vector of quantile regression parameters at the qth quantile and $Q\left(\mathrm{Ln}\ (\mathrm{W_i}) \mid X_i\right)$ denotes the qth conditional quantile of response variable given the covariates. The vector $\beta(\mathrm{q})$ is estimated by minimising:

$$\sum_{i=1}^{n}\left|LnW_i - X_i^T\beta\right|\left\{(1-q)I\left(LnW_i - X_i^T\beta \le 0\right) + qI\left(LnW_i - X_i^T\beta > 0\right)\right\}$$

with respect to $\beta$. Solutions to this minimisation problem are usually obtained using linear programming methods (Koenker and d'Orey, 1987). However, functions for fitting quantile regression now exist in standard statistical software for example in the R statistical package (library quantreg) and in STATA (command qreg), the latter being used in the estimations presented in this paper.

This approach has the additional advantage that it allows the regression coefficients associated with different covariates to vary across the centiles of wage distribution at time t. A common problem for all quantile-type fitted regression planes is that they can cross over for different values of q. This lack of monotonicity is essentially a finite sample problem and is typically due to a combination of regression model misspecification, collinearity and highly influential data values. There is extensive literature on the general case of how to monotonize fitted quantile regression planes.(Koenker, 1984; He, 1997). In this paper the choice of the log wage as the dependent variable assisted us in ensuring monotonicity.

Let us denote by the quantile at which we wish to fit the model, the quantile regression equation for the logarithm of wages at age 42 is then defined by:

$$\begin{aligned}\mathrm{Ln}(\mathrm{W_t}) = {} & \beta_0(\mathrm{q}) + \beta_1(\mathrm{q})\,\mathrm{Ln}(\mathrm{W_{t\text{-}1}}) + \beta_2(\mathrm{q})\,\mathrm{Ln}\,(\mathrm{W_{t\text{-}1}})^2 + \beta_3(\mathrm{q})\,\mathrm{Gender} + \beta_4(\mathrm{q})\,\mathrm{Gender{*}Ln}\,(\mathrm{W_{t\text{-}1}}) \\ & + \sum_{\mathrm{k=1}}^{\mathrm{p}}\beta_{\mathrm{k}}(\mathrm{q})(X_{\mathrm{kt}} - X_{\mathrm{kt\text{-}1}}) + \sum_{\mathrm{j=1}}^{\mathrm{p}}\beta_{\mathrm{j}}(\mathrm{q})Z_{\mathrm{kt\text{-}1}}\end{aligned}$$

Quantile regressions were fitted at the $10^{th}$, $25^{th}$, $50^{th}$,$75^{th}$ and $90^{th}$ percentiles. For the independent variables, *Z*, attached to the start of the period, t-1, we were mindful that using the wage at time t-1 as one of the covariates would already have controlled for the obvious determinants of wage levels up to time t-1. None the less, we considered it likely or possible, given occupation by age trajectories, that certain occupations (and possibly education levels) might have different wage growth paths. In addition, since high earners, overlapping with high occupations, are often able to command higher pay rises, we needed to capture this phenomenon. We included occupation codes at t-1 as dummies, therefore, to capture these effects.[4] In addition, we included an indicator of whether the occupation at time t-1 was mainly dominated by women, men or integrated across men and women (the reference group was mainly female dominated).[5] It was also possible that living with a partner or dependent children at base year would affect the extent to which individuals focused on their job and gaining promotion over the next 9 years. Stereotyped expectations here would be that having a partner and having a dependent child at t-1 would each motivate men to gain more promotion and wage growth, whereas they may impede women from pursuing their career. Partner and dependent child were entered as time t-1 dummy variables.

For independent variables, *X*, indicating change from time t-1 to time t, we included measures of human capital accumulation and depletion from the work histories of the individuals in the data over this 9 year period. Dummies were included for additional training spells of 3 days or more, gaining an additional educational qualification, any periods out of the labour market, or spells of part-time work, and whether there had been changes in 'jobs', to capture some of the earlier findings in the literature reviewed above.

---

[4] Education qualifications were also entered into the model, but overlapped to a considerable extent with the occupation indicators, reducing their significance while not adding anything to the model. A number of studies have argued that education affects occupation mobility, with higher mobility and promotion resulting from higher education (Harper, 1995; Sicherman and Galor, 1990). However, these earlier studies did not also include occupations. Occupations clearly encompass educational qualifications to a large extent, and are more closely related to the alternative career paths in or between organisations. Highest education dummies were also tried in the model instead of occupations, but did not perform as well and so were dropped from the final versions.

[5] Occupations were classified as mainly male or mainly female if they had a 70% share or more male (or female) employees in that occupation. Integrated occupations were the rest.

Our data on job changes did not contain information about whether individuals stayed with the same employer. However, we consider that changes of occupation category are most likely to indicate a change of employer. We included, therefore, a dummy variable for an upward change in occupation from the origin occupation at time t-1 to the destination occupation at time t. Similarly, a dummy for a downward occupational move over this period was also included. A third dummy indicated the individual reported a job change over the period, but no occupation change. Thus the omitted reference group is the group who stayed in the same occupation and did not change their job. This does not fully capture employer changes versus internal job changes, but it is a reasonable approximation. It has the added benefit of allowing us to incorporate the direction of occupational change. [6]

In addition changes in partner status and the addition of a new birth over the period were also included as another block of dummy variables. Several other variables were also tried but discarded because of their lack of significance.[7]

Models were run using the combined sample of men and women, using a dummy indicator to capture gender differences, and interaction terms with gender and wage at t-1, and wage at t-1 squared. At this point we need to point out that the interaction terms, (gender*log age 33 wage and gender*squared log age33 wage) were not significant for any of the models we fitted and therefore were not included in any of the final models.

[6] The vertical scale used is indicated in Appendix 1, and was devised by calculating the mean hourly wage of SOC 1990 major groups, and ranking them accordingly.
[7] Dummy variables were used to capture excessively large (or small) gaps between the two interviews which may have increased (or reduced) the length of the growth period. Dummies to recognise plus or minus two standard deviations from the mean inter-interview gap were not significant. We also included changes in region between the higher wage South East and other lower wage regions which were not significant and attitude measures from age 33 NCDS, but the missing values on these variables substantially reduced our sample size, by several hundred women and the same for men, so we decided to drop these variables.

Variables were entered in blocks in order to see whether one set moderated the effects of others. Blocks of variables were entered as follows; (1) gender on its own; (2) gender plus the wage at age 33 and age 33 wage squared term; (3) occupations at age 33; (4) occupational change and occupational segregation; (5) changes in human capital; and lastly (6), family partner and ill health variables. Where variables were not significant at a particular quantile, they were dropped from that quantile in the final model.

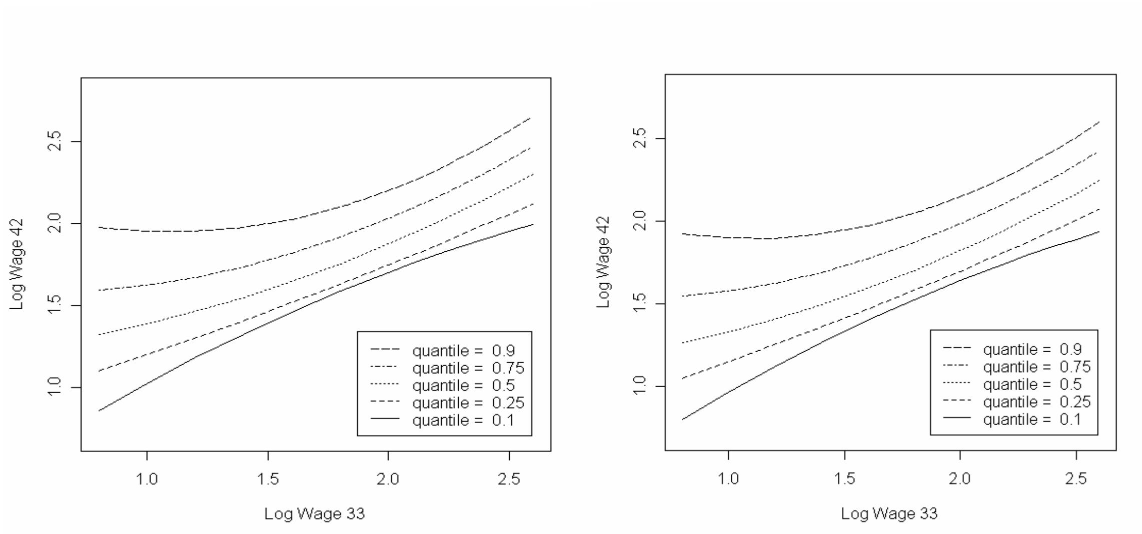
## 5. Results

Before other controls were added, the gender coefficient showed a male wage advantage at age 42 which was lowest in the 25$^{th}$ quantile at 11 per cent, followed by 16-17 per cent in the 10$^{th}$ and median quantiles, but considerably higher in the upper quantiles, 20 per cent at the 75$^{th}$ and 24 per cent at the 90$^{th}$ quantile. The results of the final models including all significant covariates are displayed in Table 1. The starting wage at age 33 was highly significant and positive. Adding the starting wage at age 33 on its own reduced substantially the male advantage in all quantiles but more at the top than the bottom quantiles. The addition of the wage squared term was also significant in all but one (25$^{th}$) quantiles, but did not add anything further to the explanation of the gender wage gap, except at the highest 90$^{th}$ quantile where the gender gap reduced by a further one percentage point.  The introduction of an interaction term between gender and age33 wage was never significant and was dropped therefore. After controlling for all the other variables, being a man gives an advantage over being a women of 5-6 per cent of the log age 42 hourly wage at all quantiles , with little variation, although the gender gap is slightly larger at the 10% than at the other quantiles. The exception is the 90$^{th}$ quantile, where the gender coefficient does not reach significance. At the top end of the wage distribution, therefore, the covariates can be argued to account completely for the gender wage gap.

The wage square term also changed the shape of the relationship between the logs of age 42 and age 33 hourly wage rates when it was first introduced by placing more emphasis in the higher quantiles on the squared age 33 wage effect and less on the wage at age 33 itself. The wage at age 33 and wage squared terms were all significant but varying in size and sign across the quantiles. These results are signaling a widening gap and wage growth at age 42 for those who had the highest wages at age 33. Since more men in this sample had higher wages at age 33 than women, this is an additional implied advantage to the men who managed to get the highest wages by age 33. The shapes of these relationships between age 33 wages and the age 42 wage are displayed in Figure 3 (a and b) using the coefficient values that emerged after other covariates were added. It is clear that the shapes of these relationships varies by quantile. For wage quantiles below the median, there is a concave relationship such that increases in age 33 wage rates are associated initially with increases in age 42 wages but they eventually start to plateau off. For quantiles above the median, a convex relationship is evident with the age 42 wage rates rising at faster rates as age 33 wages rates increased. The 75th wage quantile has started to move in the same direction, but has not quite reached the ability of those at the top wage rates, not only to stay at the top of the wage distribution, but to extend their lead, akin to the controversial ‘fat cat’ executive pay phenomenon.

**Figure 3:** Relationship between Log-wage at age 33 and Log-wage at age 42 at different percentiles for men (right plot) and women (left plot).

(a) Women (b) Men




Over and above wage rates at age 33, occupation dummies at age 33 were also significant predictors of age 42 wages across all quantiles. The addition to age 42 wages generally increased moving up through the occupational hierarchy at each wage quantile, in comparison with the reference occupation, elementary occupations. Interestingly, it is not the case that being in a higher quantile, but the same occupational category was associated with higher wage growth. This is because we have already accounted for directly the higher wage growth associated with starting off with higher wage rates. In terms of the top management, professional and associate professional occupations, wage growth was largest for those who were in the 25th or, to a lesser extent, the median quantiles, after controlling for the effect of their starting wage. However, there were some variations in rankings by occupation. At the lowest two quantiles, personal service and sales ranked above skilled trades occupations in their wage growth over the period, with personal service above sales in growth. At the median quantile and above, sales ranked

**Table 1. Quantile regression results for model of hourly wage at age 42.**

| | Q 10% | Q 25% | Median | Q 75% | Q 90% | Mean Regression |
|---|---|---|---|---|---|---|
| Constant | 0.053 (0.195) | 0.693(0.100)* | 1.115(0.085)* | 1.590(0.092)* | 2.279(0.156)* | 0.996(0.104)* |
| Gender (male=1) | 0.062 (0.027)* | 0.050(0.016)* | 0.054(0.014)* | 0.047(0.016)* | 0.055(0.030) | 0.051(0.018)* |
| **Wage at age33** | | | | | | |
| Log wage33 | 1.030 (0.179)* | 0.412(0.089)* | 0.083(0.073) | -0.218(0.079)* | -0.697(0.133)* | 0.261(0.090)* |
| Log wage33 squared | -0.118(0.040)* | 0.045(0.020)* | 0.135(0.016)* | 0.207(0.017)* | 0.315(0.028)* | 0.074(0.020)* |
| **Occupation age33** | | | | | | |
| Managerial | 0.234 (0.062)* | 0.289(0.034)* | 0.312(0.031)* | 0.297(0.034)* | 0.294(0.062)* | 0.348(0.037)* |
| Professional | 0.280 (0.063)* | 0.333(0.035)* | 0.300(0.032)* | 0.271(0.036)* | 0.294(0.066)* | 0.350(0.039)* |
| Associate professional | 0.251 (0.063)* | 0.309(0.035)* | 0.275(0.031)* | 0.271(0.035)* | 0.303(0.065)* | 0.317(0.038)* |
| Administrative-secretarial | 0.176 (0.061)* | 0.271(0.034)* | 0.224(0.031)* | 0.161(0.034)* | 0.139(0.063)* | 0.226(0.037)* |
| Skilled trades | 0.054 (0.057) | 0.112(0.032)* | 0.123(0.029)* | 0.071(0.033)* | 0.056(0.060) | 0.111(0.036)* |
| Personal service | 0.153 (0.064)* | 0.171(0.036)* | 0.133(0.032)* | 0.123(0.036)* | 0.059(0.067) | 0.150(0.039)* |
| Sales and customer service | 0.080 (0.071) | 0.117(0.041)* | 0.178(0.036)* | 0.171(0.041)* | 0.217(0.076)* | 0.168(0.045)* |
| Process, plant and machine ops | 0.027 (0.059) | 0.074(0.034)* | 0.057(0.030) | 0.037(0.034) | -0.007(0.063) | 0.042(0.037) |
| Elementary(Ref) | | | | | | |
| **Mobility changes** | | | | | | |
| Upward occupation mobility | | | 0.031(0.013)* | 0.047(0.015)* | | 0.029(0.016) |
| Downward occupation mobility | -0.167 (0.027)* | -0.131(0.016)* | -0.096(0.014)* | -0.068(0.016)* | -0.089(0.063)* | -0.117(0.017)* |
| Job change no occupation mobility | | | | | | 0.029(0.023) |
| **Segregation at age33** | | | | | | |
| Male dominated | 0.057 (0.023)* | 0.111 (0.019)* | 0.104(0.017)* | 0.131(0.019)* | 0.195(0.036)* | 0.138(0.021)* |
| Integrated | | 0.040 (0.018)* | 0.050(0.017)* | 0.069(0.019)* | 0.121(0.034)* | 0.064(0.020)* |
| Female | | | | | | |

| | | | | | | |
|---|---|---|---|---|---|---|
| dominated(Ref) | | | | | | |
| **Human capital changes** | | | | | | |
| Training by 42 | 0.084 (0.020)* | 0.067(0.011)* | 0.055(0.010)* | 0.038(0.012)* | 0.077(0.022)* | 0.065(0.013)* |
| 1 spell of non-work | -0.177 (0.039)* | -0.154(0.022)* | -0.101(0.020)* | -0.084(0.022)* | -0.024(0.041) | -0.109(0.024)* |
| 2+ spells non-work | -0.195 (0.068)* | -0.218(0.040)* | -0.206(0.035)* | -0.157(0.039)* | -0.180(0.072)* | -0.196(0.043)* |
| Part time spell by 42 | -0.123 (0.062)* | -0.107(0.037)* | | | | -0.059(0.041) |
| New qual by 42 | | | 0.039(0.013)* | | | 0.031(0.017) |
| Limiting illness by 42 | -0.062 (0.023)* | -0.037(0.013)* | | | | -0.034(0.014)* |
| Not partnered at 33 | | | | 0.039(0.014)* | | 0.021(0.022) |
| Pseudo/R square,N=3558 | 0.262 | 0.315 | 0.350 | 0.356 | 0.344 | 0.480 |

* significant at 95% confidence level. ( standard errors in parentheses)

higher in wage growth than both personal service and skilled trades jobs and even above administrative secretarial jobs at the 75th and 90th quantiles.

Benefits to wage growth from upward occupational mobility were only evident for those in the median and 75th quantiles whereas the bad effects on wages of downward occupational mobility by age 42 were evident across all quantiles, largest at the 10th and 25th quantiles. Benefits of being in a male dominated occupation occurred at all quantile levels, were higher than the benefits of working in an integrated occupation, and progressively increased from the lowest to the highest quantile. There was a wage growth benefit from working in an integrated occupation, which also increased moving to higher quantiles. However, it was not significant at the 10th quantile, possibly because integrated occupations are not common and gender segregated jobs more common at the low wage end of the wage spectrum.

Having additional training by age 42 was associated positively with wage growth at all quantiles, but gaining a new educational qualification only had a significant positive contribution to wage growth at the median quantile. Spells out of work had the expected depressing effect on wage growth, multiple spells having worse effects. The depressing effect of one spell out of work was highest at the lowest quantile, and lowest at the highest quantile. This is also an interesting finding. Having a higher wage appears to be a protective factor against wage depreciation from one but not from multiple spells out of work. Having a spell working part time had a negative effect on wage growth when in the 10th or 25th quantile, but not at higher quantiles. It may be the case that part-time work is not a disadvantage when taken in higher paid jobs, although the lack of significance of part-time work at higher quantiles is also consistent with their being few part-time jobs in higher paid positions. Having a limiting longstanding illness between ages 33 and 42 was also associated with a significant negative effect on wage growth only at the lower quantiles. It may be the case that those who have such illnesses are only able to keep working in lower paid, and possibly less demanding jobs.

Of the demographic and family variables, very few were significant at any quantile level, the exception being those who were not partnered at age 33 had slightly higher (4%) wage growth, and only when they were at the $75^{th}$ wage quantile. This may be an effect for unpartnered women, since unpartnered men are known to do worse in their wages than partnered men (Akerloff, 1998). It is notable and contrary to stereotype expectations, that the gap between men's and women's wage growth at this stage of their careers, admittedly a select group of women, was not explained by the presence or absence of family commitments.

More of the benefits of using a quantile model to estimate the coefficients are apparent in comparison with the single regression model estimated on the full set of covariates at the mean and displayed in the final column of Table 1. Although the overall mean gender coefficient is a reasonable representation of the quantile effects, the coefficients on the age 33 wage and wage squared terms and its change of shape would have been hidden in a single regression on the same covariates. Similarly, on the basis of a single regression we would have concluded that there is no effect on age 42 wages of experiencing upward mobility, of having a new qualification by age 42 or of having any part-time work experience since age 33. We would also have concluded, mistakenly, that downward occupational mobility, being in a male dominated or integrated occupation, having spells out of work all had a constant effect across the wage distribution.

## 5. Conclusions

In this paper we have examined wage change between age 33 and 42 of the group of men and women selected because they were reasonably close comparators in their labour market experience. At the starting point, age 33, these women and men differed in their hourly wages, with men having the advantage, despite the fact that these women were more highly educated than the men. Over this period of the 1990s, where economic growth took over from a recession in the early part of the decade, the gender wage ratio moved further against this groups of NCDS women. Clearly these men's wages grew more strongly through this decade, especially at the top of the distribution. Our analyses point to the main reasons for NCDS employed men's superior wage growth over this decade of mid life: It is their advantaged starting point both at higher wage rates and in occupations with higher wage growth over the decade. Modelling these effects at different quantiles, revealed the much stronger positive effect of the squared wage term at higher quantiles. To some extent these results support the view that men and women are being paid unequal wages for equal work in the same occupations. However, our occupational categories are broad, so it is not clear from our data that men and women are working along side each other in the same jobs, even though they are in the same occupational categories. Occupational segregation also helps to explain men's stronger wage growth where they benefit increasingly in the higher quantiles from working in male dominated occupations. In part this is due to occupational choices differing between men and women usually earlier in their careers. This is one of the results that became apparent only by using a quantile approach. What was notable is that domestic responsibilities and partnership status at age 33 and changes in these by age 42 did not contribute any explanation to the variations in wage growth and the gender wage gap between these men and women over their 30s. However, such factors may have helped to explain the different wage and occupational starting points of these men and women.

Other advantages of having modelled the quantiles are that it became evident that having a new qualification by age 42 helped wage growth of those in the median wage range and having any part-time work experience since age 33 only penalized those at the bottom of the wage distribution. In addition, experiencing upward mobility, rather than having no significant effect at the mean, was seen to be significant at the median and 75$^{th}$ quantiles but not at other wage levels. Obviously upward occupational mobility is not possible in the top most occupation. But the results may also signal, for those with wages below the median, that there is no wage growth advantage from alternate strategies of job changing to a higher occupation, versus staying in the same job. Quantile regression also helped us to see that experiencing downward occupational mobility or having spells out of work have greater percentage wage penalties for those at the bottom end of the wage distribution than for those at the top end. From other studies, we know that the 1990s was a period of differential opportunities for pay rises at the top and bottom end of the wage distribution, heightened insecurity and chance of real wage loss for men who stayed in the same job, but also a time of greater opportunity for large pay rises for those who changed jobs. This study suggests that pay rises from changing jobs may be more common in the upper half of the wage distribution, and that pay falls also typically resulted from job changes that involved downward occupational moves. However, our study also shows that the same beneficial opportunities were not so available to women. Overall the quantile modeling reinforced the view that ‘to those who have, more will be given’ and less will be taken away.

## References

Akerloff, G. (1998) ‘Men without children’, *Economic Journal*, 108, pp.257-309.

Anderson, T. Forth, J. Metcalf, H. and Kirby, S. (2001) *The Gender Pay Gap*. Report to the Women and Equality Unit, London: Women and Equality Unit, DTI.

Arulampalan, W. Booth, A.L. and Bryan, M.L. (2006) 'Is there a glass ceiling over Europe? Exploring the gender pay gap across the wages distribution', *Industrial and Labor Relations Review*, 60 (2), pp.

Becker, G. (1975) *Human Capital*, second edition, Boston: National Bureau of Economic Research.

Breckling, J. and Chambers, R. (1988). M-quantiles. *Biometrika*, 75, 761-771.

Blackaby, D, Booth, A. and Frank, J (2005) 'Outside Offers and the Gender Pay Gap: Empirical Evidence from the UK Academic Labour Market', *Economic Journal,* 115, F81-F107,

Booth, A. and Francesconi, M. (1999) 'Job mobility in 1990s Britain: Does gender matter? *Research in Labor Economics*, 19, pp.173-89.

Booth, A., Francesconi, M. and Garcia-Serrano, (1999) 'Job tenure and job mobility', *Industrial and Labor Relations Review*, 53(1) pp.43-70.

Booth, A., Francesconi, M. and Frank, J. (2003) 'A sticky floors model of promotion and pay and gender', *European Economic Review*, 47 (2), pp.99-126.

Butcher,T. (2005) 'The hourly earnings distribution before and after the National Minimum Wage', *Labour Market Trends*, Vol.113, No.10, pp.427-35.

Chambers, R. and Tzavidis, N. (2006). M-quantile models for small area estimation. *Biometrika*, 93, 255-268.

Davies, H. and Peronaci, R. (1997) 'Male Wages and Living Arrangements', Birkbeck College Discussion Paper in Economics 5/97.

Dex, S. (1985) *The Sexual Division of Work*, Brighton: Wheatsheaf.

Dex, S. (2003) *Families and work in the 21$^{st}$ century*, York: Joseph Rowntree Foundation.

Ermisch, J. and Wright, R. (1993) ‘Wage offers and full-time and part-time employment of British Women’, *Journal of Human Resources*, 28, pp.111-133.

Equal Pay Task Force (2001) *Just Pay.* Manchester: Equal Opportunities Commission.

Francesconi, M. (2001) ‘The determinants and consequences of promotions in Britain’, *Oxford Bulletin of Economics and Statistics*, 63 (3) pp.279-310.

Greenhalgh, C. (1980) Male-female wage differentials: is marriage an equal opportunity? *Economic Journal*, 90 751–75.

Hakim, C. (1998) *Social Change and Innovation in the Labour Market,* Oxford, Oxford University Press

Harper, B. (1995) ‘Male occupational mobility in Britain’, *Oxford Bulletin of Economics and Statistics*, 57 (3), pp.349-69.

Hawkes, D. and Plewis, I. (2006) “Modelling Non-Response in the National Child Development Study”, *Journal of the Royal Statistical Society, Series A.*

He, X. (1997). Quantile curves without crossing. *The American Statistician*, 51, 186-92.

Joshi, H. (2005) ‘Gender and Pay: some more equal than others’ in A. Heath, J. Ermisch and D. Gallie (eds) *Understanding Social Change*. British Academy Centenary Volume, Oxford: Oxford University Press, pp. 151 – 185.

Joshi, H., Makepeace, G. and Dolton, P (2007) ‘More or less unequal? Evidence on the pay of men and women from the British Birth Cohort Studies’, *Gender, Work and Organization,* Vol.14. No.1, pp.37-55.

Joshi, H. and Paci, P. (1998) *Unequal pay for women and men: evidence form the British Birth Cohort Studies.* Cambridge, Mass: MIT Press.

Joshi, H., Paci, P. and Waldfogel, J. (1999) ‘The wages of motherhood: better or worse?’ *Cambridge Journal of Economics*, 23, 5, 543-564.

Kingsmill, D. (2001) *Women’s Employment and Pay.* London: Cabinet Office and DTI, Available online at http://www.kingsmillreview.gov.uk/ Last consulted 25 September 2006.

Koenker, R., and Bassett, G. (1978): Regression Quantiles. *Econometrica*, 46, 33-50.

Koenker, R. (1984). A note on L-estimators for linear models. *Statistics and Probability Letters*, 2, 323-25

Koenker, R., and d'Orey, V. (1987).Computing Regression Quantiles. *Applied Statistics*, 36, 383-93.

Makepeace, G., Dolton, P. and Joshi, H. (2004) Gender earnings differentials across individuals over time in British Cohort Studies, 1991 and 2000. *International Journal of Manpower*. Vol 25 no 3/4, 251-263 .

Manning, A. (1996) 'The Equal Pay Act as an experiment to theories of the labour market', *Economica*, 663: pp.191-222.

Manning, A. and Petrongolo,B. (2004) *The part time pay penalty*, Report to the Women and Equality Unit, London: Women and Equality Unit, DTI.

Manning, A. and Swaffield, J. (2005) 'The Gender Gap in Early-Career Wage Growth'. CEP Discussion Paper 0700, London School of Economics, http://econ.lse.ac.uk/staff/amanning/work/gender.html

Millward, N. and Woodland, S. (1995) Gender Segregation and male/female wage differentials, in Humphries, J. and Rubery, J. (eds.) pp. 221-244.

Mincer,J. (1974) *Schooling, Experience and Earnings,* New York: Columbia University Press.

Mincer, J.and Polachek, S. (1974) 'Family investments in human capital: earnings of women', *Journal of Political Economy*, Vol.82 (2) pp.S76-S108.

Nickell,S. Jones,P. and Quintini,G. (2002) 'A picture of job insecurity facing British men' in *Economic Journal*, Vo.112, No.476, pp.1-27.

Paci, P. Joshi, H. and Makepeace, G. (1995) 'Pay gaps facing men and women born in 1958: differences within the labour market' in Humphries, J. and Rubery, J. (eds) *The Economics of Equal Opportunity,* Manchester: Equal Opportunities Commission, 87–112.

Purcell, K., P. Elias and N. Wilton (2006) *Looking Through the Glass Ceiling: a detailed investigation of the factors that contribute to gendered career inequalities.* Liverpool: UK Higher Education European Social Fund/ downloadable from www.warwick.ac.uk/go/glmfSicherman,N. and Galor,O. (1990) 'A theory of career mobility', *Journal of Political Economy*, Vol.98, pp.169-92.

Women and Equality Unit, WEU (2004) The Gender pay gap in the UK, Gender Pay Gap Factsheet, London: Women and equality Unit. Dowloadable PDF from http://www.womenandequalityunit.gov.uk/research/factsheets/gender_paygap_2004.doc.

Women and Work Commission (2006) *Shaping a Fairer Future.* London: Women and Equality Unit, Department of Trade and Industry.

Wright, R.E. and Ermisch, J.F. (1991) Gender discrimination in the British labour market: a reassessment. *Economic Journal,* 101,406, 508–52.

Zabalza, A. and Arrufat (1985) 'The extent of sex discrimination in Great Britain', in Zabalza, A and Zannatos, Z. (eds) *Women and Equal Pay: the Effect of Legislation on Female Employment and Wages in Britain* Cambridge: Cambridge University Press

**Appendix Table A1: Means and standard deviations of variables**

| Descriptives of variables included in model | Men | | Women | |
|---|---|---|---|---|
| | Mean | SD | Mean | SD |
| Difference in LOG WAGES BETWEEN 42 AND 33 | 0.135 | 0.415 | 0.123 | 0.327 |
| Managers and Senior Officials | 0.218 | 0.413 | 0.168 | 0.374 |
| Professional | 0.113 | 0.317 | 0.150 | 0.357 |
| Associate Professional and technical | 0.107 | 0.310 | 0.186 | 0.389 |
| Administrative and secretarial | 0.066 | 0.248 | 0.298 | 0.458 |
| Skilled Trades | 0.214 | 0.410 | 0.020 | 0.140 |
| Personal service | 0.076 | 0.265 | 0.081 | 0.273 |
| Sales and Customer service | 0.039 | 0.193 | 0.039 | 0.193 |
| Process, plant and machine operatives | 0.118 | 0.323 | 0.048 | 0.215 |
| Elementary | 0.048 | 0.215 | 0.010 | 0.097 |
| Upward Mobility * | 0.243 | 0.429 | 0.280 | 0.450 |
| Downward Mobility * | 0.188 | 0.391 | 0.196 | 0.397 |
| No change in mobility status | 0.569 | 0.497 | 0.524 | 0.498 |
| Had any change in job but not occupation level | 0.075 | 0.264 | 0.079 | 0.270 |
| Male Dominated Occupation | 0.708 | 0.455 | 0.166 | 0.372 |
| Integrated Occupation | 0.227 | 0.419 | 0.336 | 0.473 |
| Female dominated occupation | 0.065 | 0.246 | 0.498 | 0.500 |
| Done any work related training for 3 or more days aged 42 | 0.474 | 0.499 | 0.455 | 0.498 |
| Worked Part Time from age 33 to age 42 | 0.011 | 0.105 | 0.064 | 0.245 |
| Experienced one period of non-working between age 33 and age 42 | 0.066 | 0.248 | 0.084 | 0.278 |
| Experienced two periods of non-working between age 33 and age 42 | 0.021 | 0.144 | 0.020 | 0.140 |
| Worked continuously from age 33 to age 42 | 0.913 | 0.282 | 0.896 | 0.306 |
| New qualification by age 42 | 0.149 | 0.356 | 0.231 | 0.422 |
| Had partner at 33 | 0.832 | 0.374 | 0.676 | 0.468 |
| Had dependant child at age 33 | 0.639 | 0.480 | 0.435 | 0.496 |
| Had no partner at age 33 and had a partner by age 42 | 0.080 | 0.271 | 0.110 | 0.313 |
| Had a partner at age 33 and no partner at age 42 | 0.066 | 0.248 | 0.096 | 0.294 |
| Had more children by age 42 | 0.304 | 0.460 | 0.165 | 0.377 |
| Had longstanding limiting illness at age 42 | 0.238 | 0.426 | 0.267 | 0.443 |
| Log wage at 33 | 2.223 | 0.414 | 2.061 | 0.394 |
| Log wage at 42 | 2.357 | 0.510 | 2.184 | 0.436 |
| Sample Size | 2606 | | 952 | |

*The occupational ranking for upward or downward moves from age 33 to age 42 was derived from 9 SOC major 1990 occupations ranked according to occupational mean wages: at the top end was professionals followed by manager; associate professional; administrative-clerical; skilled trades; process and plant operatives; personal and protective; sales; other elementary. An upward move was set to 1 if any occupation at age 42 was a higher level occupation then their occupation at age 33 and a downward move scored 1 if their occupation at age 42 was a lower level occupation compared to their occupation at age 33.