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Departamento de Economía
Universidad Carlos III de Madrid
Calle Madrid, 126
28903 Getafe (Spain)
Fax (34) 916249875

Estimating Union Wage Effects and the Probability of Union Membership in the U.K During 1991-2003[1-2]

Georgios Marios Chrysanthou
Department of Economics
Universidad Carlos III de Madrid
28903, Madrid, Spain
E-mail: gchrysan@eco.uc3m.es

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Abstract

Using a dynamic model of unionism and wage determination we find that the unobserved factors that influence union membership also affect wages. We observe a significant decline in trade union membership persistence during the period under analysis. We find that UK trade unions still play a non-negligible, albeit diminishing, role in wage formation. While unions were unable to establish a wage premium for male members during the two periods considered, the female union wage effect stood at (19.4%, 17.6%) during (1991-1996, 1997-2002) respectively. The endogeneity correction procedure employed yields a discernible pattern of the union wage effect relative to OLS and fixed effects thus, refuting the pessimistic conclusions reached by Freeman and Medoff (1982) and Lewis (1986) that endogeneity correction methodologies do not contribute to our understanding of the union wage effect puzzle.

Key Words: union membership persistence, union wage effects, unobserved heterogeneity, dynamic model of unionism and wage determination.
JEL classification: C31,C33, J31, J51

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1. INTRODUCTION: THE JOINT DETERMINATION OF UNION STATUS AND UNION WAGE EFFECTS

In the past decade few studies were dedicated to the topic of trade union wage differentials in the UK. Most authors became disinterested in the topic in that trade unions are currently viewed as having little relevance in the process of wage formation.

The successive anti-union (1980, 1982, 1988, 1990) Employment Acts and the consequent dramatic decline in aggregate trade union membership and recognition led to a fall in unions' bargaining power and a significant decline in union wage premia (see for instance Stewart, 1995).

Furthermore, as the simultaneous determination of union membership status and union wage differentials complicates the identification of the union wage impact, some authors (e.g. Swaffield, 2001; Blanchflower and Bryson, 2010) disregarded the potential pitfalls of employing a *selection on observables* approach and ignoring the problem of endogeneity.

It is hard to conceptualise a realistic scenario in which the unobserved determinants of the unionisation decision would not affect wages. We cannot readily ignore to encompass how the unobserved individual heterogeneity underlying the union membership decision is rewarded (see Robinson, 1989a; Vella and Verbeek, 1998).

We therefore estimate union wage differentials employing a *selection on unobservables* methodology and longitudinal data from the British Household Panel Survey (BHPS) spanning the period of (1991-2003).

The general consensus by many researchers in the past was that cross-sectional OLS analyses of the union wage effect are contaminated by the selectivity of union sector employees inflating the union wage effect (see Abowd and Farber, 1982; Freeman, 1984).

While there was little disagreement that union membership status is not exogenous (e.g. Freeman, 1984; Duncan and Leigh, 1985; Robinson, 1989a) authors such as Freeman and Medoff (1982) and Lewis (1986) have reached the pessimistic conclusion that there is no discernible pattern to the estimates of the union wage impact, many were considered to be suspiciously high or low, and endogeneity correction methodologies have contributed little to our understanding of the union wage differential *puzzle* (Robinson, 1989a, p.640).

Upon summarising the then existing literature Robinson (1989a) concludes that the outcome is conflicting: endogeneity correction methods (such as Inverse Mills Ratio and Instrumental Variables) produced an upward adjustment as opposed to OLS, whereas longitudinal (differencing) methods produced a downward adjustment (Robinson, 1989a, p.640).

Though, some researchers employing longitudinal data sets attribute the resulting reduction in the union wage effect to fixed effects of higher quality workers present within the unionised sector this is not consistent with studies employing IMR or IVE methods to deal with such effects (Robinson, 1989, p.658).

Estimation of union wage differentials using longitudinal data requires controlling for the endogeneity of union membership status. Fixed effects and Instrumental Variables estimators (e.g. Hausman and Taylor, 1981) assume that this endogeneity is individual-specific and fixed and are thus restrictive in their treatment of unobserved heterogeneity (see Robinson 1989a,b; Vella and Verbeek 1999a).

Thus, we adopt a version of the estimation methodology offered by Vella and Verbeek (1998, 1999b) that explicitly identifies the different sources of endogeneity of union membership status.

We wish to establish whether UK trade unions still play a role in wage formation following the introduction of the successive Employment Acts targeted towards weakening their bargaining strength. We are interested in estimating the degree of true structural persistence of union membership. Further, we are interested in exploring the role of unobserved individual heterogeneity in the unionisation decision and the manner in which it is rewarded. Finally, we wish to investigate whether the economic sorting structure governing the entry into the two sectors provides a discernible pattern of the endogeneity corrected estimates relative to the *uncorrected* estimates of the union wage effect.

The empirical results indicate that trade unions in the UK still play a non-negligible, albeit diminishing, role in wage formation. We observe a significant decline in trade union membership persistence during the period under analysis. While unions failed to establish a union wage differential for their male members, the estimated female union impact still remains significant. Finally, the endogeneity correction procedure employed yields a discernible pattern of the estimated union wage impact relative to OLS and fixed effects. This is in line with Robinson (1989a) and refutes the pessimistic conclusions reached by Freeman and Medoff (1982) and Lewis (1986) in their influential surveys.

The rest of the paper is organised as follows. Section 2 outlines the econometric model; Section 3 the estimation procedure; Section 4 analyses the estimation results; and finally Section 5 concludes.

2. A DYNAMIC MODEL OF UNIONISM AND WAGE DETERMINATION

Equation (1) outlines the primary wage equation and assumes that individuals sort themselves into their preferred sector (union/non-union) on the basis of wages which are determined by observed and unobserved attributes and their respective prices. The potential wage corresponding to individual i employed in sector j, in time period t is given by $w_{j,it}$. The non-unionised and unionised sectors are denoted by $j = \{0, 1\}$ respectively, β is an unknown parameter vector and x_{it} is the conventional vector of personal and industrial characteristics which is also inclusive of time dummies. The unobserved random components of the employee's wage are given by $(\alpha_{j,i}, \varepsilon_{j,it})$ and the usual error component structure assumes $\alpha_{j,i} \sim iidN(0, \sigma_\alpha^2)$ and $\varepsilon_{j,it} \sim iidN(0, \sigma_\varepsilon^2)$:

$$\begin{aligned} w_{j,it} &= \beta'_{j,t} x_{it} + \alpha_{j,i} + \varepsilon_{j,it} \\ t &= 1, ..., T;\ i = 1, ..., N;\ j = \{0, 1\} \end{aligned} \tag{1}$$

Employment within a unionised establishment is also contingent on the employer's willingness to hire him/her (see Abowd and Farber, 1982). The major limitation of the estimation methodology employed here is that it does not sufficiently control for employer characteristics while on the other hand, individual employees' attributes are allowed to be an integral part of the employer's decision making process. While in the estimated models presented employer attributes are captured through the industrial classification dummies and establishment size controls these are not adequate in order to assign any specific effects purely to unobserved heterogeneity.

The dynamic reduced form model depicting the decision of an individual to join either the unionised or non-unionised sector is provided in equation (2). The benefits of employment within the unionised sector are captured by the latent variable U^*_{it}. The union membership status of an individual i in period t, is indicated by the dummy variable U_{it}.

The unknown parameters to be estimated are $(\gamma_1', \gamma_2)'$ and the composite error term ν_{it} captures the unobserved individual heterogeneity underlying the union membership decision and is decomposed into an individual-specific component ϑ_i and an individual time-specific effect η_{it}. The logarithm of the gross average hourly wage rate (log of weekly wage divided by usual paid hours including overtime) is denoted by w_{it}:

$$\begin{aligned} U^*_{it} &= \gamma_1' x_{it} + \gamma_2 U_{i,t-1} + \vartheta_i + \eta_{it} \\ U_{it} &= I(U^*_{it} \rangle 0) \\ w_{it} &= w_{j,it} \;\; if \;\; U_{it} = j \end{aligned} \tag{2}$$

Denote $ej_{,it} = \alpha_{j,i} + \varepsilon_{j,it}$ and $\nu_{it} = \vartheta_i + \eta_{it}$, let ν_i be a T vector of ν_{it} and $x_i = [x_{i1}, ..., x_{iT}]'$. Assuming that:

$$\nu_i \mid x_i \sim iidN(0, \sigma^2_\vartheta ii' + \sigma^2_\eta I) \tag{3}$$

$$E\{ej_{,it} \mid x_{i,}\nu_i\} = \tau_1 \nu_{it} + \tau_2 \bar{\nu}_i \tag{4}$$

where $\bar{\nu}_i = \frac{1}{T}\sum_{t=1}^{T} \nu_{it}$, i is a T dimensional vector of ones and (τ_1, τ_2) are unknown constants to be estimated.

Expression (3) enforces normality and a strict error components structure on the reduced form model for union membership and precludes any form of autocorrelation in η_{it} while equation (4) permits heteroskedasticity and autocorrelation in $\varepsilon_{j,it}$ but imposes the strict exogeneity of x_{it}.[3]

Potential seniority and non-pecuniary benefits can be sufficiently strong motives for individuals to remain within the unionised sector, irrespective of wage changes, and this introduces state dependence in the model (see Vella and Verbeek, 1998). The inclusion of a lagged union membership status variable in the reduced form model prevents the error components from incorrectly capturing the dynamics which should be credited to lagged union membership.[4]

The random components $(\alpha_{j,i}, \theta_i)$, $(\varepsilon_{j,it}, \eta_{it})$ in equations (1,2) denote the individual-specific and the individual/time-specific effects respectively. It is assumed that these are independently and identically distributed drawings from a multivariate normal distribution, where every effect is potentially correlated with its counterpart, of the same dimension, in the other equation. More specifically the four covariances $(\sigma_{j,\alpha\theta}, \sigma_{j,\varepsilon\eta})$ are allowed to be non-zero.[5] These covariances indicate that the random components in the wage equation are potentially correlated with the random components in the union membership equation and this is

[3] Note that testing for non-normality in the reduced form model for union membership can be quite difficult computationally.

[4] It is not possible to include dynamics coming through the lagged dependent variable in the wage equation. Arellano *et al* (1997) propose an alternative estimator, constrained to models with Tobit types of censoring, that permits for lagged latent dependent variables to enter both the primary and reduced form equations linearly.

[5] The covariances between the effects in the union/ non-union wage equations are not specified, whereas all remaining covariances are set to zero.

precisely what produces the potential endogeneity of union membership status in the primary equation.

The covariances convey valuable information about the form of sorting into the two sectors (see Vella and Verbeek, 1999a). Note that the θ_i are constructed so that their average value for union employees is positive while their average value for non-union employees is negative. For tractability assume that the endogeneity is taken to operate purely via the individual-specific effects $(\alpha_{j,i}, \theta_i)$. In the case that either covariance between (α, θ) is non-zero then the unobserved factors that determine union membership influence wages as well (and similarly if either covariance between $(\varepsilon_{j,it}, \eta_{it})$ is non-zero).

A hierarchical sorting structure requires that both covariances are positive $(\sigma_{1,0}>0)$ so that individuals with high values of θ are, on average, the best employees in terms of their endowment of unobserved productivity, irrespective of whether they are located in the union or non-union sector (and *vice-versa*).

A comparative advantage (or positive sorting structure) requires that employees perform differently in the two sectors and sort themselves appropriately $(\sigma_{1,0}<0)$. This implies a negative association between the relative productivity in the two sectors and demands that the contribution of unobserved heterogeneity raises wages in both sectors (i.e. $\sigma_{1,\alpha\theta}>0, \sigma_{0,\alpha\theta}<0$).[6]

Note that solely a degenerate hierarchical structure, imposing perfect correlation between sector-specific skills, can meet the strict and restrictive requirement of the equality of the two covariances imposed by either of the Instrumental Variables or the restricted Control Function estimators. A comparative advantage structure is precluded *a priori* (see Vella and Verbeek, 1999a).

To estimate the union wage differential we enforce the restriction that the returns to observed characteristics are both time and sector invariant. The wage equation (*eq.*1) then becomes:

$$\begin{aligned} w_{it} &= \beta' x_{it} + \delta U_{it} + e_{it} \\ e_{it} &= U_{it}(\alpha_{1,i} + \varepsilon_{1,it}) + (1 - U_{it})(\alpha_{0,i} + \varepsilon_{0,it}) \end{aligned} \tag{5}$$

3. ESTIMATION PROCEDURE

Following Vella and Verbeek (1998, 1999b) we start with equation (5) which is made conditional on the t-dimensional vector U_i, and the vector of exogenous variables x_{it}:

$$E(w_{it} \mid x_{it}, U_i) = \beta' E(x_{it} \mid x_{it}, U_i) + \delta E(U_{it} \mid x_{it}, U_i) + E(\alpha_{j,i} \mid x_{it}, U_i) + E(\varepsilon_{j,it} \mid x_{it}, U_i) \tag{6}$$

Estimation of the reduced form model for union membership (*eq.*2), provides the estimates of the unobserved individual heterogeneity (see Appendix II). This is a dynamic random effects Probit model with a likelihood function:

$$\prod_i \int \prod_t \Phi\left(\frac{\gamma' \Psi_{it} + \theta_i}{\sigma_\eta}\right)^{U_{it}} \Phi\left(-\frac{\gamma' \Psi_{it} + \theta_i}{\sigma_\eta}\right)^{1-U_{it}} \frac{1}{\sigma_\theta} \phi(\theta/\sigma_\theta) d\theta \tag{7}$$

where, $\gamma = (\gamma_1', \gamma_2)'$, $\Psi_{it} = [x_{it}, U_{i,t-1}]$, and (Φ, ϕ) correspond to the cumulative probability and density functions of the standard Normal distribution.

[6] Note that $\sigma_{1,0}$ cannot be estimated directly.

The inclusion of the lagged union membership variable as a regressor in ($eq.$2) gives rise to the problem of initial conditions (refer to Heckman, 1981a). The initial conditions problem occurs when the initial value of the dependent variable is correlated with unobserved individual heterogeneity. The presence of individual-specific effects θ_i clearly invalidates the assumption of exogeneity of union membership status in the first period.

The initial conditions problem cannot be readily ignored since the random effects maximum likelihood estimator in its standard form will be inconsistent (see Heckman, 1981a,b). Further, ignoring the correlation between individual-specific effects ϑ_i and the initial conditions will overstate the degree of state dependence.

We employ Wooldridge´s (2005) solution to the initial conditions problem. This involves modelling the distribution of the unobserved effect conditional on the initial value and the observed history of strictly exogenous explanatory variables (see Appendix I).

The last two conditional expectations in ($eq.$6) are estimates of the unobserved heterogeneity and can be expressed (see Appendix II) as:

$$\begin{aligned} E(\alpha_{j,i}|x_{it}, U_i) &= \sigma_{j,\alpha\theta}\left\{\frac{T}{\sigma_\eta^2 + T\sigma_\theta^2}E(\bar{\nu_i}|\ x_{it}, U_i)\right\} = \sigma_{j,\alpha\theta}C_i \\ E(\varepsilon_{j,it}|x_{it}, U_i) &= \sigma_{j,\varepsilon\eta}\left\{\sigma_\eta^{-2}E(\nu_{it}|\ x_{it}, U_i) - \frac{T\sigma_\theta^2}{\sigma_\eta^2(\sigma_\eta^2 + T\sigma_\theta^2)}E(\bar{\nu_i}|\ x_{it}, U_i)\right\} = \sigma_{j,\varepsilon\eta}C_{it} \end{aligned} \tag{8,9}$$

The endogeneity correction terms (C_i, C_{it}) defined in equations $(8, 9)$ above, are added as additional terms in the equation of primary interest to be estimated jointly with (β', δ) in the second step from conditional moment restrictions such as least squares based on equation (5).[7]

Under the null hypothesis of no endogeneity ($\sigma_{j,\alpha\theta} = \sigma_{j,\varepsilon\eta} = 0$), the conventional standard errors can be used. Otherwise, the standard errors should be adjusted for heteroskedasticity and the inclusion of the endogeneity correction terms (see Newey, 1984; Vella and Verbeek, 1999b).

4. ESTIMATION RESULTS

We employ British Household Panel Survey (BHPS) data and construct two balanced panels of employees aged between 16-65 during (1991-1996, 1997-2002). These consist of a full-time male employees' sample and a full sample of female employees (full-time and part-time). Part-time male employees are excluded since the small gains in terms of sample size are more than outweighted by the costs of a potential increase in the heterogeneity of the male samples. The female samples can provide a comparison group against the male sample that could potentially suffer from selectivity bias. Note that the former is also prone to sample selection bias caused by the labour market participation decision (see Swaffield, 2001, p.439).

We opted to split the panels into two distinct time periods in order to compare the end of the Conservative´s two decades of anti-union legislation and see whether anything actually changed under New Labour that came into office in 1997.

Descriptive statistics for the set of variables and the endogeneity correction terms are given in Tables (1, 7-10) respectively (see Appendix IV).

[7] The estimation procedure of the endogeneity correction terms is provided in Appendix III.

The unconditional male union wage differential is approximately 4.5% and 4% in the (1991-1996) and (1997-2002) samples, respectively, while, the corresponding unconditional female union wage differentials are approximately 14% and 11.9%. Further, the descriptive statistics Tables reveal that females earn approximately 20.6% less than their male counterparts during (1991-1996) while, they earn approximately 14.4% less during (1997-2002).

Male union membership during (1997-2002) has fallen by approximately 15.3% compared to the (1991-1996) male membership level. The female union membership rate on the other hand, has actually risen by approximately 9.5% so that by the last cross-section of the survey female rates converged to the level of male membership rates at a marginally higher percentage.

Note that, an individual is taken to be a trade union member if he/she has responded positively to the question "Are you currently a member of: Trade Unions" in the Social and Interest Group Membership section of the BHPS. Unfortunately, this question was only asked every other year after the fifth wave (1995-96) of the survey since the data depositors believe that there is not a lot of movement in and out of organisations and therefore, it was not felt necessary to ask this every year. A further union membership variable which is available in the BHPS, "Member of workplace union", can be used as a proxy for waves six (1996-97), eight (1998-99), ten (2000-01) and twelve (2002-03). This includes "in-house" staff associations, but excludes employers' organisations.

This introduces a degree of discontinuity in our measure of trade union membership, but, nevertheless when one wishes to undertake a longitudinal analysis using BHPS data this is the only alternative available.

Prior to embarking on the analysis of the estimated results some important issues need to be addressed. Regarding the issue of identification, the non-linear mapping from the reduced form union membership variables to the endogeneity correction terms identifies all parameters in the wage equation (5).

The exclusion of lagged union membership status from the empirical counterpart of ($eq.5$) identifies the equation as long as γ_1 differs from zero. Note that it is assumed that the long-term advantages of union employment, whilst generating persistence of union membership status, do not have a significant impact on wages.

It can be argued that, while the lagged value of union membership status affects an individuals' unionisation decision it does not have any significant effect on the current wage. This occurs in that union membership status may capture movement costs that are not specific to union employment. Workers are therefore assumed to change union membership status only if they change jobs. Furthermore, the long-term advantages of union employment, whilst generating persistence of union membership status are not expected to have a significant impact on wages and therefore lagged union membership status is expected to have a minor effect on current wages (see Vella and Verbeek, 1998, p.167).

Note that we exclude wages from the set of explanatory variables in the union membership models. Wages can be viewed as a collective good by the individual therefore casting doubts on whether they can be considered as a determinant of his/her unionisation probability (see Booth, 1986). This occurs in that, individuals can always free ride on union bargained wages that apply to all employees within a covered establishment irrespective of their union membership status.

Wages can be included in the union membership models to detect whether individuals do free-ride and an insignificant coefficient can then validate the public goods theory of union wages given that earnings are not proxying any omitted variables. Nevertheless, even if wages and union status are not determined simultaneously wages might be acting as a proxy for omitted variables that are simultaneous such as job security and pension provisions and this would produce biased results. Since wages can be either a complement or a substitute of union negotiated non-pecuniary benefits the direction of the bias cannot be determined *a priori*, thus rendering the interpretation of the resulting coefficients on wages problematic (Booth, 1986, p.43).

Due to the statistical significance of the endogeneity correction terms (refer to Tables 3-6) the standard errors in the wage models should be adjusted for heteroskedasticity and the inclusion of the correction terms. We use nonparametric (pairs) Bootstrap by resampling observations with replacement to estimate the standard errors of the parameters in the reduced form models. Concerning the wage models we use wild Bootstrap for inference robust to heteroskedasticity of unknown form since it provides refinements for the linear regression model in the presence of heteroskedasticity (see Flachaire, 2005).[8,9]

Higher-order terms of the latent effects, squared correction terms and their interactions with union status, were included in the models in order to detect non-normality (see Pagan and Vella, 1989) and these are reported when normality is an issue.

Lastly but not least, it should be noted that all estimated models reported in Tables (2-6) exclude occupational controls. Occupational status, being a measure of ability, could be potentially endogenous and can contaminate the conclusions concerning the role of unobserved individual heterogeneity (see Vella and Verbeek, 1998, pp.168-169).

Including occupational controls reduced the coefficients on the fixed individual effects in most estimated models suggesting that a component of unobserved heterogeneity underlying union status is correlated with occupational classification. Further, the inclusion of occupational controls was found to inflate the union wage differentials. We therefore opted to exclude occupational controls from the set of regressors.[10]

4.1. The Persistence of Trade Union Membership

The reduced form estimates presented in (Table 2) suggest that union membership remains persistent even after controlling for the unobserved effect.[11,12]

[8] In both estimating equations we use between 210-250 Bootstrap replications and these are generally adequate for standard error estimation (see Efron and Tibshirani, 1993).

[9] We use the Rademacher distribution in the wild Bootstrap DGP. Davidson and Flachaire (2001) demonstrate that it always produces better results than Mammen´s (1993) distribution. Both methodologies gave similar results. In the case of the (1991-1996) female estimates we use Mammen´s DGP as the Rademacher distribution gave inconclusive results in the case of the unrestricted estimates.

[10] While the reduced form models contain interactions between University and Vocational Qualifications and the Public Administration and Education industrial classification these were excluded from the primary wage equations as in many cases they exacerbated non-normality.

[11] With the exception of the second period female estimates whereby, lagged membership status is only significant at the not so stringent 10% significance level.

[12] The presence of x_i in expression (AI$_3$) implies that we are not able to identify the coefficients on time-constant explanatory variables in x_{it}. Including time-constant explanatory variables in x_{it} only increases the explanatory power as it is not possible to separately identify the partial effect of time-constant variables from their partial correlation with the unobserved effect (see Wooldridge, 2005, p.44).

Exogeneity of the initial conditions is rejected since there is a considerable correlation between the unobserved individual heterogeneity and the initial condition. This occurs in that, the coefficients on the initial value of union membership appear in all estimated models with particularly strong and statistically significant effects that are greater in magnitude as opposed to the coefficients on the lagged value of membership status.

According to the predicted probability ratios (see Table 2) a male worker with a given set of observable and unobservable attributes is $(2.9, 1.6)$ times as likely to remain a union member during (1991-1996, 1997-2002) respectively. Similarly, a female worker is approximately $(1.8, 1.2)$ times as likely to remain a union member in the current year if she had been so in the previous year.[13]

While the fall in persistence in the male samples was around 45.2% the corresponding female decline was approximately 32%. This occurs in that though the male union membership rate declined throughout the period concerned (refer to Table 1) the respective female rate, followed an increasing path.

Conclusively then, the extent of state dependence in union membership status declined significantly during the period under analysis. This is in line with our expectations following the introduction of a series of anti-union legislation (1980, 1982, 1988, 1990 Employment Acts) and its resulting impact.

Furthermore, the role of habit persistence (i.e. the behavioural effect of remaining in unions due to experiencing unionisation in the past) diminished relative to the impact of unobserved heterogeneity since the proportion of the total error variation attributed to unobserved individual heterogeneity, ρ, is markedly higher in the second period estimates (see Table 2).

4.2. The Wage Regressions Under Hierarchical Sorting

To determine the economic sorting structure consistent with the data we begin with the hierarchical sorting estimates that restrict the endogeneity correction terms to be invariant to sector.

In all estimated models under hierarchical sorting the selection terms (individual-specific and time-variant effects) were found to be jointly statistically significant and this is suggestive of selectivity bias. Therefore, union membership and wages are determined simultaneously and ignoring this gives biased estimates of the union wage effect.

With regard to the (1991-1996) male estimates under hierarchical sorting, the negative and statistically significant coefficient on the individual-specific correction term suggests that employees who receive lower wages, upon conditioning on their attributes and in the absence of trade unions, are those more likely to be union members (refer to Table 3). Concerning the female estimates, the statistically significant and negative coefficients on both correction terms in (1997-2002) and the time-variant correction term in the (1991-1996) estimates (Tables 5, 6, respectively) are also indicating that lower paid individuals have a higher propensity towards union membership *ceteris paribus*.

However, concerning the (1997-2002) male estimates (Table 4) the positive and statistically significant coefficients on both correction terms, imply that male employees who receive higher wages, controlling for their characteristics and in the

[13] Note that the degree of state dependence in male union membership appears to be more pronounced regarding both periods under analysis. This is not surprising given that the male samples employed solely consist of full-time employees.

absence of unions, are those more likely to be trade union members.

The statistically significant coefficients on the time-variant individual effects in the female estimates for both periods (Tables 5, 6) and the (1997-2002) male estimates (Table 4) indicate that, fixed effects estimation is inappropriate as the time varying endogeneity is not eliminated and continues to contaminate the resulting estimates. Note that even in the case whereby unobserved heterogeneity is individual-specific and time-invariant the fixed effects estimator imposes the invariance of heterogeneity rewards to sector and therefore, it is fairly restrictive in its treatment of unobserved heterogeneity (see Vella and Verbeek, 1998).

In all reported models the estimated union effect under fixed effects assumptions is notably low. With the exception of the second period male estimates, whereby under hierarchical sorting we get a statistically insignificant wage differential, fixed effects estimates provide the lower bound of the union wage effect as opposed to the *uncorrected* OLS and the endogeneity corrected estimates. This is consistent with the general consensus in the union literature that the longitudinal differencing estimates of the union wage effect produce a downward adjustment (e.g. Freeman, 1984; Robinson, 1989a; Jakubson, 1991).

The outcome that higher paid employees are less likely to seek union membership (apart from the second period male estimates) is in line with insider-outsider theories (Lindbeck and Snower, 1986) suggesting that groups of highly skilled employees can be seen as acting as a *de facto* union on its own since they cannot be rapidly and costlessly replaced (see Blanchflower el al, 1990). It is also consistent with the argument that the standardisation of wage rates, via the bargaining process, implies that workers with high degrees of human capital would be less prone to union membership as it would entail reduced human capital premia (see Abowd and Farber, 1982).

The finding concerning the (1997-2002) male estimates could be attributed to highly skilled male employees within the heavily unionised public sector which are more likely to be union members (see Chrysanthou, 2008; Blanchflower and Bryson, 2010).

However, the presence of the higher-order endogeneity correction terms due to non-normality has the implication that unobserved heterogeneity impacts negatively on the wages of union members. The dominant effect stems from the negative coefficient on the higher-order time-invariant correction term indicating diminishing returns to unobserved heterogeneity. It is plausible that the standardisation of wages in the union sector results in reduced human capital premia for the best employees in terms of their endowment of unobserved productivity (see Abowd and Farber, 1982).

The reduced form estimates (Table 2) reveal that during (1997-2002) the role of habit persistence diminished relative to the impact of unobserved heterogeneity. Therefore, regarding the male estimates it seems plausible that the higher proportion of highly skilled employees remaining in unions during the second period do so due to unobserved determinants such us solidarity, age of establishment and several other factors that we fail to account for as opposed to maximising their earnings.

While on the basis of individual behaviour it seems reasonable to argue that less productive employees might be more prone towards union membership in order to appropriate some share of monopoly rents, it appears improbable that profit maximising employers would hire them (see Abowd and Farber, 1982; Vella and Verbeek, 1998, p.173). Since in modelling wage determination and the union membership decision we do not explicitly incorporate the employers' role, at a minimum we ought to investigate the impact on the estimates upon the relaxation of the assumption that unobserved heterogeneity is equally valued in each sector.

4.3. The Wage Regressions Under Unrestricted Sorting

To test whether the random components are differentially rewarded in the two sectors we modify the models by allowing the endogeneity correction terms to vary by sector.

To investigate how unobserved individual heterogeneity is rewarded in each sector more rigorously we need to examine the average wage contribution of the endogeneity correction terms.

The male unrestricted sorting estimates during (1991-1996: Table 3) are consistent with a comparative advantage sorting specification. The negative coefficient on the individual-specific effect indicates that employees that are the recipients of lower wages, upon conditioning on their attributes and in the absence of trade unions, are more likely to be union members. The positive coefficient on the interaction between the fixed effect and union membership, however, implies that such individuals perform relatively better in the union sector. In other words, employees perform differently within the two sectors and sort themselves appropriately so that the average contribution of unobserved heterogeneity to the wages of union and non-union members is positive (see Table 7). Therefore, those with a relatively higher propensity towards union membership benefit most from sorting themselves into the union sector.

Regarding the (1997-2002) male estimates (see Table 4) all endogeneity correction terms are statistically insignificant. Thus, we reject the unrestricted sorting structure estimates in favour of the hierarchical sorting estimates.

Finally, considering the female estimates for both periods (Tables 5,6) given the statistical insignificance of the fixed individual effect the dominant effect determining the sorting structure consistent with the estimates stems from the time-variant effects. The average contribution of the time-variant effects to the wages of union members is negative whereas the corresponding contribution to the wages of non-union members is positive. Therefore unobserved heterogeneity impacts negatively on the wages of union members, irrespective of the sector they sort themselves into. Hence, we cannot reject the hierarchical sorting specification.

While the restriction that the random components are equally rewarded in the two sectors is only rejected in the (1991-1996) male estimates we do not believe that estimation of the remaining models by Instrumental Variables or restricted Control Function estimators is appropriate. Primarily since the structure of sorting into the union and non-union sectors cannot be known *a priori* we see no reason why one should preclude a comparative advantage sorting structure by employing these estimators.

Further, one should bear in mind that solely a degenerate hierarchical structure, imposing perfect correlation between sector-specific skills, can meet the strict and restrictive requirement imposed by Instrumental Variables and variants of the restricted Control Function estimators (see Vella and Verbeek, 1999a).

4.4. The Estimated Union Wage Effects

According to the economic sorting structure supported by the estimates UK trade unions were unable to establish a union wage effect for their male members.

However, regarding the female portion of the labour market the union wage effect is estimated to be approximately (19.4% , 17.6%) during (1991-1996, 1997-2002) respectively (refer to Tables 5, 6). The significant contribution of unobserved individual heterogeneity renders the total union wage differential highly variable across individuals.

Note that, it is likely that the paucity of employer controls might be biasing upwards the union wage impact since unionised establishments tend to pay more for reasons not directly linked to union membership (see Blanchflower and Bryson, 2010).

The outcome that unions were unable to establish a union wage differential for their male members is not surprising given the successive legislative changes targeted towards weakening the bargaining strength of UK trade unions (for a detailed account refer to Stewart, 1995). The overall effect of the (1980, 1982, 1988 and 1990) Employment Acts was that the ensuing period was an era of a dramatic decline in aggregate trade union membership and union recognition in the UK.

Nevertheless, the estimated female union wage differentials imply that unions still play a non-negligible role in wage formation even though we observe a decline of approximately 9.3% in the female union wage effect during the second period under analysis.

The outcome that female employees still receive substantial union wage differentials is consistent with the expectation that unions raise the wages of those who would have been relatively lower paid in their absence. In fact, the descriptive statistics (in Table 1) reveal that females earn approximately 20.6% and 14.4% less than their male counterparts during (1991-1996, 1997-2002), respectively.

Further, provided that unionism is assumed to be a normal good an employee's demand for unionism is positively related to his/her union wage differential (see Chang and Lai, 1997, p.121). Therefore lower paid employees, in the absence of unions, for discriminatory reasons are more likely to be prone towards union membership (see for instance Heywood, 1990). This is in line with the fact that, in the period under analysis, while male union membership has been declining female union membership has been rising (refer to Table 1).

The estimated union wage effects are comparable to the recent estimates of Blanchflower and Bryson (2010) using data from the Labour Force Survey (LFS). We opted to compare the results of this study to the estimates of Blanchflower and Bryson (2010) in that, they use a similar time period and UK data while they do not account for the endogeneity of union status.

The corresponding *uncorrected* OLS union wage differentials of Blanchflower and Bryson (2010), using personal and industrial controls, are approximately (6.5%, 4%) for males and (19.1%, 13.8%) for females during (1993-1999) and (2000-2006), respectively (see Blanchflower and Bryson, 2010, p.101).

One could argue that the evident continuous decline in UK union wage differentials explains the relatively higher endogeneity corected estimates of the female union wage effects (19.4%: 1991-1996, 17.6%: 1997-2002) from the present study that uses data from an earlier period. Therefore, we should not be particularly preoccupied with the endogeneity of union membership status.

However, while in the case of the female union wage effects the results from this study are comparable to the estimates of Blanchflower and Bryson (2010) the discrepancy in the male union wage impact highlights the necessity to control for the endogeneity of union membership status.

OLS differentials denote the difference between the average wage of those that are actually in the union sector and the average of those employees that are located in the non-union sector *ceteris paribus.* Endogeneity corrected differentials on the other hand come from a random assignment to sector (see Robinson, 1989a, p.660).[14]

According to the *orthodox* view profit maximising establishments within the unionised sector will select the best employees thus inflating the OLS estimate of the union wage differential (e.g. Abowd and Farber, 1982; Freeman, 1984). However, this will only result in an upward bias for OLS under a hierarchical sorting structure of omitted ability when the average contribution of unobserved individual heterogeneity to the wages of union members is positive as is the case with the (1997-2002) male estimates in this study.

Under hierarchical sorting, the average contribution of unobserved heterogeneity to the wages of male union members during (1991-1996) and female members during both periods is negative so that they receive lower wages irrespective of the sector they are located in. Therefore, this reverses the direction of the OLS bias from positive to negative.

Recall that, the union wage impact under fixed effects provides the lower bound among the statistically significant estimated union wage effects. Conclusively then, the endogeneity correction procedure employed yields a discernible pattern of estimates of the union wage differential relative to OLS and fixed effects. Hence, it does contribute to our understanding of the union wage effect *puzzle*. This conflicts with Freeman and Medoff (1982) and Lewis (1986) and is in line with Robinson (1989a).

5. SUMMARY AND CONCLUSIONS

Using data from the BHPS (1991-2003) we employ a dynamic model of union membership status and wage determination to estimate the wage impact of UK trade unions. The findings reveal that the unobserved factors that influence union membership also affect the wage impact of unions. Therefore, we cannot fail to account for the endogeneity of union membership status.

Though trade union membership remains persistent even after controlling for the unobserved effect, we observe a significant decline in union membership persistence during the period under analysis.

While UK trade unions were unable to establish a union wage effect for their male members, the female union wage differential is estimated to be approximately (19.4% , 17.6%) during (1991-1996, 1997-2002) respectively. The significant contribution of unobserved individual heterogeneity renders the total union wage differential highly variable across individuals.

The outcome that unions were unable to establish a union wage differential for their male members is in agreement with our expectations following the successive (1980, 1982, 1988 and 1990) Employment Acts targeted towards weakening the bargaining strength of UK trade unions.

However, the estimated female union wage differentials imply that unions still play a non-negligible role in wage formation though we observe a decline of approximately 9.3% in the estimated wage effect during the second period under analysis. The result that female employees still receive substantial union wage differentials

[14] Note that there are no individuals who are randomly assigned. Random assignment refers to controlling for the unobservables underlying the unionisation decision which are correlated with the wage rate (see Vella and Verbeek, 1999a, p.474).

is consistent with the expectation that unions raise the wages of those who would have been relatively lower paid in their absence.

While our estimated female wage differentials are comparable to the recent *uncorrected* estimates of Blanchflower and Bryson (2010), the results concerning the male union wage effects highlight the necessity to account for the endogeneity of union membership.

Employees receiving lower wages, upon conditioning on their attributes and in the absence of trade unions, are those more likely to be union members (excluding the 1997-2002 male estimates). This is in accordance with insider-outsider theories (Lindbeck and Snower, 1986) indicating that groups of highly qualified employees do not need to rely on unions to gain higher wages since they can be valuable assets to employers (see Blanchflower *el al*, 1990). It is also consistent with the argument that the standardisation of wage rates through the bargaining process suggests that those with higher levels of human capital would be less likely to join unions as it would entail reduced human capital premia (Abowd and Farber, 1982).

The models for union status determination show that the role of habit persistence diminished relative to the impact of unobserved heterogeneity during the period under analysis. Therefore, regarding the male estimates it seems plausible that the higher proportion of highly skilled employees remaining in the union sector in the second period, do so due to unobserved determinants such us solidarity, age of establishment and several other factors that we fail to account for as opposed to maximising their earnings.

We conclude that the endogeneity correction procedure employed yields a discernible pattern relative to OLS and fixed effects and therefore contributes to our understanding of the union wage differential *puzzle*. This refutes the pessimistic conclusions reached by Freeman and Medoff (1982) and Lewis (1986) and is in agreement with Robinson (1989a).

The obvious future research direction is to use the new cross sections of the BHPS in order to estimate the union wage impact during (2003-2009). Further, given the outcome regarding the second period male estimates it could be worth exploring the wage impact of UK trade unions in the public and private sectors using matched employer-employee data.

6. APPENDIX I: HANDLING THE INITIAL CONDITIONS PROBLEM

Union membership decision is modelled using the dynamic Random effects Probit specification given in equation (AI_1):

$$U_{it}^{*} = \gamma_1' x_{it} + \gamma_2 U_{i,t-1} + \vartheta_i + \eta_{it};\ \eta_{it} \sim N(0, \sigma_\eta^2) \qquad (AI_1)$$

Adopting the Mundlak (1978)-Chamberlain (1984) specification we can allow for a correlation between ϑ_i and the time means of the observed time-varying characteristics taking the form of $\vartheta_i = \overline{x}_i' \varkappa + \kappa_i$. Substituting this expression for ϑ_i in equation (AI_1) we arrive at specification (AI_2) where it is assumed that $\kappa_i \sim iidN(0, \sigma_\kappa^2)$ and is independent of (x_{it}, η_{it}) for all i and t:

$$U_{it}^{*} = \gamma_1' x_{it} + \gamma_2 U_{i,t-1} + \overline{x}_i' \varkappa + \kappa_i + \eta_{it};\ \eta_{it} \sim N(0, \sigma_\eta^2) \qquad (AI_2)$$

We employ Wooldridge's (2005) solution to the initial conditions problem. This specifies a distribution of unobserved individual heterogeneity conditional on the initial condition instead of obtaining the joint distribution of all outcomes of the endogenous variables.

We begin by specifying the distribution of the unobserved effect as:

$$\vartheta_i | U_{i1}, x_i \sim N(\kappa_0 + \kappa_1 U_{i1} + x_i' \kappa_2, \sigma_\kappa^2);\ x_i = \{x_{i1}, ..., x_{iT}\} \qquad (AI_3)$$

where the $(1xT)$ row vector x_i contains all non-redundant explanatory variables in all periods under consideration.

The density $D(U_{i1}, ..., U_{iT} | U_{i1} = U_1, x_i = x, \kappa_i = \kappa)$ is given by:

$$\prod_{t=1}^{T} \{\Phi(\gamma_1' x_t + \gamma_2 U_{t-1} + \kappa_0 + \kappa_1 U_{i1} + x' \kappa_2 + \kappa)^{y_t} . [1 - \Phi(\gamma_1' x_t + \gamma_2 U_{t-1} + \kappa_0 + \kappa_1 U_{i1} + x' \kappa_2 + \kappa)]^{1-y_t}\} \qquad (AI_4)$$

To find the joint distribution of $(U_{i2}, ..., U_{iT} | U_{i1} = U_1, x_i = x)$ we need to integrate out κ_i. Integrating (AI_4) against the Normal (0, σ_κ^2) gives the likelihood function in expression (AI_5) which is identical to the structure of the standard Random effects Probit model with the only difference that the explanatory variables at time t are $\{z_{it} \equiv (1, x_{it}, U_{i,t-1},\ U_{i1},\ x_i)\}$:

$$\int_{\mathbb{R}} \prod_{t=1}^{T} \{\Phi(\gamma_1' x_t + \gamma_2 U_{t-1} + \kappa_0 + \kappa_1 U_{i1} + x' \kappa_2 + \kappa)^{y_t} . \qquad (AI_5)$$
$$[1 - \Phi(\gamma_1' x_t + \gamma_2 U_{t-1} + \kappa_0 + \kappa_1 U_{i1} + x' \kappa_2 + \kappa)]^{1-y_t}\} (1/\sigma_\kappa) \phi(\kappa/\sigma_\kappa) d\kappa$$

To obtain the magnitudes of state dependence the average partial, marginal, effects $(\hat{p}_j - \hat{p}_0)$ and predicted probability ratios $(\hat{p}_j / \hat{p}_0)$ were estimated using the following counter-factual probabilities that take U_{t-1} to be fixed at 0 and 1 and are evaluated at $x_{it} = \overline{x}$ (see Stewart, 2007):

$$\begin{aligned} \hat{p}_j &= \frac{1}{N} \sum_{i=1}^{N} \Phi\{(\hat{\gamma}_1' \overline{x} + \hat{\gamma}_2 + \hat{\kappa}_1 U_{i1} + \overline{x}_i' \hat{\kappa}_2) \sqrt{1 - \hat{\rho}}\};\ \hat{p}_0 = \frac{1}{N} \sum_{i=1}^{N} \Phi\{(\hat{\gamma}_1' \overline{x} + \hat{\kappa}_1 U_{i1} + \overline{x}_i' \hat{\kappa}_2) \sqrt{1 - \hat{\rho}}\} \\ \rho &= cor(v_{it}, v_{is}) = \frac{\sigma_{\kappa_i}^2}{\sigma_{\kappa_i}^2 + 1};\ t,\ s = 2, ..., T;\ t \neq s \end{aligned} \qquad (AI_6)$$

7. APPENDIX II: DERIVATION OF THE ENDOGENEITY CORRECTION TERMS

Following Vella and Verbeek (1998, 1999b) we use the assumption of joint normality to derive the expectation of $e_{j,it}$ in $(eq.5)$ conditional on the vector ν_i and the vector of exogenous variables x_{it}. Employing the standard formulae for the conditional expectation of normally distributed vectors:

$$\begin{aligned} E(\alpha_{j,i}|x_{it},\nu_i) &= \sigma_{j,\alpha\theta}\left\{\frac{T}{\sigma_\eta^2+T\sigma_\theta^2}\bar{\nu_i}\right\} && (\mathrm{AII}_1) \\ E(\varepsilon_{j,it}|x_{it},\nu_i) &= \sigma_{j,\varepsilon\eta}\left\{\sigma_\eta^{-2}\nu_{it}-\frac{T\sigma_\theta^2}{\sigma_\eta^2(\sigma_\eta^2+T\sigma_\theta^2)}\bar{\nu_i}\right\} && (\mathrm{AII}_2) \end{aligned}$$

To obtain the conditional expectations, given the t-dimensional vector U_i, we replace ν_i in $(eq.\mathrm{AII}_1,\mathrm{AII}_2)$ by their conditional expectations given U_i. Using the first law of iterated expectations:

$$\begin{aligned} E(\nu_{it}|x_{it},U_i) &= E(\vartheta_i+\eta_{it}|x_{it},U_i)=E_{\vartheta_i}[\vartheta_i+E\{\eta_{it}|x_{it},U_i,\vartheta_i\}] && (\mathrm{AII}_3) \\ &= \int_{-\infty}^{\infty}[\vartheta_i+E(\eta_{it}|x_{it},U_i,\vartheta_i)]f(\vartheta_i|x_{it},U_i)d\vartheta_i \end{aligned}$$

where $f(\vartheta_i|x_{it},U_i)$ is the conditional density of ϑ_i.

Given the assumption of the strict exogeneity of x_{it} the conditional distribution of ϑ_i is:

$$f(\vartheta_i|x_{it},U_i)=\frac{f(U_i|x_{it},\vartheta_i)f(\vartheta_i)}{f(U_i|x_{it})} \qquad (\mathrm{AII}_4)$$

Substituting expression AII_4 into AII_3 we arrive at:

$$\begin{aligned} E(\nu_{it}|x_{it},U_i) &= \frac{1}{f(U_i|x_{it})}\int_{-\infty}^{\infty}[\vartheta_i+E(\eta_{it}|x_{it},U_i,\vartheta_i)]f(U_i|x_{it},\vartheta_i)f(\vartheta_i)d\vartheta_i = && (\mathrm{AII}_5) \\ &= \frac{1}{\int f(U_i|x_{it},\vartheta_i)f(\vartheta_i,x_{it})d\vartheta_i}\int_{-\infty}^{\infty}[\vartheta_i+E(\eta_{it}|x_{it},U_i,\vartheta_i)]f(U_i|x_{it},\vartheta_i)f(\vartheta_i)d\vartheta_i \end{aligned}$$

where we have used $f(U_i|x_{it})=\int f(U_i|x_{it},\vartheta_i)f(\vartheta_i,x_{it})d\vartheta_i$ and $E\{\eta_{it}|x_{it},U_i,\vartheta_i\}$ denotes the cross-sectional generalised Probit residual (see Gourieroux *et al*, 1987) obtained from the first step estimates of the reduced form model:

$$E(\eta_{it}|x_{it},U_i,\vartheta_i)=(2U_i-1)\sigma_\eta\left\{\frac{\phi\{(2U_i-1)(\gamma'\Psi_{it}+\theta_i)/\sigma_\eta\}}{\Phi\{(2U_i-1)(\gamma'\Psi_{it}+\theta_i)/\sigma_\eta\}}\right\} \qquad (\mathrm{AII}_6)$$

8. APPENDIX III: ESTIMATION OF THE ENDOGENEITY CORRECTION TERMS

The term in the denominator in expression AII_5 is the likelihood contribution for individual i. Given the parameter estimates from the reduced form model $\delta_1 = (\gamma_1, \gamma_2, \sigma_{\vartheta_i})$ we can approximate expression AII_5 using quadrature methods or simulation (numerical integration).

To estimate the simulated counterpart of AII_5 we obtain R draws of ϑ_i^r from $f(\vartheta_i|\sigma_{\vartheta_i}) = N(0, \sigma_{\vartheta_i})$ and compute the corresponding log-likelihood for individual i conditional on the draw:

$$f(U_i|x_{it}, \vartheta_i^r) = \prod_{t=1}^{T} f(U_{it}|x_{it}, \vartheta_i^r) \tag{AIII$_1$}$$

To provide a better coverage of the integrals we use randomised Halton draws since the asymptotic properties of simulation-based estimators are obtained under the assumption of randomness.[15] Supplementing these with antithetic draws we induce a negative correlation over observations thus, further improving the coverage (see Train, 2003; Cappellari and Jenkins, 2006). The procedure is repeated R times and averaging over these replications we obtain the simulated log-likelihood function:

$$f(U_i|x_{it}, \vartheta_i^r) = \ln \frac{1}{R} \sum_{r}^{R} f(U_i|x_i, \vartheta_i^r) \tag{AIII$_2$}$$

Estimators obtained by maximising likelihoods that are approximated by simulation techniques are known as MSL estimators and have the same large-sample properties as maximum likelihood, if the number of repetitions used to approximate the integral grows at a higher rate than the square root of the number of observations in the sample (see Drukker, 2006, p.153). We use 500 randomised Halton draws and 500 antithetic draws, $R = 1000$, and given the sample sizes $\sqrt{N}/R \longrightarrow 0$.

Given the estimates from the reduced form model we can simulate the expression for $E(\nu_{it}|x_{it}, U_i)$.

Taking again R draws from $f(\vartheta_i|\sigma_{\vartheta_i})$ expression AII_5 is approximated by:

$$\tilde{\nu_{it}} = \frac{1}{\frac{1}{R}\sum_{r}^{R} f(U_i|x_{it}, \vartheta_i^r)} \frac{1}{R} \sum_{r}^{R} [\vartheta_i^r + E(\eta_{it}|x_{it}, U_i, \vartheta_i^r)] f(U_i|x_{it}, \vartheta_i^r) \tag{AIII$_3$}$$

The individual specific means $E(\bar{\nu_i}|\ x_{it}, U_i)$ can be computed using: $\bar{\nu_i} = \frac{1}{T_i} \sum_{t=1}^{Ti} \tilde{\nu_{it}}$. Substituting the estimates for $E(\nu_{it}|x_{it}, U_i)$ and $E(\bar{\nu_i}|\ x_{it}, U_i)$ into $(eq.8, 9)$ we obtain the endogeneity correction terms (C_i, C_{it}).

[15] Randomised Halton draws have identical properties of coverage over observations as Hatlon draws, however, they are not systematic at least in the same manner pseudorandom numbers are random (see Train, 2003, pp.234-235).

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9. APPENDIX IV: TABLES

TABLE 1

Descriptive Statistics

	1991-1996				**1997-2002**			
Gender	**Male**		**Female**		**Male**		**Female**	
Variable	**Mean**	**Std. Dev.**	**Mean**	**Std.Dev.**	**Mean**	**Std. Dev.**	**Mean**	**Std.Dev.**
Log of Hourly Wage	**2.073**	**0.464**	**1.746**	**0.472**	**2.264**	**0.466**	**1.995**	**0.461**
Trade Union Membership	**0.4099**	**0.492**	**0.3173**	**0.465**	**0.3471**	**0.476**	**0.3474**	**0.476**
Log (1+Potential Experience)	3.223	0.634	3.166	0.715	3.558	0.456	3.544	0.467
Marital Status	0.692	0.462	0.683	0.465	0.665	0.472	0.642	0.479
Full-Time Employment	_	_	0.672	0.469	_	_	0.709	0.454
Maternity Leave	_	_	0.011	0.104	_	_	0.013	0.113
Black (Caribbean, African, Other)	0.001	0.035	0.005	0.068	0.002	0.047	0.006	0.079
Asian (Indian, Pakistani, Chinese, Other)	0.014	0.116	0.008	0.090	0.013	0.113	0.013	0.111
Other Ethnic Minority Group	0.009	0.093	0.005	0.068	0.007	0.080	0.005	0.072
Inner/ Outer London and R of South East	0.300	0.458	0.300	0.458	0.269	0.444	0.282	0.450
South West	0.097	0.296	0.070	0.255	0.094	0.291	0.095	0.293
Midlands	0.167	0.373	0.158	0.365	0.192	0.394	0.172	0.377
Scotland	0.078	0.269	0.097	0.296	0.078	0.268	0.089	0.285
Wales	0.053	0.224	0.047	0.211	0.056	0.230	0.050	0.219
North West	0.110	0.313	0.112	0.316	0.102	0.303	0.111	0.315
North East	0.150	0.358	0.175	0.380	0.170	0.376	0.160	0.366
East Anglia	0.044	0.205	0.041	0.198	0.040	0.195	0.041	0.198
University Degree or Higher	0.158	0.365	0.109	0.312	0.169	0.375	0.133	0.340
HND, HNC, Teaching	0.082	0.274	0.074	0.263	0.087	0.281	0.078	0.268
A Levels	0.239	0.426	0.155	0.362	0.267	0.442	0.225	0.418
O Levels or CSE	0.333	0.471	0.426	0.495	0.345	0.475	0.409	0.492
Fair, Poor, V Poor Self-Assessed Health	0.157	0.364	0.199	0.399	0.181	0.385	0.206	0.404
Workforce >500	0.213	0.410	0.156	0.363	0.206	0.405	0.173	0.379
Workforce 100-499	0.303	0.460	0.228	0.419	0.297	0.457	0.213	0.409
Workforce 25-99	0.271	0.445	0.275	0.447	0.248	0.432	0.271	0.445
Workforce <25	0.213	0.409	0.340	0.474	0.249	0.432	0.343	0.475
Industrial Classification Dummies								
Agriculture, Forestry & Fishing	0.016	0.125	0.006	0.080	0.006	0.078	0.005	0.067
Energy and Water Supplies	0.051	0.220	0.007	0.085	0.025	0.155	0.007	0.083
Extraction of Minerals & Manufacture of Metals	0.052	0.221	0.018	0.134	0.053	0.225	0.014	0.116
Metal Goods, Engineering & Vehicles Industries	0.156	0.363	0.050	0.219	0.138	0.345	0.030	0.171
Other Manufacturing Industries	0.122	0.328	0.067	0.250	0.151	0.359	0.062	0.241
Construction	0.044	0.204	0.006	0.076	0.055	0.227	0.006	0.079
Distribution, Hotels & Catering (Repairs)	0.116	0.320	0.177	0.382	0.122	0.327	0.185	0.388
Transport & Communication	0.091	0.288	0.030	0.171	0.105	0.307	0.036	0.187
Banking & Finance	0.134	0.341	0.153	0.360	0.143	0.350	0.154	0.361
Public Administration, Education, Other	0.219	0.414	0.484	0.500	0.202	0.401	0.502	0.500
Occupational Classification Dummies								
Professional Occupations	0.124	0.330	0.107	0.310	0.098	0.297	0.106	0.308
Managers & Administrators	0.180	0.384	0.097	0.296	0.196	0.397	0.120	0.325
Associate Professional & Technical	0.108	0.311	0.124	0.329	0.110	0.312	0.141	0.349
Clerical & Secretarial	0.099	0.298	0.333	0.471	0.081	0.273	0.302	0.459
Craft & related	0.182	0.386	0.026	0.161	0.202	0.401	0.021	0.143
Personal & Protective Service	0.069	0.253	0.117	0.322	0.059	0.235	0.131	0.338
Sales	0.033	0.179	0.076	0.266	0.035	0.183	0.090	0.287
Plant & Machine Operatives	0.152	0.359	0.037	0.188	0.158	0.365	0.030	0.170
Other Occupations	0.054	0.227	0.081	0.273	0.063	0.243	0.059	0.235
Number of Observations	**4818**		**5172**		**5538**		**5760**	

Source: BHPS (1991-2003), ISER, Essex, SN:4967, June 2004

TABLE 2

Dynamic Random Effects Probit Models of Union Membership

	Male 1991-1996		Male 1997-2002		Female 1991-1996		Female 1997-2002	
Variable	**Coef.**	**z**	**Coef.**	**z**	**Coef.**	**z**	**Coef.**	**z**
Lagged Trade Union Membership	**1.386**	**8.21**	**0.568**	**2.60**	**0.940**	**6.24**	**0.298**	**1.78**
Trade Union Membership (1)	**1.877**	**5.83**	**3.676**	**3.62**	**1.932**	**6.03**	**3.234**	**5.97**
Log(1+Potential Experience)	0.027	0.32	0.045	0.20	0.098	1.22	0.024	0.10
Marital Status	-0.026	-0.09	0.261	0.61	-0.168	-0.77	0.268	0.74
Mean(Marital Status)	0.140	0.47	-0.244	-0.54	0.555	2.27	0.350	0.82
Full-Time Employment	_	_	_	_	0.109	0.58	0.038	0.17
Mean(Full-Time Employment)	_	_	_	_	0.268	1.25	0.381	1.05
Maternity Leave	_	_	_	_	-0.034	-0.08	-0.147	-0.33
Mean(Maternity Leave)	_	_	_	_	1.280	1.20	1.183	0.60
Asian(Indian, Pakistani, Chinese, Other)	-0.222	-0.76	1.102	1.49	0.715	1.87	-0.052	-0.08
Black(Caribbean, African, Other)	-0.281	-0.04	-0.926	-0.19	-0.545	-0.84	1.694	1.33
Other Ethnic Minority Group	0.191	0.29	-1.493	-1.45	-0.658	-0.34	1.425	1.24
Inner/ Outer London and R of South East	-0.282	-2.26	-0.324	-1.05	-0.202	-1.27	-0.421	-1.33
South West	-0.072	-0.42	0.368	0.99	-0.034	-0.18	-0.430	-1.05
Scotland	-0.394	-2.22	0.152	0.42	0.320	1.53	0.385	0.94
Wales	0.340	2.01	0.741	1.70	0.333	1.47	0.690	1.64
North West	0.010	0.06	0.294	0.85	-0.153	-0.83	0.057	0.14
North East	-0.196	-1.28	-0.043	-0.14	0.316	1.79	0.339	0.91
East Anglia	-0.397	-1.74	0.057	0.12	0.007	0.03	-0.697	-1.40
Public, Educ, Other& University Qual/Higher	0.085	0.22	0.323	0.48	0.727	1.62	1.793	1.85
Public, Educ, Other& Vocational Qualifications	0.287	0.91	0.462	0.70	1.285	1.70	0.135	0.18
University Degree or Higher	-0.374	-1.49	-0.617	-1.49	-0.580	-1.55	-0.710	-0.79
HND, HNC, Teaching	-0.049	-0.23	-0.808	-1.65	-0.650	-1.03	0.680	1.04
A Levels	0.003	0.02	0.171	0.63	0.099	0.59	0.475	1.26
O Levels or CSE	0.069	0.57	0.204	0.70	0.024	0.21	0.300	0.84
Workforce >500	0.418	2.57	0.896	3.35	0.440	2.63	0.660	2.60
Workforce 100-499	0.295	1.95	0.973	3.88	0.348	2.39	0.547	2.39
Workforce 25-99	0.151	0.91	0.543	2.47	0.152	1.16	0.188	0.90
Agriculture, Forestry & Fishing	-0.965	-2.21	-2.554	-0.43	-8.660	-5.69	-1.170	-0.11
Extraction of Minerals & Manufacture of Metals	-0.235	-0.71	-1.305	-1.72	0.267	0.40	-0.734	-0.68
Metal Goods, Engineering & Vehicles Industries	-0.643	-2.69	-1.086	-1.53	-0.848	-1.27	-1.117	-1.06
Other Manufacturing Industries	-0.102	-0.40	-1.180	-1.57	-0.287	-0.47	-0.929	-0.92
Construction	-0.480	-1.50	-1.092	-1.33	0.647	0.20	-0.254	-0.10
Distribution, Hotels & Catering (Repairs)	-1.125	-3.87	-1.504	-2.00	-0.235	-0.40	-0.919	-0.98
Transport & Communication	-0.058	-0.22	-0.772	-0.97	0.082	0.12	-0.767	-0.76
Banking & Finance	-0.383	-1.32	-1.126	-1.45	0.144	0.24	-0.799	-0.85
Public Administration, Education, Other	0.291	1.07	-0.272	-0.35	0.217	0.36	-0.156	-0.17
Constant	-1.803	-4.17	-2.157	-1.63	-3.182	-4.37	-2.787	-2.08
ρ	**0.395**	**8.63**	**0.671**	**8.91**	**0.498**	**10.32**	**0.743**	**18.31**
Average Partial Effect	**0.396**		**0.088**		**0.259**		**0.046**	
Predicted Probability Ratio	**2.909**		**1.595**		**1.792**		**1.218**	
Log-Likelihood	**-1050.74**		**-1079.59**		**-1325.87**		**-1475.37**	

Notes

1. Standard error estimates provided above obtained using (211, 210, 242, 217) "Nonparametric Bootstrap" replications respectively
2. Time dummies included in all models

Source: BHPS (1991-2003), ISER, Essex, SN:4967, June 2004

TABLE 3

Wage Regressions (1991-1996), Males

Variable	OLS Coef.	t	Fixed Effects Coef.	t	Hierarchical Coef.	z	Unrestricted Coef.	z
Trade Union Membership	**0.042**	**3.48**	**0.038**	**2.89**	**0.086**	**4.96**	**0.004**	**0.25**
Potential Labour Market Experience	0.012	6.79	0.066	15.54	0.008	4.26	0.008	4.30
Squared Experience	0.000	-4.46	0.000	-3.95	0.000	-2.28	0.000	-2.69
Marital Status	0.175	14.40	0.005	0.27	0.155	12.02	0.160	12.38
Asian(Indian, Pakistani, Chinese, Other)	-0.020	-0.44	_	_	-0.014	-0.36	0.005	0.11
Black(Carribean, African, Other)	-0.187	-1.23	_	_	-0.137	-1.20	-0.158	-1.40
Other Ethnic Minority Group	-0.040	-0.68	_	_	0.050	0.86	0.055	0.98
Inner/ Outer London and R of South East	0.235	14.21	0.037	0.56	0.210	11.34	0.207	11.35
South West	0.104	4.80	0.100	1.19	0.018	0.74	0.040	1.65
Scotland	-0.013	-0.55	-0.083	-0.67	-0.042	-1.63	-0.030	-1.19
Wales	-0.006	-0.23	0.030	0.31	-0.045	-1.84	0.005	0.19
North West	0.081	3.87	-0.183	-2.13	0.049	2.16	0.078	3.36
North East	0.083	4.37	0.222	2.15	0.089	4.31	0.042	2.02
East Anglia	0.071	2.48	0.000	0	0.000	-0.01	0.008	0.24
University Degree or Higher	0.559	24.35	0.159	1.47	0.519	19.00	0.499	18.74
HND, HNC, Teaching	0.391	15.81	0.046	0.39	0.357	14.82	0.370	15.29
A Levels	0.273	15.04	0.203	2.42	0.260	12.13	0.261	11.36
O Levels or CSE	0.153	9.28	0.084	1.21	0.161	8.65	0.166	8.89
Workforce >500	0.178	10.26	0.031	1.84	0.181	9.67	0.148	7.90
Workforce 100-499	0.126	7.94	0.031	2.07	0.118	7.14	0.106	5.84
Workforce 25-99	0.088	5.60	0.024	1.72	0.073	4.23	0.062	3.39
Agriculture, Forestry & Fishing	-0.472	-9.34	-0.254	-3.23	-0.460	-8.85	-0.442	-8.71
Extraction of Minerals & Manufacture of Metals	-0.232	-6.92	-0.075	-1.5	-0.213	-6.31	-0.203	-6.48
Metal Goods, Engineering & Vehicles Industries	-0.266	-9.49	-0.108	-2.4	-0.243	-8.01	-0.224	-7.66
Other Manufacturing Industries	-0.279	-9.65	-0.087	-1.88	-0.241	-7.25	-0.259	-8.39
Construction	-0.251	-7.07	-0.131	-2.68	-0.254	-7.09	-0.182	-5.19
Distribution, Hotels & Catering (Repairs)	-0.372	-12.32	-0.191	-4.21	-0.329	-9.16	-0.382	-11.06
Transport & Communication	-0.318	-10.58	-0.147	-2.99	-0.287	-9.35	-0.303	-9.53
Banking & Finance	-0.056	-1.94	-0.155	-3.56	0.003	0.08	-0.016	-0.50
Public Administration, Education, Other	-0.195	-7.30	-0.129	-2.83	-0.196	-6.76	-0.147	-5.27
C_i	_	_	_	_	**-0.030**	**-2.75**	**-0.044**	**-2.74**
C_{it}	_	_	_	_	**-0.015**	**-1.03**	**-0.010**	**-0.56**
C_i.(Union Membership)	_	_	_	_	_	_	**0.056**	**2.54**
C_{it}.(Union Membership)	_	_	_	_	_	_	**0.076**	**2.64**
Constant	1.403	32.50	0.425	3.94	1.539	29.52	1.564	33.28
Adj R^2	0.375		_		0.355		0.347	

Notes

1. Hierarchical/ Unrestricted Standard Errors obtained using 250 "Wild Bootstrap" replications
2. All models, except Fixed effects, are inclusive of time dummies

Source: BHPS (1991-2003), ISER, Essex, SN:4967, June 2004

TABLE 4

Wage Regressions (1997-2002), Males

	OLS		Fixed Effects		Hierarchical		Unrestricted	
Variable	Coef.	t	Coef.	t	Coef.	z	Coef.	z
Trade Union Membership	**0.059**	**5.13**	**0.054**	**3.85**	**0.014**	**0.78**	**0.099**	**5.49**
Potential Labour Market Experience	0.014	6.09	0.069	11.70	0.008	3.55	0.009	3.71
Squared Experience	0.000	-7.10	0.000	-3.21	0.000	-4.78	0.000	-4.51
Marital Status	0.212	20.42	0.086	4.94	0.215	20.84	0.192	18.33
Asian(Indian, Pakistani, Chinese, Other)	-0.157	-3.62	_	_	-0.185	-3.52	-0.176	-3.14
Black(Carribean, African, Other)	0.155	1.48	_	_	0.020	0.16	0.041	0.36
Other Ethnic Minority Group	-0.025	-0.41	_	_	-0.128	-1.77	0.009	0.12
Inner/ Outer London and R of South East	0.171	11.72	0.094	1.24	0.165	10.11	0.151	9.16
South West	0.128	6.69	0.081	0.67	0.121	5.27	0.147	6.67
Scotland	0.023	1.13	-0.061	-0.37	-0.006	-0.29	0.023	0.94
Wales	-0.017	-0.72	0.087	0.90	-0.010	-0.50	-0.035	-1.64
North West	0.031	1.63	-0.078	-0.92	0.030	1.53	0.028	1.38
North East	0.005	0.33	0.037	0.58	0.012	0.62	-0.014	-0.75
East Anglia	0.027	1.02	0.064	0.61	0.010	0.40	0.020	0.72
University Degree or Higher	0.542	28.08	0.172	1.78	0.531	23.77	0.577	26.27
HND, HNC, Teaching	0.398	18.19	0.298	2.77	0.383	16.51	0.423	17.84
A Levels	0.230	13.79	0.086	1.00	0.222	12.65	0.253	13.75
O Levels or CSE	0.150	9.61	-0.055	-0.63	0.139	8.17	0.171	10.21
Workforce >500	0.185	12.22	0.030	1.81	0.182	10.55	0.167	9.63
Workforce 100-499	0.115	8.24	0.024	1.72	0.096	6.25	0.090	5.91
Workforce 25-99	0.040	2.90	0.007	0.53	0.010	0.64	0.020	1.40
Agriculture, Forestry & Fishing	-0.670	-9.70	-0.109	-1.53	-0.643	-7.47	-0.756	-9.90
Extraction of Minerals & Manufacture of Metals	-0.115	-3.09	-0.001	-0.02	-0.124	-3.83	-0.148	-4.36
Metal Goods, Engineering & Vehicles Industries	-0.227	-6.73	-0.055	-1.29	-0.207	-6.62	-0.258	-8.51
Other Manufacturing Industries	-0.271	-8.07	-0.004	-0.10	-0.211	-6.35	-0.281	-8.96
Construction	-0.207	-5.54	-0.007	-0.16	-0.136	-3.84	-0.218	-6.10
Distribution, Hotels & Catering (Repairs)	-0.396	-11.48	-0.085	-1.93	-0.367	-10.77	-0.410	-12.94
Transport & Communication	-0.290	-8.47	0.027	0.58	-0.249	-7.59	-0.287	-9.22
Banking & Finance	-0.067	-1.97	-0.071	-1.59	-0.043	-1.18	-0.092	-2.60
Public Administration, Education, Other	-0.223	-6.85	-0.099	-2.24	-0.191	-6.23	-0.246	-7.75
C_i	_	_	_	_	**0.040**	**1.90**	**-0.002**	**-0.08**
C_{it}	_	_	_	_	**0.029**	**2.04**	**-0.008**	**-0.46**
C_i.(Union Membership)	_	_	_	_	_	_	**-0.057**	**-1.53**
C_{it} .(Union Membership)	_	_	_	_	_	_	**-0.030**	**-1.16**
C^2_i	_	_	_	_	**-0.134**	**-4.55**	_	_
C^2_{it}	_	_	_	_	**0.020**	**2.29**	_	_
Constant	1.656	31.92	0.013	0.10	1.847	33.61	1.810	33.39
Adj R^2	0.419		_		0.400		0.406	

Notes

1. Hierarchical/ Unrestricted Standard Errors obtained using 250 "Wild Bootstrap" replications
2. All models, except Fixed effects, are inclusive of time dummies

Source: BHPS (1991-2003), ISER, Essex, SN:4967, June 2004

TABLE 5

Wage Regressions (1991-1996), Females

	OLS		Fixed Effects		Hierarchical		Unrestricted	
Variable	**Coef.**	**t**	**Coef.**	**t**	**Coef.**	**z**	**Coef.**	**z**
Trade Union Membership	**0.128**	**10.54**	**0.013**	**1.02**	**0.177**	**9.22**	**0.174**	**8.39**
Potential Labour Market Experience	0.005	2.96	0.067	16.28	0.000	0.21	0.004	2.41
Squared Experience	0.000	-3.86	0.000	-5.07	0.000	-1.62	0.000	-2.91
Marital Status	0.071	6.06	0.022	1.24	0.082	6.78	0.060	4.89
Full-Time Employment	0.103	8.54	-0.109	-7.51	0.096	7.13	0.099	6.72
Maternity Leave	-0.340	-6.92	-0.403	-12.28	-0.377	-3.35	-0.257	-3.24
Asian(Indian, Pakistani, Chinese, Other)	0.076	1.31	_	_	0.093	1.95	0.166	2.90
Black(Carribean, African, Other)	0.085	1.12	_	_	0.125	2.13	0.003	0.06
Other Ethnic Minority Group	-0.020	-0.26	_	_	0.042	0.44	-0.010	-0.10
Inner/ Outer London and R of South East	0.204	12.58	0.146	2.34	0.204	10.58	0.204	11.38
South West	0.020	0.85	0.010	0.11	0.055	2.02	0.028	1.09
Scotland	0.081	3.83	0.424	4.52	0.068	2.98	0.053	2.29
Wales	0.070	2.58	0.165	0.78	0.063	2.32	0.082	3.15
North West	0.077	3.87	0.232	1.67	0.075	3.30	0.077	4.12
North East	-0.009	-0.50	0.092	0.81	-0.025	-1.28	-0.017	-0.94
East Anglia	-0.021	-0.74	0.831	4.13	-0.043	-1.37	-0.035	-1.10
University Degree or Higher	0.543	24.17	0.055	0.49	0.545	24.58	0.571	25.42
HND, HNC, Teaching	0.418	17.43	-0.004	-0.05	0.403	15.62	0.441	16.42
A Levels	0.213	11.00	-0.009	-0.15	0.175	9.19	0.254	12.01
O Levels or CSE	0.128	8.76	-0.056	-1.05	0.112	8.24	0.143	10.71
Workforce >500	0.191	11.67	0.031	1.72	0.166	8.98	0.124	7.12
Workforce 100-499	0.148	10.30	0.038	2.49	0.141	8.86	0.111	7.15
Workforce 25-99	0.079	5.92	0.016	1.30	0.082	5.22	0.038	2.66
Agriculture, Forestry & Fishing	_	_	_	_	_	_	_	_
Extraction of Minerals & Manufacture of Metals	0.082	1.44	0.113	1.72	0.005	0.08	-0.111	-1.85
Metal Goods, Engineering & Vehicles Industries	-0.017	-0.34	0.081	1.55	-0.085	-1.40	-0.183	-3.50
Other Manufacturing Industries	-0.111	-2.39	0.016	0.31	-0.172	-2.67	-0.291	-5.42
Construction	0.054	0.68	0.081	1.12	-0.072	-1.01	-0.157	-2.01
Distribution, Hotels & Catering (Repairs)	-0.199	-4.54	0.029	0.62	-0.292	-4.67	-0.402	-7.91
Transport & Communication	-0.013	-0.26	0.021	0.35	-0.112	-1.57	-0.254	-4.22
Banking & Finance	0.107	2.42	0.048	0.98	-0.004	-0.07	-0.117	-2.30
Public Administration, Education, Other	0.029	0.69	0.048	1.03	-0.082	-1.36	-0.188	-3.80
C_i	_	_	_	_	**0.002**	**0.16**	**0.025**	**1.19**
C_{it}	_	_	_	_	**-0.110**	**-8.40**	**-0.081**	**-4.32**
C_i.(Union Membership)	_	_	_	_	_	_	**-0.080**	**-2.27**
C_{it} .(Union Membership)	_	_	_	_	_	_	**0.026**	**0.84**
Constant	1.108	21.47	0.080	0.80	1.367	19.31	1.423	24.42
Adj R^2	0.403		_		0.407		0.408	

Notes

1. Hierarchical/ Unrestricted Standard Errors obtained using 250 "Wild Bootstrap" replications
2. All models, except Fixed effects, are inclusive of time dummies

Source: BHPS (1991-2003), ISER, Essex, SN:4967, June 2004

TABLE 6

Wage Regressions (1997-2002), Females

	OLS		Fixed Effects		Hierarchical		Unrestricted	
Variable	Coef.	t	Coef.	t	Coef.	z	Coef.	z
Trade Union Membership	**0.115**	**10.49**	**0.048**	**3.81**	**0.162**	**9.25**	**0.149**	**8.25**
Potential Labour Market Experience	-0.003	-1.54	0.065	10.44	-0.003	-1.18	-0.004	-1.59
Squared Experience	0.000	-0.12	0.000	-2.15	0.000	-0.43	0.000	-0.23
Marital Status	0.046	4.46	0.069	3.67	0.027	2.22	0.034	2.89
Full-Time Employment	0.067	6.00	-0.065	-4.64	0.071	5.58	0.056	4.71
Maternity Leave	-0.236	-5.59	-0.239	-7.66	-0.239	-3.27	-0.232	-3.36
Asian(Indian, Pakistani, Chinese, Other)	-0.169	-3.91	_	_	-0.238	-4.87	-0.170	-3.27
Black(Carribean, African, Other)	-0.077	-1.26	_	_	-0.089	-3.19	-0.092	-2.87
Other Ethnic Minority Group	-0.081	-1.22	_	_	-0.026	-0.33	0.018	0.21
Inner/ Outer London and R of South East	0.189	12.89	0.038	0.53	0.197	11.91	0.209	13.56
South West	0.062	3.21	0.037	0.36	0.062	2.97	0.077	4.22
Scotland	0.089	4.49	0.317	1.94	0.078	3.45	0.118	5.23
Wales	0.060	2.48	0.130	0.87	0.080	2.80	0.109	4.07
North West	0.049	2.66	0.080	0.53	0.043	2.07	0.035	1.75
North East	0.019	1.15	0.006	0.07	0.038	2.03	0.012	0.69
East Anglia	-0.020	-0.75	-0.029	-0.18	-0.020	-0.76	-0.047	-1.83
University Degree or Higher	0.500	25.16	0.411	3.22	0.534	25.62	0.498	25.01
HND, HNC, Teaching	0.343	15.71	0.335	2.41	0.343	13.90	0.365	14.62
A Levels	0.180	10.88	0.163	1.34	0.194	10.20	0.215	11.92
O Levels or CSE	0.147	10.01	0.100	0.93	0.148	9.92	0.151	9.77
Workforce >500	0.181	12.42	0.050	2.62	0.193	12.27	0.174	10.95
Workforce 100-499	0.156	11.61	0.051	3.24	0.166	11.62	0.179	11.70
Workforce 25-99	0.092	7.43	0.032	2.39	0.097	7.18	0.087	6.21
Agriculture, Forestry & Fishing	-0.205	-2.25	0.011	0.11	-0.289	-4.48	-0.090	-1.32
Extraction of Minerals & Manufacture of Metals	-0.030	-0.43	0.056	0.69	-0.058	-0.91	0.126	2.18
Metal Goods, Engineering & Vehicles Industries	-0.130	-2.05	0.055	0.73	-0.199	-3.73	-0.102	-1.94
Other Manufacturing Industries	-0.202	-3.37	0.030	0.42	-0.244	-4.71	-0.136	-2.87
Construction	-0.127	-1.52	0.075	0.73	-0.176	-2.94	-0.034	-0.58
Distribution, Hotels & Catering (Repairs)	-0.342	-5.85	-0.072	-1.03	-0.384	-7.92	-0.294	-6.16
Transport & Communication	-0.085	-1.36	0.014	0.20	-0.102	-1.91	-0.026	-0.48
Banking & Finance	0.008	0.14	0.043	0.63	-0.038	-0.80	0.085	1.87
Public Administration, Education, Other	-0.121	-2.10	0.021	0.31	-0.180	-3.82	-0.068	-1.47
C_i	_	_	_	_	**-0.068**	**-2.96**	**-0.030**	**-0.86**
C_{it}	_	_	_	_	**-0.035**	**-2.95**	**-0.037**	**-2.61**
C_i.(Union Membership)	_	_	_	_	_	_	**-0.046**	**-0.97**
C_{it} .(Union Membership)	_	_	_	_	_	_	**0.008**	**0.38**
Constant	1.661	23.47	-0.340	-2.06	1.732	26.68	1.627	26.03
Adj R^2	0.399		_		0.414		0.419	

Notes

1. Hierarchical/ Unrestricted Standard Errors obtained using 250 "Wild Bootstrap" replications
2. All models, except Fixed effects, are inclusive of time dummies

Source: BHPS (1991-2003), ISER, Essex, SN:4967, June 2004

TABLE 7

Descriptive Statistics, Endogeneity Correction Terms (1991-1996), Males

	Union Members		Non-Union Members	
Latent Effect	Mean	Std. Dev.	Mean	Std. Dev.
C_i	0.385	0.488	-0.408	0.449
C_{it}	0.281	0.414	-0.083	0.425
C^2_i	0.386	0.658	0.368	0.481
C^2_{it}	0.250	0.615	0.187	0.666

TABLE 8

Descriptive Statistics, Endogeneity Correction Terms (1997-2002), Males

	Union Members		Non-Union Members	
Latent Effect	Mean	Std. Dev.	Mean	Std. Dev.
C_i	0.194	0.317	-0.376	0.248
C_{it}	0.261	0.447	-0.061	0.395
C^2_i	0.138	0.278	0.203	0.152
C^2_{it}	0.267	0.655	0.159	0.642

TABLE 9

Descriptive Statistics, Endogeneity Correction Terms (1991-1996), Females

	Union Members		Non-Union Members	
Latent Effect	Mean	Std. Dev.	Mean	Std. Dev.
C_i	0.396	0.474	-0.400	0.327
C_{it}	0.394	0.451	-0.089	0.400
C^2_i	0.382	0.559	0.267	0.309
C^2_{it}	0.359	0.644	0.168	0.565

TABLE 10

Descriptive Statistics, Endogeneity Correction Terms (1997-2002), Females

	Union Members		Non-Union Members	
Latent Effect	Mean	Std. Dev.	Mean	Std. Dev.
C_i	0.207	0.267	-0.284	0.207
C_{it}	0.295	0.438	-0.099	0.436
C^2_i	0.114	0.175	0.124	0.104
C^2_{it}	0.279	0.595	0.200	0.648