UNIVERSIDAD CARLOS III DE MADRID working papers

Working Paper 08-23
Economic Series (15)
May 2008

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Determinants of Trade Union Membership in Great Britain During 1991-2003[1,2]

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Abstract

We analyse the determinants of union membership in the UK using data from the BHPS (1991-2003). Employing three alternative methodologies to control for the problem of initial conditions we find that union membership remains persistent even after controlling for the unobserved effect. There is evidence of a considerable correlation between the unobserved individual heterogeneity and the initial membership status. Ignoring this overstates the degree of state dependence of union membership greatly. The extent of state dependence in union membership status is notably higher in the (1991-1996) period estimates and appears to be more pronounced in the case of male employees for the entire period under analysis. The second period estimates reveal that unobserved heterogeneity has a more prominent impact in determining future unionisation probability versus past union membership. Finally, the estimates suggest that an individual´s propensity to unionise is determined by a mixture of industrial and personal characteristics. This is at odds with earlier studies, such as Booth (1986) and Wright (1995), failing to control for unobservable effects and concluding that personal attributes do not have a significant impact on unionisation propensity.

Keywords: union membership, initial conditions, unobserved individual heterogeneity, state dependence, dynamic random effects probit models.

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JEL Classification: C23, J51

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1 The views presented in this paper are the author´s and do not reflect those of the BHPS data depositors, namely, the Institute for Social and Economic Research at the University of Essex, U.K.

2 I am grateful to Jeffrey Wooldridge for his comments and suggestions. I would also like to thank Jean Marie Linhart, Guillaume Frechette, Mark Stewart, John Bone, Keith Hartley, John Hutton and Andrew Jones for their valuable help. All errors and omissions remain my own.

1. INTRODUCTION

This paper analyses the determinants of union membership in the U.K during 1991-2003. We are interested in investigating the characteristics of the individuals who join trade unions and identify the traits that are associated with higher, as well as, lower unionisation propensity.

Further, we wish to explore whether both personal and industrial characteristics are equally important determinants of individual unionisation propensity. In addition, we aim to establish if union membership is still persistent in an era that is characterised by rapid deunionisation. In other words, we seek to determine whether individuals who have been union members is the past are likely to do so in the future and if the present disadvantageous position of unions in the UK is coupled with a fall in union persistence.

Unobserved individual heterogeneity (i.e. the unobserved individual characteristics underlying the union/non-union decision) and actual experience of union membership in the past, offer opposing justifications to the likelihood that individuals might remain in unions in the future (see Hsiao, 2003). Thus, we need to determine the relative importance of unobserved heterogeneity versus past experience of unionisation in determining future union status.

Provided that the initial value of union status is affected by past experience we need to employ an estimation methodology that accounts for the initial conditions problem. Primarily since the random effects maximum likelihood estimator in its standard form will be inconsistent (see Heckman, 1981a,b) and furthermore, because the correlation between unobservable effects and the initial value of union status will overstate the degree of state dependence. We employ three alternative methods, namely, Heckman (1981b), Orme (2001), Wooldridge (2005) to account for initial conditions and obtain the degree of true, as opposed to spurious, state dependence in union membership. Therefore, our final research task is assessing the relative merits and comparing the results obtained from these alternative estimators.

The legal framework governing collective bargaining arrangements in the United Kingdom was substantially altered during the 1980's. Stewart (1995) provides a concise account of the successive legislative changes that were targeted towards weakening the bargaining strength of UK trade unions.

The introduction of the 1980 and 1982 Employment Acts strengthened the case for claiming unfair dismissal on the basis of an employee's refusal to enter a closed shop. The 1982 Act dictated that all post-entry closed shops had to be sanctioned via a ballot of the workforce. The successive 1988 and 1990 Employment Acts prohibited all means to enforce a closed shop and thus, rendered post-entry and pre-entry closed shops, respectively, illegal. The overall effect was that the 1980s was a decade of a dramatic decline in aggregate trade union membership and union recognition in the UK (Stewart, 1995, pp.143-145).

Changes in the pattern of trade union membership and union recognition in the UK have been studied by a number of authors (e.g. Booth, 1986; Gregg and Naylor, 1993; Andrews and Naylor, 1994; Wright, 1995; Disney *et al*, 1996; Disney *et al*, 1999). Most of these studies employ cross-sectional data and are, therefore unable to control for unobservable individual-specific effects something that might well result in biased estimates of the effect of observable attributes. Further, upon failing to let the effects of the explanatory variables be time dependent the effects of establishment characteristics are particularly prone to bias (see Arulampalam and Booth, 2000, p.290).

Arulampalam and Booth (2000) using data from the National Child Development Study was the first longitudinal study of the patterns of changes in individual membership across time for a single cohort of individuals (young men) in the UK. However, the nature of their data set does not allow them to discriminate between calendar time and life-cycle effects.

Costs and benefits of potential union membership are quite hard to quantify and the relevant data are not always available. Individuals might opt to become trade union members so as to benefit from incentive private excludable goods available only to trade union members. Arulampalam and Booth (2000) summarise these as protection against unfair dismissal, discrimination due to ethnic minority group membership, grievance procedures, pension plans advice and the implementation of well-defined dismissal arrangements in recessionary periods (p.291). Naylor and Raaum (1993) note that trade union membership can be "influenced by both social custom effects and by resources devoted by management to opposing unionisation" (p.591). Farber (1983) emphasises that "workers are heterogeneous in their preference for union representation to the extent that workers of different characteristics derive different amounts of pecuniary and non-pecuniary benefits from unionisation" (p.1420).

Conclusively then, unobserved individual heterogeneity plays an important role in modelling union membership status and failing to control for this provides biased estimates. However, it should be remembered that the existing data restrictions, and in particular the paucity of employer controls, mean that we cannot attribute any specific effects entirely to unobserved heterogeneity.

Our findings provide evidence that union membership remains persistent even after controlling for the unobserved effect. The likelihood to remain in unions having experienced unionisation in the past has, however, fallen during the period under analysis. Further, failing to control for unobserved individual-specific effects overstates the true degree of union membership persistence and indicates the inappropriateness of the pooled Probit estimator that assumes away any correlation between the successive disturbances. We conclude that the relative importance of unobserved heterogeneity in determining future unionisation propensity rises during the second period as opposed to the effect of past union membership.

Heckman's (1981b) solution to the initial conditions problem is our preferred estimator on the grounds that it incorporates pre-sample information, whereas, Wooldridge's (2005) methodology is attractive due to its computational simplicity. Orme's (2001) method is found to be fairly restrictive as the assumption of a sufficiently low correlation among the initial condition and the unobserved effect is rejected by the data thus rendering the resulting estimates inconsistent.

Lastly, but not least, contrary to earlier studies failing to control for unobservables, we find that both industrial and personal characteristics are important factors in the union/non-union decision. The most prominent observed heterogeneity effect, other than lagged union status, on the probability of unionisation comes from the size of workforce that it is positively associated with union membership. Furthermore, with respect to both genders during the entire period under analysis, highly qualified individuals appear to be less prone to unionisation. However, employees in the public and education sectors seem to have a higher propensity towards unions.

Concerning the male estimates there is evidence of a widespread decline in the probability of union membership across the majority of occupations during the first period and further, the second period results reveal a prevalent deunionisation tendency across all industries except the public and education sectors. The female estimates on the other hand, display no evidence of prevailing deunionisation across any occupation. This occurs in that, contrary to the male samples, female union membership rates have been increasing during the period under analysis.

The remainder of this paper is structured as follows. Section 2 outlines data and sample selection issues; Section 3 the estimation methodology; Section 4 presents and explores the estimated results; and Section 5 concludes.

2. DATA AND SAMPLE SELECTION

The British Household Panel Survey (BHPS) originated in 1991 and follows the same representative sample of individuals over a number of time periods. The first wave of the Panel (1991-92) consisted of 5,500 households and 10,300 individuals drawn from 250 regions of Great Britain. In 1999 further samples of 1,500 households in the cases of both Scotland and Wales were added and in 2001 a sample of 2,000 households from Northern Ireland was included in the survey.

The data were split into two balanced panels (1991-1996 and 1997-2002) comprising of employees aged between sixteen and sixty five, excluding full-time students, that have participated in all of the respective six waves of the survey.[3] The ECHP[4] samples and the new Scotland, Wales and Northern Ireland samples were selected out.

The traditional practice in union literature is to use male manual full-time employees in order to estimate the determinants of union membership (see Swaffield, 2000, p.439). However, this study uses a full- time male employees' sample and a full sample of female employees (full-time and part-time inclusive of female employees on maternity leave).

Part-time male employees were excluded from the male regression sample used in this study in that 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 regression samples.

In addition, the full female sample can provide a comparison group against the male full-time employee sample that could potentially suffer from selectivity bias. According to Swaffield (2001) it might be possible, however, that the female sample is also prone to sample selection bias caused by the labour market participation decision (p.439).

Further disaggregation towards a male manual full-time sample was not performed however as it would be fairly costly in terms of sample sizes. Descriptive statistics Tables for the respective selected samples are provided in Table 1 in the Appendix and include of a set of personal and industrial characteristics.

2.1. Trade Union Membership

The dependent variable is an indicator function taking a value of one if an individual is a trade union member and a value of zero otherwise. An individual is taken to be a trade union member if he or she has responded positively to the question "Are you currently a member of: Trade Unions" in the Social and Interest

[3] Proxy respondents allocated zero weights by the data depositors, by default, were selected out.

[4] European Community Household Panel survey.

Group Membership section of the BHPS. The respective variable in the BHPS data set is "Member of trade union". 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 union membership for waves six (1996-97), eight (1998-99), ten (2000-01) and twelve (2002-03). This variable is derived through the question that was asked conditionally following a positive response to the question "Is there a trade union, or a similar body such as a staff association, recognised by your management for negotiating pay or conditions for the people doing your sort of job in your workplace?".[5]

Following this response, a positive reply to "Are you a member of this trade union/association?" is recorded as membership of workplace union which includes "in-house" staff associations, but excludes employers' organisations. This introduces a degree of discontinuity in the BHPS data regarding trade union membership as this broader definition is also inclusive of staff associations, but, nevertheless when one wishes to undertake a longitudinal analysis spanning all of the first and last six waves of the BHPS this is the only alternative available.[6]

The "Union or staff association at workplace" variable allows us to construct a measure of union coverage. However, the fact that an individual's workplace is covered by a trade union that is recognised by management for bargaining purposes does not necessarily imply that this particular individual will actually be a member of the union.

According to Andrews *et al* (1998) while at the individual level members and non-members doing the same job within the same establishment earn the same wage, when comparing across establishments trade union membership is "a closer indicator of a differential than coverage"(p.453).

Thus, to conclude this Section. The cross-sectional time-series trade union membership variable constructed for all twelve waves of the survey suffers from a certain degree of discontinuity. Nevertheless, the two alternative questions available in the BHPS make it feasible to construct a measure of union membership. The individual responses from the two alternative union membership questions and the union coverage outcomes can be compared in order to detect any potential measurement error in union membership responses (see Swaffield, 2001). However, this is beyond the scope of this paper.

The descriptive statistics (in Table 1) reveal that male union membership during (1997-2002) has fallen by approximately 15.3 percent compared to the (1991-1996) levels. The respective female union membership rate has actually risen by approximately 9.5 percent so that by the last cross-section of the survey female unionisation rates converged to the level of male unionisation rates at a marginally higher percentage.

3. ESTIMATION METHODOLOGY

The dynamic model depicting the decision of an individual to join either the unionised or non-unionised sector is outlined in equation (1) below. The benefits

[5] The resulting variable is termed as "Union or staff association at workplace".

[6] Swaffied (2001) employs a similar approach.

of employment within the unionised sector are captured by the latent variable y_{it}^*. The union membership status of an individual i in period t, is indicated by the dummy variable y_{it}.

The unknown parameters to be estimated are $(\gamma, \beta)'$ and x_{it} is a vector of exogenous explanatory variables (personal and industrial characteristics). The composite error term ν_{it} captures the unobserved individual heterogeneity underlying the union membership decision and is decomposed into an individual-specific component ε_i and an individual time-specific effect u_{it}:

$$\begin{aligned} y_{it}^* &= x_{it}'\beta + \gamma y_{i,t-1} + \nu_{it} \\ y_{it} &= I(y_{it}^* \rangle 0) \\ \nu_{it} &= \varepsilon_i + u_{it};\ t = 1, ..., T;\ i = 1, ..., N \end{aligned} \tag{1}$$

3.1. Potential Limitations

The combined effect of the 1988 and 1990 Employment Acts, that precede the introduction of the BHPS in 1991, was the effective outlaw of both post-entry and pre-entry closed shops (see Stewart, 1995). This has the implication that individuals undertake their unionisation decision on the basis of wages, individual preferences and non-pecuniary benefits without coercion.

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 obvious drawback, however, of adding the lagged union membership variable as a regressor is that it gives rise to the problem of initial conditions which is discussed in greater detail in Section 3.2.

Of course, becoming a union member is also conditional upon the employer's hiring decision (see Abowd and Farber, 1982) and the primary deficiency of the union membership model employed is that it ignores the role of the employer in determining union status. Despite the fact that employer attributes are captured through the industrial classification dummies and the establishment size controls these are not adequate in order to assign any specific effects purely to unobserved heterogeneity (see Vella and Verbeek, 1998, p.164). Further, as has been already mentioned in the previous Section, the BHPS limits our ability to control for potential establishment age effects on union membership.[7]

Another potential shortcoming of the model concerns free riding on union bargained wages that apply to all employees within a covered establishment irrespective of their union membership status. Booth (1986) argues that wages can be viewed as a collective good by the individual therefore casting doubts on whether wages can actually be considered as a determinant of an individual's unionisation probability.

The collective good argument does not invalidate the theoretical prediction that union density is potentially positively associated with higher union wages. At the margin an individual's union membership decision might not be affected by the level of wages set by collective bargaining. Primarily, this occurs because the individual does not consider that his decision will make an important difference to the aggregate union density in his sector. What is more, the "marginal" individual can always free-ride (see Booth, 1986, p.44 and pp.58-59).

Some readers might consider the exclusion of wages from the sectoral choice model of union membership as inappropriate. Wages can be included in the union

[7] Via the establishment "age effect" impact on union recognition (see Stewart, 1995).

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. However, according to Booth (1986) this scenario is highly unlikely (p.44).

Vella and Verbeek (1998) suggest that the lagged value of union membership status affects an individuals' unionisation decision while it does not have any significant effect on the current wage. It is argued 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 (p.167).

The joint determination of union status and wages renders the coefficients of a single equation model inclusive of wages biased and inconsistent. Even if wages and union status are not determined simultaneously, which is rather improbable, 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). A two-step methodology such as the one offered by Vella and Verbeek (1998, 1999b) provides a far more attractive alternative and this paper focuses on the first stage estimation of reduced form models of union membership determination.

3.2. Estimation Procedure and the Problem of Initial Conditions

Modelling any stochastic process with structural dependence among time-ordered outcomes requires initialising the process (Heckman, 1981a, p.118). Initial conditions do become irrelevant asymptotically as T gets infinitely large although, in the case of short panels as is the case in this study the problem cannot be overlooked since $T = 6$ and asymptotics instead rely on large sample sizes, N. 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 ε_i clearly invalidates the assumption of exogeneity of union membership status in the first period of the survey.

The initiation of the stochastic process determining union membership has been in operation prior to initiation of the BHPS in 1991. This occurs since a large fraction of individuals in the samples used were labour market participants before 1991. Thus, the initial value of union membership cannot be taken to be exogenous. The conditional probability that an individual will become a union member in the future is a function of past experience.

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 ε_i and the initial conditions will overstate the degree of state dependence.

State dependence and individual heterogeneity offer "diametrically opposite" explanations of the notion that those individuals who have experienced an event in the past are more likely to do so in the future (Hsiao, 2003, p.216).

Considering otherwise identical individuals it is possible that those who have experienced unionisation in the past will amend their preferences determining future propensity to unionise (e.g. via potential seniority and non-pecuniary benefits pertaining to union membership). This is an entirely behavioural effect.

Alternatively, it is possible individuals differ in specific unmeasured variables that affect their propensity to unionise while they are not influenced by experiencing unionisation *per se*. In the event whereby such variables are correlated over time, and are not appropriately controlled for, past experience may turn out to be a determinant of the individuals' future propensity to unionise since it acts as proxy for the temporally persistent unobservables. This is what Heckman (1981a, 1981b) terms as "spurious state dependence" as opposed to "true (structural) state dependence" occurring in the former scenario.

Union membership decision is modelled using the dynamic Random effects Probit specification given in equation (2). The random effects formulation was chosen since in the case of dynamic models with large N and small T fixed effects produce inconsistent estimates of the parameters as differencing out ε_i generates a linear regression model with lagged dependent regressors and serially correlated disturbances. Further, the Random effects Probit model is used instead of its Logit counterpart since random effects yield correlations among the successive disturbances. For this purpose, the multivariate normal distribution is more flexible than the corresponding logistic distribution which requires that all correlations are equal to 0.5 (see Maddala, 1987):

$$y_{it}^{*} = x_{it}^{\prime}\beta + \gamma y_{i,t-1} + \varepsilon_i + u_{it};\ u_{it} \sim N(0, \sigma_u^2) \tag{2}$$

The random effects model in its standard form assumes that ε_i is not correlated with x_{it}. The presence of the individual-specific time-invariant effect ε_i, however, renders the composite error term $v_{it} = \varepsilon_i + u_{it}$ temporally correlated even when the u_{it} are taken to be serially independent. Adopting the Mundlak (1978)- Chamberlain (1984) specification we can allow for a correlation between ε_i and the time means of the observed time-varying characteristics taking the form of $\varepsilon_i = \overline{x}_i^{\prime}a + \alpha_i$. Substituting this expression for ε_i in equation (2) we arrive at specification (3) where it is assumed that $\alpha_i \sim iidN(0, \sigma_\alpha^2)$ and is independent of (x_{it}, u_{it}) for all i and t:

$$y_{it}^{*} = x_{it}^{\prime}\beta + \gamma y_{i,t-1} + \overline{x}_i^{\prime}a + \alpha_i + u_{it};\ u_{it} \sim N(0, \sigma_u^2) \tag{3}$$

The individual-specific random effects framework suggests an equi-correlation, ρ, between any two successive disturbances for the same individual unit. Given that y_{it} is dichotomous, a normalisation is necessary and it is commonly assumed that $\sigma_u^2 = 1$. The resulting expression for ρ is given in equation (4):

$$\rho = cor(v_{it}, v_{is}) = \frac{\sigma_\alpha^2}{\sigma_\alpha^2 + 1}\ ;\ t, s = 2, ..., T; t \neq s \tag{4}$$

The initial conditions problem is tackled using three alternative estimation methodologies suggested by Heckman (1981b), Orme (2001) and Wooldridge (2005). Heckman's solution to the initial conditions problem approximates the reduced form marginal probability of the initial state by a Probit function which has as its argument all of the available pre-sample information on the exogenous variables (Heckman, 1981b, p.188). Orme (2001) suggests a two step methodology which is an approximation if the correlation between the initial condition and individual random effects is weak. Wooldridge (2005) proposes an alternative approach which involves modelling the distribution of the unobserved effect conditional on the initial value and the observed history of strictly exogenous explanatory variables (p.40).

3.2.1. Heckman's Estimator

Stewart (2006) provides a Stata command, -redprob-, for Heckman's estimator of the dynamic Random effects Probit model. In the spirit of Heckman (1981b) we specify a linearised reduced form for the initial observation given by equation (5) where z_{i1} denotes a vector of strictly exogenous instruments such as pre-sample information[8] affecting an individual's propensity to unionise, the vector of means $\overline{x}_i$ and x_{i1}, η_i is correlated with α_i but not with u_{it} for $t \succeq 2$:

$$y_{i1}^{*} = z_{i1}'\pi + \eta_i \tag{5}$$

In terms of orthogonal error components η_i can be expressed as:

$$\eta_i = \theta\alpha_i + u_{i1};\ \theta \succ 0 \tag{6}$$

By construction (α_i, u_{i1}) in equation (6) are independent of one another. Exogeneity of the initial conditions occurs when $\theta = 0$. Further, it is assumed that u_{i1} meets the same distributional assumptions as u_{it} for $t \succeq 2$. The linearised reduced form for sectoral choice in the initial time period is given by:

$$y_{i1}^{*} = z_{i1}'\pi + \theta\alpha_i + u_{i1};\ t = 1,\ i = 1, ..., N \tag{7}$$

The joint probability of $(y_{i1}, ..., y_{iT})$ for individual i, given α_i, suggested by Heckman's methodology is shown in (8) where Φ is the standard normal cumulative distribution function:

$$\Phi[(z_{i1}'\pi + \theta\alpha_i)(2y_{i1} - 1)].\prod_{t=2}^{T}\Phi[(x_{it}'\beta + \gamma y_{i,t-1} + \overline{x}_i'a + \alpha_i)(2y_{i1} - 1)] \tag{8}$$

Thus, expression (9) denotes the likelihood function to be maximised for a random sample of individuals. F stands for the distribution function of $\alpha^{*} = \alpha/\sigma_\alpha$ and given the normalisation adopted $\sigma_\alpha = \sqrt{\rho/1-\rho}$. Stewart (2006) provides a program for this maximum likelihood estimator whereby, assuming that α is normally distributed, the integral over α^{*} is evaluated using the Gaussian-Hermite quadrature:

$$\prod_i \int_{\alpha^{*}} \{\Phi[(z_{i1}'\pi + \theta\sigma_\alpha\alpha^{*})(2y_{i1}-1)].\prod_{t=2}^{T}\Phi[(x_{it}'\beta + \gamma y_{i,t-1} + \overline{x}_i'a + \sigma_\alpha\alpha^{*})(2y_{i1}-1)]\}dF(\alpha^{*}) \tag{9}$$

[8] The "year current labour force status began" variable in the BHPS allows us to obtain information about the industrial/occupational classification for those individuals that have entered the labour force prior to 1991 and 1997 in the two panels. Further pre-sample information variables were constructed for the highest educational qualification of those individuals aged over 25 at the initiation of the respective panels.

3.2.2. Orme's Pseudo Maximum Likelihood Estimator

Orme (2001) offers a "pseudo maximum likelihood estimator" *PMLE* procedure that rectifies the standard *MLE* for the effect of previous event history. The linear specification in terms of orthogonal disturbances in equation (6) allows for the possibility of non-zero correlation $r = corr(a_i, \eta_i)$. Orme's two-step methodology is an approximation when r is local to zero (i.e. when the correlation between the initial condition and the random effect is small).

Following Arulampalam *et al* (2000) we start with equation (3). Equation (6) is redefined as $\alpha_i = \delta\eta_i + w_i$ where by construction (η_i, w_i) are orthogonal of one another, $\delta = r\sigma_\alpha/\sigma_\eta$ and $var(w_i) = \sigma_\alpha^2(1-r^2)$. Substituting this new specification for the random effect into equation (3) we arrive at:

$$y_{it}^* = x_{it}'\beta + \gamma y_{i,t-1} + \overline{x}_i'a + \delta\eta_i + w_i + u_{it} \tag{10}$$

We begin with the linearised reduced form for the initial observation equation (5). Assuming that (α_i, η_i) are bivariate normal then $E(w_i \mid y_{i1}) = 0$ and $E(\eta_i \mid y_{i1}) = e_i$ where:

$$e_i = \frac{(2y_{i1} - 1)\varphi(z_{i1}'\pi)}{\Phi(\{2y_{i1} - 1\}z_{i1}'\pi)} \tag{11}$$

is the generalised Probit residual (see Gourieroux *et al*, 1987) obtained from equation (5). Assuming that u_{it} is independent of x_{it} we take w_i to be the typical Random effects Probit error component given that we replace η_i by its conditional expectation.

The assumption of bivariate normality of (α_i, η_i) renders the error component w_i in the second stage Random effects Probit model in equation (10) heteroskedastic since:

$$var(w_i \mid y_{i1}) = \sigma_\alpha^2(1 - r^2\zeta_i^2);\ \zeta_i = \frac{\varphi(z_{i1}'\pi)}{\sqrt{\Phi(z_{i1}'\pi)\Phi(-z_{i1}'\pi)}} \tag{12}$$

Effectively then, Orme's two stage methodology involves estimating an "artificial" Random effects Probit model which is augmented by the generalised Probit residual obtained from the first stage linearised reduced form for the initial period equation (5) under appropriate normality assumptions.

The rationalisation of Orme's approach is grounded on the assumption of r being local to zero so that $var(w_i \mid y_{i1}) \cong \sigma_\alpha^2$. Note that since union membership is a dichotomous variable the Probit model for the initial period needs the normalisation that $\sigma_\eta = 1$. Given the expression for δ in equation (10) we obtain $r = \delta/\sigma_\alpha$. Rearranging equation (4) we arrive at $\sigma_\alpha^2 \cong \rho/(1-\rho)$ and hence, an estimate of r can be obtained by the expression provided in equation (13) where δ is the coefficient on the generalised Probit residual from equation (10):

$$r = \delta\sqrt{(1-\rho)/\rho} \tag{13}$$

The "pseudo log-likelihood" to be maximised includes lagged union membership but treats y_{i1} as exogenous. Augmenting the set of regressors by e_i "provides a test of, and an approximate control for, the initial conditions problem". Locally equivalent alternative methodology guarantees that the usual t-test of the coefficient of the generalised residual gives an asymptotically valid test of $r = 0$ (Orme, 2001, p.6).

3.2.3. *Wooldridge's Conditional Maximum Likelihood Estimator*

Wooldridge's (2005) solution to the initial conditions problem 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:

$$\varepsilon_i | y_{i1}, x_i \sim N(\alpha_0 + \alpha_1 y_{i1} + x_i' \alpha_2, \sigma_\alpha^2); \; x_i = \{x_{i1}, ..., x_{iT}\} \tag{14}$$

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

The presence of x_i in expression (14) implies that we are not able to identify the coefficients on time-constant explanatory variables in x_{it} although time-constant explanatory variables can be included in x_i and this is the main disadvantage of Wooldridge's approach. Including time-constant explanatory variables in x_{it} will only increase the explanatory power of the model as it is not possible to separately identify the partial effect of time-constant variables from their partial correlation with the unobserved effect (Wooldridge, 2005, p.44).

Of course, the implication of adopting the Mundlak (1978)-Chamberlain (1984) device allowing for a correlation between ε_i and the time means of the observed time varying characteristics, $\varepsilon_i = \overline{x}_i' a + \alpha_i$, with the Heckman (1981b) and Orme (2001) approaches is that the presence of the $\overline{x}_i$ means that we cannot separately identify the effects of time-constant variables there, either.

The density $D(y_{i1}, ..., y_{iT} | y_{i1} = y_1, x_i = x, \alpha_i = \alpha)$ is given by:

$$\prod_{t=1}^{T} \{\Phi(x_t'\beta + \gamma y_{t-1} + \alpha_0 + \alpha_1 y_{i1} + x'\alpha_2 + \alpha)^{y_t} . [1 - \Phi(x_t'\beta + \gamma y_{t-1} + \alpha_0 + \alpha_1 y_{i1} + x'\alpha_2 + \alpha)]^{1-y_t}\} \tag{15}$$

To find the joint distribution of $(y_{i2}, ..., y_{iT} | \; y_{i1} = y_1, x_i = x)$ we need to integrate out α_i. Integrating (15) against the Normal (0, σ_α^2) gives the likelihood function in expression (16) 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 $\{z_{it} \equiv (1, x_{it}, y_{i,t-1}, \; y_{i1}, \; x_i)\}$. It is assumed that data are observed for each cross-sectional unit in all time periods although given specific sample selection mechanisms (16) can be employed for the subset of observations forming a balanced panel:

$$\begin{aligned} &\int_{\mathbb{R}} \prod_{t=1}^{T} \{\Phi(x_t'\beta + \gamma y_{t-1} + \alpha_0 + \alpha_1 y_{i1} + x'\alpha_2 + \alpha)^{y_t} . \\ &[1 - \Phi(x_t'\beta + \gamma y_{t-1} + \alpha_0 + \alpha_1 y_{i1} + x'\alpha_2 + \alpha)]^{1-y_t}\} (1/\sigma_\alpha) \phi(\alpha/\sigma_\alpha) d\alpha \end{aligned} \tag{16}$$

Essentially then, what Wooldridge (2005) suggests is that by adding y_{i1} and x_i as additional explanatory variables in each time period under consideration and using a computationally easy standard Random effects Probit software we can estimate $(\beta, \gamma, \alpha_0, \alpha_1, \alpha_2, \sigma_\alpha^2)$.

To sum up, the three alternative methodologies can be employed to test whether the initial value of union status is exogenous. If our expectation that the initial conditions are endogenous is verified then the three estimators can provide consistent results. Finally, controlling for the correlation between unobservable effects and the initial condition will provide the degree of true as opposed to spurious state dependence in union membership. Orme (2001) and Wooldridge (2005) provide simpler solutions to the initial conditions problem that are computationally less demanding to Heckman's (1981b) estimator. The estimates and relative performance of the three solutions are discussed in the following Section.

4. ESTIMATES

We focus on the estimates obtained using the Heckman (1981b) and Wooldridge (2005) estimation methodologies. Concerning Orme's (2001) two-step methodology, the estimates of the correlation between the initial condition and the random effect, r, were all in excess of 0.8. In the light of this, we cannot ignore the inherent heteroskedasticity of the residual component, w_i, in the second stage Random effects Probit model (*eq*.10) as it produces inconsistent parameter estimates. The estimated results are therefore, not reported here.

The requirement that r is local to zero under Orme's (2001) estimator is quite a stringent condition. Arulampalam *et al* (2000) looking at unemployment persistence obtain significantly lower values of r. We would normally expect union membership to be far more persistent than unemployment, in that the former entails long term benefits whereas long term unemployment reduces the employment probability of an individual. Therefore, the high correlation between the initial value of union status and the random effect found here is not surprising.

The remaining estimated models of union membership employing Heckman's (1981b) and Wooldridge's (2005) estimators are provided in Tables (2-5) in the Appendix. In all reported models the null under the Wald statistic[9] for multiple exclusion restrictions is rejected and hence, all inclusive covariates are jointly statistically significant. Further, the estimated parameters generally possess the theoretically predicted coefficients.

The reported Random effects Probit estimates were arrived at employing the adaptive Gauss-Hermite quadrature method to compute the log likelihood and its derivatives since the large group sizes and the sizeable within group correlations, ρ, in all of the estimated models made the non-adaptive quadrature approximation inaccurate. The models were estimated by increasing the number of quadrature points to 30 which is equivalent to increasing the degree of the polynomial approximation (see Butler and Moffitt, 1982).[10]

The major advantage of Wooldridge's estimator is its computational simplicity which reduces estimation time substantially compared to Stewart's (2006) command used to implement Heckman's estimator. Wooldridge's estimator for $t > 1$ was extended by a standard Probit estimator for $t = 1$ to enable comparability with Heckman's estimator, although the latter still produced a superior log-likelihood for all models. This occurs in that Wooldridge's estimator fails to incorporate pre-sample information on the basis of which Heckman's solution to the initial conditions problem is our preferred estimation methodology. However, the estimates obtained using both approaches are reported here as this reinforces the argument that the correlation between the initial condition and the unobserved heterogeneity

[9] LR for the pooled Probit estimates.

[10] The quadrature checks undertaken revealed that the coefficient estimates were nearly invariant to the quadrature point variation. The polynomial approximation with 30 quadtrature points is therefore sufficiently accurate and the estimated results can be interpreted with confidence.

results in an over-statement of the extent of state dependence in union membership, more effectively.

Stewart's -redpprob- command reports a pooled Probit model for $t > 1$, the estimates of the reduced form model for the initial observation[11] and the dynamic Random effects Probit estimates[12] for $t > 1$. The pooled Probit estimator treats the whole sample as a large cross-section therefore assuming away all cross period correlation. This restrictive Probit estimator provides an initial consistent estimate of the parameters although it is inefficient (see Maddala, 1987) and the respective estimates are only provided here to demonstrate the overstatement of the degree of state dependence. The null hypothesis under the likelihood-ratio test for ρ was rejected in each of the estimated Random effects Probit models. This implies that $\rho \neq 0$ and thus the successive error components for the same individual unit are correlated. In the light of this, the reported pooled Probit estimates for $t > 1$ are rejected in favour of the Random effects Probit estimates. Consequently, we cannot fail to accommodate the role of unobserved characteristics in modelling the unionisation decision of individuals.

Regarding both the Heckman and Wooldridge estimators, the proportion of the total error variation attributed to unobserved individual heterogeneity, ρ, was significantly higher in the (1997-2002) dynamic Random effects Probit models for both genders. Further, the unobserved heterogeneity in the female random effects estimates generally constitutes a notably greater proportion of the unexplained variance of the composite error term than it does in the corresponding male estimates.[13] This could signal our failure to use a sufficiently rich set of explanatory variables in order to explain the union determinants of a far more heterogeneous sample that is inclusive of part-time employees which are absent from the male samples.

The higher intra-panel correlation in the second period estimates suggests that unobserved attributes, underlying individual unionisation propensity, have a greater impact in the respective period. Failure to control for the evident correlation among the unobservables will lead to spurious state depedence (see Heckman, 1981a,b). Since state dependence and unobserved heterogeneity offer opposing rationalisations of future individual propensity towards unionisation, union persistence should have a diminishing effect in the second period estimates; this is what we explore in the following Section.

4.1. The Persistence of Trade Union Membership

Turning to the exogeneity of the initial conditions, in the case of the Heckman estimator this requires that $\theta = 0$. It is evident that exogeneity[14] is strongly rejected in all models with the exception of the male (1991-1996) model where it is rejected at the not so stringent 10% level of significance. Hence, the union status

[11] The estimates of the reduced form model for the initial period are not reported here as they are not of direct interest.

[12] The reported Random effects Probit estimates are inclusive of either a set of industrial or occupational controls as the there is a certain degree of overlapping among the two classifications. Unless one of the two sets of controls was excluded, the maximum likelihood functions using the -redprob- command would not converge due to this collinearity. The same exclusions were made in the Wooldridge estimates for comparison purposes.

[13] Apart from the (1997-2002) Heckman estimator where ρ was marginally higher in the model for male employees.

[14] Testing for $\theta = 0$ must take into account that it lies on the boundary of the parameter space (see Stewart 2006) .

of employees at the beginning of each time period under analysis is correlated with unobservable individual-specific characteristics.

With regard to the Wooldridge estimator, the coefficients on the initial value of union membership enter all of the estimated models with particularly strong and statistically significant effects that are much greater in magnitude than the coefficients on the lagged value of union status. This indicates that there is a considerable correlation between the unobserved individual heterogeneity and the initial condition. This correlation becomes even more pronounced in the (1997-2002) estimates for both genders compared to the respective (1991-1996) estimates and it is slightly more accentuated in the male models.

The coefficient on the lagged value of trade union membership enters all of the estimated models with a generally strongly statistically significant coefficient, an outcome that indicates that there is positive state dependence in union membership status even after controlling for the unobserved effect. The prevalent correlation between the initial condition and the unobserved individual heterogeneity renders the pooled Probit estimates inconsistent and the consequent over-statement of the extent of state dependence, across all models, is quite clear.

The pooled Probit and Random effects Probit estimates employ different normalisations. The former uses $\sigma_\nu^2 = 1$ and therefore gives an estimate of γ/σ_ν whereas the latter employing $\sigma_u^2 = 1$ reports γ/σ_u. For comparison purposes the random effects estimates have to be multiplied by the factor $\sqrt{1-\hat{\rho}} = \sigma_u/\ \sigma_\nu$ so that they are converted into γ/σ_ν (see Arulampalam, 1999). The rescaled coefficients[15] on lagged union membership status are reduced even further compared to the "inflated" pooled Probit estimates and in the case of both female (1997-2002) models they are driven to statistical insignificance.

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 y_{t-1} to be fixed at 0 and 1 and are evaluated at $x_{it} = \overline{x}$ (see Wooldridge, 2006; Stewart, 2007):

$$\hat{p}_j = \frac{1}{N}\sum_{i=1}^{N} \Phi\{(\overline{x}'\hat{\beta} + \gamma_j + \overline{x}_i'\hat{a})\sqrt{1-\hat{\rho}}\};\ \hat{p}_0 = \frac{1}{N}\sum_{i=1}^{N} \Phi\{(\overline{x}'\hat{\beta} + \overline{x}_i'\hat{a})\sqrt{1-\hat{\rho}}\} \quad (17)$$

The average partial effects and predicted probability ratios provided at the bottom of Tables (2-5) in the Appendix are an estimate of state dependence of union membership. The Heckman and Wooldridge estimators generally produce similar magnitudes while the respective pooled Probit values are substantially "inflated".

The predicted probability ratios in the case of the Heckman and Wooldridge (1991-1996) estimates suggest that a male worker with a given set of observable and unobservable attributes is 3.1 and 2.4 times, respectively, as likely to be a union member at period t if he had been so at $t-1$. The corresponding probability ratios from the (1997-2002) estimates are 1.3 and 1.6.

A female worker possessing a given set of observable and unobservable characteristics is approximately 2 and 1.7 times as likely to be a union member at period t if she had been so at $t-1$ according to the respective predicted probability ratios

[15] {1.249, 0.988; 0.256, 0.326} for the (1991-1996; 1997-2002) male models for the Heckman and Wooldridge estimators respectively. Similarly, the corresponding values in the female models are {0.663,0.625; 0.090,0.151}.

from the Heckman and Wooldridge (1991-1996) estimates. Similarly, the respective ratios from the (1997-2002) estimates suggest that a female worker is 1.1 and 1.2 times as likely to remain a union member.

Conclusively then, the extent of state dependence in union membership status is markedly higher in the (1991-1996) estimates as opposed to the (1997-2002) estimates concerning both genders. This outcome implies that for British employees during the period under analysis the probability of remaining a union member has declined. This is in line with our expectations following the introduction of the successive (1980, 1982, 1988, 1990) Employment Acts. The resulting reduction in UK trade union bargaining strength has effectively redefined British labour relations by taking away power from unions and "transferring it" to management.

Further, the degree of state dependence in male union membership appears to be more pronounced than the respective female dependence in both time periods under consideration. This is not surprising as the male samples employed consist solely of full-time employees and moreover, male labour market participation is not as discontinuous as it is for female employees.

The prevalent correlation among the initial value of union status, in both time periods, and the unobservable characteristics implies that failure to control for the correlated unobservables overestimates the relative importance of past unionisation experience as a determinant of future union propensity. In fact, our findings indicate a more prominent role for the unobserved determinants underlying union propensity in the (1997-2002) estimates as opposed to past unionisation experience.

4.2. Observed Individual Heterogeneity (Full-Time Male Employees)

Our final research question is assessing the effect of observed characteristics on individual unionisation propensity. We wish to identify the traits that have the most prominent impact on the probability of union membership. Moreover, we are interested in establishing whether using longitudinal data and controlling for the unobserved effects, unlike earlier studies such as Booth (1986) and Wright (1995), implies that both personal and industrial characteristics affect the sorting decision of individuals into unions.

The male estimates for both periods under analysis (refer to Tables 2, 3) indicate that both personal and industrial characteristics affect individual propensity towards unionisation. Thus, controlling for the unobservables and the use of longitudinal data invalidate the assertion that it is mainly the industrial attributes that impact on the union/non-union decision.

The most prominent industrial characteristic effect, during the entire period under consideration, stems from workforce size that appears to be positively associated with union membership. This is in line with the notion that larger establishments are more likely to enjoy market power and thus offer a greater scope for unionisation as there are more quasi-rents to be bargained over.[16] However, establishment size could be acting as a proxy for union recognition (see Disney *et al*, 1996).

The first period estimates (in Table 2) indicate a widespread decline in the probability of union membership across the majority of occupations and further, that individuals in white collar occupations display a greater degree of union aversion. The second period results (see Table 3) reveal a prevalent deunionisation tendency across all industries except the public and education sectors that do not

[16] Hirsch and Addison (1986) provide a comprehensive review of a number of studies establishing this.

seem to share the fate of the former stronghold of trade unions, namely, the British manufacturing industry.

The estimates are consistent with insider-outsider theories (see Lindbeck and Snower, 1986; Blachflower et al, 1990) suggesting that highly qualified employees are less likely to rely on unions in order to extract concessions from employers. It is also plausible that the consequent standardisation of wages within the unionised sector, resulting in reduced human capital premia, impacts negatively on the union propensity of highly qualified employees (see Abowd and Farber, 1982).

Highly skilled individuals within the heavily unionised public and education sectors are, however, found to be more prone to unions with regard to the entire period under analysis. This occurs in that the public and education sectors in the UK still remain relatively heavily unionised and further, they are associated with comparatively greater proportions of highly skilled employees.[17] It is plausible that given the high percentage of employees that are union members in the aforementioned sectors, that non-union members are the most marginalised employees in the sector (see Blanchflower and Bryson, 2007, p.1).

4.3. Observed Individual Heterogeneity (Female Employees)

We anticipate that male and female employees have different propensities towards unionisation. On one hand, the theoretical prediction is that females will have a lower propensity to unionise due to their discontinuous labour market participation. On the other hand, provided that female labour market participants receive, on average, lower wages for discriminatory reasons it is likely that they are more prone towards unions (see, for example, Heywood, 1990).

Concerning the observed individual characteristics' impact on the probability of unionisation, as with the male estimates, the results (provided in Tables 4, 5) suggest that individual propensity towards unionisation is determined by both industrial and personal characteristics.

Out of the industrial attributes the most pronounced impact on the probability of union membership comes from establishment size and it is found to be positively related with union propensity. As with the male estimates, this invalidates the *institutionalist* view that large establishments act as if they were unionised and subsequently offer higher wages in order to avoid unionisation.[18]

Married females appear to be more likely to be union members- an outcome that is consistent with the "exit-voice" scenario suggesting that individuals with family obligations might seek greater job security through unionisation (see Booth, 1986, pp.46-47). Moreover, this could be the result of the active trade union policies tailored to the needs of the female portion of the labour force in an effort to attract more female members in an era of declining male union membership (see Budd and Mumford, 2004). Further, the estimates suggest that part-time employees have a lower probability to be union members possibly owing to their discontinuous labour market participation. It is plausible that costs of union membership might surpass the potential benefits for this group.

[17] The data reveal that the male union membership percentages in the "Public Administration, Education, Other" industrial classification were (59.91, 57.97) during (1991-1996) and (1997-2002) respectively. The corresponding figures for females are (45.85, 49.03) and, as with the male membership rates, they are well above the average union membership rates in the respective samples.

[18] This would reverse the positive association between union membership and the size of establishment (see Brown and Medoff, 1989).

Highly qualified females are found to have a lower unionisation propensity thus, validating Abowd and Farber's (1982) prediction. On the other hand, as with their male counterparts, highly skilled females in the public and education sectors are more likely to be union members. This is consistent with our expectation that high degrees of human capital within the heavily unionised public and education sectors are positively related to unionisation propensity.

Finally, given the observed increase in the female union membership rate, the estimated results display no tendency of prevailing deunionisation across any occupation contrary to the male estimates.

5. SUMMARY AND CONCLUSIONS

In this paper we study the determinants of trade union membership in the UK during 1991-2003. The use of longitudinal data allows us to ascertain whether individual union membership at any point of time is a persistent phenomenon or instead a random process. By employing three alternative methodologies to control for the problem of initial conditions the results suggest that trade union membership is quite persistent even after controlling for the unobserved effect.

There is evidence, however, of a considerable correlation between the unobserved individual heterogeneity and the initial condition and failure to control for this overstates the degree of state dependence of union membership substantially. Failure to control for the correlated unobservables overestimates the relative impact of past unionisation experience as a determinant of future union propensity. This highlights the inappropriateness of the pooled Probit estimator that assumes away any correlation among the successive disturbances.

Heckman's (1981b) estimator is our preferred solution to the initial conditions problem since it incorporates pre-sample information and thus produces a superior log-likelihood to Wooldidge's (2005) estimator extended by a standard Probit estimator for the first period to enable comparability. The comparative advantage of Wooldridge's (2005) estimator, however, lies in its computational simplicity. Orme's (2001) methodology restriction that the correlation between the initial condition and the random effect is sufficiently low was rejected by the data and this provides inconsistent parameter estimates.

The extent of state dependence in union membership status is notably higher in the (1991-1996) period estimates as opposed to the (1997-2002) estimates: an outcome that implies that the probability of remaining a union member, during the entire period under analysis, in the U.K has declined. This is in line with our expectations following the introduction of the successive (1980, 1982, 1988, 1990) Employment Acts and the consequent reduction in UK trade union bargaining strength.

Further, the degree of state dependence in male union membership appears to be more pronounced than the respective female dependence in both time periods under consideration. This is not surprising since the male samples employed in this study consist of full-time employees only and moreover, male labour market participation is not as discontinuous as it is for female employees.

We conclude that during the second period under analysis the role of habit persistence (i.e. the behavioural effect of remaining in unions due to experiencing unionisation in the past) has diminished relative to the impact of unobserved heterogeneity in moulding individual propensity towards unions.

Hence, unobserved individual heterogeneity plays a critical role in modelling union membership status and failing to control for this gives biased estimates. Arulampalam and Booth (2000) make a similar statement and note that this is, of course, nothing new. It is consistent with social custom union theories that propose a plethora of union membership determinants such as commitment to unions and solidarity which are generally not observed by the econometrician (see Arulampalam and Booth, 2000, p.290 and p.308).

However, this is also consistent with the paucity of employer characteristics in our estimations. It should be restated that while employer attributes are captured through the industrial classification and establishment size controls these do not suffice so as to assign any specific effects purely to unobserved heterogeneity (see Vella and Verbeek 1998; Arulampalam and Booth, 2000).

In all respects it is plausible, that given the diminishing importance of trade unions in British industrial relations, those individuals remaining in unions during the second period do so due to factors such as personal or political beliefs which we fail to control for.

The observed heterogeneity estimates suggest that an individual's propensity to unionise is determined by a mixture of industrial and personal characteristics that have a differential impact on male and female propensities. This is at odds with earlier studies, such us Booth (1986) and Wright (1995), generally employing cross-sectional data (hence failing to control for unobservable individual specific effects) and suggesting that it is mainly the industrial characteristics that typically have a significant impact on the propensity to unionise.

Establishment size enters all of the estimated models with the most prominent observed heterogeneity effect, other than lagged trade union status, and is positively associated with union membership. This is in line with the notion that larger establishments are more likely to enjoy market power and hence offer a greater scope for unionisation since there are more quasi-rents to be bargained over. However, establishment size could be acting as a proxy for union recognition (see Disney *et al*, 1996).

Concerning the (1991-1996) male union membership estimates it is evident that there is a widespread decline in the probability of union membership across the majority of occupations and it appears that employees in white collar occupations display a greater degree of union aversion. This is consistent with insider-outsider theories (Lindbeck and Snower, 1986) indicating that highly skilled individuals would be less prone to rely on unions in order to extract concessions from employers as they cannot be rapidly nor costlessly replaced. Furthermore, this outcome is in line with the notion that highly qualified employees might be less likely to seek union membership in that the consequent standardisation of wages results in reduced human capital premia (see Abowd and Farber, 1982).

On the other hand, highly skilled male employees in the heavily unionised public and education sectors are more likely to be union members. This result is reinforced by the respective (1997-2002) estimates that further reveal that the public sector does not seem to share the fate of the British manufacturing industry in terms of rapid deunionisation. Therefore, while concerning the private sector educational attainment appears to be negatively associated with individual unionisation propensity, this association becomes positive with respect to the public and education sectors. Given the combination of a greater proportion of highly qualified individuals and the high percentage of union members in the public and education sectors, it is plausible that non-union members are the most marginalised in the sector (see Blanchflower and Bryson, 2007, p.1).

There is evidence that married females are more prone to unions and this is consistent with the notion that individuals with family responsibilities might seek greater job security through unionisation (see Booth, 1986, pp.46-47). Moreover, this outcome could stem from the active union policies designed to attract more female members in an era of declining male union membership (see Budd and Mumford, 2004). Furthermore, part-time female employees appear less likely to be in unions possibly due to their discontinuous participation in the labour market.

Highly qualified females seem to be less likely to be union members- an outcome that verifies Abowd and Farber's (1982) model prediction. However, as with the male estimates, highly skilled females within the public and education sectors have a higher propensity towards unionisation according to the observed heterogeneity estimates for both periods. This reflects the high representation of women in both the teaching profession and public administration which still remain two of the most heavily unionised industries in the UK. Finally, given the increasing female unionisation rates throughout the period under analysis, the female estimates show no evidence of prevalent deunionisation across any occupation unlike the male estimates.

Future work could entail looking at the individual responses from the two alternative membership questions and the union coverage outcomes in the BHPS in order to detect any potential measurement error and its impact on the estimates of union membership determinants. Further, another interesting extension is evaluating the sensitivity of the Heckman and Wooldridge estimators to the normality assumption on the unobserved individual effects by employing a discrete mass point distribution to model the unobserved heterogeneity (see Stewart, 2007).

References

Abowd, J. M. and Farber, H. S. (1982). 'Job Queues and the Union Status of Workers', *Industrial and Labor Relations Review*, Vol. 35, pp. 354-367.

Andrews, M. and Naylor, R. A. (1994). 'Declining Union Density in the '80s: What Do Panel Data Tell us?', *British Journal of Industrial Relations*, Vol. 32, pp. 413-431.

Andrews, M., Stewart, M. B., Swaffield, J. and Upward, R. (1998). 'The Estimation of Union Wage Differentials and the Impact of Methodological Choices' *Labour Economics*, Vol. 5, pp. 449-474.

Arulampalam, W. (1999). 'A Note on the Estimated Coefficients in Random Effects Probit Models', *Oxford Bulletin of Economics & Statistics*, Vol. 61, pp.597-602.

Arulampalam, W. and Booth, A. L. (2000). 'Union Status of Young Men in Britain: A Decade of Change', *Journal of Applied Econometrics*, Vol. 15, pp. 289-310.

Arulampalam, W., Booth, A. L. and Taylor, M. P. (2000). 'Unemployment Persistence', *Oxford Economic Papers*, Vol.52, pp.24-50.

Berndt, E., Hall, B., Hall, R. and Hausman, J. (1974). 'Estimation and Inference in Nonlinear Structural Models', *Annals of Economic and Social Measurement*, pp. 653-665.

Blanchflower, D. G. and Bryson, A. (2007). 'The Wage Impact of Trade Unions in the UK Public and Private Sectors', *IZA*, Discussion Paper, No.3055.

Blanchflower, D. G., Oswald, A. J. and Garrett, M. D. (1990). 'Insider Power in Wage Determination', *Economica*, Vol. 57, pp. 143-170.

Booth, A. L. (1986). 'Estimating the Probability of Trade Union Membership: A Study of Men and Women in Britain', *Economica*, Vol. 53, pp. 41-61.

Brown, C. and Medoff, J. (1989). 'The Employer Size-Wage Effect', *Journal of Political Economy*, Vol. 97, pp.1027-1059.

Budd, J. and Mumford, K. (2004). 'Trade Unions and Family Friendly Work Policies in Britain', *Industrial and Labor Relations Review*, Vol. 57, pp. 204-222.

Butler, J. and Moffitt, R. (1982). 'A Computationally Efficient Quadrature Procedure for the One Factor Multinomial Probit Model', *Econometrica*, Vol. 50, pp. 761 -764.

Chamberlain , G. (1984). 'Panel Data', *Handbook of Econometrics*, Ch.22, Griliches Z. and Intrilligator M., editors, North Holland, Amsterdam.

Disney, R., Gosling, A. and Machin, S. (1996). 'What Has Happened to Union Recognition in Britain', *Economica*, Vol.63, pp. 1-18.

Disney, R., Gosling, A., McCrae, J. and Machin, S. (1999). 'The Dynamics of Union Membership in Britain - A Study Using the Family and Working Lives Survey', *Department of Trade and Industry*, Research Report No.3.

Farber, H. S. (1983). 'The Determination of the Union Status of Workers', *Econometrica*, Vol. 51, pp.1417-1438.

Frechette, G. R. (2002). '*Superxt: to solve convergence problems associated with random-effects estimators: Version 1.0.0*'.

Gould, W., Pitblado, J. and Sribney, W. (2003). 'Maximum Likelihood Estimation with Stata', *Stata Press*.

Gourieroux, C., Monfort, A., Renault, E. and Trognon, A. (1987). 'Generalised Residuals', *Journal of Econometrics*, Vol. 34, pp.5-32.

Gregg, P. and Naylor, R. A. (1993). 'An Inter-Establishment Study of Union Recognition and Membership in Great Britain', *Manchester School*, Vol. 61, pp.

367-385.

Heckman, J. J. (1981a). 'Statistical Models for Discrete Panel Data', in Manski, C. F., and McFadden, D. L., Editors, '*Structural Analysis of Discrete Data and Econometric Applications*', Cambridge: The MIT Press.

Heckman, J. J. (1981b). 'The Incidental Parameters Problem and the Problem of Initial Conditions in Estimating a Discrete Time-Discrete Data Stochastic Process', in Manski, C. F. and McFadden, D. L., Editors, '*Structural Analysis of Discrete Data and Econometric Applications*', Cambridge: The MIT Press.

Hernández Quevedo, C., Jones, A. M. and Rice, N. (2004). 'Reporting Bias and Heterogeneity in Self-Assessed Health. Evidence from the British Household Panel Survey', *ECuity III Working Paper*, No.19.

Heywood, J.S. (1990). 'Who Queues for a Union Job?', *Industrial Relations*, Vo.29, pp. 99-127.

Hirsch, B. and Addison, J. (1986). 'The Economic Analysis of Unions', *Allen and Unwin*.

Hsiao, C. (2003). 'Analysis of Panel Data', *Econometric Society Monographs*, Camrbidge University Press, 2nd ed.

Lindbeck, A. and Snower, D. (1986) 'Wage Setting, Unemployment and Insider-Outsider Relations', *American Economic Review*, Papers and Proceedings, 76, pp. 235-239.

Maddala, G. S. (1987). 'Limited Dependent Variable Models Using Panel Data', *Journal of Human Resources*, Vol. 22, pp.307-338.

Mundlak, Y. (1978). 'On the Pooling of Time Series and Cross Section Data', *Econometrica*, Vol. 46, pp.69-85.

Naylor, R. A. and Raaum, O. (1993). 'The Open Shop Union, Wages and Management Opposition', *Oxford Economic Papers*, Vol. 45, pp. 589-604.

Newey, W. K. (1987). 'Efficient Estimation of Limited Dependent Variable Models with Endogenous Explanatory Variables', *Journal of Econometrics*, Vol. 36, pp. 231-250.

Orme, C. D. (2001). 'Two Step Inference in Dynamic Non-Linear Panel Data Models', School of Economic Studies, *University of Manchester*, mimeo.

Stewart, M. B. (1995). 'Union Wage Differentials in an Era of Declining Unionization', *Oxford Bulletin of Economics & Statistics*, Vol. 57, pp.143-166.

Stewart, M. B. (2006). '-redprob- A Stata Program for the Heckman Estimator of the Random Effects Dynamic Probit Model', *University of Warwick*, mimeo.

Stewart, M. B. (2007). 'The Inter-related Dynamics of Unemployment and Low-Wage Employment', *Journal of Applied Econometrics*, Vol. 22(3), pp. 511-531.

Swaffield, J. (2001). 'Does Measurement Error Bias Fixed-Effects Estimates of the Union Wage Effect?', *Oxford Bulletin of Economics & Statistics*, Vol. 63, pp. 437-457.

University of Essex, *Institute for Social and Economic Research*, British Household Panel Survey; Waves 1-12, 1991-2003 [computer file]. Colchester, Essex: UK Data Archive [distributor], SN: 4967, June, 2004.

Vella, F. and Verbeek, M. (1998).'Whose Wages Do Unions Raise? A Dynamic Model of Unionism and Wage Rate Determination for Young Men', *Journal of Applied Econometrics*, Vol. 13, pp.163-183.

Vella, F. and Verbeek, M. (1999b). 'Two-step Estimation of Panel Data Models with Censored Endogenous Variables and Selection Bias', *Journal of Econometrics*, Vol. 90, pp. 239-263.

Wooldridge, J. M. (2005). 'Simple Solutions to the Initial Conditions Problem in Dynamic, Nonlinear Panel Data Models with Unobserved Effects', *Journal of Applied Econometrics*, Vol. 20, pp.39-54.

Wooldridge, J. M. (2006). 'Unobserved Heterogeneity and Estimation of Average Partial Effects', *Department of Economics, Michigan State University*, mimeo.

Wright, P. (1995). 'Union Membership and Coverage: An Econometric Study Using the SCELI Dataset', *International Journal of Manpower,* Vol.16, pp.53-59.

Appendix

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.**
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
(Public Admin/Educ/Other)*(University/Higher Qual)	0.061	0.240	0.079	0.270	0.052	0.222	0.101	0.301
(Public Admin/Educ/Other)*(Vocational Qualification)	0.034	0.182	0.052	0.222	0.024	0.152	0.052	0.221
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 Services	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 (1991-1996), Males

	Pooled Probit		Heckman		Wooldridge	
Variable	**Coef.**	**z**	**Coef.**	**z**	**Coef.**	**z**
Lagged Trade Union Membership	2.562	43.04	1.731	10.38	1.331	10.33
Trade Union Membership (1)	_	_	_	_	2.164	8.46
Log(1+Potential Experience)	0.117	1.85	0.349	2.80	0.068	0.61
Marital Status	0.039	0.20	-0.047	-0.21	-0.046	-0.20
Mean(Marital Status)	0.060	0.29	0.289	1.11	0.238	0.89
Black(Caribbean, African, Other)	-0.096	-0.37	0.161	0.36	-0.516	-1.11
Asian(Indian, Pakistani, Chinese, Other)	0.431	0.63	0.169	0.19	-0.266	-0.23
Other Ethnic Minority Group	0.119	0.37	0.369	0.68	-0.040	-0.07
Inner/ Outer London and R of South East	-0.094	-1.06	-0.203	-1.35	-0.203	-1.28
South West	-0.146	-1.26	-0.351	-1.75	-0.183	-0.90
Scotland	-0.269	-2.03	-0.439	-1.97	-0.366	-1.58
Wales	0.171	1.17	0.362	1.49	0.407	1.56
North West	0.095	0.87	0.291	1.51	0.053	0.28
North East	-0.071	-0.70	-0.015	-0.09	-0.221	-1.22
East Anglia	-0.183	-1.16	-0.284	-1.07	-0.253	-0.91
(Public Admin/Educ/Other)*(Univ/Higher Qual)	0.604	3.86	1.133	4.26	0.727	2.85
(Public Admin/Educ/Other)*(Voc Qualification)	0.584	2.85	0.940	2.83	0.882	2.48
University Degree or Higher	-0.079	-0.51	-0.142	-0.58	-0.088	-0.34
HND, HNC, Teaching	0.083	0.50	0.237	0.86	0.146	0.50
A Levels	0.057	0.58	0.180	1.10	0.119	0.69
O Levels or CSE	0.086	0.99	0.178	1.23	0.170	1.12
Fair, Poor, V Poor Self-Assessed Health	0.143	1.29	0.153	1.20	0.156	1.15
Mean(Fair,Poor,V Poor Self-Assessed Health)	-0.264	-1.60	-0.306	-1.28	-0.281	-1.11
Workforce >500	0.441	4.69	0.673	5.22	0.545	3.93
Workforce 100-499	0.377	4.36	0.507	4.42	0.413	3.24
Workforce 25-99	0.215	2.41	0.252	2.19	0.208	1.63
Professional Occupations	-0.560	-3.36	-0.831	-3.37	-0.777	-2.90
Managers & Administrators	-0.648	-4.24	-1.028	-4.31	-0.817	-3.30
Associate Professional & Technical	-0.446	-2.74	-0.747	-3.06	-0.526	-2.02
Clerical & Secretarial	-0.307	-1.94	-0.553	-2.37	-0.387	-1.55
Craft & related	-0.260	-1.79	-0.343	-1.65	-0.232	-1.00
Personal & Protective Service	0.028	0.17	-0.104	-0.42	0.233	0.87
Sales	-0.607	-2.56	-0.894	-2.77	-0.541	-1.56
Plant & Machine Operatives	-0.317	-2.15	-0.480	-2.21	-0.499	-2.11
Time Dummy 1993	_	_	_	_	-0.002	-0.01
Time Dummy 1994	-0.035	-0.45	-0.047	-0.53	-0.037	-0.34
Time Dummy 1995	0.057	0.72	0.040	0.45	0.075	0.68
Time Dummy 1996	0.133	1.67	0.140	1.56	0.211	1.89
Constant	-1.782	-6.42	-2.373	-4.89	-2.198	-4.59
ρ	_	_	0.479	4.11	0.449	7.15
θ	_	_	3.140	1.82	_	_
Average Partial Effect	0.800		0.462		0.375	
Predicted Probability Ratio	9.007		3.138		2.449	
Log-Likelihood	-1142.13		-1524.70		-1066.65	

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

TABLE 3

Dynamic Random Effects Probit Models of Union Membership (1997-2002), Males

	Pooled Probit		Heckman		Wooldridge	
Variable	**Coef.**	**z**	**Coef.**	**z**	**Coef.**	**z**
Lagged Trade Union Membership	2.412	42.41	0.697	4.47	0.569	4.67
Trade Union Membership (1)	_	_	_	_	3.680	11.50
Log(1+Potential Experience)	0.017	0.24	0.033	0.10	0.044	0.25
Marital Status	0.038	0.20	0.263	0.95	0.263	0.97
Mean(Marital Status)	-0.053	-0.27	-0.036	-0.09	-0.245	-0.77
Black(Caribbean, African, Other)	0.436	1.68	1.289	1.97	1.104	1.77
Asian(Indian, Pakistani, Chinese, Other)	0.058	0.11	1.197	1.52	-0.934	-0.59
Other Ethnic Minority Group	-0.123	-0.37	-1.612	-2.47	-1.481	-1.39
Inner/ Outer London and R of South East	-0.130	-1.54	-0.221	-0.88	-0.326	-1.50
South West	0.077	0.71	0.384	0.83	0.368	1.32
Scotland	0.062	0.51	0.512	1.73	0.151	0.49
Wales	0.345	2.63	1.633	4.60	0.732	2.24
North West	0.205	1.95	1.636	3.25	0.295	1.09
North East	-0.009	-0.10	0.525	1.64	-0.041	-0.18
East Anglia	0.034	0.24	0.718	1.84	0.061	0.17
(Public Admin/Educ/Other)*(Univ/Higher Qual)	0.403	2.35	0.798	2.13	0.319	0.84
(Public Admin/Educ/Other)*(Voc Qualification)	0.223	1.03	1.645	3.12	0.472	0.96
University Degree or Higher	-0.448	-3.30	-1.175	-2.69	-0.617	-1.88
HND, HNC, Teaching	-0.381	-2.65	-1.644	-3.52	-0.817	-2.27
A Levels	-0.010	-0.10	-0.021	-0.05	0.168	0.69
O Levels or CSE	0.000	0.00	-0.069	-0.19	0.204	0.88
Fair, Poor, V Poor Self-Assessed Health	-0.102	-1.03	-0.135	-0.99	-0.154	-1.13
Mean(Fair,Poor,V Poor Self-Assessed Health)	0.009	0.06	-0.347	-1.02	0.123	0.39
Workforce >500	0.604	6.76	1.168	6.26	0.902	5.37
Workforce 100-499	0.615	7.37	1.169	6.06	0.974	6.17
Workforce 25-99	0.356	4.05	0.641	3.67	0.548	3.46
Agriculture, Forestry & Fishing	-1.187	-2.30	-2.613	-1.97	-2.580	-2.20
Extraction of Minerals & Manufacture of Metals	-0.796	-3.70	-1.481	-2.83	-1.303	-3.16
Metal Goods, Engineering & Vehicles Industries	-0.941	-4.73	-1.157	-3.26	-1.086	-2.85
Other Manufacturing Industries	-0.826	-4.21	-1.157	-3.31	-1.189	-3.21
Construction	-0.876	-3.97	-1.276	-3.29	-1.104	-2.67
Distribution, Hotels & Catering (Repairs)	-1.122	-5.34	-1.767	-4.73	-1.500	-3.72
Transport & Communication	-0.475	-2.37	-0.645	-1.66	-0.770	-1.97
Banking & Finance	-0.837	-4.23	-1.343	-3.67	-1.129	-2.85
Public Administration, Education, Other	-0.351	-1.79	-0.199	-0.53	-0.280	-0.72
Time Dummy 1999	_	_	_	_	-0.359	-3.03
Time Dummy 2000	0.031	0.41	0.002	0.02	-0.178	-1.51
Time Dummy 2001	-0.031	-0.40	-0.031	-0.29	-0.202	-1.70
Time Dummy 2002	0.112	1.41	0.230	1.95	0.056	0.45
Constant	-1.102	-3.17	-1.615	-1.12	-2.149	-2.60
ρ	_	_	0.865	26.76	0.672	16.72
θ	_	_	1.164	6.03	_	_
Average Partial Effect	0.767		0.095		0.089	
Predicted Probability Ratio	9.737		1.310		1.592	
Log-Likelihood	-1258.97		-1595.28		-1078.95	

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

TABLE 4

Dynamic Random Effects Probit Models of Union Membership (1991-1996), Females

	Pooled Probit		Heckman		Wooldridge	
Variable	**Coef.**	**z**	**Coef.**	**z**	**Coef.**	**z**
Lagged Trade Union Membership	2.095	37.92	1.083	10.15	0.909	9.01
Trade Union Membership (1)	_	_	_	_	2.072	10.70
Log(1+Potential Experience)	0.082	1.67	0.132	1.25	0.095	0.98
Marital Status	-0.115	-0.75	-0.120	-0.65	-0.157	-0.83
Mean(Marital Status)	0.352	2.09	0.621	2.57	0.589	2.50
Full-Time Employment	0.037	0.30	0.109	0.72	0.138	0.90
Mean(Full-Time Employment)	0.261	1.80	0.755	3.28	0.297	1.39
Maternity Leave	0.026	0.10	-0.049	-0.17	-0.074	-0.25
Mean(Maternity Leave)	0.857	1.34	2.765	2.05	1.638	1.35
Black(Caribbean, African, Other)	0.317	1.03	0.682	0.97	0.554	0.87
Asian(Indian, Pakistani, Chinese, Other)	-0.010	-0.03	0.447	0.57	-0.496	-0.72
Other Ethnic Minority Group	-0.171	-0.42	-0.483	-0.51	-0.690	-0.75
Inner/ Outer London and R of South East	-0.148	-1.75	-0.337	-1.86	-0.117	-0.70
South West	-0.071	-0.57	-0.188	-0.68	-0.055	-0.22
Scotland	0.215	2.03	0.524	2.22	0.397	1.85
Wales	0.258	1.98	0.641	2.17	0.395	1.48
North West	0.012	0.12	0.090	0.40	-0.098	-0.47
North East	0.254	2.86	0.608	3.03	0.384	2.12
East Anglia	-0.079	-0.51	-0.347	-1.00	0.046	0.15
(Public Admin/Educ/Other)*(Univ/Higher Qual)	0.809	3.91	1.369	3.87	0.840	2.45
(Public Admin/Educ/Other)*(Voc Qualification)	1.207	4.04	2.147	4.29	1.641	3.37
University Degree or Higher	-0.407	-1.96	-0.315	-0.83	-0.298	-0.82
HND, HNC, Teaching	-0.621	-2.20	-0.773	-1.60	-0.671	-1.43
A Levels	0.141	1.42	0.434	2.05	0.282	1.45
O Levels or CSE	0.089	1.18	0.252	1.54	0.147	0.98
Fair, Poor, V Poor Self-Assessed Health	0.121	1.39	0.129	1.24	0.122	1.15
Mean(Fair,Poor,V Poor Self-Assessed Health)	-0.253	-1.89	-0.357	-1.42	-0.247	-1.07
Workforce >500	0.340	4.23	0.534	4.02	0.402	3.02
Workforce 100-499	0.247	3.31	0.335	2.73	0.291	2.36
Workforce 25-99	0.183	2.62	0.209	1.91	0.139	1.26
Professional Occupations	-0.119	-0.80	-0.293	-1.07	-0.435	-1.62
Managers & Administrators	-0.359	-2.60	-0.599	-2.42	-0.556	-2.30
Associate Professional & Technical	-0.021	-0.16	0.046	0.19	-0.052	-0.22
Clerical & Secretarial	-0.212	-1.92	-0.434	-2.01	-0.322	-1.56
Craft & related	-0.006	-0.04	0.112	0.32	0.044	0.13
Personal & Protective Service	-0.028	-0.23	-0.028	-0.13	0.064	0.29
Sales	-0.356	-2.49	-0.679	-2.57	-0.430	-1.68
Plant & Machine Operatives	-0.297	-1.74	-0.369	-1.16	-0.327	-1.06
Time Dummy 1993	_	_	_	_	0.166	1.66
Time Dummy 1994	-0.109	-1.50	-0.127	-1.46	-0.024	-0.24
Time Dummy 1995	0.085	1.18	0.113	1.30	0.216	2.11
Time Dummy 1996	0.257	3.59	0.401	4.47	0.532	5.06
Constant	-2.086	-9.33	-3.080	-6.33	-3.160	-7.04
ρ	_	_	0.625	13.05	0.527	11.60
θ	_	_	1.098	6.54	_	_
Average Partial Effect	0.704		0.247		0.244	
Predicted Probability Ratio	6.407		1.964		1.712	
Log-Likelihood	-1446.08		-1758.06		-1331.85	

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

TABLE 5

Dynamic Random Effects Probit Models of Union Membership (1997-2002), Females

	Pooled Probit		Heckman		Wooldridge	
Variable	**Coef.**	**z**	**Coef.**	**z**	**Coef.**	**z**
Lagged Trade Union Membership	1.995	38.46	0.235	2.26	0.298	2.92
Trade Union Membership (1)	_	_	_	_	3.303	12.95
Log(1+Potential Experience)	-0.015	-0.26	-0.176	-0.73	0.005	0.03
Marital Status	-0.010	-0.06	0.193	0.86	0.212	0.95
Mean(Marital Status)	0.257	1.50	0.710	2.15	0.441	1.55
Full-Time Employment	-0.067	-0.61	-0.022	-0.15	0.012	0.09
Mean(Full-Time Employment)	0.288	2.23	1.074	3.40	0.400	1.59
Maternity Leave	-0.025	-0.11	-0.148	-0.48	-0.131	-0.42
Mean(Maternity Leave)	0.226	0.41	1.689	1.04	0.854	0.54
Black(Caribbean, African, Other)	0.059	0.28	0.233	0.47	-0.112	-0.16
Asian(Indian, Pakistani, Chinese, Other)	0.843	2.04	4.272	2.92	1.973	1.50
Other Ethnic Minority Group	0.536	1.69	1.899	2.52	1.496	1.56
Inner/ Outer London and R of South East	-0.211	-2.83	-0.989	-3.83	-0.422	-1.89
South West	-0.148	-1.50	-0.532	-1.66	-0.449	-1.54
Scotland	0.168	1.75	0.498	1.52	0.426	1.48
Wales	0.322	2.70	1.303	2.96	0.728	2.02
North West	0.015	0.17	-0.322	-0.83	0.038	0.14
North East	0.128	1.57	0.353	1.34	0.365	1.54
East Anglia	-0.280	-2.03	-1.090	-2.26	-0.751	-1.85
(Public Admin/Educ/Other)*(Univ/Higher Qual)	1.110	6.00	3.006	5.96	2.324	4.98
(Public Admin/Educ/Other)*(Voc Qualification)	0.472	2.50	1.005	1.44	0.476	1.03
University Degree or Higher	-0.565	-3.02	-1.059	-2.14	-0.976	-1.98
HND, HNC, Teaching	-0.107	-0.63	0.249	0.30	0.602	1.32
A Levels	0.127	1.46	0.682	2.43	0.579	2.24
O Levels or CSE	0.089	1.15	0.247	1.01	0.370	1.60
Fair, Poor, V Poor Self-Assessed Health	0.017	0.21	-0.042	-0.38	-0.126	-1.13
Mean(Fair,Poor,V Poor Self-Assessed Health)	-0.079	-0.65	-0.267	-0.64	0.039	0.13
Workforce >500	0.430	6.01	0.781	4.67	0.705	4.70
Workforce 100-499	0.251	3.73	0.649	4.38	0.555	4.02
Workforce 25-99	0.187	2.96	0.297	2.32	0.221	1.79
Professional Occupations	0.080	0.58	0.307	1.10	0.119	0.42
Managers & Administrators	-0.312	-2.49	-0.266	-1.03	-0.305	-1.16
Associate Professional & Technical	-0.001	-0.01	0.176	0.70	0.115	0.45
Clerical & Secretarial	-0.181	-1.67	-0.023	-0.09	-0.154	-0.63
Craft & related	-0.137	-0.70	-0.109	-0.27	-0.146	-0.35
Personal & Protective Service	-0.069	-0.59	0.178	0.68	0.154	0.59
Sales	-0.308	-2.39	-0.421	-1.49	-0.304	-1.09
Plant & Machine Operatives	-0.236	-1.39	0.007	0.02	-0.204	-0.57
Time Dummy 1999	_	_	_	_	-0.672	-6.30
Time Dummy 2000	0.360	5.56	0.550	5.98	0.230	2.22
Time Dummy 2001	-0.302	-4.45	-0.142	-1.53	-0.494	-4.58
Time Dummy 2002	0.387	5.94	0.701	7.31	0.380	3.60
Constant	-1.638	-6.35	-2.767	-2.62	-3.275	-4.27
ρ	_	_	0.853	42.57	0.744	27.09
θ	_	_	0.729	8.19	_	_
Average Partial Effect	0.672		0.036		0.047	
Predicted Probability Ratio	4.157		1.080		1.214	
Log-Likelihood	-1766.94		-1980.20		-1485.09	

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