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Trade Unions and Unpaid Overtime in Britain

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NON-TECHNICAL SUMMARY

Trade unions have traditionally focused on working time aspects and have campaigned extensively against long working hours in Britain. An aspect of this long working hours "culture" is the phenomenon of unpaid overtime, which has increased in extent and importance in Britain and other industrialized countries during the last decades. This paper examines the relationship between unionization and unpaid overtime in Britain. We suggest that the impact of union status on the amount of unpaid overtime will depend on the nature of the firm and the sector in which it operates.

Analysis of the first seventeen waves of the British Household Panel Survey (BHPS) confirms these hypotheses. In the for-profit, non-caring sector of the economy, being covered by a trade union in the workplace leads to less unpaid overtime, probably because unions protect employees from employer coercion and negotiate standardized reward and promotion procedures which provide no long-term incentives to work extra hours. On the other hand, in the non-profit, caring sector, union members work more unpaid overtime than covered non-members. Evidence is presented in favour of a specific pro-social ethos of union members: being a union member is associated with a higher probability of belonging to any other social or interest group organization in Britain. It appears the pro-social attitudes of union members lead them to donate extra working time in the non-profit, caring sector, while union coverage has no effect.

The above results are important for two different reasons. First, they enhance our understanding of what unions do in the contemporary British labour market, contributing to the broad literature of union effects on various aspects of the employment relationship. Second, they can form the basis of further research on the overall attitudes and beliefs of union members, something that has been ignored in the economic analysis of the union membership decision.

Trade Unions and Unpaid Overtime in Britain

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Abstract

In this paper we use British Household Panel Survey data to examine the relationship between unionization and unpaid overtime in Britain. The findings indicate that in the for-profit, non-caring sector of the economy, union covered employees supply fewer unpaid overtime hours than non-covered ones due to union protection and the weakening of economic incentives caused by union bargaining. On the other hand, in the non-profit, caring sector, union members offer more unpaid extra hours than covered non-members because of their specific pro-social motivations. Additional evidence is presented that confirms that union members are actually characterized by a specific pro-social ethos.

JEL Classification: J22, J51
Keywords: trade unions; unpaid overtime; pro-social motivation

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1. Introduction

The 26th of February 2010 was declared by trade unions in Britain as the *Work Your Proper Hours Day*, "the day when the average person who does unpaid overtime would start to get paid if they did all their unpaid overtime at the start of the year" (TUC, 2010). An annual campaign by British unions has been taking place for some years now with the aim to raise awareness among employees and the broad public of the incidence and extent of unpaid extra hours and the possible harmful effects it can have on employees' overall welfare. These campaigns are, of course, consistent with the traditional interest of trade unions in working hours and their "fair remuneration", as well as their recent contribution to the overall political discourse in Britain concerning long working hours and "work-life balance".[1]

Unpaid overtime has increased in the UK during the last decades, though it seems that it has come to a halt lately. Campbell and Green (2002), using UK Labour Force Survey (LFS) data, report a rising proportion of employees working extra hours for no pay, from 12.8 percent in 1983 to 23.7 percent in 1998. Also using the LFS, we calculate that the incidence of unpaid overtime has stabilized at over 20 percent since then, while the average usual weekly hours of unpaid overtime for those workers that do unpaid extra time has fluctuated around 6.5-7 hours since the '90s.[2] The main occupations reporting large amounts of unpaid overtime are managers and professionals, though one can detect significant amounts in other occupational categories as well (e.g. sales occupations; see Section 3). The apparent increase in the incidence of unpaid work in the last decades is not only a feature of the UK labour market. Observers in Australia are also concerned with the extent of unpaid overtime and its relatively recent rise (Campbell, 2005), while in Japan it has been linked to *karōshi* or "death by overwork", a legally recognized work-related cause of death in the country since the '80s (*The Economist*, 2007).

The aim of this paper is to empirically examine the relationship between unionization and unpaid overtime, using data from the first 17 waves (1991-2007) of the British Household Panel Survey (BHPS). The relationship between unionism and the extent of unpaid overtime working is a relatively empirically unexplored area in the literature of trade union effects on labour market outcomes (see e.g. Booth, 1995, for a survey). This seems odd at first sight given the importance of

[1] See Kodz *et al.* (2003) for a report on long working hours in the UK sponsored by the government and Bunting (2005) for a popularized account; see also Walsh (2010) for a more general consideration of the "work-life balance" issue.
[2] The numbers reported here refer to fulltime employees. Both incidence and hours of unpaid overtime are lower for part-time workers.

the subject for a more complete understanding of workers' welfare and the availability of relevant data, at least for the developed economies of Western Europe, North America and Australasia. Also, trade unions themselves show interest in this issue and in working time aspects more generally, as the opening paragraph demonstrated for the British case. *All Work and No Pay* (Unison, 2003) and *Something for Nothing* (Fear and Denniss, 2009) are characteristic slogans/titles of union-related research on unpaid overtime in Britain and in Australia respectively. They are indicative of the interest that trade unions show in the impact of extra unpaid hours on the welfare of their members and employees in general.

The structure of this paper is as follows: the following section describes and comments briefly on some papers that have empirically examined the determinants of unpaid overtime, building on various motivations and theoretical approaches. Section 3 outlines our own theoretical framework, while section 4 describes the testable predictions of our hypotheses in conjunction with the econometric methods that are employed to estimate models of unpaid overtime. In this section we also describe our data and comment on the distribution of unpaid overtime across various job characteristics. In section 5 we present the results and discuss our findings in detail, while section 6 presents various checks of the sensitivity and robustness of our results. Finally, section 7 concludes.

2. Related Literature

A number of recent papers have studied the overall determinants of unpaid overtime. Bell and Hart (1999) use UK Labour Force Survey data to study the correlates of unpaid overtime. Based on various theoretical hypotheses including, for example, uncertainty over task completion times, teamwork and the gift-exchange hypothesis, they emphasize the importance of unpaid overtime in the UK labour market, especially for more high ranking occupational groups. Their results show a positive relationship between unpaid overtime hours and higher wages, team leadership status and lower productivity, while they find a negative association between union membership and unpaid overtime. Moreover, they show that accounting for unpaid overtime hours in standard estimated log-wage equations significantly decreases the returns to education, experience and tenure.

Gregg *et al.* (2008) view unpaid extra hours as donated labour, driven by a pro-social motivation of employees in the non-profit, "caring" (education, health and social work) sectors of the economy. Using the BHPS, they provide evidence that supports their claims. They find that employees in the non-profit, caring sector are more likely to offer unpaid overtime than comparable employees in the for-profit, caring sector. Moreover, their analysis shows that employees select themselves into the non-profit and for-profit sectors according to their propensity to supply unpaid overtime. As regards the trade union effect, in their pooled OLS model of unpaid overtime both the coverage and membership coefficients are negative and statistically significant, while in their fixed effects model only the coverage coefficient remains so, though it is dropping in magnitude. We discuss their paper below in more detail, since a part of our argument builds on their theoretical framework.

Van Echtelt *et al.* (2007) relate the phenomenon of unpaid overtime working with intra-firm arrangements in the broad context of the so-called “post-Fordist” or “high-performance” workplace and human resource practices. Using Dutch data from 30 organizations, they find evidence that employees in post-Fordist workplaces (captured by proxies of workplace practices reported by the firms’ management) supply more unpaid overtime hours than comparable employees in other workplaces. However, the authors acknowledge that the specific channels through which this result operates are not clarified by their results.

Based on a different theoretical framework, Engellandt and Riphahn (2005) are interested in the behavioural effects of different types of employment contracts. They postulate that workers in temporary contracts should exert more effort than comparable employees in a permanent employment status. This behaviour is the result of the career concerns of temporary workers who generally view their current status as a “stepping stone” towards a more favourable (in terms of remuneration and general employment terms) permanent employment contract. Unpaid overtime (along with work attendance/absence) is used by the authors as a measure of observable effort. Hence, the scope of their study is not a theoretical explanation and empirical investigation of the determinants of unpaid overtime, but an examination of the incentive effects that a temporary employment status might have on employees. Using data from the Swiss Labour Force Survey, they find evidence confirming their hypothesis: temporary employees are significantly more likely to offer unpaid extra hours than comparable permanent ones.

The analysis of *paid* overtime, on the other hand, has received more attention in the literature and, especially for our specific interest in this paper, its relation with trade union bargaining has been studied in detail.[3] For the US, Trejo (1993) finds that union coverage is positively related with the prevalence of overtime premium pay and negatively related with the incidence and hours of paid

[3] See Hart (2004) for a detailed treatment of (mainly) paid overtime, covering both theoretical and empirical aspects.

overtime. Kalwij and Gregory (2005) using UK data from the New Earnings Survey show that while overtime increases when standard hours are reduced and decreases when the wage rate is increased, unionization is a weak determinant of paid overtime incidence and hours. In general, though these findings have their own theoretical and empirical importance, paid overtime is a quite distinct phenomenon to unpaid overtime. In the words of Campbell (2005; p. 6), "[p]aid overtime is transparent, but unpaid overtime is more opaque, and in the latter case it is always useful to look closely at the basis of the exchange between employer and employee". In the next section, we try to explicitly theorize on the determinants of unpaid overtime, with a specific focus on the channels through which unionization and union status can affect its prevalence.

It is apparent that empirical work is lacking concerning the impact of trade unions on unpaid overtime. The data that we use provide rich information and are very suitable for this purpose. We utilize the arguments and findings in the related literature as a basis in order to construct our own distinct arguments and testable hypotheses and this will be obvious in the overall discussion that follows. Our aim is to point to the specific channels through which the union effect operates and to look for differential effects that depend on the nature of the firm and the industry in which the employees work. This is also an aspect where we depart from all previous papers described above that estimated pooled models of unpaid overtime, without considering possibly different determinants based on, for example, industry or sector.

3. Theory and Hypotheses

The incidence of unpaid overtime hours undertaken by employees in modern workplaces cannot easily be explained by reference to the simple labour-leisure choice model of economic theory. The fact that additional effort in the form of working time is undertaken without direct remuneration contradicts any simple textbook definition of the labour process, such as that of Boeri and van Ours (2008, p. 3): "In order to be in the labour market, there must be an exchange of a labour service for a wage". In view of that, in this section we outline some reasons that can explain the phenomenon of unpaid overtime and we hypothesize on how trade union activity and union membership status can affect those reasons and, consequently, employees' behaviour.

Why, then, do employees offer extra working hours if there is no direct payment for these? First, explicit or implicit coercion from the part of employers can make the worker supply extra hours in fear of dismissal or other related retaliating behaviour. Second, and assuming absence of any form of coercion, workers may have career concerns and expectations that they think can be successfully advanced through exercising more effort in the form of unpaid extra hours. In this way, employees view unpaid work in the present as a form of investment for higher future rewards (Campbell and Green, 2002; Anger, 2005; Pannenberg, 2005). This incentive is stronger in occupations or workplace settings where output or performance cannot be directly observed and measured. Employers in such cases can judge individual performance by reference to observable effort such as working hours and, then, make decisions about promotions or pay rises. The individual employee knows that and offers more, and even unpaid, hours in return for future rewards (van Echtelt *et al.*, 2007, pp. 42-43).

Third, employees may not only be motivated by *extrinsic* rewards in their decision about offering extra unpaid hours. *Intrinsic* motivations may as well drive worker's behaviour (Serra *et al.*, 2010). In this respect, unpaid overtime can be seen as a kind of pro-social behaviour arising as a result of a social service ethos in industries and workplace settings where such behaviours can be relevant, as is the case in education and healthcare (Gregg *et al.*, 2008). According to Gregg *et al.* (2008), a not-for-profit character of the organization in which the employee works is also crucial for such behaviours to manifest themselves. They rely on the following reasoning concerning the importance of the not-for-profit character of the organization: employees' donated labour (i.e. unpaid overtime) in a for-profit setting will be expropriated by the employer; hence the preferred outcome (quality of patient care, in their example for the health sector) will not be achieved. Knowing this, the employees will "decide not to donate their labour in the first place" (*ibid.*, p.2).

Having outlined the above possible reasons for employees to undertake unpaid extra hours, we turn now to the question of interest: how can trade union organization and membership affect the incidence and amount of unpaid overtime? In theory, a plausible answer is that it depends on the reason unpaid work is observed and this is the main argument of the paper. The first two reasons above should be weakened under trade unionism. The managerial prerogative and arbitrariness within the workplace context is moderated under trade union presence, while career concerns are not as effective in inducing extra work as in the non-union sector, due to the standardization of pay and more regulated, collective decisions over pay rises or personnel promotions within the unionized firms (Metcalf *et al.*, 2001). For example, objective rules for pay determination such as seniority are more likely to be followed in unionized establishments than in non-unionized ones. In the latter, individualistic rules like merit or performance related pay is more usually the norm (Zangelidis, 2008). Moreover, the fact that unions lower pay dispersion means that the incentive for working longer, even unpaid, hours is weaker in unionized settings (Bell and Freeman, 2001; Campbell and Green, 2002).[4] Thus, unpaid extra work that results from these two reasons should be *lower* for union members.

On the other hand, the relationship between union membership and unpaid extra work that is the result of pro-social or altruistic behaviour by the individual is, at first sight, theoretically ambiguous. However, unobserved individual characteristics can play an important role in a more complete understanding of this relationship. Expanding on this, consider the union membership decision. Workers join the union for protection and advancement of their materialistic working conditions, such as pay and fringe benefits.[5] But workers also join a voluntary organization like a trade union for reasons having to do with their overall social values, beliefs and ideology (see e.g. Deery and De Cieri, 1991, for the case of Australian union members; see also, Adams, 1974, for an early account). Apart from economic organizations, trade unions are also social organizations related with specific social values and beliefs. For example, it is well established in the empirical literature of union membership determinants that workers with more left-wing views are more likely to be union members, *ceteris paribus* (Schnabel, 2003). If these values and beliefs are such that also lead to more pro-social behaviour, and if unpaid overtime or donated labour can be thought as a form of such behaviour in specific industries and organizations, as Gregg *et al.* (2008) argue, then union members can be expected to be *more likely* to offer unpaid overtime hours.

Two important points must be clarified here. First, we expand the concept of pro-social or altruistic behaviour by assuming membership in a voluntary organization like a trade union as being driven to some extent by attitudes and beliefs that are also consistent with a pro-social motivation. That is also in line with some studies in Britain that treat membership in a trade union as a measure of "associational social capital" (see e.g. Warde *et al*., 2003). The crucial point here is to show empirically that union members are actually characterized by such a distinctive ethos compared to non-members.

Second, we follow Gregg *et al.* (2008) in hypothesizing that pro-social motivation and the consequent behaviour in the public sector can only be expressed within firms in the broader "caring"

[4] Bryson and Forth (2010) present recent evidence for Britain which shows that "the 'sword of justice' effect whereby unions compress pay differentials remains" (*ibid*., p. 14). They document a rise in pay dispersion among non-unionized employees relative to comparable unionized ones since the early '90s.
[5] See Booth (1995) for an overview of the empirical literature on the trade union impact on wages and other labour market outcomes.

sector (i.e. education, health and social work). Hence, in contrast with other studies of public sector motivation or ethos (see e.g. John and Johnson, 2008), we distinguish the "caring" sector from the overall non-profit sector. In this way we can more easily rationalize our hypotheses concerning the union impact on unpaid overtime incidence. In firms in the for-profit, non-caring sectors, union membership status should be associated with *less* unpaid overtime because employer coercion and employee's career concerns are mitigated under trade unionism. On the other hand, in firms in the non-profit, caring sectors of the economy (where also employer coercion and career concerns are weaker, irrespective of union presence or not[6]), union membership should be correlated with *more* unpaid overtime because of the distinctive ethos of the union members that we hypothesize here. The implicit assumption, of course, is that this kind of pro-social behaviour in the form of unpaid overtime is relevant only in the caring sectors of the economy and in non-profit organizations. We, thus, assume there is no such behaviour in the first group of firms (for-profit, non-caring).

4. Econometric Methods and Data

Methods and Testable Predictions

We can summarize our hypotheses here and discuss them in relation to the econometric methods we use and the results we expect to find. The dependent variable we use in our analysis is the amount of *unpaid overtime hours* an employee works each week. Since this variable is censored at zero, a Tobit model is the appropriate choice (see Verbeek, 2004, pp. 218-227 and pp. 377-378 for more details)[7]:

$$y_{it}^* = x_{it}'\beta + \alpha_i + \varepsilon_{it} \tag{3.1}$$

and

$$\begin{aligned} y_{it} &= y_{it}^* \quad if \quad y_{it}^* > 0 \\ y_{it} &= 0 \quad if \quad y_{it}^* \leq 0 \end{aligned} \tag{3.2}$$

[6] For example, contingent pay systems such as "individual payment-by-results" schemes are traditionally less widespread in the public sector in Britain; see Pendleton *et al.* (2009) for a detailed descriptive analysis using the successive Workplace Industrial-Employment Relations Surveys (WIRS/WERS).

[7] In the final section of this paper, we report results from different modelling choices as a test of the robustness of our main results.

where $i = 1,...,N$ and $t = 1,...,T_i$, y^* is the underlying latent variable (e.g. desired unpaid hours or propensity to offer unpaid overtime hours), y is the observed variable of unpaid overtime hours, x is a vector of explanatory variables including trade union status, β a conformable vector of coefficients and ε_{it} is assumed to be *NID* (0, σ_ε^2) and uncorrelated with x. In order to estimate the model by maximum likelihood, we also have to make a distributional assumption for the unobserved heterogeneity component, α_i. The way we treat α_i is crucial for the testing of our hypotheses that were outlined in the previous section.

The individual heterogeneity component α_i represents factors influencing unpaid overtime that differ across individuals and are constant over time. These are unobserved and can bias the estimates of the model's coefficients if they are correlated with the observed variables included in *x* and are not properly accounted for in the estimation procedure. α_i can include factors such as an individual taste towards hard work, a personality type relevant to unpaid overtime working and/or, following our discussion in the previous section, values and beliefs related to a pro-social motivation and a specific social service ethos.

Throughout the paper, we model unpaid overtime separately for two different samples of employees, one referring to workers in for-profit, non-caring sector workplaces and the other to those in non-profit, caring sector workplaces.[8] We exclude workers in the non-profit, non-caring and for-profit, caring sectors of the British economy. The theoretical reasoning outlined in the previous section lies behind this choice: we did not offer explicit hypotheses concerning the relationship between unpaid overtime and union status in these excluded sectors. An account of the behaviour and incentives of workers in these sectors related to unpaid overtime working is more complicated, given the assumptions we have made and the hypotheses we want to test. Hence, the aim of this paper is to give an explanation about the above relationship in these two specific kinds of workplaces, the for-profit, non-caring and the non-profit, caring ones.

[8] We describe below how we identify these two samples in the BHPS by the specific questions included in the survey.

The empirical strategy we follow consists of three consecutive steps: first, we run pooled Tobit models for the two samples. In these models, we ignore the unobserved heterogeneity component.[9] The likelihood contribution of each observation in this case is given by:

$$l_{it}\left(\beta,\sigma_{\varepsilon}^{2}\right)=\left\{1-\Phi\left(\frac{x_{it}'\beta}{\sigma_{\varepsilon}}\right)\right\}^{d_{it}}\left\{\frac{1}{\sqrt{2\pi\sigma_{\varepsilon}^{2}}}\exp\left[-\frac{1}{2\sigma_{\varepsilon}^{2}}\left(y_{it}-x_{it}'\beta\right)^{2}\right]\right\}^{(1-d_{it})} \tag{3.3}$$

where d_{it} = 1 if $y_{it}=0$ and d_{it} = 0 if $y_{it}>0$. The two constituent parts of (3.3) were derived from the contribution of each observation to the likelihood, which either equals the probability of $y_{it}=0$ or the density of y, given that y is positive, times the probability of $y_{it}>0$.[10] Thus, to take the pooled estimates $(\hat{\beta},\hat{\sigma}_{\varepsilon}^{2})$, we maximize the following log-likelihood function with respect to $(\beta,\sigma_{\varepsilon}^{2})$:

$$\ell\left(\beta,\sigma_{\varepsilon}^{2}\right)=\sum_{i}\sum_{t}\ln l_{it}\left(\beta,\sigma_{\varepsilon}^{2}\right) \tag{3.4}$$

Note that we completely ignored the unobserved heterogeneity component here. The aim is just to detect the correlation between union status and unpaid overtime hours, having controlled for various demographic and job characteristics. From the discussion in the previous section, we expect a negative coefficient on the union status variable in the model for the for-profit, non-caring sample and a positive union coefficient in the non-profit, caring sample. In the for-profit sector, union status is expected to be associated with a lower amount of unpaid overtime hours since coercion and career concerns are mitigated under trade unionism, while the pro-social motivation of employees does not play a role in such workplaces. In the non-profit sector, we hypothesize that union members are characterized by a social service ethos that leads them to undertake more unpaid overtime than similar non-unionized workers.

Second, we use the panel character of our dataset to explicitly model the unobserved heterogeneity component. In this way, we are trying to propose a causal interpretation of the union effect on unpaid overtime, having dealt with the unobserved characteristics that are an important part of our overall argument. The negative effect of union status is expected to also be found in the for-profit, non-

[9] In the pooled Tobit model, we relax the assumption we made about the distribution of ε_{it} and we allow for correlation of the error terms within each individual. For this reason, we use cluster-robust standard errors in all estimated pooled models.

[10] See Verbeek (2004, pp.220-221) for the details on how we algebraically end up with (3.3) starting from these initial probabilities in the log-likelihood function.

caring sample once we control for unobserved individual effects. Getting rid of these effects should not drive the impact of unionism to zero, since we hypothesize that the reduction in unpaid overtime is a direct result of union bargaining and policies in the workplace. On the other hand, the union effect is expected to be zero in the non-profit sample following this second modeling strategy. This latter result follows from our discussion in the previous section, where we postulated that the positive relationship between membership and unpaid overtime in the non-profit, caring sector is due to unobserved individual characteristics (pro-social values and beliefs) that are now accounted for.

A *correlated random effects* (RE) Tobit specification (following Mundlak, 1978, and Chamberlain, 1982) instead of a simple RE Tobit approach will be used for modeling the α_i here, since our discussion points to a correlation between the union status of employees and unobserved individual values and beliefs that can bias the union coefficient in the pooled specification of the non-profit, caring sector model. The other constant unobserved factors mentioned above can also bias the coefficients in the for-profit model as well. The correlated RE model is in some sense a middle way between the simple RE and a fixed effects model (or a specification where the constant unobserved components are added as dummy variables for each individual and are directly estimated). The latter cannot be implemented because of the non-linear nature of the Tobit model and the incidental parameters problem (see Wooldridge, 2002, p. 484). The correlated RE model relaxes the simple RE model's assumption of zero correlation between the constant unobserved heterogeneity and the independent variables and imposes a specific form on this correlation.

Hence, in this approach, we assume the following specific form of relationship between the individual heterogeneity component α_i and the explanatory variables:[11]

$$\alpha_i = \overline{x}_i'\gamma + u_i \tag{3.5}$$

The $\overline{x}_i$'s are the individual means of time-varying regressors, while u_i is normally distributed with mean zero and variance σ_u^2 and assumed uncorrelated with the *x*'s. Assuming that the unobserved heterogeneity is of this form, we take account of possible correlations between the independent variables and the error term that would bias both the pooled and the simple random effects estimates.

[11] For other empirical applications of the correlated RE model, see Cai (2010), who uses a (dynamic) correlated RE Tobit model for the analysis of working hours, and Taylor (2006), who models job satisfaction through a correlated RE ordered probit model.

Note that this method is simply a random-effects Tobit model where the means of time-varying regressors are included in the model as independent variables.

In the correlated RE Tobit model just described, the likelihood contribution of the T_i observations of individual i is given by:

$$l_i = \int_{-\infty}^{+\infty} \prod_t \left\{ \left\{ 1 - \Phi\left(\frac{x'_{it}\beta + \bar{x}'_i\gamma + u_i}{\sigma_\varepsilon} \right) \right\}^{d_{it}} \left\{ \frac{1}{\sqrt{2\pi\sigma_\varepsilon^2}} \exp\left[-\frac{1}{2\sigma_\varepsilon^2} \left(y_{it} - x'_{it}\beta - \bar{x}'_i\gamma - u_i \right)^2 \right] \right\}^{(1-d_{it})} \right\} f(u_i)du_i \qquad (3.6)$$

where d_{it} is defined as before. The density f of u_i is given by:

$$f(u_i) = \frac{1}{\sqrt{2\pi\sigma_u^2}} \exp\left[-\frac{1}{2\sigma_u^2} u_i^2 \right] \qquad (3.7)$$

Note that u_i is unobserved (the random effect), so we have to integrate it out when specifying the joint density of $(y_{i1},...,y_{iT_i})$. Finally, the log-likelihood function is given by:

$$\ell = \sum_{i=1}^{N} \ln l_i \qquad (3.7)$$

Again, we maximize (3.7) with respect to $(\beta, \gamma, \sigma_u^2)$ and we take the respective maximum likelihood estimates.

Finally, we will run a simple linear probability model of membership in voluntary associations in order to show that union members are actually characterized by a pro-social motivation that leads to the behaviour of donating extra working time in the non-profit caring sector.

Data and Descriptive Statistics

In order to estimate the above mentioned models of unpaid overtime, we use data from the British Household Panel Survey (BHPS) covering the period 1991-2007 (BHPS waves 1-17). The BHPS is a panel dataset of a nationally representative sample of about 5,500 households (containing 10,000 individuals) that are interviewed annually since 1991. Only adults over 16 years old in each household are included in the sample. We use the original sample which covered England, Scotland and Wales, disregarding the extension samples of 1999 covering Scotland and Wales, and 2001, which added Northern Ireland. The BHPS is very suitable for our research question since it contains rich information on demographics, household structure and job characteristics that enables us to control for various factors in our unpaid overtime models.

We restrict attention in our sample to fulltime employees (≥30 hours normally worked each week) below state pension age (16-64 for men, 16-59 for women) only. We also drop workers that report more than 90 hours normally worked each week to get rid of extreme and/or invalid observations. The sector of employment is identified through the "wJBSECT" variable in the BHPS ('w' refers to the wave). Employees in the civil service, central or local government, NHS or higher education, nationalized industries and non-profit organizations are categorized as being employed in the not-for-profit (or non-profit) sector of the economy while employees working in private firms/companies are assigned a for-profit employment status. People in the armed forces are excluded from the sample. Moreover, caring sector workers are defined as those employed in the education and health/social work industries and are identified through their SIC80 industry classification up to wave 11 and their SIC92 industry classification for the remaining BHPS waves (variables "wJBSIC" and "wJBSIC92" respectively in the BHPS).

The BHPS does not explicitly record the total amount of unpaid overtime hours worked each week. The interviewee is asked to report her total usual weekly overtime hours and how many of them are paid. Thus, we derive usual unpaid overtime hours by subtracting paid overtime hours from the total overtime hours that are stated by the employee. This is the dependent variable in our models. Concerning the independent variables, we closely follow the specifications in Gregg *et al.* (2008) and Bell and Hart (1999). The regressors, thus, include (see Table A.1 in the Appendix): three dummy variables capturing different ranges of amounts of normal weekly working hours, sex, age, age squared, health status (disability), marital status and number of children in the household; eight 1-digit occupation dummies, permanent status of employment, managerial or supervisory duties in the workplace, two dummies for the number of employees in the worker's firm, whether the employee holds a second job, tenure and tenure squared; dummies capturing satisfaction with job security, whether pay includes bonuses and a dummy capturing a subjective evaluation of promotion opportunities within the firm by the employee;[12] and a series of education, region of work, industry and year dummies.

The hourly wage is not included in the estimated models because of endogeneity concerns.[13] However, because of the well-documented union-wage effect (see Blanchflower and Bryson, 2010, for recent estimates for Britain), it is important to account for the hourly wage in our models. Thus, the average by 3-digit occupation level of the log hourly wage of the employees in the sample is calculated from LFS data for each year and put into the model.[14] Additionally, the same is done for the standard deviation of the log hourly wage. This latter variable can be seen as a measure of earnings inequality within occupational groups of employees, something that can lead to more effort (and, thus, more unpaid extra hours worked), as was mentioned in the theoretical section (see Bell and Freeman, 2001).[15] In Appendix Table A.1, we report detailed information on the construction of some of the variables used in the models, as well as descriptive statistics for all of them.

The variable of interest is the union status of the employee. The BHPS asks if the individual employee is covered by a trade union or staff association in the workplace (one, thus, that determines employee's pay and working conditions through bargaining). If the answer is yes, the employee is then asked if she is a member of that union or staff association. We include two variables constructed by these questions: one that denotes union coverage and one that denotes union membership, conditional on coverage. This may reveal interesting findings concerning their impact on the amount of unpaid overtime worked and these are discussed in the following section. Recall that our hypothesis concerning the values and pro-social motivations of the employees have to do with their *membership* status.

After the above described sample selection procedure and the exclusion of person-year observations with missing information on any of the variables included in the models, we end up with 34,708

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[12] A possible objection to including the variables capturing promotion opportunities and bonus payments in the model is that these are aspects that are influenced by trade unions and it would be better to be captured by the union variables only. By including them, thus, we "over-specify" our models. However, all results that will be reported below are almost identical to those where these specific variables are excluded from the models.
[13] While higher wages can lead to more unpaid overtime (e.g. through a gift-exchange mechanism; see Bell and Hart, 1999), more unpaid overtime can also lead to higher wages following the reasoning outlined in the previous section.
[14] A similar procedure was followed by Gregg *et al.* (2008; p. 14), who use the median wage by occupation, year and age group. The LFS reports earnings data since 1993. Thus, in order not to discard the information in the first two waves of the BHPS, we created duplicates of the 1993 LFS data file also for 1991 and 1992. See the Appendix for more details.
[15] Because of the change in the recording of occupations in the LFS from the SOC90 classification to the SOC2000 from 2001 on, both earnings measures were also interacted with a post-2001 dummy and put into all models. In this way we control for any spurious effects that may be the result of this difference in the LFS reporting of occupations after 2001.

observations for the for-profit, non-caring sample and 7,028 observations for the non-profit, caring sample. The number of individuals in the former sample is 6,695 with an average of 5.2 waves per person, while in the latter it is 1,587 with an average of 4.4 waves.

Before proceeding with the presentation of the results, it is worth describing in brief our dataset and presenting a picture of the distribution of unpaid overtime across various variables that are used in the econometric analysis below. Unpaid overtime incidence in Britain has increased steadily since 1991, rising from 22% of all fulltime employees to 30% in 2007.[16] This increase has occurred in both the profit, non-caring sector and the non-profit, caring sector. The mean hours of unpaid overtime undertaken also increased from 1.8 in 1991 to 2.3 in 2007. This is the result of the growing incidence of unpaid extra hours among the employees in our sample, since the average number of hours undertaken by workers reporting supplying unpaid overtime declined from 8.5 to 7.8 in the same period. In contrast, paid overtime incidence followed the opposite trend, declining from 31% to 23% of fulltime employees in our sample. This provides some evidence on the growing importance of unpaid overtime in the contemporary British labour market, at least as portrayed by the BHPS and the specific sample we use here.

The occupational groups reporting higher incidence of weekly unpaid overtime are managers/administrators and professional workers (53% and 62% of employees respectively). However, large minorities of "lower-ranking" occupational groups work extra hours for no pay: 15% of employees in personal services' occupations and 25% in sales mentioned that they usually work unpaid overtime in a given week.

We now turn to the distribution of unpaid overtime incidence and amount by our variables of interest, i.e. union membership status and sector of employment. Table 1 reports the relevant percentages and numbers and points to a first crude confirmation of our theoretical hypotheses. Union members in for-profit, non-caring sector workplaces are less likely to report working unpaid extra hours than non-members. This difference of 15 percentage points is statistically different from zero. When working unpaid overtime, union members also supply a slightly lower amount of hours than non-members (the difference is weakly significant with a *p*-value of 0.09). On the other hand, in the non-profit caring sector, British trade union members are more likely to supply extra working hours for no pay and, when they do it, they work more hours than non-members (both differences are

[16] All results reported here are based on our final sample used in our models; we described above the process of selecting this sample. Weighting the data with cross-sectional weights does not change these patterns.

statistically different from zero at a 1% level of significance). These findings provide some first evidence in favour of our hypotheses. Of course, more sophisticated analysis is needed in order to reach a more robust conclusion and this is the subject of the following sections of the paper.

Table 1: Unpaid Overtime by Industry-Sector and Union Status

	For-Profit, Non-Caring Sector		**Non-Profit, Caring Sector**	
	Union	**Non-Union**	**Union**	**Non-Union**
Unpaid Overtime Incidence (% of employees)	13%	28%	46%	35%
Unpaid Overtime Amount in Hours (employees working unpaid overtime)	7.7	8.0	9.4	7.4
Sample Size (workers working unpaid overtime in parentheses)	7,474 (965)	27,234 (7,519)	4,757 (2,186)	2,271 (798)

Source: British Household Panel Survey, 1991-2007
Note: See text for the specific sample used to calculate the numbers reported.

5. Results

Table 2 reports the results for the estimated Tobit (pooled and correlated RE) models separately for each sector. Our main focus is on the two union coefficients in each specification. However, before going on discussing the findings concerning the union effects in more detail, it is worth first commenting on our results for the other determinants of unpaid overtime since they are of particular interest, both on their own and in relationship with the related literature on the subject.

Table 2: The Determinants of Unpaid Overtime by Sector

Dep. Variable: Unpaid Overtime Hours	For-profit, non-caring sector		Non-profit, caring sector	
	(1)	(2)	(3)	(4)
	Pooled Tobit	**Correlated RE Tobit**	**Pooled Tobit**	**Correlated RE Tobit**
Union Covered	-1.740***	-0.930***	-1.794**	0.280
	(0.335)	(0.267)	(0.788)	(0.573)
Union Member	-1.603***	-0.235	0.896*	-0.121
	(0.442)	(0.357)	(0.502)	(0.412)
Average hourly wage	7.679***	2.746***	3.684**	-0.213
	(0.654)	(0.523)	(1.719)	(1.285)
St. dev. Hourly wage	10.320***	4.114***	-5.635*	-4.486**
	(1.782)	(1.463)	(3.054)	(2.178)
Average wage*post2001	0.234	0.250	2.196*	0.403
	(0.644)	(0.500)	(1.156)	(0.858)
Average wage*post2001	-5.613**	-5.469**	-0.601	3.178
	(2.842)	(2.287)	(5.391)	(3.649)
Normal Hours 30-35	3.774***	7.675***	8.918***	11.507***
	(0.560)	(0.399)	(1.344)	(0.755)
Normal Hours 36-40	4.719***	7.759***	8.879***	10.892***
	(0.504)	(0.335)	(1.312)	(0.731)
Normal Hours 41-48	3.725***	5.706***	3.353**	5.259***
	(0.549)	(0.372)	(1.494)	(0.871)
Female	-1.122***	-0.200	1.337**	1.036**
	(0.310)	(0.313)	(0.580)	(0.504)
Age	0.565***	1.129***	0.630***	0.583***
	(0.090)	(0.116)	(0.173)	(0.198)
Age squared	-0.007***	-0.010***	-0.008***	-0.006***
	(0.001)	(0.001)	(0.002)	(0.002)
Married or Cohabiting	0.606**	0.068	-1.196**	0.029
	(0.294)	(0.265)	(0.496)	(0.440)
Number of children in hhold	-0.473***	-0.424***	-0.467*	-0.800***
	(0.157)	(0.124)	(0.270)	(0.211)
Managers & Administrators	3.684***	1.687**	5.071**	2.031
	(0.823)	(0.778)	(2.004)	(1.769)
Professionals	1.652*	0.965	6.683***	2.575
	(0.882)	(0.822)	(1.977)	(1.710)
Ass. Professional & Technical	1.367*	0.604	1.504	1.286
	(0.829)	(0.793)	(1.730)	(1.652)
Clerical & Secretarial	1.345*	0.078	1.938	0.764
	(0.739)	(0.752)	(1.587)	(1.723)
Craft & related	-3.553***	-1.241	-0.540	0.895
	(0.818)	(0.792)	(2.961)	(2.463)
Personal & Protective Services	2.454**	1.822*	0.865	-0.360
	(0.985)	(0.936)	(1.461)	(1.614)
Sales	3.469***	1.407*	-2.011	2.523
	(0.831)	(0.785)	(4.950)	(4.947)
Plant & Machine Operatives	-2.739***	-0.689		
	(0.826)	(0.780)		
Permanent	2.541***	2.565***	1.321*	1.791***
	(0.631)	(0.578)	(0.731)	(0.589)
Manager/Foreman/Supervisor	4.559***	2.633***	3.218***	1.170***
	(0.270)	(0.196)	(0.394)	(0.306)

Workplace Size 1-50	-0.349	0.005	2.805***	0.021
	(0.368)	(0.282)	(0.552)	(0.483)
Workplace Size 50-499	-0.013	0.210	0.983*	0.497
	(0.335)	(0.254)	(0.565)	(0.466)
Holding Second Job	-0.622	-0.291	-0.668	-0.326
	(0.422)	(0.339)	(0.535)	(0.395)
Tenure in Years	-0.193***	-0.157***	0.006	-0.029
	(0.053)	(0.042)	(0.092)	(0.067)
Tenure squared	0.003*	0.001	-0.006	-0.000
	(0.002)	(0.002)	(0.004)	(0.003)
Promotion Opportunities	0.882***	0.347*	0.470	0.341
	(0.242)	(0.177)	(0.380)	(0.274)
Pay Includes Bonus	1.120***	0.446***	-0.250	-0.116
	(0.212)	(0.158)	(0.715)	(0.489)
Dissatisfied with Security	-0.797*	-0.935***	0.253	0.386
	(0.422)	(0.322)	(0.891)	(0.535)
Neither Sat. nor Disat. With Sec.	-0.302	-0.184	-0.000	0.171
	(0.200)	(0.158)	(0.369)	(0.270)
Health Limits Type/Amount of Work	-0.233	-0.248	-1.865**	-1.625***
	(0.465)	(0.370)	(0.756)	(0.520)
Degree	6.676***	2.573*	7.390***	-0.353
	(0.681)	(1.321)	(1.656)	(2.694)
Further Education	4.307***	1.756*	4.977***	-0.931
	(0.624)	(1.030)	(1.588)	(2.632)
A Levels	4.016***	1.436	3.192*	-1.406
	(0.669)	(1.096)	(1.830)	(2.773)
O Levels	2.157***	1.101	0.235	-3.527
	(0.665)	(1.119)	(1.735)	(2.762)
Other Qualifications	1.507*	0.666	1.365	0.545
	(0.791)	(1.346)	(2.092)	(4.214)
Constant	-50.022***	-44.545***	-47.293***	-40.146***
	(2.654)	(3.903)	(5.253)	(6.068)
Observations	34708	34708	7028	7028
Log-likelihood	-39028.316	-36329.156	-12473.171	-11555.541

Notes: asterisks refer to results from two-tailed tests of the null hypothesis that the coefficient is equal to zero (* H_0 rejected at the 10% significance level; ** at 5%; *** at 1%); in the pooled specifications standard errors (in parentheses) are robust to clustering at the individual level; all specifications also include region, industry and year/wave dummies; in the correlated RE Tobit models, the individual means of time-varying regressors (except for the year/wave dummies) are included as well.

Starting from the estimates for the for-profit, non-caring sector model presented in column (1), male, older, more educated and married or cohabiting employees offer significantly more unpaid overtime hours, while the number of children in the household has a negative impact on the amount of unpaid extra hours worked. The relationship between age and unpaid overtime hours worked is estimated to be concave. Female workers may undertake fewer unpaid hours because of their weaker preferences for career progression and/or a gender-biased allocation of family responsibilities. The importance of family responsibilities is also apparent from the statistically significant and negative coefficient of the number of children in the household.[17] Bell and Hart (1999) and Gregg *et al*. (2008) report similar findings concerning the effect of the presence of children in the household. The concave relationship between age and unpaid hours might reflect the weakening of incentives to exert more effort after employees have reached a certain age. It can also reflect the natural deterioration of health that prevents older workers from working more hours.

Employees working long standard hours (over 48 per week) supply fewer unpaid overtime hours than employees reporting a lower amount of standard weekly hours. This is probably the result of hours constraints imposed by excessive standard weekly hours (Gregg *et al.*, 2008; p. 17). Managers and professionals, together with people working in protective and personal services and sales, appear to undertake more unpaid hours than similar employees in other occupations. The opposite is found for craft workers and machine operatives. This may capture differences in the nature of the jobs (also relevant for more educated workers) and the responsibilities they entail (as argued by Bell and Hart, 1999). The latter explanation is also consistent with the strong positive effect on unpaid overtime of having a managerial or supervisory status in the job. Employer coercion and fear of job loss, however, cannot be ruled out, especially for employees in low ranking blue-collar jobs. Tenure is negatively associated with extra hours, implying that newcomers in a firm exert more effort in order to secure their jobs and establish themselves within the firm. On the other hand, contrary to the argument and findings of Engellandt and Riphahn (2005), permanent workers are found to supply more unpaid overtime than temporary ones. This may be the result of the nature of the jobs of temporary workers included in the sample (seasonal and casual versus interns or workers on fixed-term contracts) that weakens a probable positive association between temporary work status and unpaid overtime driven by a willingness of temporary workers to secure a better remunerated permanent job.

Promotion opportunities and the existence of bonus payments in pay are positively related to hours worked. Both these results are consistent with our discussion in the theoretical section of this paper, where we noted the importance of career concerns and general economic incentives that influence

[17] There seems to be a different impact of family responsibilities on the amount of unpaid overtime hours that women work compared with their impact on the behaviour of similar male workers. In a model with a full set of interactions between sex and the presence of any children in the household (not reported here), women with children offer a significantly lower amount of unpaid extra hours than similar men with children. This provides evidence in favour of the hypothesis that there exists an unequal allocation of family responsibilities between the two sexes. The results are available from the author upon request. See Van Echtelt *et al.* (2009) for a study of overtime in Netherlands that specifically focuses on issues of gender.

the amount of unpaid work offered.[18] Nevertheless, workers that claim to be dissatisfied with their job security offer fewer unpaid extra hours than workers claiming to be satisfied. This result seems counterintuitive, but it may indicate a general disappointment of the group of dissatisfied employees with their job, something that does not give them the incentive to work any harder. Alternatively, it may reflect the vagueness inherent in such questions that capture subjective evaluations of job aspects.

Finally, the coefficients estimated for the earnings variables are consistent with our theoretical expectations. The average occupational wage is positively related to unpaid overtime hours, something that is in line with the presence of a gift-exchange mechanism/norm in the workplace (Bell and Hart, 1999). Dispersion in earnings, captured by the standard deviation of log hourly wages for the relevant occupational group, is also found to lead to more unpaid overtime. Inequality, thus, seems to strengthen the incentives of employees to supply more effort (Bell and Freeman, 2001). On the contrary, there seems to be no relationship between unpaid overtime and firm size, dual job holding and permanent health status (disability).

When controlling for unobserved heterogeneity in the correlated RE Tobit model (column (2)), most results that were just discussed still hold, although most of the coefficients decrease in absolute value. Sex, marital status and most of the one-digit occupation dummies and the education dummies are now not different from zero. This can be attributed to unobserved factors correlated with these variables that were biasing the estimates in the pooled model. However, at least in terms of direction of the effects and statistical significance, the previous discussion would be the same by reference to either of the two models.

There are some important differences between the estimates for the for-profit, non-caring sample and the non-profit, caring one that need to be stressed (columns (3) and (4)). Women in the non-profit sector supply *more* unpaid extra hours than similar men.[19] Also, career concerns and promotion incentives do not seem to matter in this sector, as implied by the insignificant estimates for the tenure variable and the binary indicators capturing promotion opportunities and the existence of bonus

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[18] There is evidence for Britain that overtime hours have a positive impact on actual promotions. Francesconi (2001), also using the BHPS, reports results that show that the number of weekly overtime hours has a significant and positive effect on promotion probabilities. However, he uses the total amount of overtime hours, not distinguishing between paid and unpaid extra hours.

[19] The coefficient of the number of children in the household is negative and statistically significant as in the for-profit sector. Results from an estimated model with a full set of interaction terms (as described in footnote 17) show that women with any children in this sector *do not* supply significantly different amounts of extra unpaid hours than similar men with children.

payments. This is in line with a less individualistic regulation of employment relations in the public and broad non-profit sector when compared with the profit sector. Only managers and professionals seem to offer more unpaid hours than machine operatives and employees in other occupations, though these effects fall to zero when accounting for unobserved heterogeneity in the correlated RE model. Employees with health problems work fewer unpaid overtime hours than similar healthy employees, while the same is true for married/cohabiting people (again, this effect is not found in the RE model). Workplace size has an effect on unpaid overtime in this sector, as employees in smaller workplaces work more unpaid hours. Finally, earnings dispersion (proxied by the standard deviation of the occupational hourly wage) has a negative effect on unpaid extra hours worked. This is consistent with the view that economic incentives are weaker in the non-profit sector. Additionally, notions and norms of fairness might be more strongly held by employees in the non-profit sector than in the for-profit one. Hence, higher inequality in earnings can have the opposite effect on effort in the former sector as employees perceive these norms to be violated by higher earnings dispersion.[20]

We now turn to the union variables that are the focus of this study. The results presented in Table 2 provide evidence consistent with our hypotheses. Starting from the for-profit, non-caring sector, being covered by a union in the workplace reduces the amount of unpaid overtime hours the employee works each week. The coverage coefficient is large and statistically different from zero. Once we control for unobserved heterogeneity in the correlated RE Tobit specification, the coverage coefficient reduces in absolute value, but it is still negative and significant. This provides evidence for a causal effect of trade unions on unpaid overtime hours. Covered employees are protected from employer coercion by their union, while their extrinsic incentives to supply extra unpaid hours are weakened because of union bargaining and the standardization of intra-firm reward procedures. The reduction in the absolute value of the coefficient also implies that some of the observed differential in unpaid overtime between covered and non-covered employees is the result of unobserved characteristics that lead covered workers to supply lower amounts of unpaid extra hours. Indeed, the coefficient on the individual means of union coverage in the correlated RE specification is negative (not reported). Hence, the negative association between coverage and unpaid extra hours is overstated in the pooled model.

[20] The above discussion pointed to substantial differences in the determinants of unpaid overtime working between the two sectors. Statistically, Chow-type tests of the equality of estimated coefficients in the two sectors rejected the null hypothesis of the appropriateness of pooling the two samples at the 1% level of significance.

A somehow different result is found for the membership coefficient. In the pooled Tobit specification, membership is negatively related with unpaid overtime hours, but this correlation disappears once we account for individual heterogeneity in the correlated RE specification (again, the coefficient of the individual means of membership in this specification is negative). We can interpret this in the following way: union members in the for-profit, non-caring sector are characterized by unobserved traits that lead them to offer lower amounts of unpaid overtime than comparable covered non-members and not covered employees. These traits also induce them to join the union in their workplace in the first place. A possible characterization of these traits can be a more confrontational attitude against the demands of the management of the firm regarding working hours. Because of these attitudes, employees that become union members are less likely than simply covered non-members to offer unpaid overtime hours. Hence, the negative effect of coverage is strengthened by the negative effect of membership (conditional on coverage) in the pooled specification. Recall that in the for-profit, non-caring sector any probable pro-social ethos of union members should not affect the amount of unpaid overtime (i.e. donated labour) worked.

Turning now to the non-profit, caring sector, the pooled Tobit results reveal an opposite impact of coverage and membership on unpaid overtime hours. Covered non-members supply significantly fewer unpaid extra hours than non-covered employees, while covered members supply significantly more hours than covered non-members. In the correlated RE specification, neither effect is statistically different from zero. These results are also consistent with our hypotheses: union members in the non-profit caring sector are characterized by a social service ethos that leads them to work extra hours for no pay; this is not true for covered non-members. The same values that lead members to join unions and donate their labour do not characterize covered employees who have not become union members. These "free-riders" appear to offer fewer unpaid hours than members and non-covered employees, because of a probable taste against hard work and a lack of pro-social beliefs and attitudes (free-riding cannot be considered an "altruistic" behaviour, of course, in the workplace context). In the correlated RE model, these unobserved factors are taken into account and the coefficient of coverage becomes statistically insignificant.[21] Again, the estimated coefficients for the individual means variables in the correlated RE model (not reported) confirm the opposite bias in the two estimates in the pooled model: the coefficient on individual mean coverage is negative

[21] Note also that the role of unions as protective institutions that also weaken career incentives through their impact on the internal organization of the firm is not as important in the non-profit sector as in the for-profit one (see discussion in the theoretical section and footnote 6). This is an additional reason for the insignificant coverage coefficient in the correlated RE model.

(though not significant at a conventional level of significance), while that on mean membership is positive and statistically different from zero.

Union members in the non-profit sector, however, do not supply more unpaid hours than non-covered employees. The negative coefficient of coverage is cancelled out by the positive membership one but it is not “reversed”.[22] Nevertheless, this observation does not invalidate our overall argument for two main reasons: first, union coverage in the non-profit, caring sector of our sample is 91% (and union density is 68%), meaning that non-covered employees represent a very small proportion of all employees in the sector and, hence, covered non-members comprise a more important comparison group for union members. Second, we cannot really claim something about the attitudes towards union membership of non-covered employees. Membership, in the absence of union recognition in the workplace, is usually not a choice for British workers (Green, 1990; Gregg and Naylor, 1993). This again makes any comparison between union members and non-covered employees infeasible and nonsensical.

To get a picture of the magnitude of the union effect on unpaid overtime, we calculate the *absolute* and *relative* effect of union coverage on the probability of working any unpaid overtime and on the expected number of unpaid hours worked for a representative employee. We use the estimates of the correlated RE Tobit model for the for-profit, non-caring sample (column (2) in Table 2), the only model where a statistically significant union effect is found when we account for the bias caused by unobserved heterogeneity in the pooled models.

The representative employee that we consider in our calculations is a 37-year old married woman with two children who is living in London in 2007. She is contracted to work 36-40 weekly hours, holds a clerical job in wholesale or retail trade, has A-levels, has a permanent contract, does not hold a managerial or supervisory position in her firm and does not hold a second job. She also works in a medium-sized workplace (50-499 employees), has worked for 6 years in her firm, judges her job as having opportunities for promotion and is satisfied with its security. Finally, her pay does not include bonuses and does not have a disability affecting the type and amount of her work.[23]

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[22] Based on the estimates of the pooled Tobit model in the non-profit sector, we calculated the probability of undertaking unpaid overtime and the expected number of unpaid hours worked for a representative employee (by assigning specific values to our independent variables). The difference between the probabilities of working unpaid overtime and the difference between the expected number of unpaid extra hours for a representative employee that is a union member compared with an identical one that is not union covered were not statistically different from zero.

[23] We also assign the sample means as the representative values for the two earnings measures and the individual mean variables. The unobserved heterogeneity component is set equal to zero.

Our calculations indicate that this representative employee, when covered by a trade union in her workplace, has a 9.6 percent probability of working any unpaid overtime, while an identical employee not covered by a union has a higher probability, equal to 12.2 percent. Hence, the absolute negative effect of coverage on the probability of supplying unpaid extra hours is 2.6 percentage points that corresponds to a proportionate (relative) reduction in unpaid overtime incidence of 21 percent. At the same time, the covered representative employee is expected to supply 0.31 hours of unpaid overtime each week compared with 0.41 hours for an identical not covered worker. This absolute difference of 0.10 hours corresponds to a 24 percent reduction in weekly unpaid overtime hours that is caused by union coverage. Both estimates reveal a substantial impact of trade unions on unpaid overtime for employees in the for-profit, non-caring sector, operating through the protection they offer against employer coercion and the weakening of individual extrinsic incentives.

Social Behaviour of Union Members

We have, thus, found evidence consistent with our hypothesis concerning the role of unions in the for-profit sector and the pro-social motivation of union members in the non-profit, caring sector. The next step is to additionally show that union members are actually different than non-members in their altruistic motivations. To this end, in Table 3 we report OLS results of a model explaining membership in voluntary associations.

Table 3: Membership in Social and Interest Groups (LP model)

Coverage coefficient	0.011 (0.012)
Membership coefficient	0.079*** (0.013)
Sample mean of dependent variable	0.51

Notes: R^2 is 0.08; sample size is 27,428 observations; asterisks refer to results from two-tailed tests of the null hypothesis that the coefficient is equal to zero (*** H_0 rejected at 1%); the sample includes fulltime employees only; OLS results, robust to clustering standard errors in parentheses; regression also includes controls for political party supported, sex, age, age squared, marital status, number of children, 1-digit occupation, profit sector, log hourly wage, health, region and year; see Appendix Table A.2 for full results

The dependent variable in this model is a dummy taking the value of one if the individual is a member of any of various social groups such as a political party, a parents' association, an environmental group etc.[24] The rationale for this is that becoming a member of a social group or organization is a behaviour consistent with individuals with specific pro-social beliefs and attitudes. Alternatively, membership in such groups can enhance people's "social capital" and general social concerns (see, *inter alia*, Putnam, 1995; Warde *et al.*, 2003; and Hall, 1999). Hence, if it is found that union members are more likely than non-members to belong in such groups, then we have evidence of the distinctive values that characterize union members and can explain their behaviour concerning donated labour in the non-profit, caring sector.

Indeed, the results presented in Table 3 show that, once we control for various demographic, job and political attitudes characteristics, union members are more likely to also be members of other social and interest groups. The membership coefficient is statistically different from zero and quite large in magnitude. Note importantly that, on the other hand, covered non-members are equally likely to belong to a social group as non-covered employees. A union member, thus, is almost 8 percentage points more likely to be a member of a social organization than a similar covered non-member or a similar non-covered worker. Hence, it is only union members that show this distinctive social activity. This is in line with our findings in Table 2 concerning the non-profit, caring sector and the estimated coverage and membership coefficients, and provides additional evidence for our overall argument.[25]

6. Sensitivity Checks

The results reported above confirm our hypotheses that were put forward in sections 3 and 4 concerning the impact of unionization on unpaid overtime. The purpose of this section is to discuss the robustness of our conclusions to various changes in the sample, methodology and the estimation procedures we follow.[26]

[24] Details on the construction of this variable and what it measures are given in the Appendix, at the bottom of Table A.3. This table reports descriptive statistics for all variables included in the model, while Appendix Table A.2 reports the full results for the estimated model.
[25] In order to check if there is any difference in the social behaviour (as captured by membership in voluntary social groups) of union members employed in the two different sectors of our analysis, we estimated separate models for the employees in the for-profit, non caring workplaces and those in the non-profit, caring ones (results available upon request). In both models the results were similar to those reported in Table 3 for the whole sample. The union membership coefficients were significantly positive and large, while their difference was not statistically different from zero. This finding indicates that people that become union members are actually characterized by a specific pro-social ethos, irrespective of their sector of employment. The subsequent behaviour of donating their labour, however, depends crucially on the profit/non-profit character of the firm in which they are employed, as we discussed in the theoretical section above and as the results in this section indicate.
[26] All results reported in this section are available from the author upon request.

First, we included in our estimations the extension samples for Wales and Scotland that were added in the BHPS since 1999.[27] In this way we no longer base our estimations on the original sample of the BHPS but, on the other hand, we increase somehow our sample size. The estimation results for our models are almost identical with those reported in Table 2 for the original BHPS sample that we use here. Second, we estimated simple linear models of unpaid overtime hours for the two sectors. This modeling procedure ignores of course the censoring in our dependent variable. However, it enables us to use a linear fixed effects model (FE) and remove the unobserved heterogeneity component, without having to assume a specific form of it as in the correlated RE Tobit model. The signs and statistical significances of the estimates obtained from the linear pooled OLS and FE models are identical to those reported in Table 2 for the pooled Tobit and the correlated RE Tobit models respectively. In the for-profit, non-caring sector, both the coverage and membership coefficients are negative and statistically different from zero in the pooled OLS models, while in the FE model only the coverage coefficient retains its significance (and reduces in absolute value). Similarly, in the non-profit, caring sector, the coverage and membership coefficients have opposite signs and are statistically significant in the pooled OLS model, while none is statistically different from zero in the FE model.

Third, an alternative measure of union membership can be constructed from the BHPS data and used instead of the measure we included in our models above. This comes from the BHPS variables concerning social and interest group membership (the ones also used to construct the membership in social groups' variable above; see the Appendix for more details). The correlation between the two union membership measures is quite high (correlation coefficient equal to 0.83). However, by using this second measure we take a lower union density in our (pooled for-profit and non-profit) sample (27% versus 29.3%), probably because this measure does not include membership in staff associations (Swaffield, 2001). Moreover, the union membership variable originally used in our models relates strongly to our union coverage variable, since the questions used to construct them are consecutive in the BHPS questionnaire (see discussion in Section 4 and in the Appendix). Finally, a drawback of this alternative membership variable is the fact that it is only available in eleven out of the seventeen BHPS waves that we use.

[27] We did not include the extension sample covering Northern Ireland that was incorporated in the BHPS in 2001. The reason for this is that the variable we used to construct the occupation dummies (based on the Standard Occupation Classification 1990) is not available for the Northern Ireland sample.

Using this second union membership variable and keeping the rest of our specifications intact, we re-estimated the models in Table 2. The results are similar to the ones reported there, where the original union membership variable is used. The only difference is that now the union membership coefficient in the non-profit, caring sector, while still positive, is statistically different from zero only at a 20% level of significance. This may be the result of the substantively smaller sample size (4,542 versus 7,028 in the original model) that prevents an accurate estimation of the union membership effect. This coefficient also reduces substantially in the correlated RE Tobit model.[28] Due to the larger sample size and the direct relationship between the union coverage and membership measures, we judge our baseline results that use the original membership measure as more preferable.

Finally, as Gregg *et al.* (2008; p. 10) also note, a possible objection to the definition of "caring" used in this paper is that the industry-wide definition ignores the diversity of occupations within each industry. For this reason, we follow Gregg *et al.* (2008) and we identify caring occupations within caring industries. Hence, we now include in the non-profit, caring sample only managers, natural scientists, health and teaching professionals, health associates, social welfare associates and childcare employees (identified in the BHPS by the 1990 SOC variable) that work in the caring industries that have been used up to now (education, health and social work). Re-estimating the pooled Tobit and correlated RE Tobit models for this new sample reveals the same pattern of results as in the second panel of Table 2. The union coverage and membership coefficients again obtain opposite signs in the pooled Tobit model. The union membership coefficient is now equal to 0.760, which is similar in magnitude with the coefficient reported in Table 2 (though insignificant now due to a larger standard error resulting from the smaller sample size). Also, both coefficients are very small and insignificant in the correlated RE model.

7. Conclusion

Trade unions have traditionally focused on working time aspects and have campaigned extensively against long working hours in Britain. An aspect of this long working hours "culture" (Bunting, 2005) is the phenomenon of unpaid overtime, which has increased in incidence and importance in Britain (and other industrialized countries) during the last decades. This paper tried to establish the relationship between unionization and unpaid overtime in Britain. We hypothesized on the impact of

[28] Estimating linear models (pooled OLS and FE) of unpaid overtime hours, we find the same results concerning signs and statistical significances as in Table 2. The union membership coefficient in the pooled OLS model for the non-profit, caring sector is positive and now statistically significant at a 10% level of significance.

union membership upon the amount of unpaid extra hours supplied and we noted a probable differential relationship, depending on the sector of employment.

Using data from the first seventeen waves of the BHPS, we found evidence consistent with our hypotheses: in the for-profit, non-caring sector, being covered by a trade union in the workplace is associated with a lower amount of unpaid overtime hours due to the mitigation of employer coercion and the standardization of reward and promotion procedures resulting from union presence and bargaining. On the other hand, in the non-profit, caring sector, union members supply more unpaid extra hours than covered non-members. The latter is the result of specific pro-social beliefs and motivations that characterize union members and lead to the behaviour of donating working time in this kind of workplaces. Indeed, when we account for unobserved heterogeneity in the estimation, the positive relationship between membership and unpaid overtime in this sector disappears. Evidence in favour of a specific pro-social ethos of union members is also presented: being a union member is associated with a higher probability of belonging to any other social or interest group or organization in Britain.

The above results are in general robust to changes in the sample, methodology and estimation procedures used and they are important for two different reasons: first, they enhance our understanding of what unions do in the contemporary British labour market, contributing to the broad literature of union effects on various aspects of the employment relationship (wages, employment, working hours etc.). Second, they can form the basis of further research on the overall attitudes and beliefs of union members, something that has been ignored in the economic analysis of the union membership decision.

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APPENDIX

Table A.1: Variables and Descriptive Statistics

	For-Profit, Non-Caring Sector				Non-Profit, Caring Sector			
Variable	**Mean**	**Std. Dev.**	**Min**	**Max**	**Mean**	**Std. Dev.**	**Min**	**Max**
Unpaid Overtime Hours	1.945	4.298	0	25	3.781	5.97	0	25
Union Coverage	0.362		0	1	0.910	0.286	0	1
Union Membership	0.215		0	1	0.677		0	1
Average Log Hourly Wage	2.078	0.348	1.027	3.108	2.246	0.34	1.149	3.085
St. Dev. Log Hourly Wage	0.396	0.075	0.030	1.122	0.360	0.072	0.127	0.828
Average Wage*Post2001	0.844	1.098	0	3.1070	0.955	1.173	0	3.085
St.Dev. Wage*Post2001	0.145	0.191	0	0.848	0.137	0.169	0	0.696
Hours 30-35	0.169		0	1	0.319		0	1
Hours 36-40	0.662		0	1	0.619		0	1
Hours 41-48	0.104		0	1	0.033		0	1
Hours >48 (r)	0.065		0	1	0.029		0	1
Female	0.319		0	1	0.723		0	1
Age	36.532	11.489	15	64	39.890	10.451	16	64
Age Squared	1466.645	888.390	225	4096	1700.461	837.731	256	4096
Married	0.716		0	1	0.707		0	1
Number of Children	0.537	0.902	0	7	0.471	0.820	0	4
Permanent	0.969		0	1	0.946		0	1
Manager/Foreman/Supervisor	0.412		0	1	0.489		0	1
Workplace Size 1-50	0.444		0	1	0.408		0	1
Workplace Size 50-499	0.388		0	1	0.301		0	1
Workplace Size >=500 (r)	0.168		0	1	0.291		0	1
Second Job	0.068		0	1	0.116		0	1
Tenure	5.307	5.849	1	51	5.714	5.658	1	41
Tenure Squared	62.378	155.088	1	2601	64.662	134.838	1	1681
Promotion Opportunities	0.524		0	1	0.575		0	1
Pay Includes Bonus	0.446		0	1	0.056		0	1
Dissatisfied with Security	0.072		0	1	0.059		0	1
Neither Sat. Nor Dissat. With Sec.	0.380		0	1	0.303		0	1
Satisfied with Security (r)	0.548		0	1	0.638		0	1
Health Limits Type/Amount of Work	0.062		0	1	0.069		0	1
Managers/Administrators	0.201		0	1	0.081		0	1
Professionals	0.062		0	1	0.346		0	1
Ass. Professional & Technical	0.091		0	1	0.279		0	1
Clerical & Secretarial	0.181		0	1	0.093		0	1
Craft & related	0.166		0	1	0.018		0	1
Personal & Protective Services	0.043		0	1	0.148		0	1

Sales	0.069	0	1	0.001	0	1
Plant & Machine Operatives	0.146	0	1	0.004	0	1
Other Occupations (r)	0.042	0	1	0.029	0	1
Degree	0.129	0	1	0.372	0	1
Further Education	0.288	0	1	0.405	0	1
A-levels	0.151	0	1	0.070	0	1
O-levels	0.219	0	1	0.095	0	1
Other Qualifications	0.091	0	1	0.019	0	1
No Qualifications (r)	0.122	0	1	0.039	0	1
South East	0.207	0	1	0.188	0	1
South West	0.094	0	1	0.042	0	1
East Anglia	0.045	0	1	0.034	0	1
East Midlands	0.093	0	1	0.073	0	1
West Midlands	0.088	0	1	0.088	0	1
Northwest	0.110	0	1	0.116	0	1
Yorkshire	0.093	0	1	0.095	0	1
North	0.060	0	1	0.082	0	1
Wales	0.048	0	1	0.055	0	1
Scotland	0.077	0	1	0.117	0	1
London (r)	0.085	0	1	0.109	0	1
Agriculture & Fishing	0.012	0	1			
Mining	0.006	0	1			
Manufacturing	0.351	0	1			
Electricity, Gas & Water	0.018	0	1			
Construction	0.054	0	1			
Wholesale & Retail Trade	0.176	0	1			
Hotels & Restaurants	0.042	0	1			
Transport & Communication	0.085	0	1			
Financial Intermediation	0.079	0	1			
Real Estate & Business Activities	0.140	0	1			
Public Administration & Defence	0.002	0	1			
Education				0.461	0	1
Health & Social Work				0.539	0	1
Social & Personal Services	0.030	0	1			
Private Households & Extra-Territorial (r)	0.007	0	1			
Wave 1 - 1991	0.068	0	1	0.062	0	1
Wave 2 - 1992	0.060	0	1	0.058	0	1
Wave 3 - 1993	0.056	0	1	0.057	0	1
Wave 4 - 1994	0.057	0	1	0.054	0	1
Wave 5 - 1995	0.059	0	1	0.060	0	1
Wave 6 - 1996	0.063	0	1	0.062	0	1
Wave 7 - 1997	0.066	0	1	0.062	0	1

Wave 8 - 1998	0.067	0	1	0.064	0	1
Wave 9 - 1999	0.057	0	1	0.055	0	1
Wave 10 - 2000	0.066	0	1	0.060	0	1
Wave 11 - 2001	0.064	0	1	0.060	0	1
Wave 12 - 2002	0.059	0	1	0.053	0	1
Wave 13 - 2003	0.056	0	1	0.061	0	1
Wave 14 - 2004	0.046	0	1	0.053	0	1
Wave 15 - 2005	0.051	0	1	0.059	0	1
Wave 16 - 2006	0.053	0	1	0.062	0	1
Wave 17 - 2007 (r)	0.051	0	1	0.059	0	1

Notes: Sample size is 34,708 for the for-profit sector and 7,028 for the non-profit one; (r) denotes the reference category for each set of dummies in the models; in the non-profit sector, "plant & machine operatives" are included in the "other occupations" category and together form the reference group for the set of occupation dummies; see below for details on constructing some of the variables.

Notes on Variables in Table A.1: Most of the variables used in the analysis and their derivation are self-explanatory. However, for some of them, we will give some details of how we constructed them and what they measure. Note first that proxy responses in all variables were excluded from the analysis. The selection of the final sample is described in the main text.

<u>Unpaid Overtime Hours</u>: The BHPS asks first about normal weekly working hours: "*Thinking about your (main) job, how many hours, excluding overtime and meal breaks, are you expected to work in a normal week?*" (variable "wJBHRS"). From this question, the "Hours" dummies used here are constructed. The BHPS then proceeds with a question about total overtime hours: "*And how many hours **overtime** do you usually work in a normal week?*" (variable "wJBOT"). Paid overtime hours are given in the next question: "*How much of that overtime is usually **paid** overtime?*" (variable "wJBOTPD"). We derive unpaid overtime hours by subtracting paid overtime hours from the total overtime hours stated. Note that we drop from the analysis employees with more than 90 normal weekly hours and/or more than 30 total overtime hours and/or more than 26 paid or unpaid overtime hours. All exclusions belong to the top 0.5 percent of the distribution of the relevant variable.

<u>Union Coverage and Membership</u>: The union coverage variable in the BHPS ("wTUJBPL") is derived from 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?*". If the answer in this question is "yes", then the membership question (variable "wTUIN1") follows: "*Are you a member of this trade union/association?*". Hence, we code membership as zero if the answer in the coverage or in the membership question is "no". Both questions were not asked in waves 2, 3 and 4 to employees still in the same job as in the previous year. In order not to lose observations, we replaced these missing union data with the answers given in the previous wave/year if the employee had not changed his job. The relevant variable for job continuity is named "wJBBGLY".

<u>For-Profit and Non-Profit Sectors</u>: In the text, we note how we coded the profit and non-profit sector from the answers in the relevant variable "wJBSECT". Again, this question was not asked in waves 2, 3 and 4 to employees holding the same job as in the previous year. We followed the same procedure as with the union variables in order not to lose observations.

Earnings Measures: As mentioned in the main text, measures of earnings are derived from the Labour Force Surveys, 1993-2007. The quarterly surveys are pooled for each year in order to have an annual sample. For 1991 and 1992, we duplicated the 1993 file in order not to lose the respective waves in the BHPS. We kept in the samples only fulltime employees (usual basic/contracted hours equal or larger than 30) aged 18-64 (18-59 for women). We also dropped employees reporting more than 90 usual weekly basic hours. Then, we trimmed the distribution of paid overtime hours to exclude the top 1 percent of observations and that of the gross weekly pay (variable "grsswk" in the LFS files) to exclude the bottom and top 1 percent of observations. The hourly wage for each individual is then calculated as *hourly wage = (gross weekly pay) / [basic hours+1.5*(usual paid overtime hours)]*. The hourly wages are deflated to 2005 prices by using the Consumer Price Indices reported by the Office for National Statistics and we then calculate the natural logarithm of them. To end up with our measures included in the models, the average and the standard deviation of the log hourly wage for each 3-digit occupation is calculated. The Standard Occupational Classification (SOC) of 1990 is used for the LFS data until 2000 and the SOC 2000 for the remaining years. As we described in the text, both measures were interacted with a post 2001 dummy to take account in the estimated models of this change in the occupation recording in the LFS.

Promotion Opportunities and Pay Includes Bonus: These are binary indicators constructed from the variables "wJBOPPS" ("*In your current job do you have opportunities for promotion?*") and "wJBONUS" ("*Does your pay ever include incentive bonuses or profit related pay?*"). Again, this question was not asked in waves 2, 3 and 4 to employees holding the same job as in the previous year. We followed the same procedure as with the union and the sector variables in order not to lose observations.

Industry Dummies and Caring and Non-Caring Industries: From Wave 11, the BHPS changes the recording of industry from the 1980 Standard Industrial Classification (SIC) to the 1992 SIC. We recoded the SIC 1980 categories in the earlier BHPS waves in order to make them correspond to the new SIC 1992 ones, following the guidelines given by the Office for National Statistics in their document "Introduction to UK Standard Industrial Classification of Economic Activities UK SIC(92)" (available at http://www.statistics.gov.uk/methods_quality/sic/). Then, caring industries are those classified as "Education" and "Health & Social Work" in the 1992 SIC.

Table A.2: Determinants of Membership in Social Groups (Full results)

Union Covered	0.011 (0.012)
Union Member	0.079*** (0.013)
Conservative	0.077*** (0.014)
Labour	0.041*** (0.013)
Liberal Democrat	0.070*** (0.015)

Other Party	0.043** (0.020)
Total Working Hours	-0.001 (0.001)
Female	-0.110*** (0.011)
Age	0.008*** (0.003)
Age Squared	-0.00007* (0.00003)
Married or Cohabiting	-0.026** (0.010)
Number of children in hhold	-0.005 (0.005)
Profit Sector	-0.043*** (0.013)
Log of Hourly Wage	0.121*** (0.011)
Health Limits Type or Amount of Work	0.027* (0.016)
Managers & Administrators	0.060*** (0.022)
Professionals	0.151*** (0.023)
Associate Professional & Technical	0.077*** (0.023)
Clerical & Secretarial	0.067*** (0.021)
Craft & related	0.045** (0.022)
Personal & Protective Services	0.034 (0.023)
Sales	0.049* (0.026)
Plant & Machine Operatives	0.037* (0.022)
Constant	-0.067 (0.061)

Notes: See notes in Table 3; region and year dummies included as well

Table A.3: Descriptive Statistics for Model of Social Group Membership

Variable	Mean	Std. Dev.	Min	Max
Member of Any Group	0.511		0	1
Union Covered	0.520		0	1
Union Member	0.346		0	1
Conservative	0.265		0	1
Labour	0.423		0	1
Liberal Democrat	0.134		0	1
Other Party	0.048		0	1
No Party (r)	0.130		0	1
Total Hours	43.202	7.995955	30	102
Female	0.394		0	1
Age	37.441	11.25595	15	64
Age Squared	1528.532	880.7067	225	4096
Married/Cohabiting	0.719		0	1
Number of Children	0.533	0.892652	0	7
Profit Sector	0.717		0	1
Log hourly Wage	2.115	0.523407	-0.59111	3.888604
Health Limits Type/Amount of Work	0.064		0	1
Managers/Administrators	0.173		0	1
Professionals	0.109		0	1
Ass. Professional & Technical	0.125		0	1
Clerical & Secretarial	0.176		0	1
Craft & related	0.125		0	1
Personal & Protective Services	0.081		0	1
Sales	0.047		0	1
Plant & Machine Operatives	0.111		0	1
Other Occupations (r)	0.053		0	1
South East	0.198		0	1
South West	0.084		0	1
East Anglia	0.038		0	1
East Midlands	0.086		0	1
West Midlands	0.089		0	1
Northwest	0.110		0	1
Yorkshire	0.095		0	1
Northwest	0.066		0	1
Wales	0.053		0	1
Scotland	0.085		0	1
London (r)	0.097		0	1

Wave 1 - 1991	0.115	0	1
Wave 3 – 1993	0.100	0	1
Wave 4 – 1994	0.097	0	1
Wave 5 – 1995	0.101	0	1
Wave 7 – 1997	0.113	0	1
Wave 9 – 1999	0.094	0	1
Wave 11 – 2001	0.107	0	1
Wave 13 – 2003	0.094	0	1
Wave 15 – 2005	0.094	0	1
Wave 17 – 2007 (r)	0.086	0	1

Notes: (r) denotes the reference category for each set of dummies in the above model; see below for details on constructing some of the variables.

Notes on regression and variables in Tables A.2 and A.3: The dependent variable in this model is "Membership in Any Social Group" and is derived from the BHPS variables "wORGMA" to "wORGMM", each coding membership in the following social or interest groups: political party, environmental group, parents association, tenants or residents group, religious group, voluntary service group, other community group, social group, sports club, women's institute, women's group and other organization. The variable "wORGMB" is not taken into account since it records membership in a trade union (an alternative variable than the one used in the base models of this paper). These variables are not available in waves 6, 8, 10, 12, 14 and 16.

Political Support Dummies: These are created from the variable "wVOTE" (available in all waves except wave 2), which is a derived variable using information from questions about voting intentions and about which party the respondent feels closest to.

Total Hours: These are total weekly working hours, adding normal hours plus any paid and/or unpaid overtime hours. The derivation of the hours' variables is described above.

Log Hourly Wage: This uses the variable "wPAYGU" which records the usual gross monthly pay of the individual. For each year, the top and bottom 0.5 percent of its distribution is excluded from the sample. The hourly wage is given by *hourly wage = [(gross monthly pay)*(12/52)] / [normal hours+1.5*(usual paid overtime hours)]*. We described above how we derived basic and paid overtime hours from the BHPS dataset. The hourly wage is deflated to 2005 prices by using the Consumer Price Indices reported by the Office for National Statistics and then logged.