FREE TO CHOOSE? DIFFERENCES IN THE HOURS DETERMINATION OF CONSTRAINED AND UNCONSTRAINED WORKERS

Mark L. Bryan

ISER Working Papers
Number 2002-28

Institute for Social and Economic Research

The Institute for Social and Economic Research (ISER) specialises in the production and analysis of longitudinal data. ISER incorporates the following centres:

- ESRC Research Centre on Micro-social Change. Established in 1989 to identify, explain, model and forecast social change in Britain at the individual and household level, the Centre specialises in research using longitudinal data.

- ESRC UK Longitudinal Centre. This national resource centre was established in October 1999 to promote the use of longitudinal data and to develop a strategy for the future of large-scale longitudinal surveys. It was responsible for the British Household Panel Survey (BHPS) and for the ESRC's interest in the National Child Development Study and the 1970 British Cohort Study

- European Centre for Analysis in the Social Sciences. ECASS is an interdisciplinary research centre which hosts major research programmes and helps researchers from the EU gain access to longitudinal data and cross-national datasets from all over Europe.

The British Household Panel Survey is one of the main instruments for measuring social change in Britain. The BHPS comprises a nationally representative sample of around 5,500 households and over 10,000 individuals who are reinterviewed each year. The questionnaire includes a constant core of items accompanied by a variable component in order to provide for the collection of initial conditions data and to allow for the subsequent inclusion of emerging research and policy concerns.

Among the main projects in ISER's research programme are: the labour market and the division of domestic responsibilities; changes in families and households; modelling households' labour force behaviour; wealth, well-being and socio-economic structure; resource distribution in the household; and modelling techniques and survey methodology.

BHPS data provide the academic community, policymakers and private sector with a unique national resource and allow for comparative research with similar studies in Europe, the United States and Canada.

BHPS data are available from the Data Archive at the University of Essex
http://www.data-archive.ac.uk

Further information about the BHPS and other longitudinal surveys can be obtained by telephoning +44 (0) 1206 873543.

The support of both the Economic and Social Research Council (ESRC) and the University of Essex is gratefully acknowledged. The work reported in this paper is part of the scientific programme of the Institute for Social and Economic Research.

Acknowledgement:
I would like to thank Stephen Jenkins for much helpful advice during the drafting of this paper. I am also grateful to Alison Booth and to seminar participants at the Research School of Social Sciences (RSSS), Australian National University (ANU), and the Melbourne Institute of Applied Economic and Social Research, University of Melbourne, for useful comments. All errors are mine.

Readers wishing to cite this document are asked to use the following form of words:

> **Bryan, Mark L. (December 2002) 'Free to choose? Differences in the hours determination of constrained and unconstrained workers', *Working Papers of the Institute for Social and Economic Research*, paper 2002-28. Colchester: University of Essex.**

For an on-line version of this working paper and others in the series, please visit the Institute's website at: http://www.iser.essex.ac.uk/pubs/workpaps/

Institute for Social and Economic Research
University of Essex
Wivenhoe Park
Colchester
Essex
CO4 3SQ UK
Telephone: +44 (0) 1206 872957
Fax: +44 (0) 1206 873151
E-mail: iser@essex.ac.uk
Website: http://www.iser.essex.ac.uk

ABSTRACT

In individual surveys, large minorities of individuals typically report that they would like to change their weekly working hours at their current hourly wage. If this evidence reflects genuine constraints on individuals' choice of hours, the determinants of hours should differ between constrained and unconstrained groups. Controlling for selection by an extension of the Heckman two-step method to ordered selection and panel data, and using a sample of manual men, I find that unconstrained workers' hours are determined differently from those of constrained workers. I present evidence that local labour market conditions affect the hours of constrained but not of unconstrained workers. I also correct for the potential bias resulting from the use of observed hours to derive the hourly wage, by instrumenting it with its lagged value. The combination of ignoring hours constraints and assuming the derived hourly wage is exogenous imparts a large downward bias to estimates of the wage elasticity. I estimate the corrected uncompensated elasticity to be –0.1.

NON-TECHNICAL SUMMARY

Evidence from a range of surveys suggests that a large minority of the workforce would work a different number of weekly hours if it had a free choice and the hourly rate of pay stayed the same. The figures from the British Household Panel Survey (BHPS) used in this paper show that approximately 40% of individuals interviewed each year reply that they would like to adjust their hours, mostly in a downwards direction.

If these survey responses reflect genuine constraints on the labour supply behaviour of individuals there are several implications for labour market analysis. First, economists may have an inadequate understanding of how working hours are determined. Second, constraints on individual choice of hours may reduce overall welfare. The policy attention directed at initiatives in favour of 'work-life balance' in the UK suggests there are welfare considerations. Third, estimates of how responsive labour supply is to changes in the wage may be biased. Finally, supply-side policies aimed at changing the hours of constrained workers are unlikely to succeed.

I use data on the actual working hours and reported constraints of male manual workers from the first 9 waves of the BHPS to establish whether the hours of unconstrained individuals are determined differently from the hours of those who report being constrained. Put differently: do workers' reported constraints reveal anything about their labour supply behaviour? A negative response would cast doubt on the existence of constraints. I also look at how errors in the number of hours reported affect estimates of the reponsiveness (elasticity) of hours to the wage.

The evidence is that the determinants of working hours do differ for unconstrained and constrained individuals, lending credence to their subjective reports. A tentative finding is the state of the local labour market appears to affect the hours of constrained but not unconstrained workers. The results also point to large proportionate biases in the estimates of labour supply elasticities if constraints and measurement error are ignored. Nevertheless, all the estimates obtained are small, suggesting that the wage is not the main driver of hours worked.

1. Introduction

In the neo-classical labour supply model individuals face an exogenous wage rate and free choice of working hours. In this model, if a worker cannot choose the number of hours in the current job then competition amongst firms for labour ensures that he can find an alternative job matching his hours preferences. And yet evidence from a range of surveys suggests that a large minority of the workforce is in disequilibrium with respect to its choice of working hours. The surveys reviewed by Kahn and Lang (1996), for example, show a broadly stable pattern over time: about 60% of respondents would like to keep the same hours, just under a third would like to work more and about 10% would like to work less[1]. Results from the Panel Study of Income Dynamics (PSID) tell a qualitatively similar story[2]. In Europe there is some evidence that the pattern is reversed. In a European Union survey of 1989 quoted by Kahn and Lang far more respondents wanted to reduce their hours than increase them[3]. For example, in the UK only 12% wanted to work more against 33% who wanted to work less. These figures are similar to the results from the British Household Panel Survey (BHPS) used in this paper: approximately 40% of individuals interviewed each year in the BHPS reply that they would like to adjust their hours, mostly in a downwards direction[4].

If these survey responses reflect genuine restrictions on the labour supply behaviour of individuals there are several implications for labour market analysis. First, the existence of constraints suggests economists have an inadequate understanding of how working hours are determined (though see the discussion of constraint models below). Second, constrained hours choices may imply substantial welfare losses. The policy attention directed at initiatives in favour of 'work-life balance' in the UK suggests there are welfare considerations.[5] Third, it is likely that

---

[1] The questions generally made clear that the amount of pay received would change if hours changed (in particular, the Current Population Survey (CPS) of 1985 specified that the rate of pay would stay constant).

[2] But note the sequence of the PSID questions is somewhat complicated and the phrasing does not always make clear whether the rate of pay would stay constant. Workers who may want to work less are told they would earn less money, whereas the question about wanting more work has no reference to pay.

[3] The question stated that the rate of pay would stay the same.

[4] However, the International Social Survey Program conducted in the same year (and using a similar question) found that 24% wanted to work more and only 8% wanted to work less.

[5] See the Department of Trade and Industry website http://www.dti.gov.uk/work-lifebalance.

labour supply elasticities estimated from conventional labour supply models are biased (see *inter alia* Biddle and Zarkin, 1989). Fourth, as noted by Kahn and Lang (1991), supply-side policies aimed at changing the labour supply of constrained workers are unlikely to succeed.

The aim of this paper is to use data on the actual working hours and reported constraints of individuals from the first 9 waves of the BHPS to establish whether the hours of unconstrained individuals are determined differently from the hours of those who report being constrained. Put differently: do workers' reported constraints reveal anything about their labour supply behaviour? A negative response would cast doubt on the existence of constraints.

To answer this question, after reviewing the various theories of hours constraints in Section 2, I describe a model of hours determination and reported constraints to be estimated by a two-step method (Section 3). It is a generalisation of the model estimated by Stewart and Swaffield (1995) on one wave of data. Section 4 describes a formal test, following Ham (1982), of whether the unconstrained (desired) hours regime differs from the other regimes. I estimate the model in Section 5, applying an extension of the Heckman two-step selection correction method (described by Verbeek and Nijman, 1996) to panel data and ordered selection. The evidence is that unconstrained workers are in a distinct hours regime. In Section 6 I address a perennial concern of labour supply estimation, which is that the hourly wage is usually calculated by dividing earnings by reported hours: any measurement error in hours will impart a downward 'division bias' to the estimated wage coefficient due to the spurious correlation between the hours equation disturbance term and the wage measure. The availability of panel data enables the wage to be instrumented by its own lagged value (free of contemporaneous measurement error) to remove this bias. In Section 7 I find evidence in favour of the prediction of the model that demand side variables (in this case the local labour market unemployment-to-vacancy (U/V) ratio) affect the hours of constrained but not unconstrained manual men. Section 8 compares the calculated elasticities and implied biases which result from neglecting constraints and hours measurement error. Finally, Section 9 concludes.

2. Theories of hours constraints

Four main reasons have been advanced to explain hours constraints. First, if firms are monopsony buyers of labour then, under plausible assumptions, they will negotiate a bargain which is relatively favourable to the firm and involves higher hours and lower wages. Workers then work more hours than they would desire at the agreed wage rate (i.e. they are off their supply curves). Conversely, if workers have monopoly power (for example, through a union), it can be shown that bargaining can result in them working too few hours at the negotiated rate.

Second, constraints may be efficient responses to convexities in the production function (for example, due to start-up costs), as in the model of Card (1990) where workers face a lower bound on their weekly hours. More generally, these models predict that the hourly wage varies with the number of hours worked (Barzel, 1973 and Kinoshita, 1987) and that workers are only on their supply curves on flat portions of the wage-hours locus.

Third, in some models of long-term employment contracts, all workers are constrained in their choice of hours ex-post. This is because the contract specifies a wage which diverges from the worker's marginal product but, at the same time, efficiency requires that hours are fixed such that the marginal product equals the shadow wage. Thus the wage differs from the shadow wage and the worker is constrained.

Fourth and finally, in a dynamic matching view of the labour market the difference between actual and desired hours is one aspect of firm-worker match quality which should improve with labour market experience and mobility.

3. A model of hours determination

3.1 Hours regimes

Let the desired weekly hours of work, h_{it}^S, for individual i at time t be described by:

$$h_{it}^S = x_{1it}'\beta_1 + \eta_{1i} + \varepsilon_{1it} \tag{1}$$

where x_{1it} is a vector of explanatory variables, including the hourly wage, with associated coefficient vector β_1. Unobserved influences on desired hours are decomposed into an individual-specific, time-invariant (*permanent*) effect η_{1i} and a time-varying (*transitory*) effect ε_{1it}. This equation is the standard generic formulation for labour supply which assumes that individuals are free to choose their hours of work at a given wage rate. Individuals whose hours are described by (1) are said to be in the *unconstrained* hours regime. In the absence of hours restrictions, (1) can be estimated using data about the observed hours of the full sample of individuals to obtain estimates of the labour supply model coefficients β_1.

To allow for the possibility that some individuals may be off their supply curves (as suggested by the responses to the BHPS question), define two additional equations describing respectively the number of hours, h_{it}^U, worked by an individual who would prefer to work more hours (the *underemployment* regime):

$$h_{it}^U = x_{2it}'\beta_2 + \eta_{2i} + \varepsilon_{2it} \tag{2}$$

and the number of hours, h_{it}^O, worked by an individual who would prefer to work fewer hours (the *overemployment* regime):

$$h_{it}^O = x_{3it}'\beta_3 + \eta_{3i} + \varepsilon_{3it}. \tag{3}$$

The x_{jit} vectors (j = 2, 3) contain explanatory variables, including the wage, associated with coefficient vectors β_j. Unobserved influences on hours are captured by the permanent disturbance term η_{ji} and the transitory disturbance term ε_{jit}. The vectors x_{jit} contain not only the supply characteristics x_{1it} but also those demand characteristics which are hypothesised to influence the hours of constrained individuals, implying testable restrictions on variables which are included in (2) and (3) but omitted from (1). For example, in a monopsony model of constraints, measures of labour market conditions belong in the constrained hours equations but not in the desired hours equation. Since there is no reason to suppose that the hours of the underemployed will be determined in the same way as the hours of overemployed individuals, the coefficients and unobservables may differ between (2) and (3). Furthermore, the unobservables may be correlated with each other, as well as with the error term in (1).

The two regimes form upper and lower bounds on the hours that may be worked by an individual, implying a restriction on the coefficient vectors such that $h_i^U > h_i^O$.

3.2 Selection model

If, given their observable characteristics, individuals were equally likely to be found in the three constraint states then, using the sample responses about whether they are overemployed, underemployed or unconstrained, it would be a simple matter to estimate (1), (2) and (3) as random effects (RE) or fixed effects (FE) models on the respective subsamples and obtain unbiased estimates. If, on the contrary, individuals who generally have an above-average liking for work (a high η_{1i}) or who may like to work more from time to time for unobserved reasons (a high ε_{1it}) are systematically less likely to be unconstrained, then the estimates of β_1 will suffer from selectivity bias. To account for this, the model is extended to incorporate an equation describing the sample selection mechanism:

$$\begin{aligned} d_{it}^* &= z_{it}'\gamma + \alpha_i + \nu_{it} \\ d_{it} &= 0 \text{ if } d_i^* < \kappa_1 \text{ (underemployed)} \\ &= 1 \text{ if } \kappa_1 \le d_i^* < \kappa_2 \text{ (unconstrained)} \\ &= 2 \text{ if } \kappa_2 \le d_i^* \text{ (overemployed)} \end{aligned} \tag{4}$$

where d_{it}^* is the latent propensity to be more overemployed and less underemployed; d_{it} is an indicator variable taking values 0, 1, or 2 according to the state actually observed; z_{it} is a vector of explanatory variables associated with coefficient vector γ, and κ_1 and κ_2 are parameters (cut-off points). The z_{it} vector is expected to contain all the supply and demand side variables influencing hours – indeed, the latent variable d_{it}^* can be considered to represent the difference between desired hours h_{it}^S and constrained hours h_{it}^U or h_{it}^O. The error components α_i and ν_{it} are both assumed normally distributed and orthogonal to each other with variances σ_α^2 and σ_ν^2, where σ_ν^2 is normalised to one. Writing $u_{it} = \alpha_i + \nu_{it}$, $\text{var}(u_{it}) = 1 + \sigma_\alpha^2$ and $\text{cov}(u_{it}, u_{is}) = \sigma_\alpha^2$, $s \neq t$, so that cross-time correlation is assumed to arise via the permanent effect α_i. The selection model set out in (4) is random-effects ordered probit.

Let the variances of the error components η_{ji} and ε_{jit} (j = 1, 2, 3) be $\sigma_{j\eta}^2$ and $\sigma_{j\varepsilon}^2$ respectively. Non-random selection into a regime occurs if the errors in the selection equation (4) are correlated with the errors in the corresponding hours equation (1)-(3). I therefore allow η_{ji} and α_i to be correlated, with covariance $\sigma_{j\eta\alpha}$, and similarly let the covariance between ε_{jit} and ν_{it} be $\sigma_{j\varepsilon\nu}$. Correlation between unobservables thus arises from two sources, the permanent effect and the transitory effect.

These correlations affect the consistency of estimators of β_j. In the case of fixed effect (FE) estimation on the subsample of individuals from a particular regime, the time invariant effect η_{ji} is removed in estimation and a sufficient condition for consistency is that $\sigma_{j\varepsilon\nu} = 0$, i.e. selection operates through the permanent effect only. If this condition holds, then FE estimation removes the selection bias without any further correction. This is the reason sample selection is considered less of a concern when using panel data (Vella, 1998). In the hours constraint literature, Ball (1995) estimated a labour supply equation in first differences on a sample of unconstrained individuals, arguing that sample selection was thereby removed. However, if selection also operates through the transitory effects, i.e. $\sigma_{j\varepsilon\nu} \neq 0$, then selectivity bias remains. In the case of the random effects (RE) estimator, the requirements for consistency are stronger. For example, one sufficient condition is that $\sigma_{j\eta\alpha} = \sigma_{j\varepsilon\nu} = 0$, i.e. selection does not operate through either channel.

As noted below in Section 3.3, there is some evidence against the assumption that hours are distributed normally conditional on observables and therefore that the distribution of errors in the model (1)-(4) is jointly normal. Nevertheless the assumption will be maintained for the moment in the interest of a fuller exposition of the correction procedure. Below, as a robustness check, I re-estimate the model under the weaker assumption that the desired hours equation errors η_{1i} and ε_{1it} are simply linearly related to the selection equation errors.

3.3 Estimating strategies

Unbiased estimates of β_1, β_2 and β_3 can be obtained by estimating each hours equation (1)-(3) jointly with (4), subject to adequate identification as discussed below. However, the likelihood function includes a double integral even after some simplification (Verbeek and Nijman, 1996), making it computationally unattractive.

An alternative would be to place restrictions on the selection process. For example, the model of Stewart and Swaffield (1997) is a two-limit Tobit where the observed hours, h_{it}, of constrained individuals form exogenous censoring points on the distribution of desired hours.[6] They estimated the model on cross-sectional data but it would be feasible to extend it to panel data. In this case:

$h_{it} = h_{it}^U$ if $h_{it}^S > h_{it}^U$ (underemployed)
$h_{it} = h_{it}^S$ if $h_{it}^O < h_{it}^S < h_{it}^U$ (unconstrained)
$h_{it} = h_{it}^O$ if $h_{it}^S < h_{it}^O$ (overemployed)
Estimation of this model is feasible directly by maximum likelihood.

I adopt a second, more general approach, yielding consistent estimates, which is an extension of the two-step Heckman technique to ordered selection and panel data. It involves estimating the selection model (4) and using the estimates to generate regressors which are added to the hours equations (1)-(3) to capture the conditional expectations of the error terms. Apart from its generality, this method offers the advantage that, although it assumes normally distributed errors α_i and ν_{it} in the selection equation, the coefficient estimates of the hours equations (obtained by OLS) are more robust to non-normality of the errors η_{ji} and ε_{jit}, j = 1, 2, 3.[7] Stewart and Swaffield rejected the normality of hours in their model, remarking that this was "problematic" for their estimator. Moffitt (1999) also comments on the sensitivity of the Tobit estimator to non-normality.

4. A test of the hours constraint information

Assuming that the model of reported constraints (4) is correctly specified, then the coefficient estimates from the two stage procedure may be used to test whether individuals' hours are determined differently in the three constraint states. A common objection to the use of subjective data in economic analysis is that survey responses are highly dependent how the questions are asked and thus are not very informative of the underlying process of interest. Reports of overemployment, for

[6] Stewart and Swaffield (1995), testing down from a more general model, could not reject the restrictions implicit in the Tobit model on their sample of BHPS wave 1 data.
[7] Whilst there is evidence of a sharp spike in both actual and desired hours (Kahn and Lang, 1996), it seems plausible that the *difference* between actual and desired hours (the variable underlying the selection model) approximates to a normal distribution.

example, may be positively correlated with long hours without reflecting any actual restrictions on the economic behaviour of these respondents.

The general scheme presented above suggests two tests of whether or not the constraints reported by individuals are associated with different hours determination regimes. The first is to compare the selectivity-corrected estimates of β_1, β_2 and β_3 since under the null hypothesis that the determinants of hours are not regime specific, there should be no significant difference between the coefficient estimates. However, identification issues arise in estimating (2) and (3) because in this general scheme it is difficult to suggest variables which enter the selection model while being excluded from the constrained hours equations. For example, if constrained hours are the outcome of individual-firm bargaining then all supply and demand variables are valid explanatory variables in the h_{it}^U and h_{it}^O equations (2) and (3), and identification would be on distributional assumptions alone.[8]

On the other hand, identification of (1) is achieved by the exclusion of demand variables. Thus the second test, originally proposed by Ham (1982), is based on a comparison of two estimates of β_1, the uncorrected estimates of (1) on the entire sample and the selectivity-corrected estimates from the unconstrained sample only. Under the null hypothesis that reported constraints do not reflect different behaviour, the uncorrected estimator using the whole sample is efficient; under the alternative hypothesis that there are distinct hours regimes, the error term in (1) (estimated on the full sample) is augmented by $h_{it}^O - h_{it}^S$ or $h_{it}^U - h_{it}^S$ for constrained individuals. If this discrepancy between desired and actual hours is correlated with any of the determinants of desired hours x_{1it} (and evidence from the selection estimates reported below suggests it is), then the uncorrected estimator will be inconsistent. The selectivity-corrected estimates, on the other hand, are consistent under both the null and alternative hypotheses, assuming that (4) is an appropriate model of workers' reported constraints. A Hausman test can therefore be used to distinguish between the two hypotheses.

[8] Similar problems arise in testing the significance of demand variables in (1).

5. Two-step estimator with ordered selection and panel data

I propose a two-step method to test and correct for the ordered selection into the unconstrained regime, applying the techniques developed in the literature by Verbeek and Nijman (1996) and Vella and Verbeek (1999). Analogously to the standard Heckman model, selection bias is eliminated by incorporating two additional selection-correction regressors into the main equation. They are derived as the expected values of the error components η_{1i} and ε_{1it} conditional on the vector of all outcomes d_i. One can write:

$$E[h_{it}^S \mid d_i] = \mathrm{x}_{1it}'\beta_1 + E[\eta_{1i} \mid d_i] + E[\varepsilon_{1it} \mid d_i] \tag{5}$$

Intuitively it is necessary to condition on the selection outcome in *all* periods because of the serial correlation in the selection model error $u_{it} = \alpha_i + v_{it}$ (due to α_i): the contemporaneous value of the unobservable depends on its past and future realisations. Using assumptions made about the correlation structure of u_{it}, η_{1i} and ε_{1it}, it can be shown that the expectations of the two error terms in (5) are:

$$E[\eta_{1i} \mid d_i] = \sigma_{1\eta\alpha} / (\sigma_v^2 + T_i\sigma_\alpha^2) \sum_{t=1}^{T_i} E[u_{it} \mid d_i] \tag{6}$$

and

$$E[\varepsilon_{1it} \mid d_i] = (\sigma_{1\varepsilon v} / \sigma_v^2) E[u_{it} \mid d_i] - \sigma_{1\varepsilon v}\sigma_\alpha^2 / [\sigma_v^2 (\sigma_v^2 + T_i\sigma_\alpha^2)] \sum_{t=1}^{T_i} E[u_{it} \mid d_i] \tag{7}$$

where $\sigma_v^2 = 1$ and individual i is observed over T_i periods. It is clear that $E[\eta_{1i} \mid d_i] = 0$ if there is no correlation between the permanent effects ($\sigma_{1\eta\alpha} = 0$). The expression for $E[\varepsilon_{1it} \mid d_i]$ is more complicated and consists of a term related to the contemporaneous error in the selection equation, as well as a term related to the selection errors in all periods. Since $\sigma_{1\eta\alpha}$ does not appear in this second term, the selection equation errors in periods other than t still affect $E[\varepsilon_{1it} \mid d_i]$ even if selection only occurs through the transitory terms (i.e. $\sigma_{1\eta\alpha} = 0$) – an example of the effects of serial correlation.

To evaluate (6) and (7) an expression for $E[u_{it} \mid d_i]$ in terms of observables is required. Again this is an extension of the conditional expectation used in the

Heckman model and involves integrating over all possible values of values α_i. The general formula (as given, for example, by Verbeek and Nijman,1996) is:

$$E(u_{it} \mid d_i) = \int_{-\infty}^{\infty} [\alpha_i + E(v_{it} \mid d_i, \alpha_i)] f(\alpha_i \mid d_i) d\alpha_i \tag{8}$$

where $f(.)$ is the condition density function of α_i. The term $E[v_{it} \mid d_i, \alpha_i]$ is the generalised residual from the random effects ordered probit. The conditional density is then given by:

$$f(\alpha_i \mid d_i) = \frac{\frac{1}{\sigma_\alpha}\phi(\frac{\alpha_i}{\sigma_\alpha})\prod_{s=1}^{T_i} l_{is}(d_{is} \mid z_{is}, \alpha_i)}{\int_{-\infty}^{\infty} \frac{1}{\sigma_\alpha}\phi(\frac{\alpha}{\sigma_\alpha})\prod_{s=1}^{T_i} l_{is}(d_{is} \mid z_{is}, \alpha) d\alpha} \tag{9}$$

where l_{is} has the form of the likelihood contribution in a cross-sectional ordered probit equation. Equation (9) is derived from the rules of conditional probability: intuitively one can see the similarity to $g(v|w) = g(v,w)/\int g(u,w)\mathrm{d}u$, where $g(v,w)$ is the joint probability density of variables v and w, and $g(v|w)$ is the density of v conditional on w.

These expressions were evaluated using the estimated value of γ from the selection equation. A potential complication comes in evaluating the two integrals. Rather than Gausssian-Hermite quadrature, which is commonly employed in iterative procedures, I used a cubic spline algorithm (included in the Stata 7 software package). Though more computationally expensive in that more integration points must be used for the same accuracy, this relative inefficiency is much less of a concern when the integral only has to be evaluated once for each individual. Sensitivity checks of the interval of integration and number of points showed that the differences were negligible over a wide range. The results presented use 50 integration points over the range of –5 to +5 of the normalised integral (i.e. substituting $\alpha_i^* = \alpha_i/\sigma_\alpha \sim N(0, 1)$).

5.1 Data and estimation

The selection model (4) was estimated on a sample of data from 9 waves from the British Household Panel Survey (BHPS). The constraint indicator is constructed from the answers to the question:

"Thinking about the hours you work, assuming that you would be paid the same amount per hour, would you prefer to work fewer hours, work more hours or continue the same hours?".

The wording of this question may lead to confusion amongst individuals who are not paid by the hour and so ideally estimation would be done on hourly paid workers only in the first instance. Unfortunately the BHPS did not carry a question on whether individuals were salaried or hourly paid until wave 9, and so a sample of manual workers (who should be less likely to be salaried) was used. The sample was restricted to men to avoid the complication of modelling workforce participation. A similar sample was used by Stewart and Swaffield in their analysis based on BHPS wave 1. The summary statistics of the sample are presented in Table 1, with variable definitions listed in Table B1, Appendix B.

[Table 1 around here]

The sample consists of male manual employees aged 21-64 years who did not have second jobs and had valid observations on the variables of interest. Manual status is derived from information about socio-economic status (BHPS variable wJBSEG), summarised using 3 digit occupational codes. This level of occupational detail meant that I could drop men who were in manual occupations but likely to have had positions of responsibility (e.g. foremen) and thus less likely to have been hourly paid. A total of 5241 person-years were observed from the first 9 waves (1991-99) of the BHPS. This covered 1700 individuals observed on average over 3.1 waves. In about 55% of cases the men reported being unconstrained, whilst they were underemployed in 8% of cases and overemployed in 37% of cases.

Table 1 shows that underemployed men worked on average about 4 fewer hours per week than unconstrained men who, in turn, did about 4 fewer hours than overemployed men. Moving from underemployed men through unconstrained men to the overemployed, we see a monotonic increase in mean age and tenure, and in the likelihood of being married, trade union covered and being paid partly by bonuses. There is not, however, a substantial difference in log net weekly wage across the subgroups, but it is striking that unconstrained men had on average 32% higher weekly non-labour income than the overemployed (and 7% higher non-labour income than the underemployed). Comparing the mean values of the variables in the full sample with those of unconstrained men, we see that the differences between the subgroups tend to average out, with the exception that that weekly non-labour income is still about 10% higher for unconstrained individuals. The wage variable was derived as the net marginal overtime wage assuming an overtime premium of time-and-a-half (the actual premium applying to each individual cannot be determined from the data) and the marginal income tax rate calculated for each worker (from the UK Inland Revenue income tax bands and allowances over the period, given his wage income and marital status).[9]

Table 2 gives some idea of how reported constraints varied for individuals over the panel. The first two columns of figures repeat the overall summary statistics given in Table 1. The next two columns show the number of individuals who ever reported each of the constraint states, so that, for example, 303 (or 17.8% of) individuals reported being underemployed at some time during the observed time period, and over half of men said they were overemployed at least once. These figures are substantially larger than the overall statistics, and indicate that individuals do move between constraint states. The final column shows, for individuals observed at least twice, the percentage of time spent in each state conditional on experiencing that state at least once. For example, men who were ever underemployed reported underemployment at 39% of waves on average. The highest figure is for workers who were ever unconstrained: they said they were happy with their hours in 67% of waves on average. Again these statistics suggest a large degree of 'churning'.

[Table 2 around here]

In order to facilitate comparisons with Stewart and Swaffield (1997), I used the same functional form of the desired hours equation (1), i.e.:

$$h_{it}^{S} = x_{it}'\beta + \delta_1 \ln w_{it} + \delta_2 y_{it}/w_{it} + \eta_{1i} + \varepsilon_{1it} \qquad (10)$$

where w_{it} is the net marginal wage, y_{it} is non-labour income, x_{it} is a vector of variables influencing desired hours, and β, δ_1 and δ_2 are coefficients to be estimated. This form

[9] The wage formula is: net wage = 1.5 * (1 – marginal tax rate) * (usual gross pay per month) / [(usual standard weekly hours) + 1.5 * (usual paid overtime weekly hours)] * (12/52).

allows the labour supply curve to bend backwards. Again to facilitate comparison, the explanatory variables were the same as Stewart and Swaffield's: age and age squared, marital status and whether the job was covered by a trade union (to control for any taste differences between covered and non-covered individuals, since union status is also a demand variable).

The selection equation included all the desired hours regressors, plus demand and job characteristics likely to influence constraints. The additional variables were the unemployment/vacancy ratio in the travel-to-work area (TTWA), in order to capture labour market conditions; a dummy variable for living in inner or outer London or not, since one might expect the cost of changing jobs to be less in the densely populated labour market there; years of education; tenure (predicted to determine constraints by various contracting models); and two other dummies describing the man's employment contract: whether it covered a fixed term and whether his pay included bonus payments. By analogy with a Sargan test of over-identifying restrictions, I used a simple test to gain confidence that the variables included in the selection equations but excluded from the desired hours equation, i.e. included in z_{it} but excluded from x_{1it}, were not, in fact, incorrectly excluded from x_{1it}.[10] On this basis I rejected from the selection equation dummy variables for firm size, pay including regular increments, and evening and night work.

Table 3 reports the estimates of the selection model. They show that the probability of overemployment, rather than underemployment, rose with increasing age until about 45 years and then fell off. Several variables had a statistically significant effect, for example those receiving bonus payments were more likely to be overemployed, as were married men. Longer tenure raised the probability of overemployment. In addition to the distributional assumption, the exclusion of tenure from the desired hours equation acts as an identifying restriction. The coefficients on $\ln w_i$ and y_i/w_i imply that higher wages reduced the probability of overemployment at the sample means of wage and non-labour income. The fact that several of the

[10] This involved regressing the residuals from the selection-corrected desired hours equation on all the exogenous variables and then calculating the test statistic NR^2 where N is the unconstrained sample size and the R^2 is from this supplementary regression. The statistic was then compared to the appropriate critical value from the $\chi^2(k-1)$ distribution, where k is the number of excluded variables. The test

controls to be included in the desired hours equation had statistically significant effects in the selection model is evidence that the difference between desired and actual hours is influenced by supply characteristics. As noted above, this difference is an omitted variable in a supply equation estimated on the full sample of workers under the alternative hypothesis of distinct hours regimes. The correlation between this omitted variable and the supply characteristics x_{1it} will bias estimates of the coefficients β. The estimate of ρ is significantly different from zero, consistent with the presence of an permanent effect and justifying use of the panel correction procedure.

[Table 3 around here]

Table 4 presents summary statistics for the calculated correction terms. The permanent-effect correction term is denoted a_1 (derived from (6)) and the transitory effect term (derived from (7)) is denoted a_2. The top panel of the table shows the means of a_1 (over all 1700 individuals in the full sample) and a_2 (over all 5241 observations in the full sample). By construction, these expectations conditional on all selection outcomes d_i should sum to zero over the full sample. The means of a_1 and a_2 are very close to zero, providing a positive check on the calculation procedure. The error is presumably due to rounding. The standard deviation of a_2 is decomposed into that part due to within individual variation and that part due to between individual variation. The former is larger.

[Table 4 around here]

The bottom panel of Table 4 reports the means over the unconstrained sample only. Both means are negative, indicating that unconstrained individuals tend to have unobservable characteristics making them less likely in general to be overemployed (the permanent effect), as well as less likely than usual to be overemployed in a particular period (the transitory effect). The between and within standard deviations for a_2 show that conditional on being unconstrained, there is relatively little variation within individuals in the transitory effect, unlike in the full sample. This lends support to the view that applying a fixed effects estimator to this sub sample without any

assumes that the model is identified by the non-linearity of the joint error distribution, so that the variable exclusions over-identify the model.

further correction may not result in excessive bias (a_1 and to a large extent a_2 are swept out when deviations are taken from individual means).

Table 5 shows estimates of the hours equations. They could have been estimated using generalised least squares (GLS). However, I present OLS estimates for better comparison with the OLS results in the next section. There I argue that the wage is endogenous and propose to instrument it by its lagged value. However, since the instrument is only weakly and not strictly endogenous (it is still correlated with the lagged error), the assumptions underlying GLS (as well as the fixed effects estimator) are violated, which would lead to biased estimates. OLS, in the other hand, is consistent. The coefficients of equation (10) estimated from the unconstrained sub-sample are reported in column (1) and those estimated from the full sample reported in column (6). The estimates in column (6) are comparable with Stewart and Swaffield's OLS estimates derived from their full sample from wave 1 only. The standard errors from my larger nine-wave sample (which are adjusted to take account of arbitrary correlation between observations on the same individual) are considerably smaller as expected, but the point estimates exhibit a similar pattern. An inverted U shaped age profile in hours is evident, as is the strong effect of marital status. The wage coefficient is slightly smaller in magnitude than Stewart and Swaffield's estimate (–10.0) and the coefficient on y_i/w_i is about half the size (though the 95% confidence intervals overlap).

Columns (3) and (4) show OLS estimates with the two correction terms included in the set of regressors. The estimates suggest that the correlation between the transitory effects $\sigma_{1\varepsilon\nu}$ plays a bigger role than that between the permanent effects $\sigma_{1\eta\alpha}$. In both cases individuals who are more likely than expected, given their observable characteristics, to be overemployed tend to desire fewer hours than one would expect. Under the null hypothesis of no selection into the unconstrained sample, the standard errors in column (4) are consistent, and therefore a test of the null can be based on a *t*-test of each individual coefficient and an *F*-test of their joint significance. The coefficient on the transitory term a_2 is significant at the 5% level whilst that on the individual term is only significant at the 10% level. However, a joint test yields an F statistic of 8.0, more than twice as large as the 5% critical value

F(2,1276)=3.0, leading to a rejection of the null hypothesis of no selectivity into the unconstrained regime.

[Table 5 around here]

Under the alternative hypothesis that selection is relevant, the standard errors of the desired hours equation coefficients are no longer consistent. The covariance matrix may be adjusted using a formula which takes account of the heteroscedasticity and additional sampling error introduced into the hours equation by the use of predictions of γ and σ_α to construct a_1 and a_2, rather than their true values. An alternative procedure is to calculate an estimated covariance matrix using the bootstrap method. Bootstrapping the first stage (as well as the second) will generate variation in the predictions of a_1 and a_2 due to sampling variation of the z_{it} vectors. In general bootstrapping over both stages gives larger standard errors than bootstrapping over just the second stage and is pursued as a conservative strategy. Because there are multiple observations on individuals, the bootstrapping procedure draws individuals (clusters of individual observations) rather than single observations. Again this tends to increase the standard errors. The standard errors are reported in column (5) and do not differ very much from their counterparts in column (4).

The estimates in column (3) are consistent under both the null hypothesis that hours determination in the unconstrained sample does not differ from the other regimes as well as under the alternative hypothesis that hours determination does differ. Column (6) reports estimates derived using the full sample, which are efficient under the null but inconsistent under the alternative. They show a very strong inverted U shaped age profile and a more negative wage coefficient than the estimates in column (3). The Hausman statistic, on which the test for systematic differences between the two sets of coefficients is based, is:

$$H = \quad [b_c – b_e]' \, [\mathrm{V}(b_c) – \mathrm{V}(b_e)]^{–1} \, [b_c – b_e] \qquad (11)$$

where b_c is a k x 1 vector of estimates which are consistent under both the null and alternative, and b_e is a k x 1 vector of estimates which are efficient under the null and inconsistent under the alternative. Under the null, this Wald statistic is distributed as $\chi^2(k)$.

Column (8) shows the differences b_c–b_e and column (9) shows the square root of the difference between their estimated variances. The ratio of each pair of statistics from columns (8) and (9) then forms the square root of the Hausman statistic for that individual coefficient and is distributed as $N(0, 1)$ under the null.[11] It can be therefore be used in a *z*-test of the hypothesis for that individual coefficient. The age (squared) and wage coefficients appear to differ significantly between the two equations. These differences are qualitatively similar to those reported by Stewart and Swaffield (1997) between their simple OLS and two-limit Tobit estimates. The joint Hausman statistic calculated from (19) was 27.9. This is distributed as $\chi^2(7)$ under the null, which implies a *p*-value of 0.001.

Both the individual and joint test statistics therefore indicate that the hours regime for manual men who report being unconstrained does differ from that for overemployed and underemployed men. This result has two implications. First, it lends credence to these subjective reports. Second, since the wage coefficient appears to differ significantly when proper account is taken of constraints, it raises questions about how much bias is introduced into estimates of labour supply elasticity if constraints are ignored. This issue is explored in Section 8.

6. Measurement error

In the results presented thus far the issue of the potential endogeneity of the wage rate was ignored. The wage may be endogenous in the desired hours equation for two reasons: first, unobservable characteristics which raise wages may also be associated with the unobserved characteristics of individuals who wish to work longer hours. Second, any measurement error in hours will induce a spurious negative correlation between the calculated hourly wage and the unobserved term in the hours equation, the so-called division bias problem (Borjas, 1980). The first problem is traditionally addressed by using instruments such as education to predict wage levels or growth (in differenced hours equations). However, a review of the literature suggests that these instruments are often questionable either because it is not clear that they do not belong in the hours equation, or because they are poor predictors of the

[11] Since, for an individual test, $H \sim \chi^2(1)$, then $\sqrt{H} \sim N(0, 1)$.

wage measure. No attempt is therefore made here to correct for this source of endogeneity.

The second source is potentially more important (see the studies by Altonji, 1986 and Borjas, 1980) and more soluble with panel data. To guage the size of the problem, a simplified method based on Altonji (1986) is presented which makes use of an additional measure of the wage. It is available in wave 9 onwards and is the response to a direct question asking hourly paid workers for their hourly wage rate. Let:

$w_h = w + e_h$, and

$w_e = w + e_e$

where w_h is the reported hourly wage, w_e is the measure constructed from usual earnings and usual hours, w is the unobserved true hourly wage, and e_h and e_e are random errors uncorrelated with each other and w. The error e_h might be due to poor recall, whilst e_e is assumed to be mainly due to measurement error in hours. The assumption that the errors are additive and independent is probably quite strong, but does enable a simple evaluation of the size of the measurement error problem. Taking (co)variances:

$\text{var}(w_h) = \text{var}(w) + \text{var}(e_h)$,

$\text{var}(w_e) = \text{var}(w) + \text{var}(e_e)$, and

$\text{cov}(w_h, w_e) = \text{var}(w)$ since $\text{cov}(e_h, e_e) = \text{cov}(w, e_e) = \text{cov}(w, e_h) = 0$ by assumption.

The covariance matrix for a sample of hourly paid manual workers from wave 9 is:

	w_h	w_e
w_h	2.71	
w_e	2.32	3.73

These figures imply that $\text{var}(e_h) = 2.71–2.32 = 0.39$, which is 14% of $\text{var}(w_h)$, whereas $\text{var}(e_e) = 3.73–2.31 = 1.42$, which is 38% of $\text{var}(w_e)$. In other words nearly 40% of the variation in the calculated hourly wage appears to be noise, implying extensive mis-reporting of hours worked. Whilst this measurement error will not bias the estimates of a regression in which hours appear on the left-hand side only, the use of hours to derive the wage (a right-hand side variable) will probably give rise to a strong correlation between this calculated wage and the error term. Assuming that this result holds for the waves 1-9 data used in the estimated model, the coefficient estimates may be subject to substantial bias.

6.1 Instrumenting the wage

The panel nature of the data suggests an instrument for the calculated hourly wage: assuming that measurement error is serially uncorrelated, then last year's wage measure should be highly correlated with this year's true wage but uncorrelated with this year's measurement error. Similarly the y_i/w_i term can also be instrumented by its lagged value. To implement this method a reduced form selection equation, excluding the two endogenous terms ($\ln w_i$ and y_i/w_i), was estimated (results in Table 6). The number of observations was reduced since individuals had to be present in the preceding wave so that lagged observations were available.

[Table 6 around here]

After calculating a_1 and a_2, the two endogenous variables $\ln w_i$ and y_i/w_i were separately regressed on all the exogenous variables of the model, including their own lagged values (the instruments) and a_1 and a_2, and predictions generated for inclusion in the second stage hours equation. The estimated coefficients from these two regressions are reported in Table A1 of Appendix A. This procedure is as used by Ham (1982). The regressions were included in the bootstrap routine with the selection equation, in order that the resulting standard errors took account of the use of predicted values.

Table 7 presents the hours equations estimates using predictions of $\ln w_i$ and y_i/w_i.

[Table 7 around here]

Though not strictly comparable with the previous results because of the reduced sample size, columns (1) and (2) indicate that endogenising the wage removes the inverted U shaped age profile and reduces the magnitude of the wage coefficient, though not greatly affecting the coefficient on y_i/w_i. In columns (3) and (4), the correction terms are jointly significant (F = 6.0, *p*-value = 0.003) and individually significant, so the selection result is not affected by endogenising the wage. The results derived from the full sample and shown in columns (6) and (7) do exhibit the inverted U age profile but the point estimates are only about half the size of those in the uninstrumented equations. The wage coefficient is also substantially smaller in magnitude (-6.7 compared to –9.4 in Table 5). The individual Hausman statistics indicate the familiar pattern of significantly different coefficients. The joint test statistic is $\chi^2(7) = 20.35$.

As noted in Section 3.2, this correction method is based on the assumption that the errors in the selection and hours equations are joint normal. Following Vella and Verbeek (1999), one can relax the normality assumption in the hours equation by specifying simply that its error components η_i and e_{it} are related linearly to the selection equation errors. Equations (6) and (7) can thus be simplified to:

$$E[\eta_{1i} \mid d_i] = \tau_1 . (1/T_i) \sum_{t=1}^{T_i} E[u_{it} \mid d_i] \qquad (12)$$

and

$$E[\varepsilon_{1it} \mid d_i] = \tau_2 E[u_{it} \mid d_i] \qquad (13)$$

where τ_1 and τ_2 are parameters. Table 8 presents the results derived under this relaxed assumption. The estimates in columns (3)-(5) are very close to their counterparts in Table 7, except for the coefficients on the correction terms, which are estimates of τ_1 and τ_2, and naturally reflect the different assumed error correlation structure. Both individual and joint Hausman statistics imply a strong rejection of the hypothesis that constrained and unconstrained men are in the same hours regime, so reinforcing the preceding conclusions.

[Table 8 around here]

7. Demand-side variables and constrained workers

The results in the preceding sections indicate that the hours determination of unconstrained workers does differ from that of constrained workers. One implication, leading to an additional check on the results, is that demand-side variables should help to explain the hours of the latter group, but not those of the former. I therefore investigated the effect of the measure of local labour market conditions, the travel-to-work area U/V ratio. Table 9 reports the coefficients of hours equations which include the U/V ratio as an extra regressor, estimated over the sub-samples of unconstrained, underemployed and overemployed men. A note of caution is in order with regard to the estimates derived from the two sub-samples of constrained individuals. Since one cannot *a priori* exclude any demand variables from these equations, the estimates may be biased by omitted variables and identification may in general be weak.

Column (1) shows the estimates based on the unconstrained sample. The coefficient on U/V ratio is insignificant, which is not surprising since the tests of overidentifying restrictions had already indicated that it did not belong in the desired hours equation. The estimates for underemployed men are reported in column (2). The coefficient on U/V ratio is statistically significant (t = 2.51) and more than eight times that in column (1). The sample standard deviation of the U/V ratio is 10.7, so an increase of one standard deviation would increase expected hours by nearly 1.5 hours. By contrast, the coefficient on U/V ratio in the equation for overemployed men (column (3)) is negative, though still statistically significant (t = 3.9) and economically large: a one standard deviation increase in the U/V ratio would *reduce* expected hours by just over an hour. Notwithstanding the caveat above about the possible fragility of these estimates, they suggest that the hours of constrained individuals are influenced by the local labour market, and in different ways for the two constrained groups. Underemployed men work longer hours when labour market conditions are less favourable, possibly reflecting precautionary motives and fear of unemployment (recall that constrained hours will be determined by a combination of supply and demand factors). On the other hand overemployed men work fewer hours: their hours seem to reflect local labour demand directly. It may be noted in passing, that this effect is at variance with the monopsony explanation of overemployment: hours increase in a tight labour market, holding the wage fixed.

[Table 9 around here]

8. Labour supply elasticities

The functional form of the hours equation implies the following expressions for the wage and income elasticities:

$$\varepsilon_{wu} = (\delta_1 - \delta_2 \,.\, y/w)/h$$

$$\varepsilon_{wc} = \varepsilon_{wu} - \delta_2$$

$$\varepsilon_y = \delta_2 \,.\, y/wh$$

where ε_{wu} (ε_{wc}) is the uncompensated (compensated) wage elasticity and ε_y is the income elasticity.

Table 10 shows the calculated elasticities (evaluated at the sample means of the wage, hours and non-labour income) from the full sample without correction and from the unconstrained sample with selectivity-correction, when the wage is assumed exogenous and endogenous.

[Table 10 around here]

Whether or not the wage is assumed exogenous, estimation on the full sample of workers (assumed unconstrained) tends to exaggerate the responsiveness of hours to wages (by 20-40% for the uncompensated elasticity) and non-labour income (by 40-60%). The results suggest that the hours of constrained individuals increase more than they would like when their wages or non-labour income fall. Ignoring measurement error in the wage also seems to impart a large downward bias of about 40-60%. The preferred estimates in the final column imply a labour supply function which nevertheless bends backwards with elasticity 0.1. The compensated (substitution) elasticity is positive as predicted by theory with value 0.1. The combination of bias in the uncompensated elasticity estimates due to the assumptions that all workers are unconstrained and that the wage is exogenous is nearly 100%.

9. Conclusions

If workers' reported dissatisfaction with working hours reflects actual restrictions on their choice of hours, then unconstrained workers' hours should be determined differently from those of their unconstrained colleagues. Using a general selection model I have found evidence in support of this prediction for male manual workers, lending credence to their subjective reports of constraints. This evidence corroborates other work done on the BHPS by Stewart and Swaffield (1997), who used a cross-sectional sample and a more specific model, and Böheim and Taylor (2001), who found that adjustments in working hours were related to previous constraints.

The results suggest that policy concern over hours and 'work-life balance' may be justified, since many individuals appear unable to attain their utility maximising choice of working time. On the other hand, in the long-term contracting models mentioned in Section 2, constraints are efficient responses to asymmetric information, so the welfare implications are less clear. One drawback to the framework used here is that, while it has incorporated individual heterogeneity through the use of panel data, more dynamic aspects, such as the influence on hours of expected future wages, have not been modelled.

The results also point to large proportionate biases in the estimates of labour supply elasticities if constraints and the inherent endogeneity of derived hourly wages are ignored. Nevertheless, all the estimates obtained are small, suggesting that the wage is not the main driver of hours worked. Indeed the magnitude of the wage effect is smaller in the corrected measures. The labour supply curve bends backwards but only slightly.

Table 1
Sample summary statistics, manual men

	Underemployed		Unconstrained		Overemployed		All men	
Variable	Mean	St dev	Mean	St dev	Mean	St dev	Mean	St dev
Total weekly hours	40.63	10.43	44.90	9.48	48.67	10.16	45.96	10.09
Age	35.90	11.63	38.57	11.97	40.95	11.02	39.24	11.69
Married	0.68	-	0.74	-	0.82	-	0.76	-
Trade union covered	0.53	-	0.54	-	0.60	-	0.56	-
Log (net marginal wage)	1.65	0.35	1.71	0.34	1.69	0.34	1.69	0.34
Non-labour income/marginal wage	1.50	4.76	1.67	5.32	1.25	4.21	1.50	4.90
Non-labour income (£ / week)	7.80	27.45	8.38	25.48	6.38	20.69	7.59	23.99
Tenure (yrs)	4.83	6.25	6.35	7.46	7.57	7.69	6.68	7.50
Education (yrs)	11.74	1.41	11.63	1.30	11.52	1.14	11.60	1.25
Bonus payments (incidence)	0.33	-	0.37	-	0.40	-	0.38	-
Fixed term contract	0.03	-	0.02	-	0.01	-	0.02	-
London	0.07	-	0.05	-	0.07	-	0.06	-
U/V ratio (in travel to work area)	12.64	10.75	11.93	10.54	12.06	11.01	12.04	10.73
Underemployed	1	-	0	-	0	-	0.079	-
Unconstrained	0	-	1	-	0	-	0.549	-
Overemployed	0	-	0	-	1	-	0.373	-
N (person-years)	414		2876		1951		5241	

Notes: (1) Pooled data from BHPS waves 1-9. (2) All income variables are expressed in 1991 prices.

Table 2
Hours constraints over the panel

	Overall		Ever in state		Time in state
	Frequency	Percentage	Frequency	Percentage	Percentage
Constraint state					
Underemployed	414	7.90	303	17.82	38.87
Unconstrained	2876	54.88	1277	75.12	66.93
Overemployed	1951	37.23	885	52.06	55.63
Total	5241	100.00	2465	145.00	

Notes: (1) Number of individuals = 1700; mean number of observed waves = 3.1

Table 3
The probability of being overemployed, unconstrained or underemployed (random effects ordered probit model)

Variable	γ
Age	0.0897***
	(0.0168)
Age^2	-0.0010***
	(0.0002)
Married	0.2183***
	(0.0633)
Trade union covered	0.0444
	(0.0530)
Log (net marginal wage)	-0.2391***
	(0.0738)
Non-labour income/marginal wage	-0.0118**
	(0.0048)
Tenure	0.0163***
	(0.0039)
Education	-0.0068
	(0.0221)
Bonus payments	0.0995**
	(0.0454)
Fixed term contract	-0.2120
	(0.1435)
London	0.1727
	(0.1250)
U/V ratio	-0.0035*
	(0.0021)
κ_1	-0.1389
	(0.4176)
κ_2	2.1796***
	(0.4191)
ρ	0.4236***
	(0.0199)
N	5241
Log likelihood	-4251.4
Model significance	$\chi^2(12) = 132.66$

Notes: (1) Dependent variable: hrscon = 0 if underemployed; hrscon = 1 if unconstrained; hrscon = 2 if overremployed.
(2) Asymptotic standard errors in brackets.
(3) * significant at 10% confidence level; ** significant at 5% confidence level; *** significant at 1% confidence level
(4) The coefficient ρ is defined as $\rho \equiv \sigma_\alpha^2 / (\sigma_v^2 + \sigma_\alpha^2) = \sigma_\alpha^2 / (1 + \sigma_\alpha^2)$

Table 4
Summary statistics of calculated correction terms

Variable	Mean	Standard deviation		Observations
Full sample				
a_1	-0.00006	Overall	0.977	1700
a_2	-0.00002	Overall	0.841	5241
		Between	0.534	
		Within	0.757	
Unconstrained sample				
a_1	-0.300	Overall	0.606	942
a_2	-0.331	Overall	0.366	2876
		Between	0.409	
		Within	0.067	

Table 5
The determinants of desired hours (OLS with correction terms from random effects ordered probit model)

	Unconstrained sample (N=2876)					Full sample (N=5241)		Hausman statistics	
	b	Rob SE	b_c	Rob SE	B/S SE	b_e	B/S SE	$b_c–b_e$	$\sqrt{[v(b_c)–v(b_e)]}$
Variable	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)
Age	0.260	0.147	-0.190	0.220	0.252	0.608	0.138	-0.798	0.211
Age^2	-0.004	0.002	0.000	0.002	0.003	-0.008	0.002	0.009	0.002
Married	3.378	0.582	2.421	0.698	0.757	2.990	0.496	-0.569	0.571
TU cover	0.639	0.505	0.237	0.508	0.560	0.498	0.425	-0.261	0.364
ln(w)	-8.403	0.754	-7.450	0.819	0.828	-9.355	0.731	1.905	0.388
y/w	-0.240	0.059	-0.170	0.063	0.071	-0.253	0.051	0.083	0.050
a_1			-2.504	1.206	1.252				
a_2			-7.025	2.329	2.507				
Constant	53.463	2.609	59.320	3.389	3.901	49.517	2.424	9.803	3.057
R^2	0.12		0.13			0.12			

Note: (1) Rob SE: robust standard error; B/S SE: bootstrap standard error (2) All standard errors are adjusted for clustering on individuals. (3) Number of bootstrap replications: 100

Table 6
The probability of being overemployed, unconstrained or underemployed (random effects ordered probit model, reduced form)

Variable	γ
Age	0.0579**
	(0.0235)
Age^2	-0.0006**
	(0.0003)
Married	0.1992**
	(0.0860)
Trade union covered	0.0136
	(0.0694)
Tenure	0.0152***
	(0.0049)
Education	-0.0464
	(0.0291)
Bonus payments	0.0735
	(0.0570)
Fixed term contract	-0.2305
	(0.2198)
London	0.3536**
	(0.1680)
U/V ratio	-0.0053**
	(0.0027)
κ_1	-1.0301*
	(0.5818)
κ_2	1.4662**
	(0.5814)
ρ	0.4851***
	(0.0236)
N	3467
Log likelihood	-2712.1
Model significance	$\chi^2(10) = 55.22$

Notes: (1) Dependent variable: hrscon = 0 if underemployed; hrscon = 1 if unconstrained; hrscon = 2 if overremployed.
(2) Asymptotic standard errors in brackets.
(3) * significant at 10% confidence level; ** significant at 5% confidence level; *** significant at 1% confidence level

Table 7
The determinants of desired hours (OLS with correction terms from random effects ordered probit model, wage terms instrumented)

	Unconstrained sample (N=1897)					Full sample (N=3467)		Hausman statistics	
	b	Rob SE	b_c	Rob SE	B/S SE	b_e	B/S SE	b_c–b_e	$\sqrt{[v(b_c)–v(b_e)]}$
Variable	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)
Age	-0.082	0.183	-0.365	0.203	0.257	0.335	0.167	-0.700	0.196
Age^2	0.000	0.002	0.003	0.002	0.003	-0.005	0.002	0.007	0.002
Married	2.556	0.717	1.788	0.806	0.856	2.110	0.714	-0.322	0.473
TU cov	0.460	0.634	0.259	0.631	0.682	0.093	0.521	0.166	0.441
ln(w)	-5.288	1.360	-5.247	1.361	1.466	-6.701	1.202	1.454	0.840
y/w	-0.232	0.098	-0.207	0.096	0.100	-0.292	0.078	0.086	0.063
a_1			-3.688	1.657	1.778				
a_2			-7.831	2.691	2.815				
Constant	55.680	3.515	59.443	3.693	4.547	51.285	3.223	8.158	3.208
R^2	0.05		0.06			0.05			

Note: (1) Rob SE: robust standard error; B/S SE: bootstrap standard error (2) All standard errors are adjusted for clustering on individuals. (3) Number of bootstrap replications: 100

Table 8
The determinants of desired hours (OLS with correction terms from random effects ordered probit model, wage terms instrumented, alternative error distribution)

	Unconstrained sample (N=1897)					Full sample (N=3467)		Hausman statistics	
	b	Rob SE	b_c	Rob SE	B/S SE	b_e	B/S SE	$b_c – b_e$	$\sqrt{[v(b_c)–v(b_e)]}$
Variable	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)
Age	-0.082	0.183	-0.341	0.203	0.251	0.335	0.167	-0.676	0.188
Age^2	0.000	0.002	0.002	0.002	0.003	-0.005	0.002	0.007	0.002
Married	2.557	0.717	1.925	0.794	0.834	2.110	0.714	-0.185	0.432
TU cov	0.460	0.633	0.247	0.629	0.673	0.093	0.521	0.154	0.426
ln(w)	-5.290	1.359	-5.265	1.360	1.482	-6.701	1.202	1.436	0.867
y/w	-0.232	0.098	-0.203	0.096	0.099	-0.292	0.078	0.089	0.062
a_1			2.811	0.850	0.806				
a_2			-6.881	2.582	2.620				
Constant	55.681	3.512	59.302	3.715	4.430	51.285	3.223	8.017	3.039
R^2	0.050								

Note: (1) Rob SE: robust standard error; B/S SE: bootstrap standard error (2) All standard errors are adjusted for clustering on individuals. (3) Number of bootstrap replications: 100

Table 9
The effect of local labour market tightness on hours

	Unconstrained men (1)	Underemployed men (2)	Overemployed men (3)
U/V ratio	-0.016	0.1348**	-0.1016***
	(0.025)	(0.0537)	(0.0258)
Age	-0.358*	-0.8398	0.9308***
	(0.205)	(0.5818)	(0.2848)
Age^2	0.003	0.0068	-0.0115***
	(0.002)	(0.0068)	(0.0031)
Married	1.834**	-4.7287**	1.7752
	(0.812)	(1.8657)	(1.2487)
Trade union covered	0.303	-0.0396	-0.4564
	(0.631)	(1.3377)	(0.7884)
Log (net marginal wage)	-5.254***	-3.6502	-8.7907***
	(1.358)	(2.5788)	(1.4212)
Non-labour income/marginal wage	-0.212**	-0.6145*	-0.2839***
	(0.097)	(0.3643)	(0.1036)
a1	-3.540**	-16.0383***	5.0962**
	(1.738)	(3.5653)	(2.0043)
a2	-7.572***	-25.1705***	7.6493**
	(2.821)	(4.9889)	(3.5301)
Constant	59.500***	14.2132	37.4033***
	(3.685)	(10.7281)	(9.5021)
Observations	1897	233	1337
R-squared	0.06	0.20	0.11

Notes: (1) Dependent variable: total usual weekly hours
(2) Standard errors in brackets.
(3) * significant at 10% confidence level; ** significant at 5% confidence level; *** significant at 1% confidence level

Table 10
Elasticities of weekly hours

Elasticity	Exogenous wage		Endogenous wage	
	All men	Unconstrained men (corrected)	All men	Unconstrained men (corrected)
Uncompensated wage	-0.200	-0.160	-0.139	-0.110
Compensated wage	0.053	0.010	0.153	0.097
Income	-0.0086	-0.0058	-0.0099	-0.0070

Appendix A

Table A1
First-stage regressions of current wage and income variables on their lagged values (OLS estimates)

Dependent variable	$\ln(w_i)_t$	$(y_i/w_i)_t$
Age	0.004	-0.006
	(0.016)	(0.248)
Age^2	0.000	0.001
	(0.000)	(0.003)
Married	0.054	-0.490
	(0.055)	(0.852)
Trade union covered	0.054***	-0.230
	(0.012)	(0.185)
$\ln(w_i)_{t-1}$	0.666***	0.431*
	(0.017)	(0.261)
$(y_i/w_i)_{t-1}$	0.001	0.700***
	(0.001)	(0.017)
Tenure	0.000	-0.024
	(0.004)	(0.065)
Education	0.018	-0.013
	(0.013)	(0.208)
Bonus payments	0.021	-0.150
	(0.023)	(0.350)
Fixed term contract	-0.035	0.926
	(0.077)	(1.183)
London	0.087	0.302
	(0.102)	(1.566)
U/V ratio	-0.001	-0.019
	(0.002)	(0.024)
a1	0.017	0.085
	(0.249)	(3.844)
a2	0.033	0.384
	(0.415)	(6.392)
Constant	0.257**	0.042
	(0.103)	(1.583)
N	1897	1897
R^2	0.52	0.54

Notes: (1) Standard errors in brackets.
(2) * significant at 10% confidence level; ** significant at 5% confidence level; *** significant at 1% confidence level

Appendix B

Table B1
Definition of variables

Variable	Definition	BHPS variables used
Total weekly hours	Total weekly hours normally worked including paid and unpaid overtime	JBHRS, JBOT
Age	Age on 1st December of fieldwork year	AGE12
Married	= 1 if married or cohabiting; = 0 otherwise	MASTAT
Trade union covered(1)	= 1 if trade union or staff association at workplace; = 0 otherwise	TUJBPL
Log (net marginal wage)	= ln {(1.5 * (1 – marginal tax rate) * (usual gross pay per month) / [(usual standard weekly hours) + 1.5 * (usual paid overtime weekly hours)] * (12/52)}	PAYGU, JBHRS, JBOT, JBOTPD
Non-labour income (£ / week)	Total pension, benefit, transfer and investment income received last month	FIMNNL
Tenure	Job tenure in years	CJSTEN
Education	Number of years of education	QFACHI
Bonus payments(1)	= 1 if pay includes bonus payments; = 0 otherwise	JBONUS
Fixed term contract	= 1 if job is covered by fixed term contract; = 0 otherwise	JBTERM, JBTERM1, JBTERM2
London	= 1 if resident of inner or outer London; = 0 otherwise	REGION
U/V ratio	= unemployed stock / vacancy stock in travel to work area	-
Hours constraint indicator	= 0 if underemployed; = 1 if unconstrained; = 2 if overemployed	JBHRLK

Notes: (1) In waves 2-4 the question was only asked if the job had changed. If the job had not changed, I set these variables to their values in the previous wave. (2) All income variables are expressed in 1991 prices.

References

Altonji J. G. (1986), "Intertemporal Substitution in Labor Supply: Evidence from Micro Data", *Journal of Political Economy*, vol. 94, no. 3, pt. 2, S176-S215.

Altonji J. G. and Paxson C. H. (1986), "Job Characteristics and Hours of Work", *Research in Labor Economics*, vol. 8, part A, 1-55.

Ball L. (1995), "Intertemporal Substitution and Constraints on Labor Supply: Evidence from Panel Data", *Economic Inquiry*, 706-724.

Barzel Y. (1973), "The Determination of Daily Hours and Wages", *Quarterly Journal of Economics*, Vol. 87, No. 2, 220-238

Biddle J. E. and Zarkin G. A. (1989), "Choice Among Wage-Hours Packages: an Empirical Investigation of Male Labor Supply", *Journal of Labor Economics*, vol. 7, no. 4, 415-437.

Böheim R and Taylor M P (2001), "Actual and Preferred Working Hours", ISER working paper no. 2001-6 (April), University of Essex.

Borjas J (1980), "The Relationship Between Wages and Weekly Hours of Work: The Role of Division Bias," *Journal of Human Resource*s, Summer 1980, pp. 409-423.

Card D (1990), "Labor Supply with a Minimum Hours Threshold", *Carnegie-Rochester Conference Series on Public Policy*, 33, 137-168.

Ham J. C. (1982), "Estimation of a Labour Supply Model with Censoring Due to Unemployment and Underemployment", *Review of Economic Studies*, 49, 335-354.

Kahn S. and Lang K. (1991), "The Effect of Hours Constraints on Labor Supply Estimates", *Review of Economics and Statistics*, 605-611.

Kahn S. and Lang K. (1996), "Hours Constraints: Theory, Evidence and Policy Implications", unpublished manuscript, Boston University

Kinoshita T. (1987), 'Working Hours and Hedonic Wages in the Market Equilibrium', *Journal of Political Economy*, vol. 95, no. 6, 1262-1277.

Maddala G S (1983), Limited-dependent and qualitative variables in econometrics, *Cambridge University Press, Cambridge.*

Moffitt R A (1999), "New Developments in Econometric Methods for Labor Market Analysis", in Ashenfelter and Card (eds.), *Handbook of Labor Economics, vol. 3A*, Elsevier, 1999.

Stewart M. B. and Swaffield J. K. (1997), "Constraints on the Desired Hours of Work of British Men", *Economic Journal*, 107, 520-535.

Stewart M. B. and Swaffield J. K. (1995), 'Constraints on desired hours of work, trade unions and the length of the working week for British men', mimeo, University of Warwick.

Stewart M. B. and Swaffield J. K. (1999), "Low Pay Dynamics and Transition Probabilities", *Economica*, 66, 23-42.

Terza J. V. (1985), 'Ordinal Probit: a Generalization', *Communications in Statistics, Theory and Methods*, 14, 1-11.

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

Vella, F (1998), 'Estimating Models with Sample Selection Bias: A Survey', *Journal of Human Resources*, 33(1), 127-169

Verbeek M. and Nijman T. (1996), ‘Incomplete Panels and Selection Bias’, in Matyas L. and Sevestre P. (eds), *The Econometrics of Panel Data: a Handbook of the Theory with Applications*, second edition, Kluwer Academic Publishers, Dordrecht.

Wooldridge J. (1995), ‘Selection Corrections for Panel Data Models under Conditional Mean Independence Assumptions’, *Journal of Econometrics*, 68, 115-132.

Zabel J. E. (1992), ‘Estimating Fixed and Random Effects Models with Selectivity’, *Economics Letters*, 40, 269-272.