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TESTING LABOUR SUPPLY AND HOURS CONSTRAINTS

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Abstract

This paper provides empirical evidence on the assumption that individuals freely decide the number of hours they work at a given wage, using US data on prime age males from the National Longitudinal Survey of Youth. Two types of individuals are considered: those who change job between two consecutive periods and those who do not. We estimate an endogenous switching labour supply equation consistent with a life-cycle model under uncertainty. In this context the endogeneity of movements, ignored in previous analysis, proves to be crucial to get consistent estimates. The results confirm that individuals are constrained in the number of hours they work on a given job and that the intertemporal substitution elasticities are usually upward biased when ignoring the possibility of some groups of individuals being off their labour supply curves.

Keywords: Job changes; labour supply; wage elasticities; hours constraints.

JEL classification: J22, C33

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1. Introduction

A basic assumption of an important stream of labour supply literature is that each employer is indifferent to the number of hours his/her employees choose to work. A possible reinterpretation of this assumption would be that workers face a wide range of job options, with each job being paid at the same wage rate but demanding a different number of hours. Within this framework, the labour supply is determined by a set of personal characteristics and wages. Wages are the only job specific characteristics that affect the number of hours an individual works. Therefore, under both interpretations, workers freely choose the amount of hours they want to work at a given wage and they should be in their labour supply curves at every point in time. Several authors have criticised this assumption for excluding hours constraints and simplifying dynamics. If traditional labour supply models are misspecified when the individual faces restrictions in the hours s/he can work, policy analysis based on the individual's response according to them may be very misleading.

The aim of this paper is to provide additional empirical evidence on whether the classical life-cycle labour supply theory can satisfactorily explain changes in working hours between jobs, that is, to check whether or not individuals face hours constraints. If workers may freely vary the amount of hours supplied on a given job and labour supply largely depends on personal characteristics, we should not expect hours to vary more across jobs than within jobs. On the other hand, if labour supply depends on job-specific characteristics or the preferences of the employer on working hours play an important role, such differences in the variances of hours between jobs and within jobs will be present. Behavioural differences among those who change job voluntarily and those who are laid-off[1] are considered alongside with differences between job-movers and job-stayers.

The starting point is Altonji and Paxson (1986, 1988). Using data from the Panel Study of Income Dynamics (PSID) and the Quality of Employment Survey (QES) they find that the variance of changes in hours is between two and four times larger for those who have switched job than for those who are in the same job during two consecutive years. They also find that the data was inconsistent with a model in which hours in a given job are determined by the employer, with workers being free to switch cheaply another job offering an amount of hours equal to the individual optimal level. Their conclusion is that two structural interpretations, or a mixture of both, can be held in view of the results: either individuals are constrained in hours that, as wages, are determined by the employer or, alternatively, many non wage labour supply determinants are job specific and vary greatly across jobs. In other words, characteristics of the job held have a large influence on the amount of hours that the individuals work.

There are several shortcomings in their paper that motivate further research. First, Altonji and Paxson consider the decision of changing job as exogenous[2]. If individuals that change jobs are more likely to do so, the exogeneity assumption may lead to biased parameter estimates for the whole labour supply equation and therefore to biased estimates for the relevant variances of the changes in hours. Secondly, the PSID presents serious difficulties to identify job changes and the variable "hours" is subjected to great measurement error. Finally, more efficient estimates can be proposed using the longitudinal nature of the data.

In this paper we try to overcome the problems referred above. Instead of using the PSID, we use the National Longitudinal Survey of Youth (NLSYTH) from 1985 to 1991 for the US. This survey identifies without ambiguity movements between jobs and hours of work. Consistent and more efficient estimates under the presence of individual heterogeneity are computed, using the Generalised Method of Moments (GMM) as estimation method. Endogeneity of movements may bias estimates if not properly controlled. This method allows us to estimate and test a model with switching regimes that accounts for endogeneity of the decision to change job.

Related previous studies have address misspecification problems of traditional labour supply models under the presence of hours constraints. Abowd and Card (1989) study the covariance structure of earnings and hours changes. They also question the life-cycle model interpretation of labour supply, which implies that changes in productivity influence earnings more than hours. Their result suggests that most changes in earnings and hours occur at fixed hourly wage rates, with earnings and hours covarying proportionally.

Biddle (1988) and Ball (1990) find evidence of misspecification in life-cycle labour supply models. Intertemporal elasticities computed using constrained individuals may reflect contracting arrangements between workers and firms that move workers off their labour supply curves. Intertemporal elasticities computed using only unconstrained workers are more likely to represent workers' labour supply preferences.

Several authors have tried to incorporate hours constraints explicitly in estimation. Ham (1982) extended the traditional tobit type model for working hours by introducing censoring due to under or over employment. Dickens and Lundberg (1993), Kahn and Lang (1991) or Stewart and Swaffield (1997) use information on desired hours of work to check to which extend constraints are relevant for the US, Canada and the UK respectively. An alternative "structural" model is the hours-wage offer models where wages influence hours and hours, simultaneously, influence wages (e.g., Rosen, 1976, Biddle and Zarkin, 1989, or Tummers and Woittiez, 1991).

In this paper, we do not try to explicitly model hours restrictions. Rather we use a traditional framework of labour supply to test whether constraints in hours have any impact on the labour supply functions of prime age males. Endogeneity of movements will prove to be crucial to get consistent estimates of the parameters of interest. When ignored, upward bias in the intertemporal substitution elasticity of wages is found for the group of constrained individuals. Moreover, traditional labour supply models seem misspecified for individuals that do not change job, which casts serious doubts about the efficiency of policies based on them. Also the difference between the variances of the change in hours for workers that move and do not move jobs are downward biased under the simplifying assumption of exogeneity. Although the results contradict to some extend those of Altonji and Paxson (1986, 1988), they encompass previous findings relating misspecification of labour supply for workers rationed in the amount of hours they can work.

The rest of the paper is structured as follows. Section 2 proposes a tractable empirical model of labour supply in a life cycle context and discusses its implications under the presence of hours constraints. Section 3 presents the estimation method and some model specification tests. Characteristics of the data and the selected sample we use are discussed in Section 4. Section 5 analyses the main results and, finally, Section 6 concludes.

2. Empirical Model

This section introduces an empirically tractable model of labour supply compatible with the life-cycle theory and the presence of individual fixed effects. We present the implications of hours constraints derived from the model and the basic specification to be estimated.

2.1. Labour supply and fixed effects

Let us consider an individual with a life time horizon T, who has a utility function in period t depending on consumption, C_t, hours of work, h_t, and conditional on a set of demographic characteristics, Z_t. At period t the individual maximises his lifetime utility, that is, the discounted sum of his by period instantaneous utilities,

$$\max_{C_t, h_t} E_t \sum_{k=t}^{T} \frac{1}{(1+\rho)^k} U_k(C_t, h_t \mid Z_t) = U_t(.) + \frac{1}{(1+\rho)} E_t \sum_{k=t+1}^{T} \frac{1}{(1+\rho)^k} U_k(.) \tag{1}$$

subject to the asset accumulation constraint

$$A_t = (1+r_t)A_{(t-1)} + w_t h_t - C_t \tag{2}$$

where w_t and r_t are real wages and interest rate, respectively; A_t are the assets at the end of period t and ρ denotes the rate of time preference. The expectations operator E_t is taken over future uncertain wages and interest rates. The time dependence of the utility reflects the influence of predetermined shifter variables on life-cycle preferences[3].

Using Bellman's principle, we can define

$$V_{t+1}(A_t) = \max E_{t+1}\Big\{ \sum_{k=t+1}^{T} \frac{1}{(1+\rho)^k} U_k \Big\} \tag{3}$$

Then at period t the individual maximisation problem can be written as follows:

$$V_t = \max_{C_t, h_t, A_t} U_t + \frac{1}{1+\rho} E_t V_{t+1}(A_t) \tag{4}$$

under restriction (2).

For every period, first order conditions for an interior solution are:

$$\frac{\partial U_t}{\partial C_t} = \lambda_t \tag{5}$$

$$\frac{\partial U_t}{\partial h_t} = -\lambda_t w_t \tag{6}$$

$$\lambda_t = E_t\{\frac{1+r_{t+1}}{1+\rho}\lambda_{t+1}\} \tag{7}$$

Equation (7) implies that the individual chooses savings in such a way that his discounted expected marginal utility of wealth remains constant over time. Assuming no uncertainty about the interest rate and the discount factor, expression (7) can be written as

$$\frac{1+r_{t+1}}{1+\rho}\lambda_{t+1} = \lambda_t(1+e_{t+1}) \quad \text{where } E_t(e_{t+1}) = 0 \tag{8}$$

e_{t+1} reflects all unanticipated news gathered in period $t+1$. Expression (8) ensures λ being positive for all t and leads to the approximation

$$\ln\lambda_{t+1} \approx \ln\lambda_t + \rho - r_{t+1} + e_{t+1} \qquad t,\dots\dots T \tag{9}$$

Therefore the means of all future values of λ are revised to account for all forecasting errors at the time they are realised. So, at the start of the life-cycle the consumer sets λ_0 which takes into account all the information on future values of the variables available at that time. According to equation (7), λ_t is revised over time as new information is acquired.

Solving equations (5) to (7) for consumption and hours of work, conditional on λ for each period, the solutions are the so called Frisch or λ-constant demand functions:

$$C_t = C(\lambda_t, w_t; Z_t) \tag{10}$$

$$h_t = h(\lambda_t, w_t; Z_t) \tag{11}$$

where λ acts as a summary of between period allocations and is individual specific. It is therefore an appropriate conditioning variable although not observable in equations (10) and (11). However, assuming some restrictions about the form of within period preferences, λ can be treated as an unobservable individual fixed effect suitable of being differentiated out in the supply and demand equations, provided that these are linear in the logarithm of λ_t[4]. From this λ-constant demands we can compute λ-constant elasticities. They reflect fully anticipated movements along the wage profile[5].

We assume the familiar log-linear specification for the previous labour supply equation (see Ham, 1986, for a similar specification). Recovering individual subscripts and conditioning on wages, personal characteristics and the marginal utility of wealth, an expression for equation (11) is:

$$E(h_{ijt} | x_i, x_{it}, w_{ijt,}, \lambda_{it}) = \alpha_0 + \alpha_1' x_i + \alpha_2' x_{it} + \alpha_3 w_{ijt} + \ln \lambda_{it} \qquad (12)$$

or

$$h_{ijt} = \alpha_0 + \alpha_1' x_i + \alpha_2' x_{it} + \alpha_3 w_{ijt} + \ln \lambda_{it} + \varepsilon_{ijt} \qquad (13)$$

where h_{ijt} is the log of worked hours; w_{ijt} is the log of real wage rate; x_i is a set of labour supply determinants fixed over time (race, sex, education and the like); x_{it} is a set of time variant characteristics (marital status, number of children, non labour income); λ_{it} is the individual specific marginal utility of wealth and ε_{ijt} is a white noise error term.

Taking first differences in (13) all the characteristics fixed over time cancel out and we get an empirically tractable equation,

$$\Delta h_{ijt} = \alpha_2' \Delta x_{it} + \alpha_3 \Delta \ln w_{ijt} + \Delta \varepsilon_{ijt} + v_{ijt} \qquad (14)$$

v_{ijt} is an error component that includes e_t from equation (8), that is, all unanticipated components of wage and demographic variables. The coefficient for the log-wage, α_3, constitutes the intertemporal substitution elasticity of wages in this model. It is predicted to be positive given that an individual will work more in periods with higher wages.

2.2. Movements between jobs and hours constraints

The model presented above, in the same way that most of the conventional labour supply models, assumes that workers can freely choose the number of hours they want to work. Alternatively, hours are determined by employer preferences but workers can move at zero cost towards firms that offer the amount of hours they desire to work. There is a continuous of firms offering the whole range of possible hours of work. Both interpretations are coherent with traditional labour supply models. Wage rates and personal preferences, but not other job specific characteristics, influence primarily hours' choices. Moreover, each individual will be on his/her labour supply curve at every point in time. However, implications for mobility from both interpretations are very different.

Focusing on the first interpretation, we should not expect hours to vary more across jobs than within jobs (see Altonji and Paxson, 1986 and 1988). Finding a higher variance of the change in hours for people who move from one job to another (movers) would suggest that individuals staying in the same job (stayers) are constrained in the number of hours they can choose to work and that they are off their supply curves. Personal characteristics or wages may vary more among movers than among stayers. In this situation the variances of the change in hours would differ, but this difference should vanish once we correct hours for those more variable characteristics.

If the second interpretation is the correct one and changing jobs has no cost, the variance of the change in hours would be bigger for movers than for stayers, just because every individual has to move to change hours. To test whether this interpretation is coherent with the data, as Altonji and Paxson (1986) suggest, we can distinguish between job changes resulting from layoffs and other type of job changes. If layoffs (e.g., plant closings) are exogenous events to the individual decision process and preferences, then workers that experience a layoff will pick new jobs offering an hours' level similar to their previous job. Therefore, the variance in the change in hours, once corrected by wages and personal characteristics, should be similar for stayers and for this subset of movers.

Analytically, consider the model presented in previous section, in particular equation (14). We can distiguish two subgroups of people: the ones that change job between t and t-1 and the ones that stay at the same job.

Lets assume, as starting point, that movements are exogenous and that the variance for both subgroups is the same, that is,

$$\begin{aligned} &E(u_{ijt}S_{ijt}) = 0 \\ &E(u^2_{ijt} \mid S_{it} = 1) = E(u^2_{ijt} \mid S_{it} = 0) = E(u^2_{ijt}) \end{aligned} \tag{15}$$

where $u_{ijt}=\Delta\varepsilon_{ijt}+v_{ijt}$, from equation (14), and S_{it} is a dummy variable that equals one if the individual stayed in the same job between t and t-1. Equation (14) will hold for each subgroup of individuals, movers and stayers:

$$\Delta h^s_{ijt} = \alpha_2{}'\Delta x^s_{it} + \alpha_3(w_{ijt} - w_{ijt\text{-}1}) + u^s_{ijt} \tag{16}$$

$$\Delta h^m_{ijt} = \beta_2{}'\Delta x^m_{it} + \beta_3(w_{ijt} - w_{ij't\text{-}1}) + u^m_{ijt} \tag{17}$$

where $u^s_{ijt} = \Delta e^s_{ijt} + v^s_{ijt}$ and $u^m_{ijt} = \Delta e^m_{ijt} + v^m_{ijt}$; i denotes individual, j job and t time. m refers to movers and s to stayers. For movers, wage in period t corresponds to job j and wage in period t-1 to job j', being j different from j'.

We should expect that movers and stayers behave in the same way, so the parameters of interest must be equal for both subgroups ($\alpha_2 = \beta_2$ and $\alpha_3 = \beta_3$). In addition, if the u's are distributed with constant variance for all individuals, movers and stayers, $\text{var}(\Delta h^s_{ijt})$ may be different from $\text{var}(\Delta h^m_{ijt})$, if personal characteristics or wages are more variable for one subgroup than for the other. However the variance of the adjusted change in hours should be the same. That is:

$$\text{var}(\Delta h^s_{ijt} - \alpha_2{}'\Delta x^s_{it} - \alpha_3\Delta w_{ijt}) - \text{var}(\Delta h^m_{ijt} - \beta_2{}'\Delta x^m_{it} - \beta_3(w_{ijt} - w_{ij't\text{-}1})) \approx 0 \tag{18}$$

or, more compactly,

$$\mathrm{var}(u^{s}_{ijt}) - \mathrm{var}(u^{m}_{ijt}) \approx 0 \qquad (19)$$

Assuming homoscedasticity on the u's distribution means that movers and stayers have the same preferences for hours; movers do not have more variable preferences for hours. If our model is well specified, that is, if we have included all variables that may differ between movers and stayers, constant variance on u's would imply constant variance on both its components, ε and ν. This assumption is quite strong. Abowd and Card (1985) found some evidence against it. Even though, the point is whether all differences in the variance can be explained by differences in tastes. Altonji and Paxson (1986) found that this is not the case. We try to minimise the impact of heterogeneity by the selection of the sample as stated in Section 4, although consistent estimates under heteroscedasticity are obtained[6].

A different behaviour for stayers than for movers will be inconsistent with conventional life-cycle labour supply theory. Therefore estimation of equations (18) and (19) casts some interest by itself and is the first step on our analysis.

Estimation of both equations separately would give consistent parameter estimates under the null hypothesis of exogeneity of the movements. However, if the decision of changing job is correlated with some unobservable characteristic affecting preferences, stayers may show higher preference for stability. If so, estimates may be biased. Therefore possible endogeneity of movements has to be tested.

Nevertheless, whether S_{it} is endogenous or exogenous, estimation of

$$E(\Delta h_{ijt} | x_i, x_{it}, w_{ijt}) = E(S_{it}\Delta h^{s}_{ijt} + (1 - S_{it})\Delta h^{m}_{ijt} | x_i, x_{it}, w_{ijt}, S_{it}) \qquad (20)$$

would give us consistent estimates of the parameters of interest. If endogenous, S_{it} should be instrumented in estimation. Replacing expressions (16) and (17) in (20), the final equation to estimate is:

$$\begin{aligned} \Delta h_{ijt} &= \beta_0 + (\alpha_0 - \beta_0)S_{it} + \beta_2 \Delta x_{it} + (\alpha_2 - \beta_2)S_{it}\Delta x_{it} + \\ &\quad + \beta_3 \Delta w_{\mathrm{ijt}} + (\alpha_3 - \beta_3)S_{it}\Delta w_{\mathrm{ijt}} + u_{\mathrm{ijt}} \end{aligned} \qquad (21)$$

In the easiest case, if there are no differences in the parameters for stayers and for movers, the coefficients on S_{it} and cross products will cancel out. Under this hypothesis, to estimate only one equation with all movers and stayers will be sufficient to recover consistency whereas this is a strong restriction to impose on the data. If coefficients are different for both subgroups of individuals, equation (21) has to be estimated and requires S_{it} to be appropriately instrumented.

Identical coefficients for both groups and endogeneity of the movements are two interesting hypotheses to test. Equation (21), without imposing any restriction on the coefficients, is estimated alongside with equations (16) and (17) separately and results are compared. The estimation method and the implementation of different tests for the model specification are presented in Section 3 below.

3. Estimation Method

This section introduces the estimation method used, the Generalised Method of Moments (GMM), and compare it with Two Stage Least Squares (2SLS). We also present some tests to check for the specification of the model.

3.1 The GMM estimator

Given equations (16), (17), and (21) and a panel of individual observations, we can pool all individuals and periods and consider the sample a cross-section, if parameters remained constant during the whole observation period. This increases the sample size and it gives us consistent estimates under the maintained assumptions.

From previous section, the error term u_{ijt} has two components: ν_{ijt} and $\Delta\varepsilon_{ijt}$. The first component, ν_{ijt}, includes all unanticipated components of the explanatory variables for which there were some uncertainty at period *t*. The second component, $\Delta\varepsilon_{ijt}$, reflects a pure exogenous shock received by the individual at moment *t*. Consequently, some correlation may be present between the explanatory variables dated on *t* for which there were some uncertainty in *t-1* and the first component of the error term. Variables susceptible of such correlation are wages and other income ($E(w_{ijt}v_{ijt}) \neq 0$ and $E(\text{otinc}_{ijt}v_{ijt}) \neq 0$). Moreover, if some of these variables are generated endogenously in the model (as it certainly happens with wages) they may also be correlated with the contemporaneous exogenous shock ($E(w_{ijt}\varepsilon_{ijt}) \neq 0$). We have to find instruments highly correlated with the explanatory variables but uncorrelated with the error term. In doing so, the longitudinal nature of the data is useful. The same endogenous variables lagged two or more periods are valid instruments. Arellano-Bond type instruments are used in the estimation, then exploiting all possible past information. A description of this type of instruments is presented in The Appendix.

The Generalised Method of Moments (GMM) gives consistent estimates of the parameters of interest and more efficient than the simple IV estimator. In this particular case, the IV would be equivalent to Two Stage Least Squares estimator (2SLS).

Lets consider a general equation to be estimated,

$$y_i = x_i'\vartheta + u_i \qquad i = 1 \ldots N \tag{22}$$

alongside with a set of instruments, z_i, correlated with the explanatory variables but uncorrelated with the error term ($E(z_i u_i) = 0$). z_i is a vector of r instruments, where the number of instruments in greater than the number of parameters to estimate, k.

Using an i.i.d. sample of N individuals, an estimate of the parameters of interest is c, the solution to

$$b_N(c) = 1/N\sum_{i=1}^{N} z_i(y_i - x_i'c) = 1/N\,Z'(y - Xc) \tag{23}$$

As the number of instruments is greater than the number of parameters to estimate, there is no unique solution for (23). Therefore, an estimator of ϑ, should minimise:

$$b_N(c)'A_N b_N(c) = (y - Xc)'ZA_N Z'(y - Xc)\frac{1}{N^2} \tag{24}$$

where A_N is a matrix of weights and $Z = \begin{bmatrix} z_1' \\ . \\ z_N' \end{bmatrix}$, is the matrix of instruments. If

$A_N = \left[\dfrac{Z'Z}{N}\right]^{-1}$, the argument that minimise (24) is the familiar 2SLS estimator that would consistently estimate ϑ

$$\hat{\vartheta}_{2SLS} = [(X'Z)(Z'Z)^{-1}(Z'X)]^{-1}(X'Z)(Z'Z)^{-1}(Z'Y) \tag{25}$$

A heteroscedasticity consistent estimate of the asymptotic variance is:

$$\frac{\hat{W}}{N} = N(X'Z(Z'Z)^{-1}Z'X)^{-1}X'Z(Z'Z)^{-1} \left(\frac{1}{N}\sum_{i=1}^{N}\hat{u}_i^2 z_i z_i'\right)(Z'Z)^{-1}Z'X(X'Z(Z'Z)^{-1}Z'X)^{-1} \tag{26}$$

However, under the null hypothesis of heteroscedastic errors, 2SLS is not optimum among the GMM class, given the set of instruments Z (see Arellano and Bond, 1991). Therefore, when $E(u_i|z_i) = \sigma_i$, A_N is optimally chosen as $A_N = (\sum_i \hat{u}_i^2 z_i z_i')^{-1}$, where $\hat{u}_i$ are the residuals from a consistent first stage estimation as in (25). That would give us the more efficient estimate among the GMM class for which $E(z_i u_i) = 0$[7],

$$\tilde{\vartheta}_{GMM} = [(X'Z)(\sum\hat{u}_i^2 z_i z_i')^{-1}(Z'X)]^{-1}(X'Z)(\sum\hat{u}_i^2 z_i z_i')^{-1}(Z'Y) \tag{27}$$

Estimates presented in the Section 5 correspond to the type of robust estimators defined by (27). Standard errors are computed using (26).

3.2. Testing model specification

Endogeneity of the explanatory variables in equations (16), (17) and (21) does not have to be assumed. Along with endogeneity of movements in equation (21) it can be tested. This section introduces two types of specification tests: firstly an endogeneity test is derived and secondly, a test for the adequacy of the instruments is presented.

Lets start with the endogeneity test. Two estimators, both consistent under some null hypothesis, should yield similar sets of estimates. Consider the general equation (22). Let Z_1 and Z_2 be two sets of instruments, where $Z_2 = \{Z_1, Z^*\}$, that is, Z_2 contains the same instruments than Z_1 plus an extra set of instruments, Z^*, that may be correlated with the error term.

With Z_1 we can construct a GMM estimator as in (27) and we will get a consistent estimate, $\hat{\vartheta}_1$, of the true parameter vector ϑ. The use of Z_2 may yield more efficient estimates only if it is uncorrelated with the error term. Applying the GMM method, with Z_2, we obtain an estimate, $\hat{\vartheta}_2$, of the parameter vector ϑ_2. The relevant test can be written as:

H_0: Z^* is exogenous with respect to the error term.

H_1: Z^* is not exogenous with respect to the error term.

Under the null hypothesis $\vartheta_2=\vartheta_1$. The test can be written more specifically using the estimates of these parameters as,

$$H_0 : \hat{\vartheta}_2 = \hat{\vartheta}_1$$
$$H_1 : \hat{\vartheta}_2 \neq \hat{\vartheta}_1$$

and a Wald test for the difference of both estimates can be implemented, using the covariance matrix for the vector $(\hat{\vartheta}_2, \hat{\vartheta}_1)$.

When both estimators are consistent but only one is efficient under the null hypothesis, that test reduces to the Hausman test (Hausman, 1978), where the covariance matrix is $\text{var}(\hat{\vartheta}_2, \hat{\vartheta}_1) = \text{var}(\hat{\vartheta}_2) + \text{var}(\hat{\vartheta}_1)$. However, in our case, $\hat{\vartheta}_2$ does not need be efficient under the null, therefore this simplification can not be used. Mroz (1987), following White (1982), provides an estimator for the covariance matrix that does not rely on either the normality or homoscedasticity of the disturbances and it takes into account the correlation between the two sets of instruments. Therefore, the final test we will use has the following form:

$$W = (\hat{\vartheta}_2 - \hat{\vartheta}_1)'(\text{var}(\hat{\vartheta}_2 - \hat{\vartheta}_1))^{-1}(\hat{\vartheta}_2 - \hat{\vartheta}_1) \longrightarrow \chi_k^2 \qquad (28)$$

where k is the number of parameters of interest and therefore the number of restrictions we are imposing. The actual form for the covariance matrix appears in The Appendix. W is used to test endogeneity of movements as well as endogeneity of wages and other income.

Once endogeneity is tested, the adequacy of the instruments used should also be tested. For doing so a Sargan test is implemented. When the number of instruments, r, is greater than the number of parameters to estimate, k, the constraints implied by the econometric specification can be tested. Estimation of ϑ in (22) equals zero k linear combinations of the r ortogonality conditions defined by (23) over the sample. Therefore, there must be r-k linear combinations that are approximately equal to zero but are not zero. The following test can be derived from previous observation,

$$NS(\hat{\vartheta}) = (Z'(Y - X\hat{\vartheta}))'(\textstyle\sum_{i=1}^{N} \hat{u}_i^2 z_i z_i')^{-1}(Y - X\hat{\vartheta})Z \longrightarrow \chi^2_{r-k} \qquad (29)$$

A small value for this test indicates that the instruments used are accepted, in the sense that the restrictions they impose are close to zero. A variation of this test can be used to test additional instruments. Consider two sets of instruments, Z_1, that contains r_1 instruments, and Z_2, that contains r_2. Z would be the union of both sets and will have $r=r_1+r_2$ instruments. Using Z_1, an estimate $\hat{\vartheta}_1$ of ϑ is obtained; using Z_2 an estimate $\hat{\vartheta}_2$; and using Z an estimate $\hat{\vartheta}$. If the restrictions implied by the use of Z_1 have been accepted via their corresponding Sargan test, $NS_1(\hat{\vartheta}_1)$, or they are considered valid a priori, a statistic, *SDT* (incremental Sargan test), can be constructed to test the additional restrictions implied by the use of Z_2:

$$STD = NS(\hat{\vartheta}) - NS_1(\hat{\vartheta}_1) \longrightarrow \chi^2_{r-r1} \equiv \chi^2_{r2}$$

4. Data

Distinguishing between movers and stayers requires detailed information on job changes as well as on personal characteristics and wages during a reasonably long period. This is the type of information that the National Longitudinal Survey of Youth (NLSY) collects. The NLSY is a longitudinal survey conducted by the US Bureau of the Census and NORC-University over a population of 12,686 young men and women who were among 14 and 22 years old in 1979, when first interviewed. At the beginning, military youth and civilian Hispanics, black and economically disadvantaged white youth were oversampled, but since 1985 the military oversample disappeared and 1643 individuals of the oversampled civilian population were dropped out in 1991. The last available wave is from 1991.

Using the NLSY had some advantages over the use of the PSID. The last one presents serious difficulties to identify job changes: it generally provides no employer codes that uniquely identify jobs. Researchers must rely on reported tenure to infer job changes. Measurement error in tenure responses can lead to incorrect inferences[8]. In addition, the variable hours refer to the full calendar before the interview. Then if a job change occurs in this period, hours can refer to a mixture of hours worked on two sequential jobs. To avoid this problem, Altonji and Paxson try to construct a variable that measures unambiguously either within jobs or between job hours' changes using the hours changes over a three year gap.

Alternatively, in the NLSY all variables refer to an employer code. It provides information for, at most, five jobs by individual and year: starting and finishing dates of contract, worked hours, hourly rate of pay, tenure and so. That allows us to construct a complete work history for each individual and job changes can then be identified. Hours can in this case be measured without ambiguity.

An additional advantage of the NLSY is that it collects information for a particular cohort of the population, those who were born between 1957 and 1965, which are more likely to have the same preferences (discount rates, for example), reducing possible heterogeneity. To concentrate on prime age males, highly attached to the labour market, allow us to disregard participation decisions.

Information for the current or most recent job is more detailed than for other jobs (e.g., hours per year worked), therefore all variables will refer to this job. Then, for example, the variable hours91 is the number of hours worked in the current or most recent job during 1991. A stayer is defined as a worker who has not changed his current or most important job between two consecutive interviews. Movers can be either laid-off or can change job for any other reason.

The period chosen for the analysis is seven years, short enough to assume stability on the parameters of interest and long enough to be able to reach some conclusions. We select a subsample of males continuously interviewed between 1985 and 1991. The subsample contains individuals that reported positive hours of work and wages for every year and that gave valid answers for the rest of variables included in the analysis. In principle, as we want to test hours restrictions, workers holding more than one job at a time in any year will not be considered. Self-employed individuals are neither used in the estimation. That leaves us with a sample of 974 individuals per year. Although we have seven years of data, due to the nature of instruments we are forced to use, discussed in section 3, only five years, from 1987 to 1991, are finally included in estimation. Data for 1985 and 1986 will be used to instrument endogenous variables.

A total of 4870 individual observations constitute the final sample. Of those, approximately 78% are stayers (3806), 17.5% movers not laid-off (851) and 4.4% movers that experience a layoff (213). It is interesting to point out that 35.5% of the sample did not change job during the whole period 1985-1991, and an extra 25% did it only once. This may indicate either that everyone is in his labour supply curve or that moving jobs has some costs.

Along with hourly wage rates, other variables included in the analysis are those which can influence hours of work but change over time. Following the traditional labour supply theory such variables should relate personal characteristics rather than job characteristics. Therefore changes on health status, number of children, marital status and other income are selected. The Appendix provides a detailed explanation of the construction and definition of the relevant variables. Their mean and standard deviation for the selected sample are shown in Table A.1.

Although the survey collects alternative measures for hours, we choose hours per year because it is the more flexible. An individual may be constrained in the number of hours he works per day but in compensation he can have days or weeks off.

[Figures 1 and 2 around here]

Figure 1 shows clearly that dispersion of the change in hours among stayers is far below dispersion among movers. This graph represents the distribution of the change in yearly worked hours over the sample. It distinguishes among the three types of workers we are studying, namely stayers, movers that were laid-off and movers that were not laid-off. Actually, 48% of stayers did not change hours at all while only 12.7% of movers did not (13.75% if they quit the job and 8.45% if they were laid-off).

From this figure, dispersion for laid-off movers is higher than for stayers. Although this finding supports the idea of job changes having some cost, some caution should be taken given that no correction has been made for wages or personal characteristics. Figure 2 presents the corresponding distribution of the changes in the hourly rate of pay over the sample. Wages are also more variable among movers (of both types) than among stayers. Therefore correcting for factors such as wages is necessary before reaching any conclusion.

5. Empirical results

Estimates of the change in hours equations for movers and stayers separately are presented in Table 1 (equations (16) and (17)). Among movers, different equations are also estimated for those who experienced a layoff and for those who move voluntarily, in spite of the reduced sample size of the first subset. Explanatory variables for the change in hours are: two dummies accounting for change in health status (uphlth and dwhlth), change in marital status, change in the number of children, change in other income and change in hourly rate of pay.

Due to the error structure in (16) and (17) valid instruments for wages are dated two periods before current period. That is why the data on 1986 and 1985 cannot be used as additional observations in estimation. Some broad occupational dummies are also included as instruments (professional and non-manual workers, as defined in The Appendix). Occupation is a variable highly correlated with wages but it is also likely to be endogenous. Then lagged values of these dummies are used for estimation purposes. Arellano and Bond type instruments are used, exploiting all possible instruments for every year (see The Appendix for a detailed description). The change in wages for 1991 is instrumented with the hourly rate of pay (in levels) for 1989, 1988, 1987, 1986, 1985; change for 1990 is instrumented with 1988, 1987, 1986, 1985's hourly rate of pay, and so on.

Other income is considered exogenous in this specification. As the bottom line of Table 1 shows, the null hypothesis of exogeneity can not be rejected for any of the four subgroups considered, using a χ^2 test as defined by (28). However, lagged values of other income are also included as additional instruments, jointly with the current change in the variable. Applying the incremental Sargan test, we do not reject the validity of this additional set in any of the cases and some precision is gained in the estimates.

[Table 1 around here]

Most of the variables in Table 1 are unsurprisingly no significant. In general, children and marital status are never very significant for men. About the variable other income, it is necessary to remember that the oldest individual in the sample is 34 years old. Changes in health status are marginally significant for some of the groups. If there is some improvement in health, individuals tend to work more than in the previous year, especially if they are stayers or movers who were laid-off.

Regarding the wage coefficient, lets first concentrate on the subsamples of movers. As it should be expected, a positive coefficient for wages is obtained. The higher the wage of this year is with respect to the previous, the more hours the individual works this year with respect to the previous. However the coefficient is very imprecise, especially when we split the sample of movers in the two subgroups. For stayers also a positive wage elasticity is implied. As we will see from estimation of equation (21), equality of coefficients for movers and stayers would be rejected.

First column of Table 3 shows that the variance of the change in hours is greater for movers (both laid-off and not laid-off) than for stayers. In that column, variances are not corrected by the possible variation of other variables, such as wages. Estimation of equations (16) and (17) enables us to adjust the change in hours for the change in observable factors. Column two in the same table shows that even adjusting by those factors that can vary more for movers than for stayers, the difference in the variance of the change in hours between movers and stayers is still high and significant. Our estimates are a little bit higher than those found by Altonji and Paxson (1986), but also the sample we select is slightly different: we concentrate only on prime age males. The variance in hours is 5.5 times bigger for all movers than for stayers (4.8 times if they where not laid-off and 6.4 times if they experience a layoff). This suggests that some constraint in the hours an individual can work must exist or that the hours of work are influenced by other factors apart from personal characteristics.

The hypothesis of movers having more variable preferences and the possibility of moving jobs at no cost do not seem to be coherent with the data if we compare the behaviour of stayers and movers that were laid-off. As explained in previous section, we should expect that if layoffs are exogenous, as it seems quite plausible, stayers and laid-off behave in a similar way. However, this is not the case when comparing the variances of their change in hours of work. Laid-off have an even greater variance, both unadjusted and adjusted, than the rest of movers, as in Altonji and Paxson (1986) results. Some caution has to be taken with this result due to the small sample size for laid-off.

Estimates in Table 1 are consistent only under the restrictive assumptions that movements between jobs are exogenous and that the error term is equally distributed across individual types. These are strong assumptions that should be tested because there could be some sample selection bias on the estimates. To do so we estimate equation (21) under endogeneity and exogeneity of the movements and compare the results.

Estimates of equation (21) under both hypotheses are presented in Table 2. Only movers that were not laid-off are included in the estimation, although inclusion of the 213 laid-off does not change results, only reinforces them as can be seen in Table A.2 in the Appendix.

[Table 2 around here]

First, we discuss the specification of both hypotheses. Columns of Table 2 are estimated for other income exogenous and wages endogenous. Testing exogeneity of income we get a statistic W=7.881 if movements exogenous and W=2.634 if movements endogenous, both of them following a χ^2 distribution with 12 degrees of freedom. Therefore we can not reject exogeneity of income in any of the specifications. Instruments for the change in other income and its cross product with the dummy Stay include, along with the current change, all possible lagged values and the product of those and the instrument for the variable Stay. Testing this additional set of instruments, they are not rejected[9].

Conversely, wages can not be considered as exogenous (from *W*(wages) in Table 2) in any of the specifications. They are instrumented with two periods lagged values, two occupational dummies and the product of the instrument chosen for Stay and all previous instruments.

The only difference between the two columns of Table 2 is that in the first one, the variable Stay is exogenous, and therefore instrumented by itself. Second column considers Stay endogenous and it is instrumented with its one period lagged value. Stay equals 1 if the individual did not change job between *t-1* and *t* and thus $Stay_{t-1}$ equals 1 if the worker did not move between *t-2* and *t-1*. Therefore, even if Stay is endogenous, $Stay_{t-1}$ would be correlated with ε_{t-2}, but most probably uncorrelated with ε_{t-1} or any of the components of u_t, being a valid instrument for Stay[10].

From Table 2, results under both specifications are quite different. In the specification that assumes exogeneity of movements, hours respond to wage changes more for movers than for stayers ($\beta_1 = 0.519$ and $\alpha_1 = 0.041$), but both elasticities are positive. Hours constraints affecting stayers would be coherent with this finding[11].

The variable Stay is positive and highly significant, indicating that the change in hours is, in mean, bigger for stayers than for movers. However, although significant, its effect is rather small. None of the remaining variables has any significant effect on the change in hours of work. When we compare these results with the separate equations for each group, we see that they are quite different, suggesting that the assumption of the error term being homoscedastic for both groups is quite strong. Therefore estimates in Table 1 will be most probably biased.

When we allow for endogeneity of movements, results change quite significantly. Testing exogeneity of movements produces a test W=181.9 distributed as a χ^2 with 33 d.o.f. The null hypothesis of exogeneity is clearly rejected. Therefore specification with movements endogenous is preferred and the estimation of the separated equations is not valid any more.

First of all, the coefficient on wages for stayers implied by the second column of Table 2 is α_1=-0.137 although it is not significantly different from zero[12]. These results are contrary to those found by Altonji and Paxson (1986). They are similar to those of Ham (1986), where he estimates a labour supply equation similar to ours and allowing for hours constraints in the form of unemployment or underemployment to be endogenous. He found that this misspecification of the labour supply is consistent to different tests about the functional form, the omission of variables and the error structure assumed in the hours equation. The negative intertemporal substitution elasticity implied by the coefficient could show that those restricted individuals are off they labour supply function. In that case, labour elasticities may better be reflecting contracting arrangements between workers and firms[13].

Imposing equal coefficients for movers and for stayers, an intertemporal elasticity estimate of 0.359 is found. This figure is in line with previous results for the US. MaCurdy (1981) found an elasticity of 0.15 for white married men using a sample of the PSID from 1968 to 1977. Altonji (1986) got a higher elasticity (0.267) by including age in the regression. More recently Zabel (1997) replicated Altonji's result for married white men aged 25 to 45 in 1968, using the PSID from 1968 to 1987. Our estimated elasticity is somehow larger, but the selected sample we use is also slightly different: prime age males, married and unmarried, may show a higher wage elasticity. However, given the results when we split the sample between movers and stayers (Table 2), this restriction seems quite strong.

Regarding the rest of the parameters, the coefficient for the dummy Stay changes sing and its effect is bigger than in previous specification. That implies that, on average, individuals tend to work more when they change jobs, suggesting that stayers are constrained to work less hours that they wish[14]. Therefore, the change in hours is now bigger for movers than for stayers, as it should be expected under hours constraints. Other income becomes now marginally significant, at least for movers, and it positively affects the number of hours worked. The remaining variables do not have any effect as in previous specifications.

[Table 3 around here]

Implications of using an invalid specification on the variance of the change in hours can be seen in Table 3. Column 3 gives the estimation for the variance of the change in hours adjusted according to the final valid specification (movements endogenous). We see that the difference for the variance of the change in hours between movers and stayers was underestimated when using the separated equations or considering movements exogenous. The variance for stayers is smaller and the one for movers is higher than under the simplifying assumptions. The variance for movers is around 6.3 times higher than for stayers, which is well over the results obtained by Altonji and Paxson (1986), around 2 to 4 times higher for movers than for stayers.

As mentioned above, results do not change by the use of the whole set of movers (including laid-off). If anything, they become reinforced (see Table A.2), which suggest that laid-off do not behave at all as stayers. That corroborates the hypothesis that job changes are not free for individuals and therefore is not the case that they can move jobs and can remain in their labour supply curve. Column 4 of Table 3 shows the variance of the change in hours obtained when using the whole sample of movers: this is around 6.8 times bigger for movers than for stayers.

6. Conclusions

This paper provides empirical evidence on the assumption that individuals freely decide the number of hours they work at a given wage, using US data on prime age males. If this assumption holds, the behaviour of individuals who change jobs should be similar, once personal characteristics have been corrected for, to the behaviour of individuals that remain in the same job. In other words, the variance of the change in hours for individuals that move should be statistically equal to the variance of the change in hours for individuals that stay. If this assumption does not hold, differences in the variance of the change in hours would appear higher for movers than for stayers. This is because they move, among other reasons, to adjust the number of hours they work.

Taking advantage of the longitudinal nature of the data and by using GMM, we estimate a labour supply equation with endogenous switching regimes, which is consistent with a life-cycle labour supply model under uncertainty. In this context, the endogeneity of movements, ignored in previous empirical work, is crucial to get consistent estimates.

We find that the variance of the change in hours is more than six times higher for movers than for stayers once we correct for wages and other characteristics changing over time and across jobs. This figures are somehow higher than previous estimates (see Altonji and Paxson, 1986). Invalid specification of the model (i. e., the assumption of homogeneity between movers and stayers or the assumption of exogeneity of the movements) leads to downward biases on the estimated variance. Therefore the data suggest that individuals face hours constraints in a given job and change job to change hours of work, among other reasons. On average those individuals that change jobs tend to work more hours. That is consistent with stayers working less than they wish.

Taking into account the behaviour of stayers and laid-off workers, data does not seem compatible with a model in which individuals can change jobs at no cost. If that was the case and layoffs are considered exogenous, stayers and laid-off should behave similarly. However, the behaviour of the laid-off is more similar to the behaviour of other movers than to the stayers. Some caution has to be taken with this result due to the small amount of individuals that were laid-off during the sampling period.

In this paper, we have focused on the group of individuals that are more likely constrained in the number of hours they work. Individuals holding more than one job were not considered here. It remains to be said whether the apparently constrained individuals may take second jobs to avoid their hours constraints, although this is out of the scope of this study.

The results suggest that, at least for prime age males, the standard approach to estimate labour supply functions, which does not take into account the possibility of some groups of individuals being off their labour supply curves, leads to estimate an equation that is misspecified. Previous positive and quite large estimates for intertemporal labour supply elasticities using similar models (Zabel, 1997, MaCurdy, 1981, or Altonji, 1986) were found under the assumption of no constraints in the amount of hours the individual works. In general, wage elasticities for the constrained group are overestimated if the possibility of constraints is not considered. It is specially important to allow for endogeneity of the movements across jobs. It seems quite possible that a more complex labour supply model incorporating job characteristics or employer preferences would reflect better what is detected in the data. These findings are in the line of those of Ham (1986), Biddle (1988) or Ball (1990) and cast some doubts about the effectiveness of policy analysis based on traditional labour supply models.

NOTES

[1] Layoffs are considered to be exogenous events.

[2] They use weighted two-stage least squares to correct for heteroscedasticity associated with the fact that the variance of the error component depends on whether or not the job changes. In the first stage, they instrument wages to correct for its possible measurement error. However they do not correct for the possible inconsistency of the parameter estimates due to the endogeneity of job changes.

[3] All variables are individually indexed. For simplicity of the exposition the individual subscript is dropped.
[4] Heckman and MaCurdy (1980) assumed additive separability of leisure and consumption in the direct utility function while Browning, Deaton and Irish (1985) relaxed this assumption.
[5] See Blundell and Walker (1986) for a detailed discussion of this type of demand functions.
[6] Here, again, the distinction between laid-off movers and quitters may help: workers that experimented an exogenous layoff should have the same preferences than stayers. Therefore if differences are found also between these two groups they can not be just attributable to heterogeneity
[7] Other set of instruments can be more efficient. That is why (27) is efficient *only* conditional on the set of instruments used.
[8] See Brown and Light (1992) for detailed discussion.
[9] With an incremental Sargan test, S1=23.57, distributed as a χ^2 with 30 d.o.f., we can not reject the use of these additional instruments. This figure corresponds to the specification that considers endogenous movements.
[10] An exogeneity test for $Stay_{t-1}$ was performed and gave a value of W=3.061. The statistic is distributed as a χ^2 with 31 d.o.f, and therefore the null hypothesis of exogeneity of the instrument could not be rejected.
[11] This is the opposite that Altonji and Paxson (1986) found. For the more clear measure of change in hours (based on a three year gap period), the coefficient for stayers is more or less of the same range than ours but it is the same for movers. For some other measures they found a negative coefficient for movers.
[12] A Wald test of H_0: $\alpha_1=0$, gave a value W=1.229 distributed as a χ^2 with a d.o.f.
[13] See Biddle (1988).
[14] Kahn and Lang (1992) and Dickens and Lundberg (1993) found similar evidence using desired and actual hours of work.

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7. Appendix

A: Variable description

Dwhlth: dummy that equals 1 if at period *t-1* the individual did not have health limitations but he has them at *t*.

Uphlth: dummy variable that equals one if individual had some health limitations at *t-1* but he does not have any at period *t*.

ΔChildren: dummy that equals one if the individual did have a child between *t-1* and *t*.

ΔMarital Status: dummy that equals one if the individual married between *t-1* and *t*. Dummy for marital disruptions did not make any difference in estimates.

Wage: real rate of pay per hour in the current or most recent job for every year, deflated by the USA IPC for all items (constant at 1991 prices). It is a created variable constructed from the answers to two questions and the variable of hours relevant in each case. Q1: How much do you usually earn at that job? Q2: Was that per day, per hour, per week or what?

Around 40% of the individual answered per hour in Q2, 15% per week and 20% per year for every year. Then it can be subject to some measurement error.

Hours/year: variable constructed as hours per week times weeks per year worked in the current or most recent job.

Otinc: annual before taxes income from other sources but paid work, in the household, deflated by the USA IPC for all items. We created this variable as the sum of a set of variables referring other household sources of income:

- Annual income from UC benefits for the respondent or his wife.
- Annual income from child support for the respondent or his wife.
- Annual income from AFDC for the respondent or his wife.
- Annual income from SSI for the respondent or his wife.
- Annual income from veteran benefits, workers' compensation or disability for the respondent or his wife.
- Annual income from welfare for the partner (not wife).
- Annual income from welfare for other family members.
- Wife or partner's income (gross) from wages or farm/business.
- Annual income from regular sources for other family members.
- Hours per year are hours per week times weeks per year.
- Income from other sources (interest, dividends,...).

Occupation: two occupational dummies are constructed as additional instruments for wages. Ocup1 equals one if the individual is a professional and Ocup2 equals 1 if he is a non-manual worker. The 70 Census (three-digit decomposition) is used.

Stayer: individual that stayed in the same job between period *t-1* and *t*.

Stay: dummy variable that equals one if the individual worked for the same employer in his current or most important job between two consecutive periods, *t-1* and *t*.

Mover: individual that changed his most important or recent job between *t-1* and *t*.

Layoff: individual that changes job because, as stated by himself, he was either laid-off (62.91%) or the plant where he worked closed (10.8%) or he was discharged or fired (26.29%).

B: Some technical questions

B.1. Instruments Structure

Assume we only have information from 1988 to 1991. Arellano-Bond type instruments use all past information in the more efficient possible way. Therefore, due to the structure of the error term in Section 2, endogenous observations of year *t* would be instrumented with observations of the variable dated *t*-2 and backwards. That is, observations corresponding to 1991, would be instrumented by observations in 1989 and 1988, in this case, and observations corresponding to 1990 would be instrumented with the variables dated in 1988. Lets consider wages. Valid instruments in this example are w1, w2, w3, constructed as below.

T	N	Δw	w1	w2	w3
91	1	$(w_{91}-w_{90})_1$	$(w_{89})_1$	$(w_{88})_1$	0
91	2	$(w_{91}-w_{90})_2$	$(w_{89})_2$	$(w_{88})_2$	0
.	.	.	.	.	.
91	N	$(w_{91}-w_{90})_N$	$(w_{89})_N$	$(w_{88})_N$	0
90	1	$(w_{90}-w_{89})_1$	0	0	$(w_{88})_1$
90	2	$(w_{90}-w_{89})_2$	0	0	$(w_{88})_2$
.	.	.	.	.	.
90	N	$(w_{90}-w_{89})_N$	0	0	$(w_{88})_N$

where T and N represent respectively time and individual to which the observation corresponds. Δw is the variable to be instrumented.

B.2. Covariance matrix for correlated estimators

Section 3 introduced a test for the possible endogeneity of some of the explanatory variables. Given the form of the objective functions from equation (24) and from Mroz (1987), the covariance matrix of a vector $(\hat{\vartheta}_1, \hat{\vartheta}_2)$ has the following expression:

$$\operatorname{var}(\hat{\vartheta}_1, \hat{\vartheta}_2) = A^{-1}B'CBA^{-1}$$

$$A = \begin{bmatrix} X'Z_1(Z_1'Z_1)^{-1}Z_1'X & 0 \\ 0 & X'Z_2(Z_2'Z_2)^{-1}Z_2'X \end{bmatrix}$$

where $$B = \begin{bmatrix} X'Z_1(Z_1'Z_1)^{-1} & 0 \\ 0 & X'Z_2(Z_2'Z_2)^{-1} \end{bmatrix}$$

$$C = \begin{bmatrix} \sum_{i=1}^{N} \hat{u}_{1i}^2 z_{1i} z_{1i}' & \sum_{i=1}^{N} \hat{u}_{1i}\hat{u}_{2i} z_{1i} z_{2i}' \\ \sum_{i=1}^{N} \hat{u}_{1i}\hat{u}_{2i} z_{1i} z_{2i}' & \sum_{i=1}^{N} \hat{u}_{2i}^2 z_{2i} z_{2i}' \end{bmatrix}$$

Table 1: Change in hours equation by individual type

	Stayers	Movers	M.not layoff	M.layoff
Intercept	0.017	-0.066	-0.023	-0.135
	(0.006)	(0.024)	(0.026)	(0.054)
Δwage	0.229	0.386	0.242	0.468
	(0.106)	(0.169)	(0.172)	(0.326)
Uphlth	0.088	0.387	0.166	1.430
	(0.063)	(0.171)	(0.148)	(0.969)
Dwhlth	-0.053	0.034	0.104	-0.915
	(0.045)	(0.319)	(0.277)	(0.727)
Δchildren	0.011	-0.008	-0.033	0.018
	(0.013)	(0.061)	(0.062)	(0.155)
Δmarital status	0.012	0.033	0.031	0.102
	(0.020)	(0.058)	(0.061)	(0.177)
Δotinc	0.0002	-0.008	-0.010	0.008
	(0.001)	(0.009)	(0.010)	(0.019)
Observations	3806	1064	851	213
R^2	0.077	0.078	0.039	0.104
Sargan test	55.64	38.94	39.23	39.03
(p-value)	(0.041)	(0.473)	(0.459)	(0.152)
W (Income)	0.002	2.502	2.206	0.493
(p-value)	(1.000)	(0.927)	(0.948)	(0.152)

Notes: Estimation Method: GMM. Standard errors in brackets.

Instruments include a constant, Δmarital status, Δchildren, dwhlth, uphlth and Δotinc plus Arellano-Bond type instruments with lagged values of otinc and wages dated *t-2* and backwards (see Appendix for an explanation of this type of instruments), and two occupational dummies (as defined in the Appendix) dated in *t-1*.

None of the movers experience a layoff during 1988. Instruments used in estimation for this subgroup are adjusted by this fact. W(Income): Exogeneity test for income.

Table 2: Change in hours pooling individuals (excluding layoffs).

	Stay exogenous	Stay endogenous
Intercept	-0.044	0.181
	(0.026)	(0.072)
Stay	0.055	-0.194
	(0.027)	(0.080)
Δwage	0.492	0.518
	(0.165)	(0.273)
Δwage x Stay	-0.452	-0.654
	(0.203)	(0.298)
Δotinc	-0.011	0.051
	(0.010)	(0.036)
Δotinc x Stay	0.011	-0.057
	(0.010)	(0.040)
Δmarital status	0.025	-0.206
	(0.064)	(0.283)
Δmarital status x Stay	-0.002	0.257
	(0.066)	(0.346)
Uphlth	0.157	0.183
	(0.150)	(0.307)
Uphlth x Stay	-0.073	-0.098
	(0.163)	(0.376)
Dwhlth	0.174	-0.013
	(0.295)	(0.671)
Dwhlth x Stay	-0.206	-0.001
	(0.299)	(0.739)
R^2	0.078	0.146
Sargan test	193.6	87.28
(p-value)	($5.02e^{-5}$)	(0.221)
W(Income)	7.881	2.634
(p-value)	(0.794)	(0.998)
W(wages)	350.2	30.07
(p-value)	(0.000)	(0.003)
W(Stay)	181.9	
(p-value)	($1.722e^{-32}$)	
Observations	4657	

Notes: Estimation method: GMM. Standard errors in brackets. *Common instruments to both specifications*: constant, Δmarital status, Δchildren, dwhlth, uphlth and Δotinc plus Arellano-Bond type instruments with lagged values of otinc and wages dated *t-2* and backwards, and two occupational dummies dated in *t-1*. *Specific instruments for column 2*: Stay instrumented with lagged value of Stay, $Stay_{t-1}$, and cross products StayxVariable, are instrumented with the cross product of $Stay_{t-1}$ and the corresponding instrument for the Variable. *Specific instruments for column 1*: Stay instrumented with itself, and cross products StayxVariable, are instrumented with the cross product of Stay and the corresponding instrument for the Variable.

W(Income): exogeneity test for income; *W*(wages): exogeneity test for wages; *W*(Stay): exogeneity test for movements.

Table 3: Variance of changes in hours

	Unadjusted	Adjusted: estimates Table 1	Adjusted: estimates Table 2, column 1	Adjusted: estimates Table 2, column 2	Adjusted: estimates Table C.1, column 2
Stayers	0.075 (0.008)	0.085 (0.008)	0.076 (0.008)	0.076 (0.008)	0.079 (0.008)
Movers	0.438 (0.037)	0.466 (0.038)	—	—	0.540 (0.045)
Not Layoff	0.402 (0.039)	0.414 (0.040)	0.454 (0.042)	0.481 (0.046)	—
Layoff	0.537 (0.096)	0.540 (0.096)	—	—	—
Difference move-stay	0.363 (0.038)	0.381 (0.039)	—	—	0.461 (0.046)
Diff. not layoff-stay	0.327 (0.040)	0.329 (0.041)	0.378 (0.043)	0.405 (0.047)	—
Diff. layoff-stay	0.462 (0.096)	0.455 (0.096)	—	—	—

Notes: Standard errors in brackets. Standard errors are computed without assuming normality: $\text{estvar}(\hat{\sigma}^2) = \frac{1/n \sum w_i^4 - \hat{\sigma}^4}{n}$, where w_i is the variable in deviations from the mean.

Variables	Stayers	Movers	Mov. not layoff	Movers layoff
Hours/year	2212.09	1854.10	1918.03	1598.72
	(410.297)	(720.410)	(711.826)	(699.120)
Hours/week	43.342	42.565	43.154	40.211
	(7.345)	(10.041)	(9.980)	(9.962)
Weeks/year	51.030	43.244	44.170	39.545
	(3.977)	(12.433)	(12.036)	(13.308)
Wage/hour	12.101	9.620	9.798	8.908
	(5.285)	(4.994)	(5.057)	(4.675)
Otinc. x 10^{-4}	1.199	1.178	1.185	1.151
	(2.738)	(2.143)	(2.180)	(1.994)
Δln(hours/year)	0.036	-0.029	0.019	-0.224
	(0.274)	(0.662)	(0.634)	(0.733)
Δln(wage)	0.030	0.054	0.075	0.016
	(0.366)	(0.497)	(0.517)	(0.477)
Δ(otinc 10^{-4})	0.092	0.171	0.145	0.277
	(3.461)	(2.459)	(2.550)	(2.059)
ΔChildren	0.124	0.105	0.113	0.075
	(0.330)	(0.307)	(0.316)	(0.264)
ΔMarital status	0.042	0.046	0.049	0.033
	(0.201)	(0.210)	(0.217)	(0.179)
Uphlth	0.013	0.022	0.025	0.009
	(0.112)	(0.145)	(0.155)	(0.097)
Dwnhlth	0.015	0.012	0.012	0.014
	(0.122)	(0.110)	(0.108)	(0.118)
Observations	3806	1064	851	213

Table A.1: Mean of the relevant variables by individual type

Note: Standard errors in brackets.

Table A.2: Change in hours pooling individuals.
(including all movers)

	Stay exogenous	Stay endogenous
Intercept	-0.079	0.148
	(0.023)	(0.060)
Stay	0.090	-0.171
	(0.024)	(0.069)
Δwage	0.574	0.570
	(0.157)	(0.241)
Δwage x Stay	-0.521	-0.773
	(0.198)	(0.286)
Δotinc	-0.007	0.071
	(0.009)	(0.036)
Δotinc x Stay	0.007	-0.080
	(0.009)	(0.040)
Δmarital status	0.048	-0.339
	(0.060)	(0.249)
Δmarital status x Stay	-0.023	0.449
	(0.062)	(0.314)
Uphlth	0.283	0.383
	(0.173)	(0.344)
Uphlth x Stay	-0.211	-0.336
	(0.184)	(0.431)
Dwhlth	0.009	-0.103
	(0.328)	(0.828)
Dwhlth x Stay	-0.054	0.057
	(0.331)	(0.985)
R^2	0.100	0.166
Sargan test	198.7	84.38
(p-value)	($1.81e^{-5}$)	(0.291)
W(Stay)	288.6	
(p-value)	(0.000)	
Observations	4870	

Notes: Estimation method: GMM. Standard errors in brackets. *Common instruments to both specifications*: constant, Δmarital status, Δchildren, dwhlth, uphlth and Δotinc plus Arellano-Bond type instruments with lagged values of otinc and wages dated *t-2* and backwards, and two occupational dummies dated in *t-1*. *Specific instruments for column 2*: Stay instrumented with lagged value of Stay, $Stay_{t-1}$, and cross products StayxVariable, are instrumented with the cross product of $Stay_{t-1}$ and the corresponding instrument for the Variable. *Specific instruments for column 1*: Stay instrumented with itself, and cross products StayxVariable, are instrumented with the cross product of Stay and the corresponding instrument for the Variable. *W*(Stay): exogeneity test for movements.

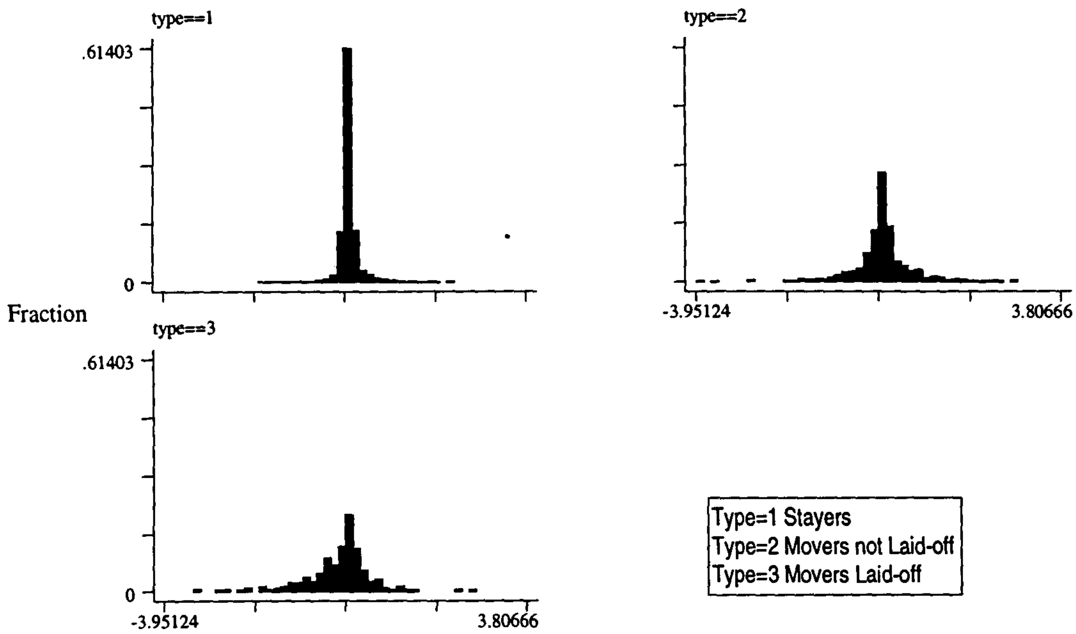

Figure 1: Histogram of change in hours per year by individual type.

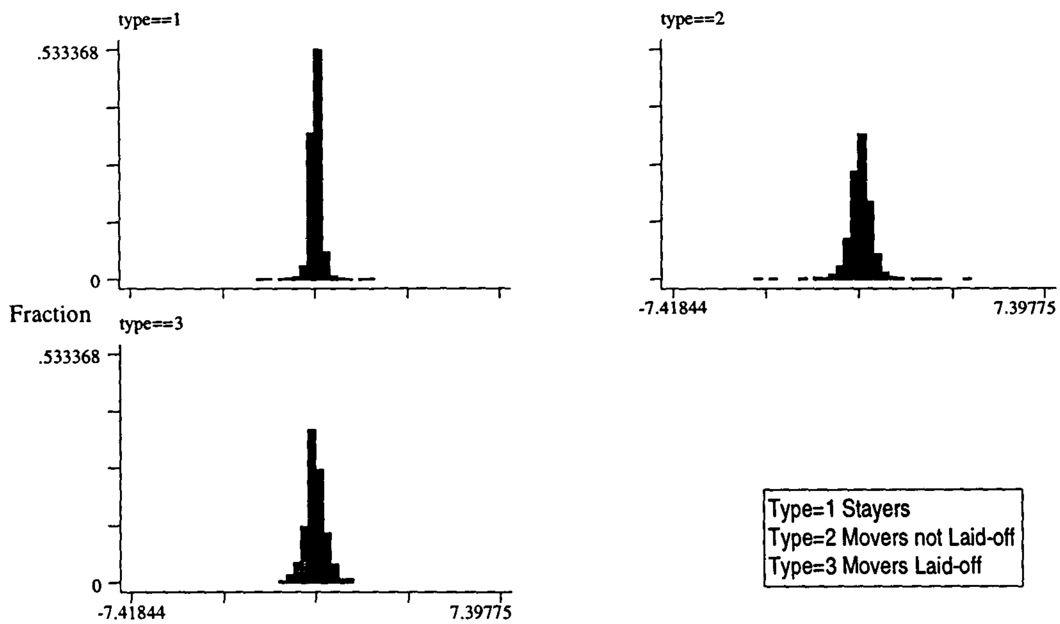

Figure 2: Histogram of change in hourly rate of pay by individual type.

Note: see Section 4 and Appendix for sample sizes and variable description.