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# Estimation of Labour Supply Functions Using Panel Data: A Survey

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# Estimation of Labour Supply Functions Using Panel Data: A Survey

by

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## Abstract

This survey aims at providing the reader with a thread through the literature on the topic of panel econometrics of labour supply, reporting also on the evaluation of the data used in these studies, and summarizing their substantive results. It documents the present trend away from models that take advantage of panel data almost exclusively in order to control for unobserved heterogeneity, towards fully dynamic models where wages become endogenous and consequently the concept of wage elasticity loses much of its appeal.

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# 1 Introduction

The econometrics of labour supply probably belongs to one of the technically most advanced fields in microeconometrics. Many specific issues such as the proper modelling of tax structures, the existence of fixed costs as well as rationing have been treated in numerous articles so that marginal gains in substantive economic insights seem low and entry costs into the field prohibitively high. Surprisingly, one of the most obvious paths for research on labour supply, the (micro-) econometric analysis of the individual's labour supply over the life cycle, has by now gained comparatively little attention. Increasing availability of panel data for many countries as well as the development of appropriate econometric techniques will make econometric studies of intertemporal labour supply behaviour using panel data not only interesting on purely theoretical grounds, they will also help to achieve a better understanding of individual retirement behaviour, the functioning of institutional settings in different countries (such as taxes, vocational training programmes, daycare for children) and the distribution of income and wealth to name only a few.

Estimation of labour supply functions using panel data has been carried out mainly in the eighties, and the number of studies reporting on such estimation is rapidly increasing. Earlier studies using panel data mainly concentrated on participation. Thus it is no surprise that the excellent surveys of Pencavel (1986), Heckman and MaCurdy (1986) and Killingsworth and Heckman (1986) hardly touched the subject. The latter survey concluded a comparison of a large number of cross-section studies with the words: "[these studies] seem to have reduced the mean and substantially increased the variance of [...] what might be called the *reasonable guesstimate* of the wage elasticity of female[1] labour supply [...]. However, [...] studies based on alternative behavioural models - notably, life cycle models, which have been used relatively little in empirical studies - are also likely to provide important insights" (pp. 196-197).

As we shall see, there is a trend away from models that take advantage of panel data almost exclusively in order to control for unobserved heterogeneity, towards fully dynamic models where wages become endogenous and consequently the concept of wage elasticity loses much of its appeal.

This survey aims at providing the reader with a thread through the literature on the topic. However, we make no claim to exhaustivity. Section 2 concentrates mainly on the theoretical aspects of the studies. Since the latter have developed in an evolutionary rather than a revolutionary fashion, that section has a strong chronological character and is largely self-contained. Since in most data at the micro level zero hours supplied can be observed Section 3 gives a brief introduction to panel econometrics for limited dependent variables. Finally, it seems worthwhile to us to supplement the survey with a section reporting on the large bulk of literature recently devoted to the evaluation of the data used in these studies. A section summarizing the substantive results precedes concluding comments.

1 Here we shall not restrict attention to female labour supply.

# 2 Theory

## 2.1 The Basic Model of Life Cycle Labour Supply

We shall not restate here the theoretical developments contained in the survey of Killingsworth and Heckman (1986, pp. 144-179) and refer the reader to them. Killingsworth and Heckman insist on the pioneering work of Mincer (1962). They show that "the distinction between permanent and transitory wages is not particularly useful from a theoretical standpoint" (p. 158) and demonstrate the usefulness of Frisch demands[2] as an alternative to the permanent-transitory distinction. They also discuss models with endogenous wages and conclude: "although much informal discussion implicitly or explicitly emphasizes the interrelationships between [...] work and wages in a life-cycle setting, rigourous analysis of such issues using formal life-cycle labour supply models with endogenous wages is still in its infancy" (p. 178). Here we will describe the models used for estimation in a selection of papers representative of the trend over the last ten years. Along the way we also give some details on the estimation techniques and on the results, illustrating the fact that econometric modelling is by no means linear: there is a feedback of estimation results on model specification.

The seminal paper, as far as empirically implementable models are concerned, is MaCurdy (1981). The assumptions retained are fairly stringent and include known life length $T$, perfect foresight and perfect credit markets, as well as constant interest and time preference rates. At time $t=0$ an individual maximizes

$$\sum_{t=0}^{T} \frac{1}{(1+\rho)^t} U[C(t), L(t)] \tag{1}$$

subject to

$$A(0) + \sum_{t=0}^{T} \frac{1}{(1+r)^t} [W(t)N(t) - C(t)] = 0 \tag{2}$$

where $C$ is consumption, $L$ leisure, $N$ hours of work ($N = \overline{L} - L$, where $\overline{L}$ denotes maximum time available in each period for allocation between leisure and market work), $r$ real interest rate, $W$ real wage,[3] $\rho$ rate of time preference, and $A(0)$ denotes initial assets.

The first order conditions (with $C \geq 0, 0 \leq L \leq \overline{L}$ and only $L \leq \overline{L}$ explicitly taken into account) include the lifetime budget restriction (2) and

$$\frac{\partial U(t)}{\partial C(t)} = \left(\frac{1+\rho}{1+r}\right)^t \lambda, \qquad t = 0, \ldots, T \tag{3}$$

$$\frac{\partial U(t)}{\partial L(t)} \geq \left(\frac{1+\rho}{1+r}\right)^t \lambda W(t), \qquad t = 0, \ldots. T \tag{4}$$

2 The uninformed reader will find a definition below.

3 The consumption aggregate is taken as numeraire in each period.

where $\lambda$ is the Lagrange multiplier of the lifetime budget restriction. The solutions are the Frisch or $\lambda$-constant demands $C[\lambda(t), W(t)], L[\lambda(t), W(t)]$ , with

$$\lambda(t) = \left(\frac{1+\rho}{1+r}\right)^t \lambda \tag{5}$$

and $\lambda$ is implicitly determined by substitution of these demand functions in (2). Thus, $\lambda$ is a function of the entire wage profile $W(t), t = 0, \ldots, T$, of the initial wealth $A(0)$, and of the interest and time preference rates $r$ and $\rho$. It is a sufficient statistic of the past and the future as far as the present decision is concerned.

Concavity implies

$$\frac{\partial C(t)}{\partial \lambda} < 0, \qquad \frac{\partial L(t)}{\partial \lambda} \leq 0, \qquad \frac{\partial^2 L(t)}{\partial \lambda^2} \leq 0,$$

$$\frac{\partial \lambda}{\partial A(0)} < 0, \qquad \frac{\partial \lambda}{\partial W(t)} \leq 0, \qquad \forall t = 0, \ldots, T. \tag{6}$$

In order to obtain an empirical model, MaCurdy specifies the following additively separable utility function for individual $i$:

$$U_i(C, L) = \gamma_{Ci}(t) C^{\beta} - \gamma_{Ni}(t) N^{\alpha}, \quad [N = \overline{L} - L], \qquad i = 1, \ldots, N. \tag{7}$$

Concavity requires $0 < \beta < 1, \quad \alpha > 1$. Heterogeneity, both observed and unobserved, is modelled through random preferences with the specification

$$\ln \gamma_{Ni}(t) = \sigma_i - u_i^*(t), \tag{8}$$

where $u_i^*(t)$ is i.i.d. with zero expectation (note that time-varying characteristics are excluded by assumption).

For an interior solution $[N > 0]$ the resulting Frisch labour supply equation is

$$\ln N_i(t) = F_i + bt + \delta \ln W_i(t) + u_i(t) \tag{9}$$

with

$$F_i = \frac{1}{\alpha - 1}(\ln \lambda_i - \sigma_i - \ln \alpha) \tag{10}$$

and

$$\delta = \frac{1}{\alpha - 1}, \quad b = \delta(\rho - r), \quad u_i(t) = -\delta u_i^*(t), \quad \rho - r \approx \ln \frac{1+\rho}{1+r}.$$

This is a linear panel model with an individual-specific effect $F_i$ which has[4] to be treated as a fixed effect because it is correlated with $W_i(t)$ via $\lambda$.[5]

4 But see the discussion of Jakubson (1988) below.

5 See equation (6) and the implicit determination of $\lambda$.

Moreover, MaCurdy considers the following linear approximation of $F_i$:

$$F_i = \underline{Z}_i'\underline{\phi} + \sum_{t=0}^{T} \gamma(t) \ln W_i(t) + A_i(0)\theta + a_i, \tag{11}$$

where $\underline{Z}_i$ denotes a vector of household characteristics, and coefficients are identical across households. Combined with the additional assumption of a quadratic form for the $\ln W$ profile, and after some algebra which is omitted here, this leads to

$$F_i = \underline{Z}_i'\underline{\phi} + \pi_{0i}\bar{\gamma}_0 + \pi_{1i}\bar{\gamma}_1 + \pi_{2i}\bar{\gamma}_2 + A_i(0)\theta + \eta_i, \tag{12}$$

with $$\bar{\gamma}_j = \sum_{t=0}^{T} \gamma(t)t^j, \qquad j = 0, 1, 2.$$

*Interpretation of Parameters*: $\delta$ is the intertemporal substitution (or $\lambda$-constant or Frisch) elasticity. It describes the reaction to an evolutionary change of the wage rate along the wage profile. It is positive since $\alpha > 1$. *Along* a profile, *evolutionary* changes take place. MaCurdy calls changes *between* profiles *parametric* or profile changes. A change $\Delta$ from a profile I to a profile II at time $s$ causes the labour supply of profile II to be lower than that of profile I in all periods $t \neq s$ because $\lambda_{II} < \lambda_I$. Equation (11) implies

$$F_{II} - F_I = \gamma(s)\Delta < 0.$$

The net effect on labour supply in period $s$, $[\delta + \gamma(s)]\Delta$, can be positive or negative. $\delta + \gamma(s)$ and $\gamma(s)$ are the usual uncompensated (own- and cross-period) elasticities and the corresponding compensated elasticities are $\delta + \gamma(s) - E(s)\theta$ and $\gamma(s) - E(s)\theta$, respectively, where $E(s)$ denotes real earnings in period $s$. If leisure is a normal good [ $\theta < 0$ ], we have

$$\delta > \delta + \gamma(s) - E(s)\theta > \delta + \gamma(s)$$

i.e., $$e_\lambda > e_u > e_A,$$

where $e_\lambda$ is the wage elasticity with constant marginal utility of wealth, $e_A$ is the wage elasticity with constant (lifetime) wealth and $e_u$ is the wage elasticity with constant (lifetime) utility.

Estimation is conducted in two stages.

*Stage 1*: (9) is estimated in first differences:[6]

$$D \ln N_{ij} = b_j + \delta D \ln W_{ij} + \varepsilon_{ij}, \qquad j = 2, \ldots, \tau, i = 1, \ldots, N.$$

6 $D$ denotes the first difference operator. Another possibility would be to use within estimation. One advantage of estimation in first differences, however, is that no strict exogeneity assumption is needed.

$\tau$ denotes the number of waves in the available (balanced) panel, and $b_j = \delta(\rho - r)_j$ is a period effect.[7] No restriction is imposed on the covariance structure of $\varepsilon$ and system estimation (2SLS and 3SLS) is used; $\ln W_i$ is treated as endogenous and instrumented, using a human capital type equation.

In this way the reactions of $N(t)$ to the *evolutionary* changes in $W(t)$ are completely described. In order to also describe the reactions of labour supply to *parametric* changes in wages, more information is needed.

*Stage 2:* Given the first stage parameter estimates, the fixed effects can be estimated using (1) - (9) as

$$\hat{F}_i = \frac{1}{\tau} \sum_{j=1}^{\tau} [\ln N_i(t(j)) - \hat{b}t(j) - \hat{\delta} \ln W_i(t(j))], \qquad (13)$$

where $t(j)$ is age in period $j$. Similar equations are constructed also for variables having means equal to the $\pi_{hi}$, and they are estimated in a system jointly with (12).

## 2.2 Tests and Relaxation of the Assumptions of the Basic Model

*Uncertainty:* We now assume uncertainty concerning wages and interest rates. Replanning for the future takes place in every period, on the basis of the new information obtained. The individual maximizes the expected discounted utility in period $t$:

$$\mathrm{E}_t \sum_{k=t}^{T} \frac{1}{(1+\rho)^{k-t}} U(k) = U(t) + \frac{1}{1+\rho} \mathrm{E}_t \sum_{k=t+1}^{T} \frac{1}{(1+\rho)^{k-t-1}} U(k) \qquad (14)$$

subject to the budget restriction

$$A(t) = (1 + r(t))A(t-1) + W(t)N(t) - p(t)C(t), \qquad (15)$$

where $A(t)$ are the assets at the end of period $t$, $p(t)$ is the price of the consumption aggregate in period $t$, and $W(t)$ and $r(t)$ now denote the *nominal* wage and interest rate.[8] Using the Bellman principle, we define

$$V(t+1) = \max \mathrm{E}_{t+1} \left\{ \sum_{k=t+1}^{T} \frac{1}{(1+\rho)^{k-t-1}} U(k) \right\},$$

with maximisation subject to the constraint (15) written at $t+1$. This is a function of $A(t)$ alone and at period $t$ the person maximizes:

7 For clarity we follow MaCurdy in distinguishing the wave $j$ of the available panel from the variable $t$ which is related to the age of the individual $i$. Note that $b_j$ could pick up other effects than simply variation in the interest rate.

8 Some end-period constraint must be introduced, like for instance the assumption of no bequest $A(T) = 0$, but the precise form of the constraint does not modify the form of the solutions.

$$V(t) = \max_{C(t),N(t)} U(t) + \frac{1}{1+\rho} \mathrm{E}_t V(t+1) \tag{16}$$

under restriction (15). If we exclude corner solutions, the first order conditions are:

$$\frac{\partial U(t)}{\partial C(t)} = \lambda(t) p(t), \tag{17}$$

$$\frac{\partial U(t)}{\partial N(t)} = -\lambda(t) W(t), \tag{18}$$

$$\lambda(t) = \mathrm{E}_t \left\{ \frac{1 + r(t+1)}{1+\rho} \lambda(t+1) \right\}. \tag{19}$$

The last equation implies that the individual decides on savings in such a way that the discounted expected utility of money remains constant (Euler equation). If we assume that there is no uncertainty about $r(t+1)$ we have

$$\mathrm{E}_t[\lambda(t+1)] = \frac{1+\rho}{1+r(t+1)} \lambda(t)$$

or

$$\lambda(t+1) = \left[ \frac{1+\rho}{1+r(t+1)} \lambda(t) \right] (1 + e(t+1)),$$

which simply defines $e(t+1)$ with $\mathrm{E}_t[e(t+1)] = 0$ and leads to the *approximation*

$$\ln \lambda(t+1) \approx \ln \lambda(t) + \rho - r(t+1) + e(t+1). \tag{20}$$

Therefore, the "fixed effects" technique remains feasible in the presence of uncertainty about the wage profile. However, the orthogonality between $e(t+1)$ and the information available at time $t$ suggests application of the Generalized Method of Moments (GMM). Exposition here has been kept fairly sketchy. See Altug and Miller (1990) for a more elaborate treatment spelling out the implications of assuming a competitive environment with complete markets.

*Within-period additive separability:* The importance of relaxing the assumption of separability between leisure and goods is indicated in Browning and Meghir (1989) who reject this assumption, testing it within a very general scheme using 1979-1984 FES data (time series of cross-sections): preferences about goods are specified in a flexible way, with conditional cost functions where no behavioural assumption concerning labour supply or participation decision is needed.[9] Here we shall be concerned only with relaxing the assumption of additive separability between the two "goods" leisure and aggregate consumption.

9 Yet their model is not cast in the life cycle framework and the implications of their study for life cycle models should be elucidated.

In section 3 we shall see that weakening this assumption is actually not as easy as it appears at first sight when working with Frisch demands on panel data. Browning et al. (1985), however, estimate the following specification in first differences

$$N_i(t) = \alpha_1(\underline{a}_i(t)) + \beta_1 \ln \tilde{W}_i(t) + \theta_1 \sqrt{\frac{p(t)}{W_i(t)}} + \beta_1 \ln \tilde{\lambda}_i(t), \tag{21}$$

$$C_i(t) = \alpha_2(\underline{a}_i(t)) + \beta_2 \ln \tilde{p}_i(t) - \theta_2 \sqrt{\frac{W_i(t)}{p(t)}} + \beta_2 \ln \tilde{\lambda}_i(t), \tag{22}$$

where "~" indicates discounting. Symmetry of the demand functions implies that $\theta_1 = \theta_2 = \theta$ and within-period additive separability is equivalent to $\theta = 0$. $\underline{a}_i(t)$ is a vector of household characteristics. Browning et al. estimate the equations separately, i.e. they do not enforce the identity $\theta_1 = \theta_2$, as would be feasible in this context since there is no adding-up restriction (in contrast with a Marshallian demand system). However, they find $\theta_1$ and $\theta_2$ to be significantly different from zero and to have opposite signs, which makes the entire specification appear questionable. Note that, although Browning et al. consider aggregate consumption, no problem arises from working with several consumption goods. Yet, durables should be given special attention, as they might be more properly treated as assets.

So far we have focussed on the preferences of an individual. In practice, however, one often prefers to work with household preferences. One of the many reasons for doing this is the impossibility of isolating individual from household consumption in survey data. Then, another assumption which is necessary for the validity of the specifications that we have considered so far is the separability of the labour supplies of the different potential earners in a household. If it holds, the earnings of the other household members can be accounted for in $A(t)$, because then the influence of hours and wages of other household members boils down to a pure income effect. Otherwise the model is mis-specified.

A problem that arises when one considers members of a household other than the head (that asymmetry is still empirically relevant) is the participation decision. However, still keeping to the situation where only the labour supply of the household head is considered, we first turn to the empirically no less relevant problem of unemployment, because it relates well to the former developments.

*Unemployment:* Certainly one of the most questionable assumptions made so far is the assumption that unemployment is voluntary. Ham (1986) produces empirical evidence against that hypothesis in the context of life cycle models (see also Ashenfelter and Ham, 1979). Ham uses the following modification of MaCurdy's model. If an additional restriction consisting of a ceiling to the number of hours worked exists, and if $T_u$ is the set of indices of the periods where this restriction holds for individual $i$ we have

$$\ln N_i(t) < F_i^* + b\,t + \delta \ln W_i(t) + u_i(t), \qquad \text{for } t \in T_u, \tag{23}$$

$$\ln N_i(t) = F_i^* + b\,t + \delta \ln W_i(t) + u_i(t), \qquad \text{for } t \notin T_u, \tag{24}$$

where $F_i^*$ corresponds to a higher value of $\lambda$ than when $T_u = \varnothing$ : the profile of expected wages at each period is lower than in the absence of unemployment periods. Therefore, (9) will yield large residuals for $t \in T_u$ if unemployment is not the outcome of a free choice. The idea is then to estimate either

$$\ln N_i(t) = F_i^* + b\,t + \delta \ln W_i(t) + \theta_1 U_i(t) + u_i(t) \tag{25}$$

or

$$\ln N_i(t) = F_i^* + b\,t + \delta \ln W_i(t) + \theta_2 H_i^u(t) + u_i(t), \tag{26}$$

where $U_i(t) = 1$ if $t \in T_u$ and 0 otherwise, and $H_i^u(t)$ denotes yearly hours of unemployment. If the assumption is correct, then $\theta_1$ (or $\theta_2$ ) will not significantly differ from zero. Otherwise one would expect negative values.

The assumption is clearly rejected for both specifications (25) and (26), as well as for other specifications allowing for uncertainty, non-linearity ( with the additional term $[\ln W_i(t)]^2$ ), non-separability (specification (21)), as well as for various assumptions on the covariance structure of the residuals. The results of these tests suggest modelling these restrictions explicitly. Lilja (1986) makes several proposals in this direction.

However, MaCurdy (1990) criticizes Ham's argument and shows that $\theta_1$ (or $\theta_2$) significant in (25) (or(26)) is compatible with voluntary unemployment caused by a lower wage offer $W_i(t)$ for $t \in T_u$ : "The reasoning underlying the testing of exclusion restrictions in labor supply functions relies on the argument that wages fully capture the influences of demand-side factors in the supply decision. This reasoning is sound but the variable identified as relevant by intertemporal substitution theory is the offer wage; and the offer wage deviates from the observed market wage if unemployment occurs at all" (MaCurdy, 1990, p.228; see also Card, 1990, who interprets Ham's findings in favour of demand-side conditions as the main determinant of observed hours).

*Accounting for the participation decision*: The prototype here is the paper by Heckman and MaCurdy (1980) which also presents the first estimation of a Tobit model on panel data. The specification does not differ much from that of MaCurdy (1981) but now the individual considered is a married woman. Separability between the leisures of husband and wife is assumed, and the specification chosen for the utility function is

$$U(t) = \gamma_C(t) C^{\beta}(t) + \gamma_L(t) L^{\alpha}(t), \tag{27}$$

with $\alpha < 1, \beta < 1$ (we have dropped the index of the individual for simplicity). The stochastic assumptions adopted are

$$\ln \gamma_L(t) = \underline{Z}(t)'\underline{\phi} + \varepsilon_1(t), \tag{28}$$

$$\ln W(t) = \underline{X}(t)'\underline{\psi} + \varepsilon_2(t), \tag{29}$$

$$\varepsilon_1(t) = \eta_1 + u_1(t), \qquad \mathrm{E}u_i(t)u_j(s) = \delta_{ts}\sigma_{ij}, i,j = 1,2,$$

$$\varepsilon_2(t) = \eta_2 + u_2(t), \qquad \mathrm{E}u_i(t) = 0,$$

where $\eta_1$ and $\eta_2$ are individual fixed effects capturing unobserved heterogeneity in the specifications of $\ln \gamma_L$ and $\ln W$.[10] (But the claim that absence of correlation over time in the $u$ s is not a strong assumption because of the free correlation between $\eta_1$ and $\eta_2$ is questionable in two ways: (i) the $\eta$ s are time independent, (ii) they are viewed as being deterministic). Identification requires exclusion restrictions between $\underline{X}$ and $\underline{Z}$. Maximization of (1) subject to (2) with this specification yields

$$\ln L(t) = f + \frac{\rho - r}{\alpha - 1}t - \underline{Z}(t)'\frac{\underline{\phi}}{\alpha - 1} + \underline{X}(t)'\frac{\underline{\psi}}{\alpha - 1} + v(t) \text{ if } L(t) \le \overline{L}, \tag{30}$$

$$= \ln \overline{L} \qquad \text{otherwise},$$

where

$$f = \frac{1}{\alpha - 1}(\ln \lambda - \ln \alpha - \eta_1 + \eta_2),$$

and

$$v(t) = \frac{1}{\alpha - 1}[-u_1(t) + u_2(t)].$$

Equations (29) and (30) are simultaneously estimated by ML, assuming normality for $(u_1(t), u_2(t))$. The fixed effects are $f$ in the hours equation and $\eta_2$ in the wage equation. The estimation can only be performed for women who worked at least once in the observed periods. Correction for the corresponding selection bias is found to have only a minor impact. Since asymptotic arguments are not justified in the time dimension (only eight waves), estimates of the fixed effects are not consistent[11] and this inconsistency leads in principle to inconsistency of all coefficients. However, (i) Heckman (1981a) performed Monte Carlo simulations for fixed effects Probit with eight waves and found that the fixed effects Probit performed well when the explanatory variables were all strictly exogenous, (ii) Tobit should perform even better because it is a combination of Probit and linear regression. The fixed effects (incidental parameters) are estimated simultaneously with the parameters of interest through alternated iteration on both subsets of parameters. Yet their economic interpretation is difficult because the influence of $f$ is mixed with that of the time invariant variables in $\underline{Z}(t)$ and the

10 $\delta_{ts}$ in (29) is the Kronecker symbol.

11 That is, for $N \to \infty$.

same holds for $\eta_2$ and the time invariant variables in $\underline{X}(t)$. Regressions of the fixed effects on those time invariant variables complete the picture and allow to reach conclusions like the following: current-period household income (exclusive of the wife's earnings) has no significant impact on labour supply, in contrast to the impact of an eight year average income (proxy for the permanent income).

Another study taking the participation decision into account is Jakubson (1988). The specification is the same as above but identification of $\underline{\psi}$ and $\underline{\phi}$ is left aside and Jakubson specifies $\underline{X}(t) \equiv \underline{Z}(t)$. The model is thus considerably simplified and takes the general multivariate Tobit form

$$y_i^*(t) = \underline{x}_i(t)'\underline{\theta} + c_i + u_i(t), \tag{31}$$

$$y_i(t) = y_i^*(t) \qquad \text{if} \quad y_i^*(t) > 0,$$

$$= 0 \qquad \text{otherwise},$$

$$\underline{u}_i \sim \mathrm{N}(0, \Sigma).$$

Jakubson presents three approaches to the estimation of (31): simple pooling, treatment of $c_i$ as a random effect taking into account the correlation with $\underline{x}_i$ (using the approach of Chamberlain, 1984) and, as before, treatment of $c_i$ as a fixed effect. For the fixed effects, the considerations above still hold, while convergence for the random effects specification is ensured even for short panels as long as their stochastic specification is correct. For details, see Section 4.

The main conclusions are: (i) the panel estimates (fixed or random effects) of the influence of children on labour supply are only about 60% of the cross-section estimates, due to the neglect of individual effects in the latter; (ii) as concerns the life cycle hypothesis, like in the Heckman MaCurdy study, current income does not have a significant influence in the fixed effects estimation, yet this does not hold true for random effects.

Disregarding the inconsistency problem associated with fixed effects here, and considering that sampling may be endogenous (one of the selection criteria being "stable marriage", see Lundberg, 1988) the fixed effects approach might seem preferable on a priori grounds. However, as we shall see in the following section, the entire specification is questionable.

### 2.3 Alternative Parameterization and Implications

Blundell et al. (1990) show that the specification of $\lambda$-constant systems where $\lambda$, or $\ln\lambda$, appears additively and therefore can be treated as an individual-specific effect, turns out to be extremely restrictive. To see this, let us write an intertemporally additive utility function as

$$\sum_{t=0}^{T}\frac{1}{(1+\rho)^t}F[U^*(t),\underline{Z}(t)], \tag{32}$$

where $F$ increases with its first argument, $U^*$ is a representation of the within-period preferences, and $\underline{Z}$ is a vector of characteristics. Thus, three elements are necessary for a complete characterization of the intertemporal preferences: $\rho$, $F$ and $U^*$. We now consider the models of MaCurdy (1981) and Browning et al. (1985). They share the form

$$g(x_i(t)) = f\{\underline{p}(t);\underline{\theta}\} + \phi \ln \lambda(t), \tag{33}$$

where $g$ and $f$ are some functions, $x_i(t)$ denotes demand for good or leisure $i$, $\underline{p}(t)$ is the price vector at $t$, and $\underline{\theta}$ and $\phi$ are parameters. Blundell et al. show that for $g() = \ln()$ and $f$ linear the within-period utility $U_t = F(U_t^*, \underline{Z}_t)$ must be either homothetic (which is totally unattractive) or explicitly additive over all leisures and goods. Therefore $F = id$ and $U^*(t) = \sum_{i=1}^{n} U_i^*(t)$. The devastating consequence is that such intertemporal preferences are completely identified (up to $\rho$) on a single cross-section, given that some variation in the wages or prices can be observed. Thus, this type of specification hardly qualifies for exploiting panel data.

Blundell et al. show that the indirect utility function corresponding to $U_t = F(U_t^*, \underline{Z}_t)$ for the specification of Browning et al. takes the form

$$V(y(t),\underline{p}(t)) = \alpha - \phi \exp\left\{-\frac{y(t)-a(\underline{p}(t))}{\phi\mu(\underline{p}(t))}\right\}, \tag{34}$$

where $y(t)$ is the total expenditure in period $t$, $a$ and $\mu$ are two price indices and $\phi$ is the parameter appearing in (33). As a consequence, the intertemporal elasticity of substitution $\Phi \equiv V_y / y V_{yy}$ is given by $\Phi \equiv -\phi\mu/y$ and therefore, since $\phi\mu < 0$, $\Phi$ decreases for wealthier households, which goes against the intuition that the possibilities for substitution should increase with wealth.

Summing up, it turns out that the requirement that $\lambda$ or a function of $\lambda$ should appear linearly in (33) imposes very strong a priori restrictions on preferences.

An alternative strategy consists in estimating the within-period preferences $U^*$ by eliminating $\lambda$, either directly between two goods or indirectly via the period budget equation, and then estimating the monotonous transformation $F$ and the time preference rate $\rho$ separately. The advantage is that no restriction on within-period preferences is implied. Panel data are not absolutely necessary for this strategy: a time series of independent cross-sections proves to be sufficient and has even some advantages in providing valid instrumental variables more easily (see Blundell and Meghir, 1990). Blundell et al. (1989) give a good example of the application of this strategy to demands for goods. Two important panel studies use this alternative.

MaCurdy (1983) proposes to directly estimate the marginal rate of substitution functions. The first-order conditions (17) and (18) give

$$\frac{\partial U(t)/\partial N(t)}{\partial U(t)/\partial C(t)} \equiv \frac{\partial U^*(t)/\partial N(t)}{\partial U^*(t)/\partial C(t)} = -\frac{W(t)}{P(t)}. \tag{35}$$

The advantage over estimating Marshallian demands is that this allows estimation of preferences that do not imply a closed-form expression for the demand functions. The estimation of (35) does not require a panel. A cross-section with enough price variation, or indeed a time series of cross-sections, is sufficient. In spite of this, MaCurdy chooses the restrictive form

$$F_i[U_i^*(t)] = \xi_i(t)\frac{[U_i^*(t)+\nu]^{\sigma}-1}{\sigma}, \tag{36}$$

$$U_i^*(t) = \gamma_i(t)\frac{[C_i(t)+\theta_C]^{\alpha_C}}{\alpha_C} - \frac{[N_i(t)+\theta_N]^{\alpha_N}}{\alpha_N}, \tag{37}$$

with

$$\xi_i(t) = \exp\{\underline{X}_i(t)'\underline{\phi} + a_i(t)\}, \tag{38}$$

$$\gamma_i(t) = \exp\{\underline{X}_i(t)'\underline{\psi} + \varepsilon_i(t)\}. \tag{39}$$

The parameters $\underline{\phi}, \underline{\psi}, \sigma, \nu, \theta_C, \theta_N,$ and $\alpha_N$ are constant across individuals and over time. This utility function is still additive, yet no longer explicitly additive, and this form of $U^*(t)$ admits several well-known special cases such as CES, addilog and Stone Geary. (Surprisingly enough, MaCurdy is not at all interested in the intertemporal elasticity of substitution in that study). There is no identification problem here since (38) and (39) are estimated in two different dimensions: (39) is estimated in the 'individual' dimension and (38) in the 'time' dimension. Equations (35) and (37) yield

$$\ln\frac{W(t)}{P(t)} = -\underline{X}_i(t)'\underline{\psi} +$$

$$(\alpha_N - 1)\ln[N_i(t)+\theta_N] - (\alpha_C - 1)\ln[C_i(t)+\theta_C] - \varepsilon_i(t), \tag{40}$$

which gives consistent estimators (on a single cross-section if desired) for $\underline{\psi}, \alpha_N, \alpha_C, \theta_N$ and $\theta_C$. Using those one can obtain $\gamma_i(t)$ by substitution of $\underline{X}_i(t)'\underline{\psi} - \varepsilon_i(t)$ from (40) into (39). Substitution of (17) into (20) gives

$$\ln\left[F_i(t+1)'\frac{\partial U^*}{\partial C}(t+1)\right] = b(t+1) + \ln\left[F_i(t)'\frac{\partial U^*}{\partial C}(t)\right] + e(t+1), \tag{41}$$

where $b(t+1) = \rho - r(t+1) + \ln[P(t+1)] - \ln[P(t)]$, i.e. the difference between the rate of time preference and the real interest rate at time $t+1$. The above specification leads to

$$\ln \frac{\partial U_i^*}{\partial C}(t+1) - \ln \frac{\partial U_i^*}{\partial C}(t) = b(t+1) - [\underline{X}_i(t+1) - \underline{X}_i(t)]'\underline{\phi} + \tag{42}$$

$$(1-\sigma)[\ln(U_i^*(t+1)+\nu) - \ln(U_i^*(t)+\delta)] + a_i(t) - a_i(t+1) + \eta(t+1).$$

Either time series or panel data contain all the information needed to estimate (42). Instrumental variables are necessary to take account of the endogeneity of $U_i^*(t)$ and $U_i^*(t+1)$, and Pagan's (1984) method of correcting the variance of the estimators is advisable here because estimated parameters are used in the construction of regressors as well as regressands in (42). Taking account of measurement errors in hours, wages or consumption would be difficult here because such errors would contaminate $\varepsilon_i(t)$ (see (40)) and would therefore produce non-linear errors in the variables in (42).

*Errors in variables* are thoroughly treated by Altonji (1986) using instrumental variables methods. Unfortunately, in order to obtain the required linearity Altonji uses a version of MaCurdy's (1981) restrictive form, i.e. an explicitly additive within-period utility function

$$U_t = \sum_{k=t}^{T} \frac{1}{(1+\rho)^k} \left[ \frac{\gamma_{Ck}}{\alpha_C} C_k^{\alpha_C} - \frac{\gamma_{Nk}}{\alpha_N} N_k^{\alpha_N} \right], \tag{43}$$

where $\gamma_{Ck}$ and $\gamma_{Nk}$ are time-varying taste modifiers. The $\lambda$-constant demands are[12]

$$\ln N_t = cst + \delta_N[\ln W_t + \ln \lambda_t + t\ln(1+\rho) - \ln \gamma_{Nt}], \tag{44}$$

$$\ln C_t = cst + \delta_C[\ln \lambda_t + t\ln(1+\rho) - \ln \gamma_{Ct}]. \tag{45}$$

Rather than estimating (44) in first differences,[13] Altonji proposes substituting $\ln \lambda_t + t\ln(1+\rho)$ out of (44) and (45). We now assume that the observations contain the measurement errors $v_{Nt}^*$, $v_{Ct}^*$ and $e_t^*$, and consist in $n_t^* = \ln N_t + v_{Nt}^*$, $c_t^* = \ln C_t + v_{Ct}^*$ and $w_t^* = \ln W_t + e_t^*$. Since $W_t$ is not directly observed but is calculated by dividing period income by $N_t$, $v_{Nt}^*$ is correlated with $e_t^*$ but neither of the two will be correlated with $v_{Ct}^*$. Thus, we obtain the model

$$n_t^* = cst + \delta_N w_t^* + \frac{\delta_N}{\delta_C} c_t^* + \delta_N \ln \frac{\gamma_{Ct}}{\gamma_{Nt}} + v_{Nt}^* - \delta_N e_t^* - \frac{\delta_N}{\delta_C} v_{Ct}^*. \tag{46}$$

12 From now on we switch from our previous convention to letting $t$ appear as a subscript, in order to alleviate notation.

13 Yet this is done for comparison.

The advantage over first differences is that the substitution using $c_t^*$ does not bring lagged wages into the equation. Even more important perhaps, the assumption about expectations that was used above to motivate estimating first differences under uncertainty is no longer necessary. Instruments are used for $w_t^*$ and $c_t^*$. The results do not differ much from MaCurdy's.

Blundell, Browning and Meghir (1989) give a good example for a less restrictive use of the alternative estimation methods mentioned above but it is limited to the demand for goods. Yet, we do not know of a study that estimates a complete system of life cycle labour supply and goods demands using panel data. Blundell and Walker estimate such a system with a cross-section and the calculation of substitution elasticities they present is based on arbitrary assumptions. The reason for this strategy is the missing of data material that would allow identification of the rate of time preference and of the monotone transformation (see equation (32)). A remedy for this shortcoming and therefore a possible solution of the problem is the combined use of various data sources: see Arellano and Meghir (1989) for a possible approach.

## 2.4 Relaxing the assumption of intertemporal separability

Although relaxing this assumption is no easy task, it is important because all the studies that test the assumption clearly reject it. If the estimation results are to be used in policy analysis, the specification must produce interpretable parameters and not merely a separability test. In this respect, it seems difficult to simultaneously model the multiple reasons that lead to the rejection of separability. Most empirical studies therefore concentrate on only one of the aspects. The modelling of partial adjustment or rational habit formation in an optimization scheme over the life cycle is such a practicable extension.

Yet before turning to structural models relaxing the intertemporal separability assumption, it is interesting to discuss the results of a VAR approach to modelling the relationship between wages and hours of work using panel data. As a prototype for this kind of approach we will focus on the study by Holtz-Eakin et al. (1988) but also refer the reader to Abowd and Card (1989).

Holtz-Eakin et al. analyse a sample of 898 males from the Panel Study of Income Dynamics (PSID) over 16 years. They estimate linear equations for wages and hours with lags of equal lengths on both wages and hours on the right hand side of each equation, and individual effects. Note that the equation on hours does not nest the simple life-cycle model of MaCurdy (1981) since the contemporaneous wage is excluded and no serial correlation is allowed. By contrast, the form of the wage equation could be justified by human capital considerations. However. attempts at interpreting these reduced form equations are not in line with the VAR approach. The model of Holtz-Eakin et al. does not a priori impose the stationarity of the coefficients over time, not even for the individual effect. The estimation strategy relies on GMM, combined with quasi-differencing along the lines of Chamberlain (1984, p. 1263) in order to eliminate the individual effect while allowing for non-stationarity. Errors in variables are easily dealt with within this linear GMM framework, but again under the restrictive assumption that they present no serial correlation. Starting with a maximum lag length of three periods (involving four lags of the original variables in the quasi-differenced equations) parameter stability is rejected for none of the two equations, and the analysis proceeds more simply with first differences. The next step concerns testing for the lag-length, and the assumption that one lag is sufficient to describe the data is rejected in no equation at the 1% level but rejected in the hours equation at the 5% level.

Furthermore, one cannot reject the assumption that lagged hours could be excluded from the wage equation. The same holds for lagged wages in the hours equation when using only one lag but not if two lags are retained (an argument in favour of nesting the non-causality test within the hypothesis about the lag length is that in this way the test statistics turn out to be asymptotically independent, which facilitates pin-pointing the reasons for rejection of the joint hypothesis). Tests for measurement error bias are constructed using internal instruments in the simple first-order autoregressive models, in order to increase the power of the test. The assumption of absence of measurement error cannot be rejected at the 5% level but there is evidence that the test may have low power in this instance. Most results are qualitatively, and, what is more surprising, quantitatively replicated on a sample from the National Longitudinal Survey (NLS). The authors conclude (p. 1393):"Our empirical results are consistent with the absence of lagged hours in the wage forecasting equation, and thus with the absence of certain human capital or dynamic incentive effects. Our results also show that lagged hours is important in the hours equation, which is consistent with the alternatives to the simple labor supply model that allow for costly hours adjustment or preferences that are not time separable. As usual, of course, these results might be due to serial correlation in the error term or a functional form misspecification". The problem of possible serial correlation in the error term is of no minor importance and in the sequel we stress the way in which it is dealt with.

Bover (1991) estimates a rational habits model in a certainty framework with a minimum amount of replanning. The salient feature of her approach is that the model specification is constructed in such a way that it allows for an explicit expression of the marginal utility of wealth $\lambda$, as a function of future wages, initial wealth, the (constant) interest rate, and preference parameters. The advantage of such an expression is that it allows a direct analysis of wealth effects on intertemporal labour supply (see Card, 1990, for the potential importance of such effects), whereas the approach of MaCurdy (1981) allows such an analysis only in a very indirect and unsatisfactory way. However, this comes at a large cost, as we shall see. In period $t$ the individual maximizes

$$\sum_{k=t}^{T}(1+\rho)^{t-k}\left[\beta_k \ln(\gamma_N+\phi N_{k-1}-N_k)+(1-\beta_k)\ln(C_k-\gamma_c)\right] \tag{47}$$

s.t.

$$\sum_{k=t}^{T}(1+r)^{t-k}(W_k N_k - C_k) = -A_t, \tag{48}$$

where the variables have the same interpretation as in equations (1) and (2), and $\phi$ measures habit persistence. The Stone-Geary specification (47) was also used by Ashenfelter and Ham (1979) in order to derive an explicit expression for $\lambda$ under perfect foresight. The novel feature here lies in the relaxation of the intertemporal separability assumption through the rational habit formation assumption (in a former paper Bover (1986) considered two alternative models, one with partial adjustment and one with myopic habit formation which did not take account of all direct and indirect influences of current labour supply on future decisions, as the rational habit formation model does, but she found all these models to be empirically indistinguishable).

Defining $N_k^* = N_k - \phi N_{k-1}$ and $W_k^* = \sum_{j=0}^{T-k}(1+r)^{-j}\phi^j W_{k+j}$ allows to rewrite (47)-(48) in the usual form of a separable intertemporal utility function with arguments $\{N_k^*, C_k\}_{k=1,...,T}$ and an additively separable intertemporal budget constraint. The corresponding Frisch demands are linear in $\lambda_t$ and the expression of the latter is obtained by substituting these into the budget constraint. The reason for the subscript $t$ in $\lambda_t$ is the replanning that takes place at each period, when the individual forms new predictions about his wage profile. The somewhat arbitrary assumption here is that each individual's future wages lie on a specific time trend, and that the individual learns more about the two coefficients of this relationship as more time passes by. This is disturbing, because if the relationship were deterministic two observations would suffice to pin it down without error, and if not we have uncertainty about future wages, whereas the derivation of $\lambda$ assumed the $W_k^*$ to be known.

This specification yields a static nonlinear model which can be exactly linearized through transformations of the exogenous variables on the one hand and of the parameters on the other hand. The error specification is of the ECM type with the unobserved heterogeneity subsumed in a time-invariant individual effect. Bover estimates the dummy variable model with unrestricted covariance for the residual error term including also time dummies and using instruments to cope with potential endogeneity and measurement error problems concerning the wage variable. These instruments should be strictly exogenous conditional on the individual effect and the instruments used seem indeed to have this property. A $\chi^2$-test of the overidentifying restrictions leads to no clear-cut rejection of of the specification. The results show a significant effect of the lagged hours on the current decision.

The approaches of Hotz, Kydland and Sedlacek (1988) (HKS) and Shaw (1989) (S) are based on similar specifications and estimation methods, and can therefore be described together. While Bover substitutes the marginal utility of money in the Euler equation with a very special assumption about the wage path, here the agents have rational expectations concerning the uncertain wage profile, and the resulting stochastic Euler equations are directly estimated with GMM. The basic difference between the two approaches lies in the kind of non-separability which is allowed for. HKS assume rational habit formation and therefore account for intertemporal non-separability of preferences, like Bover does. Analogous to her they assume that the wage path is not influenced by the hours decision, thus assuming intertemporal separability in the budget constraint. By contrast Shaw actually relaxes the latter assumption, i.e. she allows for non-separability in the budget constraint and not in the preferences. In period $t$ the individual maximizes

$$E_t \sum_{k=t}^{T} \frac{1}{(1+\rho)^{k-t}} U(Z_k, C_k) \tag{49}$$

where $$Z_k = L_k + \alpha a_k$$

with $$a_k = (1-\eta) a_{k-1} + L_{k-1} \qquad \text{for HKS}$$

but $$Z_k = L_k \qquad \text{for S},$$

and $L$ denotes leisure. The HKS specification nests intertemporal separability $(\alpha = 0)$ and the models of Johnson and Pencavel (1984) and Bover (1986, 1991), where only the labour supply of the previous period does play a role in the preferences of the current period $(\eta = 1)$.

The budget restriction is

$$A_{t+1} = (1+r_t)(A_t + W_t N_t - C_t) \tag{50}$$

in self-explaining notations, but Shaw defines $W_t$ as the product $R_t K_t$ of the human capital stock $K_t$ and its rental rate $R_t$ and choses a quadratic approximation $f$ for the relationship between $K_{t+1}$ on the one side and $K_t$ and $N_t$ on the other side, which yields the atypical earnings function

$$\frac{W_{t+1}}{R_{t+1}} = f\left(N_t, \frac{W_t}{R_t}\right). \tag{51}$$

This equation is separately estimated using IV, and the validity of that procedure may be questioned. The specification for $U$ is the Translog in both approaches. The estimation of the preference parameters is by GMM using the orthogonality conditions in the stochastic Euler equations. In order to avoid misspecification due to the potential endogeneity of wages, HKS only use the Euler equation for consumption. Since parameters $\alpha$ and $\eta$ are identified under the maintained assumption of no contemporaneous additive separability between $Z_t$ and $C_t$, this allows testing the form of the intertemporal non-separability in preferences. Moreover, a score test of the wage exogeneity is offered. HKS also explain how to cope with a certain degree of correlation between individuals through macroeconomic shocks or regional variables. Both approaches are estimated over small samples of men from the PSID (482 for HKS and 526 for S). Neither of the two studies handles measurement errors or unobserved heterogeneity, due to the high degree of non-linearity in the Euler equations. The last point, in particular, is problematic since the presence of unobserved heterogeneity can bias the conclusions about state dependence in dynamic models (e.g. Chamberlain, 1984). The theoretical setting (Euler equation) implies orthogonality between the residual at time t and all the information available up to $t-1$. Thus, in GMM estimation, all variables dated $t-1$ or earlier qualify in principle as instruments for the equation dated $t$. This implication of theory can be tested by a $\chi^2$-test of overidentifying restrictions using two sets of instruments, one being restricted to strictly exogenous instruments. HKS conduct such a test and do not reject the null of orthogonality.

HKS separately estimate the parameters for two age groups and reach the following conclusions. The estimated parameters $\alpha$ and $(1-\eta)$ are positive and well-determined and therefore intertemporal separability is rejected, and not only $L_{t-1}$ but also leisure decisions in previous years have a direct influence on current decisions. The separability between $Z_t$ and $C_t$ in the Translog utility function is also rejected, as is exogeneity of the wages. A slightly disturbing result is the negativity of the estimated rate of time preference. At first, Shaw finds the same result, yet the introduction of sufficient observable heterogeneity in the other preference parameters yields a not unreasonable value of 4.2%. Her other conclusions are as follows. The rental rate of human capital varies considerably over time and the number of hours worked has a strong influence on future wages. This result offers a possible explanation for the misspecification of the usual static earnings function. Because of the model structure and especially the fact that the non-linearity is within the budget constraint, the overall implications of the model can only be evaluated by simulation. Simulating reveals that the intertemporal elasticity of labour supply is not constant as is usually assumed in static models, but instead rises over the life cycle.

It seems that these models have been used with male rather than with female labour supply because the estimation method used does not readily extend to discrete data. Altug and Miller (1991) propose a solution to that problem. We shall not go into the details of their paper, which is still largely preliminary, but it seems worth mentioning that this is a very sophisticated and innovative study, which also considerably improves upon the treatment of aggregate shocks adopted by the two studies just mentioned. The main drawback is that estimation of the model of Altug and Miller is intricate, combining GMM with simulation of participation probabilities and iterative estimation of Euler equations, including non-parametric regressions at each iteration. In short, it requires the use of a supercomputer. Another drawback, a theoretical one, is that the model heavily relies on the assumption that actual hours of work differ from expected or contracted hours of work in a stochastic manner. While this may be attractive for some occupations (think of academics), it is much less convincing for most one can think of. To our knowledge, this is the only study of labour supply allowing for non-separability both in the preferences and in the budget constraint.

The study of Eckstein and Wolpin (1989a) shares this generality but restricts attention to the participation decision and disregards aggregate shocks. Insofar, it does not exactly fit the framework of this survey. Yet it seems possible to formulate labour supply models in a similar way, and the study exemplifies the type of precise statements that becomes possible with this degree of generality, concerning e.g. the influence of experience on labour supply: the results of Eckstein and Wolpin show that experience lowers the utility of continued work but that the effect is over-compensated by the positive influence of experience on future wages. Their approach also has the definite advantage of explicitly taking the unobserved heterogeneity into account. Estimation is based on the explicit solution of the dynamic programming problem of each individual in each iteration of a maximum-likelihood procedure.[14] The problem is

$$\max \quad E_t \sum_{k=0}^{T-t} \left( \frac{1}{1+\rho} \right)^k U(p_{t+k}, \underline{M}_{t+k}, x_{t+k}, H_{t+k-1}, S), \tag{52}$$

where $p_s$ is the participation indicator of the period $s$, $\underline{M}_s$ is the vector of the number of children in different age groups, $x_s$ is consumption, $H_{s-1}$ is the number of periods worked before the current period and $S$ is education. The budget constraint is given by

$$y_t^f p_t + y_t^m = x_t + \underline{c}'\underline{M}_t + b\, p_t . \tag{53}$$

The LHS variables are the male and female earnings and the last terms on the RHS are the costs of children and the fixed cost of participation. Therefore, there is not any credit possibility here, in contrast to the models we have discussed so far where the existence of perfect capital markets was assumed. The functional form chosen for $U$ is

$$U_t = \alpha_1 p_t + x_t + \alpha_2 p_t x_t + \alpha_3 p_t H_{t-1} + \underline{\alpha}'_4 \underline{M}_t p_t + \alpha_5 p_t S + f(\underline{M}_t), \tag{54}$$

where $f$ can remain unspecified because $\underline{M}_t$ is not a decision variable (the women are aged 39-44 years in the initial period). The sample used consists of 318 women out of the 1967-1982 NLS). The specification is not as detailed as in Shaw's study, especially the rental rate of human capital variable is not included:

$$\ln y_t^W = \beta_1 + \beta_2 H_{t-1} + \beta_3 H_{t-1}^2 + \beta_4 S + \varepsilon_t . \tag{55}$$

An interesting idea which is mentioned, yet not pursued, by the authors would be to also let the variance of $\varepsilon_t$ depend on past decisions. Given all those ingredients the decision rule at time $t$ takes the reservation wage form:

$$\begin{aligned} p_t &= 1 \quad \text{if} \quad \varepsilon_t \geq \varepsilon_t^*(H_{t-1}), \\ p_t &= 0 \quad \text{otherwise}, \end{aligned} \tag{56}$$

where the function $\varepsilon_t^*$ depends on all parameters. Measurement errors in $\ln y_t^W$ are explicitly accounted for, using the assumptions of the classical errors-in-variables model. Except for the parameters $\alpha_1$ and $\underline{\alpha}_4$ of the utility function, which cannot be separately identified from parameters $b$ and $c$ of the budget constraint, all the other parameters are identified.

14 Eckstein and Wolpin (1989b) survey the use of this approach for several economic problems and Gönül (1989) gives another example of application to the participation decision of men in the presence of layoffs and uncertain job offers.

Noteworthy results are the following: (i) the variance of earnings accounts for 85% of the error process, which suggests a certain caution with respect to the results obtained by Shaw; (ii) $\alpha_3$ is not significant and therefore the intertemporal separability of preferences cannot be rejected; "fixed effects" and "random effects" estimations on the subsample of those women who have changed their participation status at least once show that this result is not a consequence of the neglect of unobserved heterogeneity - at least as far as the parameters $\alpha_1$ and $\alpha_3$ are concerned; (iii) education does play a major role in the explanation of the changing rates of participation.

Although this approach seems very promising, it must be stressed that it very heavily relies on the rationality of the household, since it incorporates an explicit solution of the dynamic programming problem at each period. In this respect, the approach of Altug and Miller, which rests only on the estimation of first order conditions, is much less demanding although it also assumes rational expectations. At this stage it would be difficult to say which of these two approaches will prove to be more successful in practice.

# 3 Data Issues

This section is concerned with data problems in the economics of labour supply. These are relatively minor as compared to other fields of labour economics (Hamermesh, 1988), like for instance the discussion of union differentials (see Solon, 1988, 1989, for some points on self-selection and wage differentials and also Freeman, 1984, on labour market dynamics). A number of variables play a key role in theory but can actually never be measured, such as human capital, marginal value of wealth etc. Approximations are used for resolving the dilemma. Yet, as long as this kind of variable is used in a model, the latter will be almost immune to falsification. The problem here is one of economic theory, of making models operational (testable), rather than one of econometrics (see Griliches, 1986). Then there are data that can in principle be measured but are usually not collected: data sets describing the demand and supply sides of a labour market equally well, extensive data on the biographical background of individuals, etc., are not readily available. Here an evaluation is needed of what is feasible and what is likely to be fruitful, possibly along the lines of Stafford's (1986) approach, that Hamermesh (1988) labels as Schumpeterian. A further point is the quality of the available data and this is the main concern of this section. We will discuss the extent of data mismeasurement and its implications for estimation.

Here we provide an overview of empirical evidence from validation studies on measurement errors. The point in examining the data quality of various panels instead of concentrating on the classical errors-in-variables model (EVM) is that the measured errors failed to meet the assumptions of the EVM. For example, errors in earnings were found to have positive autocorrelation over two years and to be negatively correlated with true earnings (Bound and Krueger, 1989, Bound et al. 1990). This does not mean that innumerable panels have to be re-examined to evaluate their data quality: 90% of the published studies work with either the Panel Study of Income Dynamics (PSID), the Current Population Survey (CPS) or the National Longitudinal Survey (NLS) (Hamermesh, 1988). The characteristics of the measurement errors in those data sets have all been evaluated in validation studies but doubts remain concerning the representativeness of those studies and the stability of the properties found. Yet, the only alternative to validation studies is ignorance.

An error - the term "error" does not imply that a wrong answer has not been given on purpose - is defined as the deviation of interview answers of employees from the corresponding validation data source (either employer or social security data): in the sequel we shall refer to measurements from this second data source as the "true" values.

Following Bound et al., we write the true model and the observed values of the variables as

$$y = \beta x + \varepsilon, \tag{57}$$

$$x = \tilde{x} + u, \tag{58}$$

$$y = \tilde{y} + v, \tag{59}$$

where "~" indicates observed rather than true. The classical EVM assumptions, i.e. the correlation between the error-in-variables terms as well as their correlations with the variables are all assumed to be zero, lead to the following conclusions: (i) an error in the dependent variable causes a loss in efficiency; (ii) errors in the regressor lead to downward biased and inconsistent estimates, the degree of bias expressed as the ratio of the estimated to the true coefficient $b/\beta$ being proportional to the signal-to-noise ratio (or reliability)

$$\zeta = \frac{\sigma_x^2}{\sigma_x^2 + \sigma_u^2}, \qquad \text{or} \qquad \zeta = \frac{\sigma_{x_t}^2}{\sigma_{x_t}^2 + \sigma_{u_t}^2(1-\rho)/(1-r)}, \tag{60}$$

if we want to look at the consequences of estimating in first differences, assuming that both true value and measurement error are autocorrelated, with autocorrelation coefficients $r$ for $x$ and $\rho$ for $u$.

Dropping the classical assumption[15] of no correlation between $u$ and $x$, the reliability measure $\zeta$ can be generalized to

$$\zeta = \frac{cov(x,u)}{var(x)} \qquad \text{or} \qquad \zeta = \frac{cov(\Delta x, \Delta u)}{var(\Delta x)} \tag{61}$$

for levels and first differences, respectively.[16] Alternatively (Bound et al., 1990), the downward bias induced is given by $b_{u\tilde{x}}$ obtained by running the hypothetical regression

15 Since no findings on actual correlation structures of $\varepsilon$ with the other error terms or variables can be obtained the assumption of $\varepsilon$ being uncorrelated with all other variables and error terms is maintained.

16 Bound and Krueger (1989) estimate the elements of the formulae given above. Their estimation procedure is complicated by the fact that, in the tax files they use, annual earnings are truncated at the maximum taxable income.

$u = \beta x + u^*$. If there is negative correlation between $x$ and $u$ then $b_{u\tilde{x}}$ can be smaller than $\zeta$. Allowing for correlation between the dependent variable and its error term, the induced bias is proportional to $d_{v\tilde{y}}$ obtained, as above, by hypothetically estimating $v = \delta y + v^*$. These results still hold if one is working with first differences rather than with levels. Taking the autocorrelation $r$ into account, the variance of $\Delta x$ is given by $2\sigma_x^2(1-r)$ which can be larger or smaller than $\sigma_x^2$. A worst case scenario of aggravating bias by moving from levels to differences would be a highly correlated $x$ and an almost uncorrelated $u$. This would decrease the signal-to-noise ratio considerably. On the other hand, the bias from omitting a variable that is constant over time and correlated with $\tilde{x}$ is avoided and there is no way of assessing the trade-off in bias occurring when moving from levels to first differences. Griliches and Hausman (1986) give conditions for the within estimator to be less severely asymptotically biased than a difference estimator.

Reading the following paragraphs the reader ought to keep in mind that some of the reported error characteristics might be due to the particular setting of the validation study (see Bound et al., 1990). It is always a good idea to check the original literature to see how exactly the results have been obtained. Findings on the characteristics of errors affecting different variables of interest are described below.

*Annual Earnings:* These seem to be underreported in general. Mellow and Sider (1983) find that employer-reported wage exceeds employee-reported wage by 4.8% on average. Duncan and Hill (1985) measure the average absolute difference between employer and employee data to be 7%. The average absolute change was found to be larger in employer than in employee data, so error variance did not increase the variance of the employee data. Bound and Krueger (1989) found the mean reverting error, i.e. the negative correlation of the true value with the measurement error, to be larger than -0.4 in absolute terms for men in each year of their study. Employees obviously tend to state some amount between their true income and the average income of workers. This reduces bias if earnings are used as an independent variable but produces bias if earnings are used as a dependent variable. Distribution of the measurement error is unimodal and bell shaped with very heavy tails. $\zeta$ is somewhat higher than the value found by Duncan and Hill (1985), being slightly above 80%, and it increases if autocorrelations are taken care of. With first differences the reliability falls (but not significantly) because of mean reverting error and positive autocorrelation (pp.11 and 16). Questions about earnings asking for "usual" or for last week's (month's) figures tend to be worse than what is found for annual figures (Bound et al., 1990).

*Annual Hours*: Mellow and Sider (1983) find that male workers overreport hours by 3.9% whereby in 15% of the cases employer exceeds employee response and the opposite is true for 30%. Card (1987) uses these results to estimate a true variance of 0.26 out of a total variance (including error) of 0.35. Duncan and Hill detect a 10% error in absolute differences from the mean. Bound et al. find reports of "usual" hours to be of about the same quality whereas questions about last week's hours are less reliable.

*Note*: Outliers are often removed from the panel before the estimations are run. This corresponds to the assumption that values that lie outside of a certain interval around the mean are likely to be mainly due to measurement errors. However, removing all values that were farther than 5 standard deviations away from the mean, Duncan and Hill (1985) found that the reliability of the data sunk.

*Average Hourly Earnings*: The most thorough evaluation of that figure was made by Bound et al. who compared three different ways of calculating wages, all based on forming a ratio of different earnings and hours measures in order to arrive at a hourly wage measure. They asked (a) for data of the last pay period, (b) of last year and (c) for usual earnings and hours. The quality of the hours data is fairly constant across strategies, the correlation between the interview and the true values ranging from 0.60 to 0.64. The annual earnings data show a correlation of 0.81 and a reliability of two thirds. This clearly dominates the two other strategies, which produce correlations of 0.46 and reliabilities below one fourth. The hourly wages calculated by dividing *annual* figures are clearly superior to the other two measures. Duncan and Hill calculated an error in average absolute values of -12%. Here too, the removal of outliers tends to decrease the signal-to-noise ratio.

Bound et al. give detailed tables of all the observable correlations relevant for assessing the validity of the classical assumptions as well as the nature of the bias (partly taken from Bound and Krueger, 1989).

Assessing the impact of measurement error by comparing results of regressions based on noisy employee-supplied data to more reliable employer-supplied data, Mellow and Sider consider the hours/earnings complex as a dog that does not bark. However, Duncan and Hill run the following regression

$$\ln Y = \gamma_o + \gamma_T \mathrm{TNR} + \gamma_P \mathrm{PEX} + \gamma_E \mathrm{ED} + \varepsilon, \tag{62}$$

where Y denotes annual earnings, TNR is tenure, PEX is previous experience and ED is education, and find that the returns to tenure are 25% lower when earnings are measured with error, i.e. when using the figures supplied by employees rather than employers. Bound et al. find that returns to tenure are underestimated by a third and returns to schooling overestimated by a third. The effects are less clear if earnings are used as an independent variable.

*Correlation Structures:* With the correlation structure that we have in the earnings data, the optimal choice of an estimator depends not only on the type of process, but on the particular correlation coefficients. In general the more positive autocorrelation the errors have, the more of it is eliminated using first differences. Consequently, Bound et al. (1990) conclude that "first differencing is not as harmful as had been previously thought".

*Retrospective Reports:* Data problems are generally aggravated if the data were not reported in current but in later periods. Estimations based on data obtained by retrospective reporting face serious measurement error. Bound et al. (1990) find that only one third of past spells of unemployment are reported. More than one third of long spells (30 weeks and more) were not reported, the rest was seriously underreported.

The spells close to the interview were reported more accurately but less than half of them were reported at all. This might in part explain the influence of errors on earnings functions: if unemployment is underreported and "usual" incomes are reported by workers, the income effects of "unusual" unemployment, e.g. during a recession, being ignored, we have a negative covariance of earnings and tenure via the negative correlation of unemployment and tenure.

Duncan and Hill (1985) report that the difference from average annual earnings in absolute value rises from 7% for the current year to 20% for the year preceding the report. The increased variance must be due to increased error variance. The absolute difference of reported hourly earnings also rises significantly from $2.13 to $2.63 with an average hourly wage of less than $17. The error of reported annual work hours also rose from 10% to 12% in absolute terms.

*Using Demand Side Data:* In his study on data difficulties in labour supply, Hamermesh (1988) concludes that information from the demand side will have to be used in future studies because the approach of exclusively using supply data is at the point of decreasing returns. This point is explicitly stressed by many other researchers, as for example Card (1987) and Abowd and Card (1989) who claim that the covariance structure of hours and earnings implies that both are equiproportionally affected by a component that would be identified as individual productivity growth in a life cycle context. However, individual productivity growth should affect earnings far more than hours according to the life cycle theory. Therefore the authors consider the proportional movement of earnings and hours as mainly demand-driven. The same conclusion is drawn by Altonji and Paxson (1986) who arrive at their findings by estimating different hours determination models.

Panel data sets that contain data on individuals, their jobs and the industry they work in are rarely, if at all, available.[17] A remedy would be to use the available panels on individuals and add some variables on "their" industries (from other data sources). One has to be aware of the fact that by doing so one adds yet another measurement error to the list. Mellow and Sider (1983) find that detailed industry affiliation is correctly reported only in 70% to 90% of the cases. They run a regression on the job-risk/wage-compensation trade-off and find that using the correct data instead of the interview data (which in this case -CPS- contained only 15% wrong answers) leads to an increase of 40% to 50% in the coefficient.

*Job characteristics*: The use of additional job characteristics, as suggested by Altonji and Paxson (1986), bears some risk because only 57.6% of the respondents were able to identify their detailed (three digit) occupational status correctly. Duncan and Hill (1985) find that salient fringe benefits are reported quite correctly but for example eligibility for early retirement is not reported correctly in 28% of the cases.

17 A notable exception is the German Sozioökonomisches Panel.

## Appendix on data sources

*Duncan and Hill (1985):* Data from a company compared with data from interviews that are based on the Panel Study of Income Dynamics (PSID) questionnaire.

*Mellow and Sider (1983):* Data from a special supplement to the January 1977 Current Population Survey (CPS) and from the Employment Opportunity Pilot Project, both containing data from employees and their employers. Mellow and Sider state that proxies are about as reliable as self-reported data.

*Bound, Brown, Duncan and Rodgers (1990):* Data from the PSIDVS (VS indicating that a validation study has been conducted for this particular sample from the PSID), from a firm which provided data of its workers that participated in PSID, and data from the March Current Population Survey (1977 and 1978) matched to the Social Security earnings records.

*Bound and Krueger (1989):* CPS and Social Security pay-roll tax records (see above). The sample was truncated; this is believed to impose a downward bias on reliability measures. The Mellow and Sider proxy result holds here, too.

*Abowd and Card (1989):* Data from the National Longitudinal Survey (NLS), PSID and Seattle and Denver Income Maintenance Experiments (SIME/DIME) were used, outliers were removed! The validation study strategy is to compare a second data set from supposedly more reliable sources with the data obtained from the employee interviews.

*Freeman* (1984) CPS 1977, CPS 1979.

# 4 Econometrics

While the analysis of static male labour supply functions using cross section data has always been regarded as little fruitful due to the lack of variability of hours the estimation of male labour supply over the life cycle has gained more interest in the context of panel data. Since the truncation problem is often negligible for the conventional male labour supply case, most of the standard estimation techniques developed for panel data can be applied. See Chamberlain (1984) and Hsiao (1986) for excellent surveys. In this section we restrict attention to the econometrically more challenging case of non-linear panel models that are able to account for individual heterogeneity and the participation decision. Maddala (1987) gives a non-technical survey on much of the previous work on limited dependent variable models using panel data.

Let us illustrate the available estimation techniques for a simple participation model with heterogeneity. Starting from a straightforward extension of a binary choice model for panel data we have

$$y_{it}^{*} = \underline{\beta}'\underline{x}_{it} + c_i + u_{it}, \quad i = 1, \ldots, N, \quad t = 1, \ldots, T \tag{63}$$

and

$$d_{it} = \left\{ \begin{matrix} 1 & \text{if} \quad y_{it}^{*} \geq 0, & \text{(i.e. the individual works)} \\ 0 & \text{otherwise,} & \end{matrix} \right\}$$

where $y_{it}^{*}$ represents a latent variable (e.g. the difference between market wage and reservation wage) and $c_i$ an individual specific effect. Since differencing of the i-th observation in order to eliminate the individual effect is not feasible in models with qualitative or limited dependent variables, none of the standard approaches to panel data can be applied. Given fixed individual effects standard MLE yields consistent estimates only when $T$ tends to infinity. The more likely panel data situation of a small number of waves and a large number of cross-sectional observations produces an incidental parameter problem. For MLE in qualitative or limited dependent variable models, estimates of the fixed effects $c_i$ and the common slope parameters $\beta$ are not independent of each other. Thus the inconsistencies of the fixed effect estimates lead to inconsistent estimates of the slope parameters (see Chamberlain, 1984, p.1275 for a proof based on the logit specification).

The only fixed effect approach for large N and fixed T that produces consistent estimates for $\beta$ is the conditional maximum likelihood logit approach proposed by Chamberlain (1980).[18] Given uncorrelated error terms $u_{it}$, the main idea of the conditional likelihood

18 As mentioned by Maddala (1987, 317) Chamberlain shows that this extends also to the multinomial logit and log linear models.

approach is to condition on the statistics $\sum_t d_{it}$, which are sufficient for $c_i$. To illustrate, let us assume for simplicity a panel consisting of only two waves.[19] Then the probability of participation is given by:

$$P(d_{it} = 1 \mid \underline{x}_{i1}, \underline{x}_{i2}, c_i) = \frac{\exp(\underline{\beta}' \underline{x}_{it} + c_i)}{1 + \exp(\underline{\beta}' \underline{x}_{it} + c_i)}. \tag{64}$$

Hence the i-th individual's participation probabilities conditional on a change of participation status $(d_{i1} + d_{i2} = 1)$ are:

$$P(d_{i1} = 1 \mid \sum d_{it} = 1) = \frac{\exp(\underline{\beta}'(\underline{x}_{i1} - \underline{x}_{i2}))}{1 + \exp(\underline{\beta}'(\underline{x}_{i1} - \underline{x}_{i2}))}, \tag{65}$$

$$P(d_{i2} = 1 \mid \sum d_{it} = 1) = \frac{1}{1 + \exp(\underline{\beta}'(\underline{x}_{i1} - \underline{x}_{i2}))}. \tag{66}$$

Note that the conditional probabilities do not depend on the fixed effect. Defining the random variable $w_i$ for the two sequences of change in participation status:

$$w_i = \begin{cases} 1 & \text{if} \quad (d_{i1}, d_{i2}) = (1, 0) \\ 0 & \text{if} \quad (d_{i1}, d_{i2}) = (0, 1) \end{cases} \tag{67}$$

$$i \in B = \{ j \mid d_{j1} + d_{j2} = 1 \}$$

gives the following conditional log-likelihood function for the subsample of the individuals who change their participation status:

$$\ln L = \sum_{i \in B} \{ w_i \ln\left( \frac{\exp(\underline{\beta}'(\underline{x}_{i1} - \underline{x}_{i2}))}{1 + \exp(\underline{\beta}'(\underline{x}_{i1} - \underline{x}_{i2}))} \right) +$$

$$(1 - w_i) \ln\left( \frac{1}{1 + \exp(\underline{\beta}'(\underline{x}_{i1} - \underline{x}_{i2}))} \right) \}. \tag{68}$$

Since (68) has the same form as the log-likelihood in the binary logit case, standard ML-logit software packages can be applied to obtain consistent parameter estimates of $\underline{\beta}$ and estimates of the asymptotic standard errors provided (68) satisfies some regularity conditions. The latter impose mild restrictions on the fixed effects. The econometric software package LIMDEP includes a conditional ML-logit routine for panel data up to five waves which does not require a preprocessing of the original data to obtain the conditional likelihood specification. However, since the conditional ML-logit approach uses only the observations on changes in the labour force participation status, there is likely to be a substantial reduction in the numbers of observations

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19 This classical example is given by Chamberlain (1984) and Maddala (1987).

that can actually be used for estimation. Moreover, comparative static results in terms of the marginal participation probabilities (64) are not available. The conditional logit approach only allows the evaluation of the estimated change in the log odds of participation.

If one is willing to accept the random effects assumption with $v_{it} = c_i + u_{it}$ as normally distributed error term that is correlated across cross-sectional units, ML-probit (or ML-Tobit) estimation of (63) yields consistent parameter estimates. Robinson (1982) gives a proof for the Tobit model which also holds for the probit. However, simple pooling approaches yield inefficient estimates since they ignore the correlations among the errors. Given a multivariate normal distribution for $v_{it}$, MLE remains computationally tractable for small panels $(T \leq 3)$. For larger panels some authors (see Heckman and Willis, 1976, Heckman, 1981c, and Butler and Moffitt, 1982) suggest more parsimonious specifications of the error term covariance matrix in order to avoid the computation of $T$ -fold integrals. Assuming that the individual effect results from a random distribution $G$ which depends on a parameter vector $\underline{\delta}$ and is independent of the explanatory variables, the log likelihood function for the binary choice problem with normal errors becomes:

$$\ln L = \sum_{i=1}^{N} \ln \int_{-\infty}^{+\infty} \prod_{t=1}^{T} [\Phi(\underline{\beta}'\underline{x}_{it} + c)]^{d_{it}} \left[1 - \Phi(\underline{\beta}'\underline{x}_{it} + c)\right]^{1-d_{it}} \mathrm{dG}(c \mid \underline{\delta}), \tag{69}$$

where $\Phi$ denotes the standard normal distribution. Under weak regularity conditions, maximization of (69) gives consistent estimates of $\underline{\beta}$ and $\underline{\delta}$ as $N$ tends to infinity.

As already outlined in section 2, appropriate assumptions about the preference structure yield econometrically tractable decision rules for the life cycle leisure and consumption decisions. The Euler equations generated by such a model can be estimated by exploiting the orthogonality between every variable in an information set $\Omega_t$ and the error term at $t+1$ that arises from the approximation in equation (20). Generalized method of moments estimators (GMM), sometimes also referred to as non-linear instrumental variables or orthogonality conditions estimators, impose sample analogues of population orthogonality conditions implied by the regression equation (see Hansen, 1982, and White, 1982). Unlike the previously discussed approaches to the random effects model with qualitative or limited dependent variable, application of the GMM principle to panel data does not require an explicit parameterization of the temporal covariances of the errors. Thus, GMM is immune against misspecification with respect to autocorrelation and conditional heteroskedasticity given the choice of valid instruments. Avery, Hansen and Hotz (1983) apply GMM to the female labour force participation problem and Hotz et al. (1988) use the GMM principle to estimate male labour supply over the life cycle, ignoring the truncation of the dependent variable.

Assuming some functional form for the utility function leads to the $i$-th individual's Euler equation of the general form:

$$f(\underline{x}_{it}, \underline{\beta}_o) = u_{i,t+1}\,. \tag{70}$$

Since we assume rational behaviour, $u_{i,t+1}$ is orthogonal to the information set $\Omega_{it}$, and for a vector of instruments $\underline{z}_{it}$ whose elements are contained in $\Omega_{it}$ we have

$$\mathrm{E}[f(\underline{x}_{it}, \underline{\beta}_o)\underline{z}_{it}] = 0, \tag{71}$$

where E is the unconditional expectation operator. For the panel data case the population orthogonality conditions can be summed over the $T$ waves, leading to

$$\mathrm{E}\sum_{t=1}^{T}[f(\underline{x}_{it}, \underline{\beta}_o)\underline{z}_{it}] =: \mathrm{E}[\Psi(\underline{x}_i, \underline{\beta}_o, \underline{z}_i)] = 0, \tag{72}$$

where $\underline{x}_i = (\underline{x}'_{i1}, \underline{x}'_{i2}, \ldots, \underline{x}'_{iT})'$ and $\underline{z}_i = (\underline{z}'_{i1}, \underline{z}'_{i2}, \ldots, \underline{z}'_{iT})'$. This otherwise arbitrary procedure proves very useful for panel data: the main idea of GMM estimation of $\underline{\beta}_0$ is based on the fact that the sample analogue of (72) implies the following sample orthogonality conditions

$$\mathrm{O}_N(\underline{\beta}_0) = \frac{1}{N}\sum_{i=1}^{N}\Psi(\underline{x}_i, \underline{\beta}_0, \underline{z}_i), \tag{73}$$

which converge to zero as $N$ goes to infinity. A consistent estimator of $\beta_0$ can be obtained by minimizing the following quadratic criterion function based on the sample orthogonality conditions:

$$\mathrm{O}_N(\underline{\beta})'\mathrm{A}_N\mathrm{O}_N(\underline{\beta})\,, \tag{74}$$

where $A_N$ is a symmetric positive definite weighting matrix which is usually a function of sample information. The asymptotic efficiency of the GMM estimator depends on the specific choice for $A_N$ and the number and nature of the instruments chosen. If the error term in (70) does not result from an optimal decision rule under uncertainty, more instruments (e.g. leads of $\underline{z}_{it}$) can be used in order to improve the asymptotic efficiency of the estimator. Hansen (1982) shows that the GMM estimator is consistent and asymptotically normal. Because of restrictions in space we refrain from reproducing the expressions for the optimal choice of $A_N$ and the asymptotic covariances of the estimators. The software packages DPD, HOTZTRAN, LIMDEP and MOMENTS include subroutines to perform GMM estimation.

As pointed out earlier, for normally distributed error terms, the conditional likelihood approach is not feasible since the fixed effects do not vanish by conditioning. Based on the idea of Mundlak's (1978) correlated random effects model Chamberlain (1980,1984) proposed a random effects model assuming that the individual effect and the explanatory variables arise from a joint normal distribution. In Chamberlain's panel probit approach the error terms $u_{it}$ are of the general normal form:

$$(u_{i1} \ldots u_{iT})' \text{ i.i.d } N(0, \Sigma), \tag{75}$$

where the error terms are independent of the individual effects and the explanatory variables. Suppose $x_{it}$ consists for notational simplicity of only one single variable (is a scalar). The central assumption in Chamberlain's approach is that the distribution of $c_i$ conditional on $\underline{x}_i = x_{i1}, \ldots, x_{iT}$ can be specified as:

$$c_i = \delta_1 x_{i1} + \ldots + \delta_T x_{iT} + v_i \qquad v_i \mid \underline{x}_i \sim \mathrm{N}(0, \sigma_v^2). \tag{76}$$

Given the distributional assumptions (75) and (76), the regression equation (63) takes the form:

$$y_{it}^* = \beta x_{it} + (\delta_1 x_{i1} + \delta_2 x_{i2} + \ldots + \delta_T x_{iT}) + v_i + u_{it}, \tag{77}$$

with the probability of participation given by:

$$\mathrm{P}(d_{it} = 1 \mid \underline{x}_i) = \Phi(\alpha_t(\beta x_{it} + \delta_1 x_{i1} + \delta_2 x_{i2} + \ldots + \delta_T x_{iT})), \tag{78}$$

with
$$\alpha_t = (\sigma_v^2 + \sigma_{tt})^{-1/2}.$$

Thus, $y_{it}^*$ is a function of all available leads and lags of $x_{it}$ and efficient estimation can be conducted in two stages. Rewrite (78) as

$$y_i^* = \Pi \underline{x}_i + e_i \tag{79}$$

with

$$\Pi = \mathrm{diag}\{\alpha_1, \ldots, \alpha_T\} [\beta \mathrm{I}_T + \underline{\iota}_T \underline{\delta}'] \tag{80}$$

where $\underline{\iota}_T$ is a $T$-dimensional column vector consisting of ones. In the first stage, each row of $\Pi$ is estimated separately by cross-section probit (or cross section Tobit if hours can be observed), and the restrictions given by (80) are imposed in the second stage by minimum distance estimation. Identification of the parameters can be obtained by restricting one of the $\alpha$s to unity. In the panel Tobit case the $\Pi$-matrix is equal to the matrix in squared brackets in (80) and hence identification of $\beta$ and $\underline{\delta}$ is warranted.

The major advantage of Chamberlain's approach is the solution of the incidental parameter problem by assuming (76). If one is willing to accept this strong distributional assumption, Chamberlain's approach to panel probit and Tobit models reveals a number of practical advantages. Beside the computational simplicity, it allows for an unrestricted covariance matrix of the errors and robust estimates of the standard errors. Simple $\chi^2$-tests can be applied as omnibus tests for model specification. Unlike the conditional logit case, comparative statics can be performed in terms of participation probabilities for the probit case or in terms of the model parameters for the Tobit. Although from a theoretical point of view the formulation of the individual effect is not in accordance with our interpretation in section 2, from a practical point of view the usefulness of the Chamberlain approach depends on the quality of the approximation in (76). Finally, note that the Chamberlain approach is also feasible in the case of lagged dependent variables. This might provide a simple way to relax the intertemporal separability assumption.

For the sake of completeness, McFadden's (1989) method of simulated moments (MSM) should be mentioned here. It can be used as an alternative approach to estimate a panel probit model with random effects. Unlike the maximum likelihood approach to the random effects probit model, this approach can also allow for an autoregressive error structure and errors in variables. The basic idea of the MSM estimator is based on moment conditions where response probabilities are replaced by simulated response probabilities in order to avoid numerical integration. Under not too restrictive regularity conditions the estimator is shown to be consistent and asymptotically normal. To our knowledge, the MSM has not yet been applied to the estimation of labour supply functions. See also Bloemen and Kapteyn (1990), Börsch-Supan (1990) and Gouriéroux and Monfort (1989) for related approaches.

Table 1. Summary of results

| Reference/separability/funct. form | Wage effects on hours | Life cycle |
|---|---|---|
| Altonji (1986)<br>intertemporal separability<br>within period additive separability<br>*Box-Cox type* | intertemporal subst. elasticity<br>a) .014 to .07 b) .08 to .45<br>instruments for wages:<br>a): past values<br>b): human capital variables<br>consumption as $\lambda$-proxy:<br>-.11 to .17 | |
| Bover (1991)<br>intertemporal separability in constraint, not in preferences<br>*Stone-Geary* | intertemporal subst. elasticity: .8 at sample means | lagged hours significant, underlines importance of relaxation of separability assumption |
| Browning, Deaton and Irish (1985)<br>intertemporal separability in constraint and preferences<br>within period additive separability relaxed<br>uncertainty in wages and interest rates<br>*Gorman polar form* | intertemporal subst. elasticity<br>•certainty and<br>a) additive preferences<br>manual: .15 non-manual: .14<br>b) non additive preferences<br>all: .40<br>•uncertainty and b)<br>all: .40 | significance of year dummies, incompatibilities between leisure and goods equations, reduction in consumption in presence of small children and unusual hours profiles cast doubt on life cycle hypothesis. |
| Eckstein and Wolpin (1989a)<br>intertemporal separability in constraint<br>intertemporal non-separability through previous work periods in preferences<br>no within-period additive separability<br>uncertainty in wages<br>*linear utility* | no elasticities available: solution of dynamic program at each iteration of the maximum likelihood estimation procedure; impact of a change in wages can only be evaluated by means of simulation | intertemporal separability rejected. Marginal utility of wealth varies over life cycle according to interaction between wages and participation |
| Ham (1986)<br>tests for impact of unemployment using specifications of MaCurdy (1981) (MC) and Browning, Deaton and Irish (1985) (BDI) | intertemporal subst. elasticity<br>•MC specification: a)-.10 to .17<br>................................ b)-.17 to .04<br>a): wages endogenous, uncertainty<br>b): wages, unemployment, underemployment endogenous, certainty<br>•BDI specification -.01 | either workers are off their supply function or more complex models of intertemporal substitution must be considered |

and characteristics.

| Demographics and other effects | Unemployment | Data and other characteristics |
|---|---|---|
| explicit treatment of measurement errors | voluntary | PSID 1968 to 1981, 597 males, aged 25 to 49 in 1968 whose wives were younger than 63 in 1968; subsamples according to data requirements |
| children (-) | voluntary | PSID 1969 to 1977, 785 white males, aged 20 to 50 in 1968, 0<hours<5000 in all periods |
| children (+)<br>joint inclusion of numbers of children and cohort dummies leaves no room for wages | unemployed excluded from sample | FES 1970 to 77, pseudo-panel in 8 five-year age cohorts, male household heads aged 18-58 in 1970, married. Two subsamples: manual and non-manual workers |
| children, husband's wage (-)<br>experience and schooling increase disutility of work but effect on wages (+): total effect (+)<br>measurement errors account for 85% of variation in wife's wage | voluntary | NLS 1967 to 81, 318 white women aged 39-44 in 1967, continuously married, with at least 4 consecutive years of data on participation |
| test results concerning the impact of unemployment robust to a wide variety of changes in specification and estimation method | dummies on unemployment and underemployment always significant: unemployment cannot be viewed as voluntary | PSID 1971 to 79, 473 males, aged 25-50 in 1971, continuously married (same wife), poverty subsample included |

Table 1. Summary of results

| Reference/separability/funct. form | Wage effects on hours | Life cycle |
|---|---|---|
| Heckman and MaCurdy (1980, 1982)<br>intertemporal separability<br>within period additive separability<br>*Box-Cox type* | intertemporal subst. elasticity<br>-.406 | no significant impact of transitory income fluctuations, but no perfect substitutability between leisures at different periods<br>Negligible impact of correction for selectivity bias |
| Hotz, Kydland and Sedlacek (1988)<br>intertemporal separability in constraint<br>intertemporal non-separability through leisure in preferences<br>no within-period additive separability<br>uncertainty in wages / *Translog* | no elasticities available: consumption Euler equation estimated<br>indirect inference: intertemporal substitution elasticity falling over life cycle | intertemporal separability rejected; negative estimate of time preference points to remaining misspecification. |
| Hujer and Schnabel (1990)<br>intertemporal separability<br>within-period additive separability<br>*Box-Cox type* | not identified (reduced form only) | current unearned income well determined but less significant and less important in Tobit estimates with correlated random effects and unrestricted covariance for residual errors than in other estimates |
| Jakubson (1988)<br>intertemporal separability<br>within period additive separability<br>no uncertainty in estimated models<br>*Box-Cox type* | intertemporal subst. elasticity:<br>a) 1.14 b) 1.72<br>a): restrictions on wages, numbers of children and other income<br>b): restrictions on wages only | current income insignificant with fixed effects but significant with random effects; cross-section<br>estimates biased away from 0 due to omission of individual effect correlated with observed<br>variables |
| Johnson and Pencavel (1984)<br>intertemporal non-separability in constraint<br>intertemporal separability in preferences<br>within period additive separability<br>*Stone-Geary* | compensated elasticity[20]<br>a) male .129<br>female .161<br>b) ........ .149<br>a): married couples<br>b): single females | two-period model |

20 This is a lower bound for the intertemporal substitution elasticity (see page 5).

and characteristics. (continued)

| Demographics and other effects | Unemployment | Data and other characteristics |
|---|---|---|
| directly on labour supply:<br>children (-)<br>age (-)<br>through marginal value of utility<br>education (+)<br>children (-) | voluntary;<br>husband's unemployment hours lower value of wife's time at home | PSID 1968 to 75, 672 white females aged 30-65 in 1968, and subsample of 212 continuously married (same husband). For Tobit, subsample of 452 women who had worked in at least one period |
| leisure becomes less substitutable the higher the education and the more children; exogeneity of wages not clearly rejected | voluntary | PSID 1967 to 78, 482 white male household heads aged 23-52 in 1967, continuously married, with positive hours in each year. |
| schooling (+)<br>exp. (+)<br>exp. sq. (-)<br>children (-)<br>children effects stronger in panel than in cross-section estimates | voluntary | Socio-economic panel (FRG) 1984-1987<br>1182 continously married (same husband) German women aged 16-58 over periods |
| fixed and random effects give same answer:<br>children (-)<br>schooling (+)<br>experience (+)<br>$(\text{experience})^2$ (-)<br>panel effects of children 60% of cross-section effects | voluntary | PSID 1968-70-2-4, 924 white women aged 20-50 in 1968, continuously married, not in low income subsample.<br>Focus on random vs. fixed effects vs. cross-section |
| | | SIME/DIME participants (1678 couples + 1339 single females) responding in preenrollment and first two years of NIT experiment; no marital change. |

Table 1. Summary of results

| Reference/separability/funct. form | Wage effects on hours | Life cycle |
| --- | --- | --- |
| Lilja (1986)<br>several models under certainty and uncertainty, joint decisions, rationing, additive and implicit intertemporal separability, focus on Frisch demands<br>*functional forms used include Box-Cox type and BDI specifications* | intertemporal subst. elasticity<br>•certainty<br>a) m 0.38 to 0.42 f -.15 to -.14<br>b) m -.25 to -.23 f -.84 to -.25<br>c) m 0.84 to 1.14 f -.25 to -.16<br>a): no distinction between permanent and transitory wage effect<br>b): permanent component<br>c): transitory component[21]<br>•uncertainty<br>m 0.19 to 0.24 f -.83 to -.82 | results under uncertainty and rationing suggest data difficulties: need for savings or consumption data; peak hours at peak wages does not hold true for unemployed |
| Lillard (1978)<br>no explicit reference to utility: decomposition of wages and earnings in permanent and transitory components<br>LISREL application, allowing treatment of measurement errors<br>no uncertainty | coefficient of log wage in log hours equation:<br>a) -.158 to -.184 (permanent)<br>b) -.307 to -.138 (transitory)<br>c) -.160 (equality restriction) | not based on life cycle theory |
| Lundberg (1988)<br>intertemporal separability in constraint<br>intertemporal non-separability in preferences (dynamic translating)<br>family labour supply, without separability assumption<br>*linear conditional labour supply function* | intertemporal subst. elasticity[22]<br>a) male -.067 female -.018<br>b) male 0.011 female 0.018<br>c) male 0.114 female 0.031<br><br>a): no children<br>b): one child<br>c): two or more | intertemporal substitution rather than habit formation |
| MaCurdy (1981)<br>intertemporal separability<br>within period additive separability<br>*Box-Cox type* | intertemporal subst. elasticity<br>0.10 to 0.23 | |

21 Serial correlation not accounted for.

22 Own calculations at sample means, for short run reactions.

and characteristics. (continued)

| Demographics and other effects | Unemployment | Data and other characteristics |
|---|---|---|
| male female<br>•certainty<br>(+) age (+)<br>(-) unearned income (+)<br>(-) children (-)<br><br>•uncertainty<br>(n.sig.) age (+) | current period unemployment hours and expected unemployment significant | PSID 1971 to 76, without low income subsample, 631 white male household heads aged 25-55 in 1968, employed in each year, and subsample of 212 continuously married couples with working wives (selection bias taken care of) |
| schooling (+)<br>experience (+)<br>serial correl. in transitory wage .8; measurement error accounts for 6.6% (17.4%) of variation in earnings (hours) | | PSID 1967 to 73, 1041 white male household heads aged 18-58 in 1967, not in low income subsample. |
| "traditional family" rejected, "joint utility" not rejected, but strong differences across samples:<br>a): no simultaneity but positive habit formation;<br>b) and c): strong interactions in labor supply decisions;<br>Children, rather than leisure, important jointly-consumed commodity for husbands and wives in this sample | voluntary | DIME 1972 (control group), monthly information, 381 married couples . Quarterly hours worked by husband and wife during the third year are the eight dependent variables; preceding five quarters provide lags. Three subsamples according to presence and number of small children. |
| | voluntary | PSID 1968 to 77, 513 white males aged 25-47 in 1967, and subsample of 212 continuously married (same wife). |

Table 1. Summary of results

| Reference/separability/funct. form | Wage effects on hours | Life cycle |
| --- | --- | --- |
| MaCurdy (1983)<br>intertemporal separability<br>within period additive separability<br>uncertainty in wages<br>*Box-Cox type* | compensated elasticity<br>0.74 to 1.43 | |
| Shaw (1989a)<br>intertemporal non-separability in constraint through wage endogeneity<br>intertemporal separability in preferences<br>within period additive separability<br>uncertainty in exogenous variables<br>*Translog* | no elasticities available: simulation suggests rising rather than constant elasticity over life cycle | high hours early in life at low wages due to human capital formation efforts: suggests varying rates of returns to human capital |

and characteristics. (continued)

| Demographics and other effects | Unemployment | Data and other characteristics |
|---|---|---|
| | voluntary | PSID 1968 to 1981, 526 white male household heads aged 18-41 in 1968. |
| | voluntary | DIME 1972 to 1975, monthly labour supply files, preenroll-ment file, 121 married working males from the control group. |

## 5 Overview of qualitative and quantitative results

Table 1 gives informations concerning the separability assumptions and the functional forms adopted, the intertemporal substitution elasticities reported, if any, or the nearest information of that type, qualitative results concerning the validity of the life cycle hypothesis, the significant socio-demographic effects, the treatment of unemployment and corresponding results and finally some information on the data used. We shall not go through the table in detail since it was designed to be fairly self-contained. However, some comments on the overall picture may be useful. Concentrating first on the elasticities obtained both for men and for women, it is clear that the variance of the "guesstimate" is not much lower for life cycle models using panel data than for models estimated on cross sections. A look at the results reported by Ham (1986) shows that Mroz' (1987) cross section results on the great sensitivity of elasticity estimates based on a single linear labour supply specification to exclusion restrictions, choice of stochastic assumptions and estimation methods extend to panel data studies. Table 2 makes this vivid by reporting rough statistics on the distribution of reported elasticities (i) in all studies (ii) in studies using the PSID only and (iii) in studies using Box-Cox type specifications: even in the latter fairly homogeneous groups the variability is considerable. This points to the fragility of the results, and clearly more work is needed, on refinement of the economic specification, on improving the quality of data and appropriately treating measurement errors, on specification tests and relaxing distributional assumptions, on taking account of institutional restrictions on hours choice as well as on cyclical aspects of labour demand, and possibly most importantly on dynamics. Indeed the simplest explanation for the variety of elasticities is that the assumption of a constant intertemporal substitution elasticity is a misspecification.

## 6 Concluding comments

Taking stock, we can draw the following conclusions. Firstly, in our opinion, there has still been so far too little emphasis on the relaxation of ad hoc assumptions. In a way this is understandable because researchers have been busy introducing and manipulating new and sometimes complex econometric methods. Yet it is disturbing to see how popular the additively separable Box-Cox type specification has been over the decade, even in studies putting forth ideas allowing much more flexible approaches: so far, the greater flexibility of the alternative to Frisch demands consisting in separate estimation of within-period preferences and intertemporal preferences has not been used fully in labour supply studies. Secondly, there is clearly room for progress on the data issues. Given the small sample sizes and the more or less pronounced arbitrariness of the selection, most of the studies we have discussed definitely have a methodological rather than a substantive character. Moreover, the often made selection of continuously married couples is probably not exogenous with respect to labour supply decisions and Lundberg (1988) may well be over-optimistic when she says that, thanks to the use of panel data "*most* of the cross-section simultaneity between labor supply behavior and the determinants of household formation will be avoided" (p. 226, our emphasis). For the same reason, care should be taken to adapt estimation methods so as to handle

**Table 2. Statistics on reported elasticities**

| Studies | minimum | 25% | median | 75% | maximum | number |
|---|---|---|---|---|---|---|
| All | | | | | | |
| male | -.17 | 0.04 | 0.14 | 0.23 | 1.43 | 27 |
| female | -.83 | -.62 | -.14 | 0.65 | 1.72 | 9 |
| PSID | | | | | | |
| male | -.17 | -.06 | 0.07 | 0.17 | 0.45 | 14 |
| female | -.83 | -.60 | -.15 | 0.80 | 1.72 | 7 |
| Box-Cox | | | | | | |
| type | -.11 | 0.04 | 0.10 | 0.60 | 1.43 | 10 |
| male | -.41 | | 0.65 | | 1.72 | 3 |
| female[23] | | | | | | |

unbalanced panels. Thirdly, efforts to generalize dynamic structural models of participation to less restrictive aspects of labour supply, as well as efforts towards relaxing arbitrary distributional assumptions should prove extremely rewarding.

23 Only 3 values: middle of range reported instead of median.

# 7 Literature


ABOWD, J. and D. CARD (1989): "On the Covariance Structure of Earnings and Hours Changes", *Econometrica*, 57(2), 411-445.

ALTONJI, J.G. (1986): "Intertemporal Substitution in Labour Supply: Evidence from Micro-Data", *Journal of Political Economy*, 94 (3.2) S176-S215.

ALTONJI, J.G. and C.H. PAXSON (1986): "Job Characteristics and Hours of Work", in *Research in Labor Economics*, R. Ehrenberg, ed., JAI, Greenwich, Connecticut.

ALTUG, S. and R.A. MILLER (1990): "Household Choices in Equilibrium", *Econometrica*, 58, 543-570.

ALTUG, S. and R.A. MILLER (1991): "Human Capital Accumulation, Aggregate Shocks and Panel Data Estimation", *Discussion Paper 9128*, CentER, Tilburg University.

ARELLANO, M. and C. MEGHIR (1989): "Using Complementary Data Sources: An Application to Labour Supply and Job Search", *Discussion Paper 89-06*, University College London.

ASHENFELTER, O. and J. HAM (1979): "Education, Unemployment and Earnings", *Journal of Political Economy*, 87, S99-S116

ASHENFELTER, O. and G. SOLON (1982): "Longitudinal Labor Market Data-Sources, Uses and Limitations", in *What's Happening to American Labor Force and Productivity?*, W.E. Upjohn Institute for Employment Research, 109-26.

AVERY, R.B., L.P. HANSEN and V.J. HOTZ (1983): "Multiperiod Probit Models and Orthogonality Condition Estimation", *International Economic Review*, 24(1), 21-35.

BLOEMEN, H.G. and A. KAPTEYN (1990): "The Joint Estimation of a Nonlinear Labour Supply Function and a Wage Equation Using Simulated Response Probabilities", *mimeo*, paper presented at the conference *Specification Search and Robust Estimation in Micro Labor Market Models*, June 1-2, 1990, CentER, Tilburg University.

BLUNDELL, R.W. (1987): "Econometric Approaches to the Specification of Life-Cycle Labor Supply and Commodity Demand Behaviour", *Econometric Reviews*, 6(1), 147-51.

BLUNDELL, R.W. (1988): "Consumer Behaviour: Theory and Empirical Evidence - A Survey", *Economic Journal*, 98, 16-65.

BLUNDELL, R.W., and C. MEGHIR (1990): "Panel Data and Life Cycle Models", Ch. 2 in *Panel Data and Labour Market Studies*, J. Hartog, G. Ridder and J. Theeuwes (eds.), North Holland.

BLUNDELL, R.W. and I. WALKER (1986): "A Life Cycle Consistent Empirical Model of Family Labour Supply Using Cross-Section Data", *Review of Economic Studies*, 53, 539-58.

BLUNDELL, R.W., M. BROWNING and C. MEGHIR (1989): "A Microeconometric Model of Intertemporal Substitution and Consumer Demand", *Discussion Paper 89-11*, University College London.

BLUNDELL, R.W., V. FRY and C. MEGHIR (1990): "Preference Restrictions in Microeconometric Models of Life Cycle Behaviour", Ch. 2 in *Microeconometrics: Surveys and Applications*, J.P. Florens, M. Ivaldi, J.J. Laffont and F. Laisney, eds., Basil Blackwell, Oxford.

BÖRSCH-SUPAN, A. (1990): "Estimating a Multiperiod Multinomial Probit Model with Unobserved Heterogeneity and Autocorrelated Errors", *mimeo*, University of Mannheim.

BOUND, J.C., BROWN, G.J. DUNCAN and W.L. RODGERS (1990): "Measurement Error in Cross-Sectional and Longitudinal Labor Market Surveys: Validation Study Evidence", Ch. 1 in *Panel Data and Labour Market Studies*, J. Hartog, G. Ridder and J. Theeuwes (eds.), North Holland.

BOUND, J.C. and A.B. KRUEGER (1989): "The Extent of Measurement Error in Longitudinal Earnings Data: Do Two Wrongs Make a Right?", *NBER, working paper 2885.*

BOVER, O. (1986): "Some Dynamic Life Cycle Models of Labour Supply Estimated From Panel Data", *Oxford Applied Economics Discussion Papers*, 13.

BOVER, O. (1991): "Relaxing Intertemporal Seperability: A Rational Habits Moel of Labour Supply Estimated from Panel Data", Journal of Labour Economics, 9, 85-100

BROWNING, M. and C. Meghir (1989): "Testing for Separability between Goods and Leisure Using Conditional Demand Systems", *Discussion Paper 89-05*, University College London

BROWNING, M., A. Deaton and M. Irish (1985): "A Profitable Approach to Econometric Approaches to Labor Supply and Commodity Demands over the Life Cycle", *Econometrica*, 53, 503-43.

BUTLER, J.S. and R. Moffitt (1982): "A Computationally Efficient Quadrature Procedure for the One-Factor Multinomial Probit Model", *Econometrica*, 50(3), 761-64.

CARD, D. (1987): "Supply and Demand in the Labor Market", *mimeo*, Princeton University (ESEM'87).

CARD, D. (1990): "Intertemporal Labor Supply: An Assessment", *working paper* 269, Industrial Relations Section, Princeton University.

CHAMBERLAIN, G. (1980): "Analysis of Covariance with Qualitative Data", *Review of Economic Studies*, 47, 225-38.

CHAMBERLAIN, G. (1984): "Panel Data", Ch. 22 in *Handbook of Econometrics*, Vol II, Z. Griliches and M.D. Intriligator, eds., North-Holland, Amsterdam.

DUNCAN, G.J. AND HILL, D.H. (1985): "An Investigation of the Extent and Consequences of Measurement Error in Labor-economic Survey Data", *Journal of Labor Economics*, 3(4), 509-532.

ECKSTEIN, Z. and K.I. WOLPIN (1986): "On the Estimation of Labor Force Participation, Job Search and Job Matching Models Using Panel Data", in Weiss, Y. and G. Fisherson, eds. *The Theory and Measurement of Unemployment*. London: MacMillan.

ECKSTEIN, Z. and K.I. WOLPIN (1989a): "Dynamic Labour Force Participation of Married Women and Endogenous Work Experience", *Review of Economic Studies*, 56, 375-90.

ECKSTEIN, Z. and K.I. WOLPIN (1989b): "The Specification and Estimation of Dynamic Stochastic Discrete Choice Models", *Journal of Human Resources*, 24, 562-98.

FREEMAN, R.B. (1984): "Longitudinal Analyses of the Effects of Trade Unions", *Journal of Labour Economics*, 2(1), 1-26.

GöNüL, F. (1989): "Dynamic Labor Force Participation Decisions of Males in the Presence of Layoffs and Uncertain Job Offers", *Journal of Human Resources*, 24, 195-220.

GOURIéROUX, C. and A. Monfort (1989): "Simulation Based Inference in Models with Heterogeneity", *working paper 8905*, ENSAE and INSEE, Unité de Recherche, Paris.

GRILICHES, Z. (1986): "Economic Data Issues", Ch. 27 in *Handbook of Econometrics*, Vol III, Z. Griliches and M.D. Intriligator, eds., North-Holland, Amsterdam.

GRILICHES, Z. and J. HAUSMAN (1986): "Econometrics of Panel Data with Measurement Error", in *Journal of Econometrics*, 31, S.93-118

HAM, J.C. (1986): "Testing Whether Unemployment Represents Intertemporal Labour Supply Behaviour", *Review of Economic Studies*, 53(4), 559-78.

HAMERMESH, D.S. (1988): "Data Difficulties in Labor Economics", *NBER*, 2622.

HANSEN, L.P. (1982): "Large Sample Properties of Generalized Method of Moments Estimators", *Econometrica*, 50, 1029-55.

HECKMAN, J.J. (1981a): "The Incidental Parameters Problem and the Problem of initial Conditions in Estimating a Discrete Time - Discrete Data Stochastic Process and Some Monte-Carlo Evidence", in *Strutural Analysis of Discrete Data*, C. Manski and D. McFadden, eds., MIT Press, Cambridge, MA, 179-95.

HECKMAN, J.J. (1981b): "Heterogeneity and State Dependence", in *Studies in Labor Markets*, S. Rosen, ed., University of Chicago Press, Chicago, 91-139.

HECKMAN, J.J. (1981c): "Statistical Models for Discrete Panel Data", in *Strutural Analysis of Discrete Data*, C. Manski and D. McFadden, eds., MIT Press, Cambridge, MA, 114-78.

HECKMAN, J.J. and T.E. MACURDY (1980): "A Life Cycle Model of Female Labour Supply", *Review of Economic Studies*, 47, 47-74.

HECKMAN, J.J. and T.E. MACURDY (1982): "Corrigendum on A Life Cycle Model of Female Labour Supply", *Review of Economic Studies*, 49, 659-60.

HECKMAN, J.J. and T.E. MACURDY (1986): "Labor Econometrics", Ch. 32 in *Handbook of Econometrics*, Vol III, Z. Griliches and M.D. Intriligator, eds., North-Holland, Amsterdam.

HECKMAN, J.J. and R. WILLIS (1976): "Estimation of a Stochastic Model of Reproduction: An Econometric Approach", in *Household Production and Consumption*, N. Terleckyi, ed., New York, NBER.

HOLTZ-EAKIN, D., NEWEY W. and H.S. ROSEN (1988): "Estimating Vector Autoregressions with Panel Data", *Econometrica*, 56, 1371-1395

HOTZ, V.J., KYDLAND, F.E. and G.L. SEDLACEK (1988): "Intertemporal Preferences and Labour Supply", *Econometrica*, 335-60

HSIAO, C. (1986): *Analysis of Panel Data*, Cambridge University Press, Cambridge.

HUJER, R. AND R. SCHNABEL (1990): "Spezifikation und Schätzung dynamischer Arbeitsangebotsmodelle: eine mikroökonometrische Analyse mit Daten des Sozio-ökonomischen Panels", presented at the conference *Analyse von Mikrodaten als Basis wirtschaftpolitischer Entscheidungen*, Konstanz, 5-6 Oktober 1990.

JAKUBSON, G. (1988): "The Sensitivity of Labor Supply Parameter Estimates to Unobserved Individual Effects: Fixed- and Random-Effects Estimates in a Nonlinear Model Using Panel Data", *Journal of Labor Economics*, 6(3), 302-29.

JOHNSON, T.R. and J.H. PENCAVEL (1984): "Dynamic Hours of Work Functions for Husbands, Wives and Single Females", *Econometrica*, 52, 363-89.

KILLINGSWORTH, M.R. and J.J. HECKMAN (1986): "Female Labor Supply: A Survey", Ch. 2 in *Handbook of Labor Economics*, Vol I, O. Ashenfelter and R. Layard, eds., North-Holland, Amsterdam.

LILJA, R (1986): *Econometric Analyses of Family Labour Supply over the Life Cycle Using US Panel Data*, The Helsinki School of Economics, Helsinki.

LILLARD, L.A. (1978): "Estimation of Permanent and Transitory Response Functions in Panel Data: A Dynamic Labor Supply Model", *Annales de l'INSEE*, 30, 367-94.

LUNDBERG, S.J. (1988): "Labor Supply of Husbands and Wives: A Simultaneous Equations Approach", *Review of Economics and Statistics*, 70(2), 224-35.

MACURDY, T.E. (1981): "An Empirical Model of Labor Supply in a Life-Cycle Setting", *Journal of Political Economy*, 89, 1059-85.

MACURDY, T.E. (1983): "A Simple Scheme for Estimating an Intertemporal Model of Labor Supply and Consumption in the Presence of Taxes and Uncertainty", *International Economic Review*, 24, 265-90.

MACURDY, T.E. (1985): "Interpreting Empirical Models of Labor Supply in an Intertemporal Framework with Uncertainty", Ch. 3 in *Longitudinal Analysis of Labor Market Data*, J.J. Heckman and B. Singer, eds., Cambridge University Press, Cambridge.

MACURDY, T.E. (1990): "Appraising Tests of the Intertemporal Substitution Hypothesis", in *Panel Data and Labour Market Studies* , J. Hartog, G. Ridder and J. Theeuwes (eds.), North Holland.

MCFADDEN, D. (1989): "A Method of Simulated Moments for Estimationof Discrete Response Models Without Numerical Integration", *Econometrica*, 57(5), 995-1026.

MADDALA, G.S. (1987): "Limited Dependent Variable Models Using Panel Data", *Journal of Human Resources*, 22, 307-38.

MELLOW, W. AND SIDER, H. (1983): "Accuracy of Response in Labor Market Surveys: Evidence and Implications", *Journal of Labor Economics*, 1(4), 331-344.

MINCER, J. (1962): "Labor Force Participation of Married Women: A Study of Labor Supply", *Aspects of Labor Economics*, Princeton, N.J.:NBER, Princeton University Press, 63-97.

MUNDLAK, Y. (1978): "On the Pooling of Time-Series and Cross-Section Data", *Econometrica*, 46, 69-85.

NICKELL, S. (1986): "The Short-Run Behaviour of Labour Supply", in *Advances in Econometrics: Fifth World Congress*, Vol II, T.F. Bewley, ed., Cambridge University Press, Cambridge.

PAGAN, A. (1984): "Econometric Issues in the Analysis of Regressions with Generated Regressors", *International Economic Review*, 25, 221-247.

PENCAVEL, J. (1986): "Labor Supply of Men: A Survey", in O. Ashenfelter and R. Layard, eds., *Handbook of Labor Economics*, North-Holland, Amsterdam.

POTERBA, J.M. and L.H. SUMMERS (1986): "Reporting Errors and Labor Market Dynamics", *Econometrica*, 54, 1319-1338.

ROBINSON, P. (1982): "On the Asymptotic Properties of Estimators of Models Containiong Limited Dependent Variables", *Econometrica*, 50(1), 27-41.

SHAW, K. (1989): "Life Cycle Labor Supply with Human Capital Accumulation", *International Economic Review*, 30(2), 431-57.

SOLON, G. (1988): "Self-selection Bias in Longitudinal Estimation of Wage Gaps", *Economics Letters*, 28, 285-290.

SOLON, G. (1989): "The Value of Panel Data in Economic Research", in *Panel Surveys*, D. Kasprzyk, G. Duncan, G. Kalton and M.P. Singh (eds.), John Wiley & Sons, New York.

STAFFORD, F. (1986): "Forestalling the demise of Empirical Economics: The Role of Microdata in Labor Economics Research", Ch.7 in *Handbook of Labor Economics,* Vol.I, O.Ashenfelter and R.Layard eds., North-Holland, Amsterdam

WHITE, H. (1982): "Instrumental Variables Regression with Independent Observations", *Econometrica*, 50, 483-99.