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Working Paper

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Working Paper, No. 2000:14

Provided in Cooperation with:

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Suggested Citation: Hallberg, Daniel (2004) : Does Swedish Time-Use Data from 1984 and 1993 Support the Unitary Model?, Working Paper, No. 2000:14, Uppsala University, Department of Economics, Uppsala

This Version is available at:

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Does Swedish Time-Use Data from 1984 and 1993 Support the Unitary Model?*

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June 22, 2004

Abstract

This paper studies the household's decision to supply labor and tests if the so-called unitary model holds. What is subject to a test is the resulting symmetry of the Slutsky matrix, i.e., that the compensated cross-wage effects are equal and that income is pooled in the household. The test uses Swedish time budget data from 1984 and 1993. Contrary to most other studies from other countries, the conclusion is that symmetry cannot be rejected. There are, however, some evidence suggesting that some family types behave more in accordance with the unitary model than other family types, since income pooling is rejected for families with preschool children, but not for other household types.

Keywords: Time-use data, Household behavior, Slutsky matrix, income-pooling, Household labor supply.

JEL Classification: D13, J22.

*I thank Anders Klevmarken, Matz Dahlberg, Sven-Olof Daunfeldt, Lennart Flood, Magnus Wikström, Matias Eklöf and seminar participants at Uppsala and Stockholm Universities for helpful comments and suggestions on this paper. To access the data used in this article, please consult the internet address www.handels.gu.se/econ/econometrics/hus/husin.htm.

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1 Introduction

The unitary model (Becker, 1965) has been the standard tool for analysis in the study of household decision models that explain differences in time used for household work, leisure and market work. The central assumption in this model is that there is a common household preference structure and a common family budget, which implies that the household is assumed to act as one single economic agent. As is well-known, this is problematic and has been discussed by many authors, one objection being that it is incoherent with elementary theory to a priori assume that separate individuals have a common set of preferences. Equally questionable is the result that the allocation of consumption and time supplies are independent of the personal budget of a member of the household since, in this model, all incomes are pooled into the family's common budget.

Even though the unitary model can be justifiably criticized, it has been the common theoretical framework in the analysis of the family, especially in an empirical analysis of household labor supply (see, e.g., Blundell and Walker, 1986), but also for consumption behavior (Browning and Meghir, 1991, Banks, Blundell and Lewbel, 1997, and Blundell and Robin, 2000).¹ One reason seems to be that the unitary model, compared to alternative models, is somewhat less demanding in terms of the data needed. Because of the income pooling, income data at the household level is enough. Another reason is that the unitary model is easier to implement than its main competitor in the household model literature, namely the collective model. The collective model originates from the bargaining framework (Chiappori, 1988, 1992, Apps and Rees, 1997). In this setting, each household member has his or her own set of preferences and budget and contributes in the production of a household specific good.²

¹Several empirical studies have rejected the unitary model (see, for instance, Lundberg, 1988, Schultz, 1990, Thomas, 1990, Kawaguchi, 1994, Fortin and Lacroix, 1997, and Dahlberg, 1997). There are, however, also some older studies that have not rejected the model (Ashenfelter and Heckman, 1974, and Horney and McElroy, 1980).

²The collective model has been subject to empirical tests, see e.g. Browning, et al. (1994), Fortin and Lacroix (1997), and Aronsson, et al. (1999), and the prevailing conclusion seems to be that the collective model is not rejected in favor of the unitary model.

Just because the unitary model has been rejected at an aggregate level, it does not exclude the possibility that some household types more than other behave according to the unitary model. Most empirical tests have concerned the aggregate. Lundberg (1988) estimated labor supply functions for couples in a simultaneous equation system. One finding was that parents with young children act differently than couples with older children/no children. Husbands and wives with no preschool children act like two separate individuals – the labor supply decision is not simultaneous. Parents in families with young children, however, show strong interaction. Their supplies interact positively. Interesting results can be found for Sweden that contradict the previous rejections of the unitary model. Using Swedish time-use data (moreover the same data source as the present study), Aronsson, et al. (2001) found that income-pooling could not be rejected in 1984, as was the case in 1993. Using a survey study consisting of 489 face-to-face interviews with men and women in 1998, Nyman (2002) also indicated that the source of income was of little importance for personal consumption and spending. One interpretation of these results might be that Swedish men and women are co-breadwinners to a great extent. Women’s earnings gap relative their husbands is rather narrow compared to that in other countries. Swedish men and women have a relatively high degree of economic independence, pool money, and say that they share control over financial decisions to a larger extent than men and women in other countries (Ahrne and Roman, 1997, and Nyman, 1999). When studying the synchronous time use of parents, Hallberg (2002) found that parents of small children in Sweden have equally many hours of synchronous leisure as nonparents. This is contrary than what Hammermesh (2000) and van Velzen (2001) found for the US and the Netherlands, respectively. Parents in Sweden rather perform other types of activities simultaneously, compared to nonparents. Again the gender equality and the high labor force participation of women in Sweden might play an important role for the difference between Sweden and other countries.

While the focus of many studies testing implications of the unitary model (see the above studies) has been on the income-pooling hypothesis, not many

test the second of the two testable implications consistent with the model (see, e.g., Kawaguchi, 1994), namely the symmetry of the Slutsky substitution matrix. This is done in the present study with regard to the labor supplies of the husband and the wife, together with a test of income pooling. It can be noted that failing to reject symmetry indicates that the unitary model and the collective model are observational equivalent since, in the latter, the compensated cross wage effects can be symmetrical as a special case (see Chiappori, 1988 and 1992, Apps and Rees, 1997, proposition 1, pp 183). Hence, for some situations a simpler theoretical framework as the unitary model might be preferable. In order to test the collective model, one must make (non testable) identifying assumptions, i.e. find instruments for the sharing rule.

I organize this paper as follows. In Section 2, I outline a standard unitary model of the household's decision to allocate time. Section 3 presents the empirical specification. Section 4 describes the data. Empirical results are presented in Section 5, and concluding remarks are found in Section 6.

2 The model

A household has two members, a husband (m) and a wife (f). In this model, household members consume market goods, a household produced good, and leisure (cf. Gronau, 1977). In more labor supply oriented versions of this model, household production is not explicitly recognized. Although the extension with household production provides great advantages in understanding how nonmarket time is allocated between various time uses, its use in understanding the supply of market time is not beyond that of the conventional labor supply model.³ We can substitute the household production process, formulated as $Z = Z(h_m, h_f)$ ⁴, into household utility,

³See, for instance, the discussion in Killingsworth and Heckman (1986). However, as Killingsworth and Heckman noted, it is possible that the household production model provides insights into the labor force participation decision that are not apparent in the conventional model. The model shown here assumes an interior solution, i.e., that both spouses work. Gronau's (1977) household production model also modeled the labor force participation decision.

⁴For simplicity, no market goods are used in household production.

formulated as $U = U(x, Z, l_m, l_f)$, and get a combined objective function,

$$U = U(x, Z(h_m, h_f), l_m, l_f) \equiv \Psi(x, h_m, l_m, h_f, l_f), \quad (1)$$

where x is a household composite good, l_s , $s = m, f$, is leisure, and h_s is household work. The maximization problem will instead be over this objective function, subject to constraints. It is obvious that the optimal allocation between various time uses (and consumption goods) will depend on both preferences and production technologies, but the identification of the household production process is usually impossible since the home production output is generally unobserved. The empirical use of the household production model is therefore rather limited without further assumptions about the preference structure.⁵

Note also that it is sometimes comfortable to rearrange the budget constraint, $x = \tilde{w}_m n_m + \tilde{w}_f n_f + \tilde{\mu}$, and the time constraint, $T = n_s + h_s + l_s$, into the *full income constraint*,

$$F \equiv (\tilde{w}_m + \tilde{w}_f) T + \tilde{\mu} = x + \tilde{w}_m (h_m + l_m) + \tilde{w}_f (h_f + l_f) \equiv E, \quad (2)$$

where n_s , $s = m, f$, is time spent in market work, \tilde{w}_s is the marginal (after-tax) hourly wage rate, $\tilde{\mu}$ is the household after-tax nonlabor income, and T is total available time (the price of the market goods is set to 1, and \tilde{w}_s and $\tilde{\mu}$ are considered exogenous). The interpretation is that we can think of time as a commodity that is sold (in which case it is called market work) as well as consumed (in which case it is called leisure or household work). The household's full income (F) is spent entirely on goods and nonmarket time, which summed together equal the expenditure (E) of the household.

Two substantial observations can be made about the resulting demand functions. Let the vector

$$\mathbf{q} = \mathbf{q}(\tilde{w}_m, \tilde{w}_s, p, \tilde{\mu})$$

⁵Suggestions for overcoming the identification problem can be found in Gronau (1977), Graham and Green (1984). Cf. Pollak & Wachter (1975) and Kooreman & Kapteyn (1987b), Fitzgerald, Swenson and Wicks (1996).

express the derived Marshallian demand functions, where $\mathbf{q} = \{x, h_m, l_m, h_f, l_f\}$. Suppose that the after-tax nonlabor income of the household at least partly consists of individual components and is defined as

$$\tilde{\mu} = \tilde{Y}_m + \tilde{Y}_f + b,$$

where \tilde{Y}_s is the nonlabor taxable income after tax of spouse s and b is the household nonlabor (taxable or nontaxable) income (i.e. benefits), $s = m, f$. First, note that the source of nonlabor income is irrelevant for demand, hence income is pooled, which can be written as

$$\frac{\partial q}{\partial \tilde{Y}_m} = \frac{\partial q}{\partial \tilde{Y}_f}, \text{ for all } q. \quad (3)$$

Second, note that the Slutsky substitution matrix, i.e. $\left(\frac{\partial^2 E(\theta, u)}{\partial \theta_j \partial \theta_k}\right)$, is symmetric and negative definite. Here, $E(\theta, u)$ is the expenditure function of the household, θ is the price vector of \mathbf{q} and u is a fixed level of utility.

3 Empirical specification

In this section, I present an empirical specification that enables me to estimate the demand for nonmarket time (or labor supply) of the spouses and market goods and test symmetry.

Studies by Kooreman and Kapteyn (1986, 1987a), and Fortin and Lacroix (1997) showed that allowing for a certain degree of flexibility was desirable, since this permitted the labor supply functions to be backward bending in some ranges of the wage rate and forward bending in others. At the same time, the functional specification must be consistent with economic theory. The trans-log specification does not a priori impose symmetry and is quite flexible. However, for the parameters to have an economic interpretation in terms of belonging to an expenditure function, symmetry is required.⁶

Specification is based on the aggregate of the nonmarket time of the spouses, and the aggregate of market good consumption, since household

⁶For a review of the trans-log, see Chung (1994).

production technologies are not modeled separately from household preferences.⁷ The objective function of the household can be seen as a mixture of both, as in function (1). Let the aggregate of the nonmarket time of spouse s be $l_s^* = h_s + l_s$.⁸

The expenditure function is

$$E(\theta, u) = \mathbf{h}(\theta, u)' \theta,$$

where $\mathbf{h}(\cdot, \cdot)$ is the Hicksian demand vector, θ is the price vector, and u is a fixed level of utility. Given that the household acts cost minimizing, the expenditure function gives all relevant information about the objective function (1). When (1) is linearly homogeneous, the expenditure function can be written as

$$E(\theta, u) = e(\theta)u,$$

where $e(\cdot)$ is the unit expenditure function. (Below, I propose a test of the homothetic preference assumption.) I specify the log of the expenditure function as a trans-log:

$$\begin{aligned} \ln E(\theta) &= \ln e(\theta) + \ln u \\ &= \delta_{00} + \sum_{j=1}^3 \delta_j \ln \theta_j + \frac{1}{2} \sum_{j=1}^3 \sum_{k=1}^3 \gamma_{jk} \ln \theta_j \ln \theta_k + \ln u, \end{aligned} \quad (4)$$

where $\theta = \{\tilde{w}_m, \tilde{w}_f, p\}$ is the price-vector, consisting of the male and the female marginal (after-tax) wage rates and the price of market good consumption, p . Good 1 is hence the aggregate nonmarket time of the husband (l_m^*), good 2 is the aggregate nonmarket time of the wife (l_f^*), and good 3 is the aggregate market good consumption of the household (x). For this

⁷Aggregation is necessary for identification of the household's expenditure function since all nonmarket time for spouse s have the same price (\tilde{w}_s). The aggregation approach is also taken in other studies (for instance in Fortin and Lacroix, 1997).

⁸Aggregation of nonmarket time might wipe out interesting effects, however. When estimating a model of disaggregate time-use, Kooreman and Kapteyn (1987a) found relatively strong wage and income effects for some of the disaggregate activities, but much smaller for the aggregate of nonmarket time.

function, it is required that

$$\sum_j \delta_j = 1, \quad (5)$$

$$\sum_k \gamma_{jk} = 0, \quad (6)$$

$$\gamma_{jk} = \gamma_{kj}, k \neq j. \quad (7)$$

Here (5) and (6) correspond to "adding-up" restrictions, and (7) to the symmetry restriction of the Slutsky matrix.⁹

The set of expenditure share demand functions that minimizes expenditure is obtained by differentiating (4) logarithmically:

$$B_i^{(j)} = \frac{\theta_{ij} q_{ij}}{F_i} = \delta_{ij} + \gamma_{j1} \ln \tilde{w}_{mi} + \gamma_{j2} \ln \tilde{w}_{fi} + \gamma_{j3} \ln p; j = 1, 2, 3; \quad (8)$$

where q_{ij} is the quantity consumed of the j :th good, θ_{ij} is its price, and F_i is full income. Here $i = 1, \dots, N$ is introduced to signify the household. In order to allow for heterogeneity across households, I assume that the intercept δ_{ij} is of a specific form:

$$\delta_{ij} = \alpha_j + \beta_j' \mathbf{x}_i + \varepsilon_{ij}; j = 1, 2, 3; \quad (9)$$

where \mathbf{x}_i is a vector of independent demographic variables, β_j is a parameter vector, α_j is an intercept, and ε_{ij} is a stochastic component, all belonging to household i and budget share j . Here, $\varepsilon_i = \{\varepsilon_{i1}, \varepsilon_{i2}, \varepsilon_{i3}\}$ is assumed to be independent and to follow a well defined (multivariate) distribution with zero mean and variance-covariance matrix Σ . Although ε_{ij} has been introduced to represent random preference/technology variation, it is clear that it can also comprise other sources of random variation in the budget shares, such as measurement errors, optimization errors, and so on.

⁹ Adding-up follows from the fact that budget shares must sum to 1. The off-diagonal elements of the Slutsky matrix are $\frac{\partial^2 E(\theta, u)}{\partial \theta_j \partial \theta_k} = \gamma_{jk} \frac{E}{\theta_j \theta_k} + \frac{q_j q_k}{E}$, $j \neq k$. For symmetry $\frac{\partial^2 E(\theta, u)}{\partial \theta_j \partial \theta_k} = \frac{\partial^2 E(\theta, u)}{\partial \theta_k \partial \theta_j}$, $j \neq k$, it is required that $\gamma_{jk} = \gamma_{kj}$, $j \neq k$.

Inserting (9) into (8) yields

$$B_i^{(j)} = \frac{\theta_{ij} q_{ij}}{F_i} = \alpha_j + \gamma_{j1} \ln \tilde{w}_{mi} + \gamma_{j2} \ln \tilde{w}_{fi} + \gamma_{j3} \ln p + \beta_j' \mathbf{x}_i + \varepsilon_{ij}; j = 1, 2, 3. \quad (10)$$

The “adding-up” restrictions now reads

$$\begin{cases} \sum_j \alpha_j = 1, \\ \sum_j \gamma_{j1} = \sum_j \gamma_{j2} = \sum_j \gamma_{j3} = \sum_j \beta_j = \sum_j \varepsilon_{ij} = 0, \\ \gamma_{j1} + \gamma_{j2} + \gamma_{j3} = 0; j = 1, 2, 3. \end{cases} \quad (11)$$

As usual, “adding-up” implies that one share equation is redundant. I therefore drop the budget-share equation for the market goods in estimation ($j = 3$). The market good price (p) does not vary across households, which can be handled by setting it to 1 (thus interpreting market goods in terms of expenditure and male and female marginal wage rates as real marginal wage rates). This means that we have a system consisting of 2 equations to estimate.

Note that symmetry is not imposed a priori. The symmetry of the Slutsky matrix implies

$$\begin{cases} \gamma_{12} = \gamma_{21}, \\ \gamma_{13} = \gamma_{31}, \\ \gamma_{32} = \gamma_{23}. \end{cases} \quad (12)$$

It is sufficient to test $\gamma_{12} = \gamma_{21}$ (since the other two follow due to “adding-up”).

It is straightforward to test the assumption of linearly homogeneous preferences empirically.¹⁰ Note that under homothetic preferences, the indirect utility function V can be written as

$$V(\theta, F) = v(\theta) F,$$

¹⁰The assumption of homothetic preferences may not be too critical in view of the empirical results found in Fitzgerald, Swenson and Wicks (1996). They used interviews to evaluate the outputs from several home production processes so that the household production function parameters could be directly estimated. One finding was that most home production processes followed constant returns to scale.

where F is full income. Roy's identity implies that the Marshallian income share for good j is $\theta_j q_j(\theta, F) / F = \theta_j q_j(\theta) F / F = \theta_j q_j(\theta)$. Hence, if the objective function is linearly homogeneous, then the Marshallian income shares are independent of income. Since the Marshallian income shares are observationally equivalent to $B^{(j)}$, $j = 1, 2, 3$, the assumption of homothetic preferences can be tested by examining if the shares $B^{(j)}$ are independent of full income.¹¹

Income is "pooled" in the household if (3) holds for all j . Empirically, I implement this test by including second order polynomials of individual non-labor incomes after tax, \tilde{Y}_m and \tilde{Y}_f , in the equations for $B^{(1)}$ and $B^{(2)}$ (hence treating the shares as Marshallian income shares rather than expenditure shares). If income pooling holds, then $\frac{\partial B^{(1)}}{\partial \tilde{Y}_m} = \frac{\partial B^{(1)}}{\partial \tilde{Y}_f}$ and $\frac{\partial B^{(2)}}{\partial \tilde{Y}_m} = \frac{\partial B^{(2)}}{\partial \tilde{Y}_f}$.

Possible endogeneity of the after tax wage rate to hours of work is addressed by adopting a straightforward instrumental variable (IV) technique.¹² As usual, the choice of instruments is vital. Following Blomquist (1996), I use polynomials of the gross wage rate, taxable nonlabor income, and nontaxable income as instruments. Blomquist hence treated the gross wage rate as exogenous and I will do the same. The system of budget shares is only applied to two-earner households, which might introduce sample selection bias. This is corrected for by including appropriate inverse mills-ratio terms (λ_{mi} and λ_{fi}) as additional right-hand side variables and estimating the model by a two-step procedure. The work decisions of the spouses are probably correlated, so a bivariate probit is used in the first step. This procedure is quite standard, and details can be obtained from the author upon request. Because stochastic terms might correlate across equations and since I adopt IV techniques, the system will be estimated with the three-stage least squares (3SLS) estimator.

¹¹The test is implemented by including the log of full income among the right-hand side variables. In this specification, the other parameters of the share equations should not be interpreted as derived from an expenditure function, however.

¹²Full income and individual nonlabor incomes after tax may be endogenous for the same reason as the marginal (after-tax) wage rate. The tests of homothetic preferences and of income pooling (see below) deals with this by also instrumenting the full income and individual nonlabor incomes after tax.

4 Data

Data are taken from the Swedish household panel study Household Market and Nonmarket Activities (HUS)¹³. The 1984 and 1993 waves of data include time-use surveys besides the conventional surveys. The 1984 main survey was a random sample from the Swedish population at the end of 1983 consisting of 2619 individuals aged 18-74. The 1993 main interview included 4137 individuals and covered the panel from 1984 and random renewal samples from 1986 and 1993. The time-use diary was collected by telephone as yesterday recall diaries. This was done for two randomly selected days, stratified such that one day was a weekday and the other a weekend day. In two-spouse households, the measurement days were harmonized so that both spouses were interviewed concerning the same day. The number of respondents in the 1984 (1993) time-use survey distributed over questionnaires was 2552 and 2468 (3249 and 3175), of which 2438 (3009) individuals participated in both questionnaires. 1060 persons participated in both the 1984 and the 1993 time-use study.

One advantage of HUS is that both spouses have been interviewed. I restrict the sample to married and cohabiting couples with and without children and to families where both spouses participated in at least one of the time-use surveys and in the corresponding main interview. The sample is also restricted to households where both spouses are reported to be employed in the main survey and where we observe the after tax wage rates (or can derive them from earnings data and information on normal working hours reported in the main survey). In 1984 (1993), the sample consists of 354 (416) households, which contributed 672 (678) time-use observations, 352 (366) of which concerned a weekday (Mon.-Fri.).¹⁴

¹³For further details about HUS, see Klevmarken and Olovsson (1993) and Flood, Klevmarken and Olovsson (1996).

¹⁴The same household could have participated in the time-use survey in all four waves. I will, however, not explore the panel-structure of the data. With the sample restrictions used, the cross-year panel becomes too small (74 households).

4.1 Dependent variable

The market work definition includes work breaks, business travels, official job-related travels, applying for work and military service, and traveling time to and from work. The remainder time during a day is the time use for nonmarket activities. The budget shares are computed as

$$B_i^{(1)} = \frac{\tilde{w}_{mi}l_{mi}^*}{F_i}, \text{ and } B_i^{(2)} = \frac{\tilde{w}_{fi}l_{fi}^*}{F_i},$$

where l_{si}^* is total nonmarket time for spouse $s = m, f$. Full income for household i , F_i , is estimated as

$$\hat{F}_i = 24(\tilde{w}_{mi} + \tilde{w}_{fi}) + \tilde{\mu}_i. \quad (13)$$

We can consider full income as the household's hypothetical income defined by linearizing the budget around observed marginal tax rates.¹⁵

It would be preferable to have access to each household's market good consumption, x , during exactly same time interval as the time-use interview, but this is unfortunately not recorded in the data. The second-best solution seems to be to determine x residually from the full income concept, according to (13). This procedure will give a measure of consumption, but it has obvious drawbacks.¹⁶ It may therefore be more accurate to interpret x in terms of the household's real after-tax income (p is arbitrarily set to 1), rather than the goods actually consumed during 24 hours.

The marginal, after-tax, wage rates were computed using the marginal tax-rates evaluated at annual working hours. For almost everyone in the

¹⁵Because marginal (after-tax) wage rates are fixed at their observed levels, full income is only a rough approximation of the true income the household would have earned if both spouses actually did spend 24 hours in paid labor. Most individuals experience a higher marginal tax rate as they supply more labor since earnings increase, but are also paid over-time premiums over some threshold hours. Increased earnings also affect transfers.

¹⁶For example, my estimate of consumption is probably more jumpy over single days than is true consumption, since the budget is forced to add-up. The justification for the assumption that the budget constraint is satisfied is that the two days were selected randomly. Hence, the time-use observed during a day is an approximation of the average time-use on an annual basis (multiplied by 365), and the measurement error due to this approximation is additive with the expected value zero.

sample (87%) the gross wage rate was calculated as monthly earnings divided by what the respondent reported in the main survey as being his or her *normal* working hours. This procedure should, however, not invalidate the gross wage rate as an instrument since normal working hours and the time-use response (during weekdays) often deviate, sometimes substantially (especially for women).¹⁷

The household nonlabor income net of taxes ($\tilde{\mu}$) consists of the husband's and the wife's individual nonlabor taxable income after tax, \tilde{Y}_m and \tilde{Y}_f , and nontaxable family benefits, b . Individual nonlabor taxable income Y_s is defined as the sum of individual interest incomes, interest subsidies, dividends, and capital gains less capital losses, interest on debts and administrative expenses, and other income that was not income from work. I applied different definitions of \tilde{Y}_s each year, since there was a change in the taxation rules between 1984 and 1993.¹⁸ The household nonlabor income net of taxes during one day was then computed as $\tilde{\mu} = (\tilde{Y}_m + \tilde{Y}_f + b) / 365$, and full income as (13).¹⁹ For the 1984 (1993) sample, the mean and standard deviations of full income of the household during a day are 1085 SEK and 239 SEK (2587 SEK and 694 SEK).

¹⁷The correlation between normal working hours and reported working hours from the time-use survey is weak. It ranges from 0.19 to 0.28 for males in 1993 and females in 1984, respectively. Unfortunately, measurement errors in normal working hours will bias the calculated hourly wage rate (for instance, the reported normal working hours for males are accumulated at exactly 40 hours per week). Some of the deviation may also depend of not including work breaks, jobtravels, etc. in the normal work time.

¹⁸Before the 1991 tax-reform, capital income, other types of nonlabor taxable income, and earnings from work were taxed jointly, i.e., the marginal tax rate was determined by the sum of all three (after deductions). In 1984, I therefore computed the total tax-rate (based on the sum of all three incomes) and multiplied this with Y_s to generate \tilde{Y}_s . After the tax reform, capital income and earnings from work were taxed separately. Other nonlabor taxable income – that was not capital income – was still jointly taxed with earnings from work, however. For this type of income, I used the total tax-rate (based on all taxable income exclusive of capital earnings) in 1993.

¹⁹To guarantee positive good consumption, $\tilde{\mu}$ was truncated for negative numbers. This means that the time-shares (consumption share) may therefore be underestimated (overestimated) to some extent.

4.2 Explanatory variables

The variables used to control for heterogeneity are age dummies, a university degree dummy (UNIVDEG), years of schooling (YSCHOOL) and self-employment (SELFEMP), included for both the husband and the wife. I also include age difference of the spouses (AGEDIFF), whether living in a single family house (OWNHOME), the number of children in age brackets 0-2, 3-6 and 7-18 (CH0-2, CH3-6, CH7-18), If the family uses outside home day care (DAYCARE), if the time-use response concerned a Saturday or a Sunday (WEND), and Season indicator variables of time-use interview.²⁰

As instruments for the log of after tax wage rate, I use the third-order polynomial of the log of observed gross wage, nonlabor taxable income (Y_s), nontaxable family benefits (b) and their interactions. Employment probability is determined by age, age squared, university degree, number of children in the household (NCHILD), dummies indicating the youngest child's age (YCH0-2, etc.), if there is any disabled person in the family (DISABILITY), the household's nonlabor income before tax (HNLINC) and the characteristics of the spouse (hence, $HNLINC = Y_m + Y_f + b$). Estimates of the bivariate probit model is presented in the Appendix, Table 6.

Descriptive statistics of the samples used in the estimation are given in Table 5, also in the Appendix. Some comments are needed. As a result of the tax-reform in 1991, the after tax wage rate compression is much less apparent in the 1990s than in the 1980s. This is obvious from looking at the changes in the marginal tax. In my sample for 1984, men's average after-tax wage rate was actually lower than that of women. The 1984 and 1993 samples differ in some other respects as well; a greater share in the 1993 sample has obtained a university degree and more women work as self-employed. I also note that the mean age is about 3 years higher in the second year.

²⁰In 1984, DAYCARE includes private and municipal day care, preschool and after-school programs. In 1993, DAYCARE does not include private day-care and after-school programs.

5 Empirical results

5.1 Regularity, homothetic preferences, and model diagnostics

Before examining the estimation results one must check whether the estimated expenditure function is consistent with the underlying theory, as the estimated expenditure function must meet two regularity conditions: monotonicity (positivity of the fitted budget shares) and concavity. The expenditure share system was estimated separately for weekdays (Mon.-Fri.) and weekend days (Sat.-Sun.) each year. I find that for the weekday responses, nearly all observations in both years satisfy the concavity requirement; about 2.8 percent (1.4 percent) did not for 1984 (1993). Monotonicity was never violated. For weekend days, and especially in the first year, there were, on the other hand, serious problems fitting a regular expenditure function. In the 1984 sample for weekend day responses, none of the observations fulfilled concavity. For weekend days in the second year, 12.5 percent of the observations did not satisfy the concavity requirement. There were also some observations in each year (3.1 percent for 1984 and 1.6 percent for 1993) for which the fitted budget shares were negative. The conclusion from these tests is hence that the model seems acceptable for weekdays but not for weekend days. I will therefore henceforth concentrate on weekdays.²¹

By including the log of full income in the share equations, I test the assumption of a linearly homogeneous preference/technology structure. As noted above, if the preference structure is homothetic, the log of the full income variable should not contribute significantly in any of the share equations. There is, however, an indication that the underlying preference/technology structure was not homothetic in 1984. For the weekday responses, the $\chi^2(2)$ -value of the hypothesis that the log of the full income variable is jointly zero is 14.2 and 2.4 for the 1984 and 1993 samples, respectively (the critical value

²¹This may not be unexpected. Not many people take up market work during Saturdays and Sundays, which have the implication that nonmarket time (l_s^*) is censored at 24 hours. The budget share, i.e. $B_s = \frac{\tilde{w}_s}{F} l_s^*$, is not censored at any particular value, however, due to the variation arising from \tilde{w}_s and F . The model I apply, however, rests on the assumption of an interior solution.

is 6.0 at the 5 percent level). The empirical specification seem thus not quite appropriate for 1984, but nevertheless performs well for 1993 altogether.²² I have, nevertheless, chosen not to terminate the testing sequence for that year, but some caution is needed when interpreting these results.

First-stage regressions reveal that the instruments are highly correlated with the endogenous variables, see Table 1, which only include Monday-Friday responses. In general, the explanatory power with respect to the after-tax wage rates is best for the 1993 sample. One explanation for this might be the simplifications of the tax-system introduced by the 1991 tax-reform. Model diagnostics also reveal some form of heteroscedasticity and model misspecification (see RESET and HET in the table). I did not try any alternative model formulation since the budget share equations passed the RESET test.

An augmented regression test for endogeneity (Durbin-Wu-Hausman test) shows that the husband's marginal (after-tax) wage rate should be considered to be endogenous to his wife's labor supply in 1984.²³ Sargan's test of overidentification indicates, however, that the chosen set of instruments is not valid for women in 1984, i.e., that the instruments are not independent of the structural error term. However, the instrument set is valid for the other groups since they pass this test. For men's labor supply in 1984, I find that the marginal wage rates are not endogenous, given the instruments, while I can reject the hypothesis of exogenous marginal wage rates for men and women in 1993 at the 10 percent level.

These tests hence suggest that there are some problems with the chosen model specification for 1984. Some caution is therefore needed when interpreting the results for that year.²⁴

²²Tests were performed without the symmetry constraint imposed.

²³Both this test and the Sargan test were performed separately for each equation.

²⁴A suggestion – which was not explored – would be to model the average demand of nonmarket time by using a weighting procedure of the time-use responses (5/7 for weekdays and 2/7 for weekend days). This might be questionable if parameters are not stable over the two types of response days, which they seem not to be.

Table 1: First-stage diagnostics, Sargan's test for overidentification, and Durbin-Wu-Hausman test for endogeneity (Mon.-Fri. responses)

	1984, N=352				1993, N=366			
	$B^{(1)}$	$B^{(2)}$	$\ln \tilde{w}_m$	$\ln \tilde{w}_f$	$B^{(1)}$	$B^{(2)}$	$\ln \tilde{w}_m$	$\ln \tilde{w}_f$
R^2	0.33	0.30	0.60	0.64	0.28	0.33	0.75	0.81
RESET	1.03	1.25	1.97	1.68	0.99	1.14	2.10	1.52
p -value	0.43	0.12	0.00	0.00	0.50	0.23	0.00	0.01
HET	67.26	31.50	180.94	100.39	57.65	50.07	56.91	159.16
p -value	0.03	0.97	0.00	0.00	0.12	0.32	0.13	0.00
Sargan ^a	25.21	48.14			20.79	23.99		
p -value	0.15	0.00			0.24	0.12		
# overidentifying restrictions, r	19	19			17	17		
Augmented regression test for endogeneity (Durbin-Wu-Hausman test):								
$\ln \tilde{w}_m$	0.91 ^b	2.99 ^b			0.74 ^b	1.00 ^b		
$\ln \tilde{w}_f$	0.68 ^b	1.60 ^b			1.79 ^b	2.22 ^b		
F, all exogenous	0.48	7.72			2.59	2.48		
p -value	0.62	0.00			0.08	0.09		

Note: HET is the Beusch-Pagan test for heteroskedasticity. a) Sargan is $(N - k) R^2 \sim \chi^2(r)$, where N is the number of observations, k is the number of parameters is the budget share, R^2 is taken from the OLS regression of the IV residual on the instrument set, and r is the number of overidentifying restrictions. b) T -ratios of first stage residuals included in the structural model (see Davidson and MacKinnon, 1993).

Table 2: Estimated parameters and Wald test of symmetry of cross wage effects (Mon.-Fri. responses)

<i>Parameter estimate</i>	1984		1993	
γ_{11}	0.154	(0.024)	0.089	(0.020)
γ_{12}	-0.140	(0.023)	-0.176	(0.020)
γ_{21}	-0.175	(0.037)	-0.166	(0.020)
γ_{22}	0.180	(0.024)	0.159	(0.024)
<i>Hypothesis test: $\gamma_{12} = \gamma_{21}$</i>	1984		1993	
<i>Wald1</i>	0.81	[0.37]	0.09	[0.77]
<i>Wald2</i>	0.64	[0.42]	0.08	[0.78]
<i>Wald3</i>	0.62	[0.43]	0.12	[0.73]
<i>Wald4a</i>	0.78	[0.38]	0.00	[0.96]
<i>Wald4b</i>	4.96	[0.03]	0.44	[0.51]

Note: Heteroscedastic-consistent standard errors in parenthesis, p-values of null $\gamma_{12} = \gamma_{21}$ in brackets. Full estimation results given in Tables 7 and 8 in the Appendix.

a) Families with preschool children (0-6), b) families without preschool children (0-6).

5.2 Symmetry test and income pooling test

The symmetry test is performed separately for the weekday responses in each year. *Wald1* in Table 2 refers to these tests. Estimates of the γ s are given in the same table. Wald tests of symmetrical cross wage effects suggest that symmetry cannot be rejected (the critical $\chi^2(1)$ value on 5 per cent is 3.84).²⁵

It is possible that the inference from Table 2 is slightly misleading due to the heteroscedasticity introduced by the self-selection correction. A test of the joint hypothesis that all λ -terms are zero shows that the null for any

²⁵As noted above, endogeneity of marginal (after-tax) wage rates may be less of a problem for 1993 as compared to 1984, which might result in an inefficiency of the IV-estimator for the second year. Therefore, I have performed the symmetry test using estimates from a seemingly unrelated regression model (SUR), without instrumenting the net wage rates. The hypothesis of equal cross-wage effects could not be rejected in this case either.

Table 3: Wald tests of the income pooling hypothesis (Mon.-fri. responses)

	1984		1993	
Families with preschool children	1.299	[0.522]	7.521	[0.023]
Other families	1.268	[0.531]	2.690	[0.261]
All	1.110	[0.574]	7.084	[0.029]

χ^2 -values, p -values in brackets, of the joint null
 $\frac{\partial B^{(j)}}{\partial \tilde{Y}_m} = \frac{\partial B^{(j)}}{\partial \tilde{Y}_f}$, both $j = 1$ and $j = 2$. \tilde{Y}_m and \tilde{Y}_f are included as
second-order polynomials in the equations for $B^{(j)}$.

of the samples cannot be rejected (the $\chi^2(4)$ -value is 2.96 and 3.32 for the 1984 and 1993 samples, respectively, with the critical value being 9.49 on the 5 percent level). Excluding all λ -terms and re-estimating the system shows that the same conclusion still holds: the hypothesis of symmetrical cross-wage effects cannot be rejected. *Wald2* in Table 2 relates to these tests. In cross-sectional data, one might also be concerned about a general form of heteroscedasticity. The system has, therefore, been reestimated using White's (1980) heteroscedastic consistent variance-covariance matrix. *Wald3* in Table 2 refers to these tests. Again, the symmetry condition cannot be rejected.

On an aggregate level, the evidence for 1993 seem to suggest that families do not pool income. When including all family types the income pooling hypothesis test, presented in Table 3, row "All", shows that income pooling is rejected in 1993 on the 5 percent level.²⁶ The hypothesis is however not rejected on the 1 percent level, which may be a result of mixing household with differing behavior. In 1984, however, data cannot reject income pooling. Using the same data that I am using, the conclusion in Aronsson, et al. (2001) was that the data provided some empirical support for the more general collective model. They could, just as I, not reject the income-pooling

²⁶For a fraction of the sample, income was negative. The sign of the second order term (\tilde{Y}_s^2) was reversed when $\tilde{Y}_s < 0$.

hypothesis in 1984, as was the case in 1993.^{27,28}

Families with preschool children

It is realistic to assume that the presence of preschool children is an important factor for the joint labor supply decision in a household and the allocation of leisure. Therefore, it seems reasonable to estimate separate models for families with and without preschool children and test if symmetry holds. The hypothesis is then that parents with preschool children aged six years or younger have a tighter time schedule and must organize both work and home time in a different way than families without young children. The specification used so far to control for child heterogeneity only assumes a shift in the intercept – it is assumed that the same γ s applies for both groups. When separate models are estimated (for simplicity, using the same selection model into employment for both groups), I detect no violation of the symmetry condition for parents with preschool children aged six years or younger in any of the years (*Wald4a* in Table 2). However, for families without young children, I find that the symmetry of cross-wage effects does not hold for 1984 (*Wald4b* in Table 2).²⁹ This result seems to confirm Lundberg’s (1988) findings.³⁰

Estimating separate models for families with preschool children and oth-

²⁷None of the variables that Aronsson, et al. (2001) used to determine the sharing rule (difference in husband’s and wife’s years of schooling, age, marginal wage rate and nonlabor income after tax (capital income), and the relative supply of males and females in the marriage market) turned out to be significant for 1984.

²⁸Furthermore, distorting the source of income may be elements of intra-household redistribution of mortgage interest deductions in order to exhaust tax avoidance opportunities. Because of tax legislation, this was probably more frequent in the 1980s than after the tax reform.

²⁹The sample of families with small children aged 0-6 was relatively small; 103 in 1984 and 78 in 1993. In the sample of couples without small children, there are 249 observations in 1984 and 288 in 1993. Pooling 1984 and 1993 and re-running the symmetry test show that symmetry cannot be rejected for families with small children, while it can for those without. For families with small children in 1984 there were problems in fitting a regular expenditure functions: 35 percent of the observations failed to meet the concavity condition that year. But the hypothesis of homothetic preferences could not be rejected (p -value was 0.82), however. Hence, it was for the other group that year that preferences were not homothetic.

³⁰It should be noted that families where one of the parents (usually the mother) is on parental leave for a period longer than 8 weeks are not included in our samples.

ers, I find, however, that it is families without young children that seem to pool income, see Table 3. Although this result seems a bit uncertain, it stands out from the with respect to the symmetry test result above.

5.3 Parameter estimates

Table 4 presents the demand elasticities with respect to the own and cross-prices of nonmarket time and expenditure.³¹ A general conclusion is that own-price (after-tax) elasticities are nonpositive, which should be expected from basic demand theory. In fact, in all cases, the estimated own-price elasticities are negative and significantly different from zero. Male labor supply is more elastic with respect to the own wage rate than female labor supply, especially for 1993. The own-price elasticity of males is higher and more precisely estimated in 1993 as compared to 1984, while that of females remains approximately unchanged. The cross-price effects on labor supply are weak and not significantly different from zero, but somewhat higher in 1993 as compared to 1984. Nonmarket time and market good expenditure (note that the price of market goods is set to unity) seem to be substitutes, i.e., the cross-price elasticity of wage on x is positive. This may, however, be a result of how x was determined (i.e. residually determined from full income). A 1 percent increase in \tilde{w}_m or in \tilde{w}_f increased the household's level of expenditure less in 1984 than in 1993.

A comparison between the sample of parents with small children (bottom of table) and others (not shown) shows that the former group in general seems less willing to give up nonmarket time due to a wage increase than the latter. The former also had stronger cross-price elasticities. (Men in families without children in low ages had a cross-price elasticity of -0.05 and an own-price elasticity of -0.25). This also seems to confirm the results in Lundberg (1988).

From the full estimation results, which are displayed in the Appendix, Table 7 and Table 8, we can conclude that having young children 3-6 years of age shifts women's time away from the market in 1993. For other groups

³¹Since market time of s is $T - l_s^*$, the price elasticities of labor supply are the negative of those presented in that table.

Table 4: Own and cross price elasticities for nonmarket time (l^*) and market good consumption (x), and predicted budget shares (all evaluated at the sample mean)

1984							Mean predicted budget share
	\tilde{w}_m		\tilde{w}_f		p		
l_m^*	-0.19	(0.09) ^b	-0.07	(0.25)	0.26	(0.57)	0.32
l_f^*	-0.06	(0.20)	-0.16	(0.06) ^a	0.22	(0.70)	0.40
x	0.30	(0.12) ^a	0.32	(0.11) ^a	-0.62	(0.19) ^a	0.28
1993							Mean predicted budget share
	\tilde{w}_m		\tilde{w}_f		p		
l_m^*	-0.40	(0.06) ^a	-0.14	(0.26)	0.55	(0.50)	0.34
l_f^*	-0.13	(0.24)	-0.20	(0.06) ^a	0.33	(0.76)	0.36
x	0.61	(0.10) ^a	0.40	(0.10) ^a	-1.01	(0.17) ^a	0.30
Families with preschool children (1984 and 1993 pooled)							Mean predicted budget share
	\tilde{w}_m		\tilde{w}_f		p		
l_m^*	-0.19	(0.05) ^a	-0.16	(0.22)	0.36	(0.66)	0.34
l_f^*	-0.13	(0.17)	-0.13	(0.04) ^a	0.26	(0.82)	0.41
x	0.48	(0.06) ^a	0.42	(0.06) ^a	-0.90	(0.10) ^a	0.25

Note: The own price elasticity is computed as $\hat{\eta}_{ii} = \hat{\sigma}_{ii}\hat{B}_i$ and the cross price elasticity as $\hat{\eta}_{ij} = \hat{\sigma}_{ij}\hat{B}_j, i \neq j$, where $\hat{\sigma}_{ij} = 1 + \hat{\gamma}_{ij}(\hat{B}_i\hat{B}_j)^{-1}, i \neq j$, $\hat{\sigma}_{ii} = (\hat{\gamma}_{ii} + \hat{B}_i(\hat{B}_i - 1))\hat{B}_i^{-2}$, $\hat{\gamma}_{ij}$ is an estimated parameter and \hat{B}_i is the predicted budget share evaluated at the sample mean (see Chung, 1994). Standard errors in parenthesis:^a) significant on 1 percent, ^b) significant on 5 percent. Parameter estimates from separate models, obtained by iterated 3SLS with symmetry imposed.

and years I find no significant effects of children on total nonmarket time (although the point estimates in general are positive). One explanation is probably due to the aggregation of nonmarket time, since the main differences in time allocation between families with and without children are, presumably, how they split their time between household work and private leisure. This interpretation is in line what Hallberg (2002) found. More educated spouses demand less nonmarket time/work more, but the effects are not altogether attuned. Males, but not females, allocate significantly more of their time to nonmarket activities during the summer period than the rest of the year. Parents with preschool children aged six years or younger have a smoother week pattern than others (not shown). As could be expected, the former group allocates a larger share of their full income to home activities. However, I find very few significant effects for this group, which might be because data were pooled from two years.

6 Concluding remarks

In this paper, I have estimated a household labor supply model and tested the implications of the unitary model, namely the symmetry of the cross-wage effects of the spouses, and income pooling. Two Swedish time-use data sets are used, from the middle of the 1980s and the early 1990s. Contrary to the findings of many other studies, the symmetry hypothesis cannot be rejected, implying that the fundamental assumption of a common utility function of the household seems to be consistent with the data. I confirm the results in other studies that have tested income-pooling in Sweden (Aronsson, et al., 2001, and Nyman, 2002). They found indications that Swedish couples might well choose to pool their sources of income into a family budget. In addition to these studies, however, I find some indications that families with preschool children aged six years or younger act differently than others. The behavior of the former group with respect to their labor supplies seems to be less in accordance with the unitary model, as some evidence in this study suggests that the families with preschool children do not pool income, at least not in the 1990s.

There might be more of income pooling and sharing of financial decisions among couples in Sweden than among couples elsewhere. The behavior of some subgroups in society might, however, be best approximated by one type of theoretical model, while others are best described by another type of model. Since the picture is less than uniform on this matter, at least in some countries more than others, empirical studies that exploit this variation between groups (e.g., families with small children) are of interest.

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7 Appendix

Table 5: Decriptive statistics, 1984 and 1993 samples

Variable	1984				1993			
	Mean	Std.Dev	Min	Max	Mean	Std.Dev	Min	Max
<i>Mon.-Fri.:</i>	(N=352)				(N=366)			
Nonmarket time _m	16.198	4.518	2.467	24	15.900	4.158	5.833	24
Nonmarket time _f	19.030	4.250	8.833	24	18.162	4.323	5.25	24
Budget share _m	0.324	0.109	0.051	0.811	0.336	0.096	0.100	0.640
Budget share _f	0.401	0.119	0.147	0.799	0.364	0.103	0.066	0.648
<i>Sat.-Sun.:</i>	(N=320)				(N=312)			
Nonmarket time _m	23.142	2.409	8.250	24	22.739	2.871	9.500	24
Nonmarket time _f	23.087	2.500	12.167	24	23.189	2.411	12.083	24
Budget share _m	0.459	0.092	0.163	0.711	0.486	0.098	0.190	0.912
Budget share _f	0.486	0.095	0.187	0.785	0.460	0.093	0.088	0.672
	(N=672)				(N=678)			
\tilde{w}_m	21.547	5.984	7.384	64.423	55.521	24.462	22.680	422.938
\tilde{w}_f	22.897	7.098	6	115.010	51.126	12.400	4.651	107.094
marginal tax rate _m	55.958	12.299	0	84	42.635	9.848	28.530	53.480
marginal tax rate _f	43.485	11.219	0	77	34.228	6.901	26.700	53.320
gross wage (w_m)	51.534	16.746	17.308	140.550	99.030	41.779	40.385	625.000
gross wage (w_f)	41.411	13.120	10	217	78.501	20.514	6.735	213.846
AGE _M	41.676	10.409	20	66	44.670	9.982	21	65
AGE _F	38.981	9.910	20	62	42.422	9.955	20	65
OWNHOME	0.789				0.820			
CH0-2	0.104	0.315	0	2	0.083	0.275	0	1
CH3-6	0.277	0.541	0	3	0.206	0.485	0	2
CH7-18	0.854	0.925	0	4	0.683	0.905	0	4
DAYCARE	0.310				0.237			
AGEDIFF	2.695	4.340	-22	23	2.248	3.950	-11	21
WEND	0.476				0.460			
MAR-MAY	0.266				0.285			
JUN-AUG	0.223				0.181			
SEP-NOV	0.283				0.263			
AGE1830 _M	0.146				0.117			
AGE3135 _M	0.173				0.074			
AGE4655 _M	0.235				0.347			
AGE56+ _M	0.116				0.147			
UNIVDEG _M	0.137				0.209			
SELFEMP _M	0.054				0.059			
AGE1830 _F	0.223				0.150			
AGE3135 _F	0.174				0.093			
AGE4655 _F	0.190				0.350			
AGE56+ _F	0.063				0.069			
UNIVDEG _F	0.121				0.267			
SELFEMP _F	0.012				0.024			

Table 6: Bivariate probit estimates of job probability

	1984			1993		
	Mean (std.dev.)	z_m	z_f	Mean (std. dev.)	z_m	z_f
AGE _M	46.872 (13.827)	0.135** (0.056)	0.010 (0.035)	48.369 (13.913)	0.246*** (0.043)	0.030 (0.037)
AGE _M ²	2388 (1362)	-0.002*** (0.001)	0.000 (0.000)	2533 (1407)	-0.003*** (0.000)	0.000 (0.000)
YSCHOOL _M	10.639 (3.799)	0.039** (0.018)	-0.001 (0.016)	11.578 (3.804)	0.063*** (0.019)	0.033** (0.016)
UNIVDEG _M	0.104 (0.305)	1.286*** (0.292)	-0.173 (0.168)	0.180 (0.384)	-0.232 (0.169)	-0.122 (0.141)
AGE _F	43.971 (13.475)	0.125*** (0.046)	0.175*** (0.033)	45.616 (13.756)	0.010 (0.042)	0.163*** (0.037)
AGE _F ²	2115 (1253)	-0.001*** (0.000)	-0.002*** (0.000)	2270 (1320)	0.000 (0.000)	-0.002*** (0.000)
YSCHOOL _F	10.185 (3.138)	0.021 (0.025)	0.036* (0.020)	11.486 (3.395)	0.031 (0.020)	0.004 (0.018)
UNIVDEG _F	0.082 (0.275)	-0.080 (0.259)	0.256 (0.197)	0.200 (0.400)	-0.165 (0.152)	0.452*** (0.142)
DISABILITY	0.100 (0.300)	-0.316** (0.148)	-0.601*** (0.125)	0.103 (0.305)	0.000 (0.142)	-0.459*** (0.124)
NCHILD	1.012 (1.092)	-0.087 (0.084)	-0.200*** (0.063)	0.841 (1.076)	0.025 (0.086)	-0.252*** (0.067)
YCH0-2	0.091 (0.288)	0.108 (0.280)	-0.775*** (0.205)	0.111 (0.314)	0.215 (0.216)	-0.553*** (0.201)
YCH3-6	0.150 (0.358)	-0.394 (0.243)	-0.231 (0.192)	0.115 (0.319)	0.146 (0.264)	0.095 (0.204)
YCH7-18	0.314 (0.464)	-0.455** (0.206)	0.119 (0.157)	0.239 (0.427)	-0.037 (0.175)	0.147 (0.168)
HNLINC/10 ⁶	0.029 (0.041)	-14.78*** (1.641)	-3.519*** (1.152)	0.081 (0.105)	-4.259*** (0.540)	-2.342*** (0.461)
CONSTANT		-3.253*** (0.833)	-2.249*** (0.575)		-5.002*** (0.653)	-2.779*** (0.548)
ρ		-0.07			0.29	
$\chi^2(1)$ of $\rho=0$		0.8 [0.37]			21.35 [0.000]	
Log L		-1017.1835			-1118.0223	
Observations		1470			1577	
Freq. table		$z_m=0$	$z_m=1$		$z_m=0$	$z_m=1$
$z_f=0$		172	253		205	197
$z_f=1$		125	920		142	1033

Note: Robust standard errors in parenthesis, p -values in brackets.

***, **, * indicates significance on 1, 5 and 10 percent.

Table 7: Unrestricted model estimates by 3SLS for 1984, male and female budget shares of total nonmarket time as dependents (symmetry not imposed)

Parameter	Weekdays (Mon.-Fri.)				Weekend days (Sat.-Sun.)			
	Male		Female		Male		Female	
	Coef.	St.err.	Coef.	St.err.	Coef.	St.err.	Coef.	St.err.
Constant	0.276	(0.097)***	0.411	(0.125)***	0.300	(0.064)***	0.323	(0.084)***
$\ln \tilde{w}_m$	0.154	(0.024)***	-0.175	(0.037)***	0.245	(0.017)***	-0.165	(0.029)***
$\ln \tilde{w}_f$	-0.140	(0.023)***	0.180	(0.024)***	-0.188	(0.017)***	0.230	(0.017)***
age18-30 _m	0.003	(0.027)	0.049	(0.029)	0.013	(0.018)	-0.010	(0.017)
age31-35 _m	0.004	(0.018)	0.012	(0.019)	-0.010	(0.011)	-0.011	(0.010)
age46-55 _m	0.024	(0.020)	-0.004	(0.020)	-0.021	(0.013)	0.004	(0.012)
age56- _m	0.016	(0.036)	-0.003	(0.039)	-0.028	(0.020)	-0.003	(0.024)
univdeg _m	-0.006	(0.019)	0.012	(0.021)	-0.006	(0.013)	0.012	(0.012)
yschool _m	-0.001	(0.002)	0.000	(0.002)	0.001	(0.001)	-0.001	(0.001)
selfemp _m	-0.023	(0.025)	0.016	(0.019)	-0.011	(0.014)	-0.013	(0.015)
age18-30 _f	0.038	(0.027)	-0.036	(0.030)	-0.048	(0.019)**	-0.011	(0.018)
age31-35 _f	0.010	(0.019)	-0.002	(0.019)	-0.013	(0.010)	0.013	(0.010)
age46-55 _f	-0.026	(0.018)	-0.002	(0.022)*	0.010	(0.013)	0.001	(0.012)
age56- _f	-0.009	(0.044)*	0.045	(0.042)	0.000	(0.019)	0.010	(0.027)
univdeg _f	0.026	(0.019)*	0.006	(0.021)	-0.010	(0.018)	0.021	(0.012)*
yschool _f	-0.001	(0.002)	-0.006	(0.002)**	0.000	(0.001)	-0.001	(0.002)
selfemp _f	-0.051	(0.019)***	0.043	(0.043)	0.016	(0.010)	0.016	(0.015)
ownhome	0.025	(0.014)	0.006	(0.015)	0.002	(0.009)	-0.014	(0.008)*
ch0-2	0.018	(0.031)*	0.037	(0.029)	-0.022	(0.019)	0.002	(0.016)
ch3-6	0.005	(0.013)	0.005	(0.014)	-0.006	(0.008)	0.014	(0.008)*
ch7-18	-0.002	(0.007)	0.004	(0.007)	-0.006	(0.004)	-0.005	(0.005)
daycare	-0.028	(0.014)**	0.017	(0.014)	0.015	(0.011)	-0.012	(0.009)
agediff	-0.002	(0.002)*	0.001	(0.002)	0.002	(0.001)**	0.001	(0.001)
mar-may	0.007	(0.014)	0.002	(0.015)	-0.002	(0.009)	-0.007	(0.009)
jun-aug	0.041	(0.015)***	0.004	(0.014)	0.008	(0.008)	-0.001	(0.010)
sep-nov	-0.023	(0.012)*	-0.003	(0.014)*	0.002	(0.009)	0.000	(0.008)
λ_m	0.022	(0.039)*	-0.033	(0.043)*	-0.030	(0.015)**	-0.055	(0.027)**
λ_f	0.042	(0.072)*	0.000	(0.060)*	0.033	(0.025)	-0.003	(0.035)
R^2	0.40		0.42		0.66		0.67	
Observations			352				320	
All λ -terms=0: $\chi^2(4)$			2.96 [p -value=0.565]				10.06 [p -value=0.039]	

Note: Heteroscedastic-consistent standard errors in parenthesis; ***, **, * significant on 1, 5 and 10 percent.

Table 8: Unrestricted model estimates by 3SLS for 1993, male and female budget shares of total nonmarket time as dependents (symmetry not imposed)

Parameter	Weekdays (Mon.-Fri.)				Weekend days (Sat.-Sun.)			
	Male		Female		Male		Female	
	Coef.	St.err.	Coef.	St.err.	Coef.	St.err.	Coef.	St.err.
Constant	0.738	(0.108)***	0.353	(0.120)***	0.450	(0.075)***	0.445	(0.066)***
$\ln \tilde{w}_m$	0.089	(0.020)***	-0.166	(0.020)***	0.196	(0.014)***	-0.190	(0.013)***
$\ln \tilde{w}_f$	-0.176	(0.020)***	0.159	(0.024)***	-0.189	(0.013)***	0.208	(0.013)***
age18-30 _m	0.003	(0.033)	0.022	(0.034)	-0.007	(0.021)	-0.024	(0.024)
age31-35 _m	-0.035	(0.026)	-0.001	(0.027)	-0.005	(0.019)	0.002	(0.014)
age46-55 _m	0.010	(0.017)	-0.019	(0.015)	-0.003	(0.016)	-0.001	(0.011)
age56- _m	0.004	(0.024)	-0.012	(0.025)	-0.001	(0.023)	-0.010	(0.016)
univdeg _m	-0.019	(0.016)	-0.030	(0.015)**	-0.023	(0.012)*	0.005	(0.008)
yschool _m	0.001	(0.002)	0.003	(0.002)	0.001	(0.002)	-0.002	(0.001)
selfemp _m	-0.031	(0.018)*	0.017	(0.016)	0.009	(0.015)	0.017	(0.008)**
age18-30 _f	0.020	(0.029)	-0.030	(0.033)	0.027	(0.019)	0.033	(0.018)*
age31-35 _f	0.023	(0.023)	-0.026	(0.021)	0.002	(0.017)	-0.018	(0.020)
age46-55 _f	-0.026	(0.017)	0.013	(0.015)	0.008	(0.015)	-0.003	(0.007)
age56- _f	0.033	(0.032)	0.033	(0.035)	0.062	(0.026)**	0.031	(0.018)*
univdeg _f	0.030	(0.016)*	0.003	(0.018)	0.005	(0.012)	0.002	(0.011)
yschool _f	-0.005	(0.002)***	0.000	(0.002)	-0.001	(0.002)	0.000	(0.001)
selfemp _f	-0.022	(0.031)	-0.026	(0.023)	-0.021	(0.025)	-0.030	(0.019)
ownhome	-0.007	(0.012)	-0.005	(0.012)	0.003	(0.009)	-0.005	(0.006)
ch0-2	-0.021	(0.030)	0.012	(0.030)	0.021	(0.023)	0.014	(0.017)
ch3-6	0.001	(0.012)	0.044	(0.012)***	0.019	(0.008)**	-0.003	(0.010)
ch7-18	-0.009	(0.007)	-0.003	(0.007)	0.007	(0.006)	0.007	(0.005)
daycare	-0.008	(0.012)	0.006	(0.013)	-0.005	(0.010)	-0.003	(0.007)
agediff	0.000	(0.002)	0.000	(0.002)	0.000	(0.001)	0.001	(0.001)
mar-may	0.006	(0.012)	0.010	(0.011)	-0.006	(0.009)	-0.007	(0.008)
jun-aug	0.031	(0.015)**	0.013	(0.014)	-0.005	(0.012)	0.009	(0.007)
sep-nov	-0.007	(0.010)	0.001	(0.011)	-0.008	(0.009)	-0.002	(0.008)
λ_m	-0.071	(0.078)	0.068	(0.072)	-0.004	(0.040)	-0.020	(0.037)
λ_f	0.000	(0.066)	0.030	(0.071)	-0.113	(0.047)**	-0.047	(0.035)
R ²	0.30		0.35		0.63		0.73	
Observations			366				312	
All λ -terms=0: $\chi^2(4)$			3.32 [p -value=0.505]				6.92 [p -value=0.140]	

Note: Heteroscedastic-consistent standard errors in parenthesis; ***, **, * significant on 1, 5 and 10 percent.