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Abstract
We estimate a collective household model with survey data on financial satisfaction from the European Community Household Panel. Our estimates suggest that cohabitating individuals enjoy returns to scale in consumption that are towards the larger end of the range of estimates reported in the literature. They also suggest that the share of household income provided by the female partner is a significant determinant of her share of household consumption in most countries of the countries we study.

Keywords: consumption, returns to scale, collective household models

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1. Introduction

The unitary model of consumer behavior assumes the existence of a single household utility function. This sits uneasily with the methodological individualism of economics. Moreover, the unitary model has empirical implications – for example, that household demands and saving behavior are unaffected by the distribution of income within the household – that are overwhelmingly rejected by data.

There are several ways to take the multiplicity of decision makers in a household into account, including both non-cooperative and non-cooperative approaches. However, the leading approach now seems to be "collective" models, pioneered by Chiappori (1988, 1992) and Apps and Rees (1988), and recently surveyed by Vermeulen (2002). The collective approach assumes only that intra-household decisions are Pareto efficient (in particular, it does not specify a particular bargaining structure). This turns out to be enough to generate testable restrictions. Moreover, with certain restrictions on preferences, intra-household allocation can be described by a sharing rule.

The identification and estimation of the parameters of collective household models with data on household expenditures and/or (individual) labour supply is a difficult task. To date, there have been essentially two schemes for identification. The first is to assume that there is at least one assignable good. Typical candidates for an assignable good are leisure (Chiappori, 1994; Chiappori, Fortin and Lacroix, 2002), and men/women's clothing (Browning et al., 1994). To assume that observed non-market time is private consumption of leisure is very unattractive if there is home production (Apps and Rees, 1997).[1] Private consumption of men's and women's clothing is observed only if members of a couple are indifferent to each other's sartorial choices.

In a recent paper, Browning, Chiappori and Lewbel (2003) develop an alternative estimation strategy. They show that by specifying a consumption technology and sharing rule, they can identify structural parameters (of individual preferences, the consumption technology and the sharing rule), essentially by comparing the shapes of demands between of men and women, living singly and in couples. However, it turns out that finding the structural parameters that optimally rationalize the differences in demands is a highly nonlinear, computationally intensive, problem. The authors report that estimates take a long time to converge, and that there are multiple local minima. This limits the number of specification checks and tests that they can perform.

Our goal in this paper is to explore a third alternative, which exploits data that, to the best of our knowledge, has not yet been used for this purpose. In particular, we attempt to estimate a collective household model from panel data on individual subjective financial satisfaction. The basic idea, which we lay out formally below, is as follows. When two single individuals move into cohabitation, their financial resources change in two ways. First, returns to scale in consumption mean that their potential joint consumption exceeds the sum of what they could individually consume living alone. Second, unless resources are shared perfectly equally, one individual's consumption will rise by more than is implied by returns to scale, while the consumption of the other will rise by less (or could even fall). Thus, because we observe individuals of different circumstances moving in and out of cohabitation, if we assume stable (but possibly heterogeneous) individual preferences and reporting behavior, we can infer something

[1] Chiappori (1997) demonstrates that it is still possible to proceed if all home produced goods can be freely bought and sold in the market.

about the household consumption technology and sharing rule from observed changes in individual financial satisfaction.

The use of subjective survey measures of economic wellbeing has been rising in recent years. Such measures have been repeated validated by psychologists, and are believed to be a reasonably proxy for "utility". See Frey and Stutzer (2002) for a survey. We employ longitudinal data from the European Community Household Panel (ECHP), which contain the following question on subjective financial wellbeing: "How satisfied are you with your financial situation?" Responses are recorded on a 7-point scale. Schwarze (2003) uses the answers to this question in the German Socio Economic Panel (which is a component of the ECHP) to estimate equivalence scales (effectively, the returns to scale in consumption). Kuklys (2003) performs a similar exercise on the British Household Panel Survey, which is also a component of the ECHP.[2] However, neither author considers intra-household allocation (implicitly assuming that consumption is equally allocated in the household.) Bonke and Browning (2003) conduct a cross sectional analysis of this question in the Danish component of the ECHP. Their focus is intra-household allocation. They show that husbands and wives differ in their financial satisfaction and that relative income is an important correlate of within-household differences in satisfaction. This is important evidence against the income-pooling implication of the unitary model. However Bonke and Browning do not exploit the panel nature of the data or attempt to estimate a structural model, as we do in this paper.

We believe our approach adds significantly to the existing literature in several ways. First, because we use a very different kind of information than existing estimates of collective models, our estimates provide an independent check on previous results. Second, the procedure that we

[2] Kuklys is particularly concerned with estimating the costs of disability.

develop is very computationally manageable, particularly in contrast to the methodology of Browning, Chiappori and Lewbel (2003) (although at the cost of specifying a less rich household consumption technology). It also has modest data requirements. The low computational burden makes it feasible to try a variety of specifications and robustness checks, and to quickly generate estimates for a range of samples. The use of widely available data also facilitates the generation and comparison of estimates for different time periods and populations. For example, below we report estimates of returns to scale and sharing rule parameters for ten of the countries participating in the ECHP. In turn, the ability to generate estimates from a range of samples, populations or institutional settings opens up possibilities for comparative research. To illustrate, we use our returns to scale and sharing rule estimates to calculate measures of income inequality among singles and couples for ten European countries in 2001. These calculations account for within-household inequality and we contrast them with measures that fail to account for within-household inequality. We discuss other possible applications in our concluding section.

Our estimates suggest that cohabitating individuals enjoy returns to scale in consumption that are towards the larger end of the range of estimates reported in the literature. They also suggest that, in many countries, the share of household income provided by the female partner is a significant determinant of her share of household consumption. This latter result contradicts the income pooling implied by the unitary model. In our application we find that accounting for intra-household inequality results in modest increases in the Gini coefficients for the 10 countries we examine. The impact of accounting for intra-household allocation on measured inequality differs across countries, but not so much as to dramatically change the rank ordering of countries by inequality.

The outline of the rest of the paper is as follows. In the next section we describe the European Community Household Panel survey, and the sub-sample of that data that forms the basis of our empirical work. We also take an unstructured look at the financial satisfaction of men and women, living singly and in couples, and at how financial satisfaction changes with changes in living arrangements. This helps to motivate the subsequent analysis. In Section 3 we develop the structural model that we subsequently use to interpret the data. We describe, in turn, individual preferences, the household consumption technology, and intra-household allocation. Section 4 discusses some econometric issues. Section 5 presents our main results, which are country-specific estimates of returns to scale in household consumption and of parameters of the sharing rule that determines household allocation. Section 6 reports our inequality estimates and Section 7 concludes.

2. Data and Descriptive Statistics

The European Community Household Panel Survey

The European Community Household Panel survey (ECHP) is a standardized multi-purpose annual longitudinal survey providing comparable micro-data about living conditions in the European Union Member States. The December 2003 release of the ECHP data used in this paper includes eight waves spanning the 1994--2001 time period. Over 60,000 households and 130,000 adults across the European Union were interviewed at each wave.[3]

[3] The first wave covered all EU-15 Member States with the exception of Austria, Finland and Sweden. Austria joined in the second wave, Finland in the third, and Sweden in the fourth. However data for Sweden are not longitudinal, but derived from repeated cross-sections. In the periods covering the first three waves, the ECHP ran parallel to existing similar panel surveys in Germany, Luxembourg and the United Kingdom. From the fourth wave onwards, the ECHP samples were replaced by data harmonized ex post from these three existing surveys. The ECHP data were `cloned' backwards so that two versions of German, Luxembourg, and British data are available in the first three waves of the ECHP database.

The topics covered in the survey include income, employment, housing, health, and education. A harmonized (E.U.-wide) questionnaire was designed at Eurostat. The survey was implemented by "National Data Collection Units" in member states. The public-use database is derived from the data collected in each of the Member States and is created, maintained and centrally distributed by Eurostat.

Sample

Not all of the countries represented in the ECHP have data suitable for our purposes, because of exceptions to the general design rules and missing information. We study ten countries: Denmark, the Netherlands, Belgium, France, Ireland, Italy, Greece, Spain, Portugal and the United Kingdom.[4]

Our analysis is based on individuals living as a single individual or as a member of a couple (without children.) Couples may or may not be legally married; throughout we refer to an individual living as a member of a couple as "cohabiting."

A small number of individuals in same-sex couples were dropped from the data, as were individuals in households reporting zero household income (each of these restrictions eliminated less than 0.5% of the data). We also dropped observations for which there was not a usable response to the financial satisfaction question (less than 2% of the data).

A First Look at the Data

Individual respondents to the ECHP (including multiple individuals in the same households) answered the following "Financial Satisfaction" question:

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[4] We dropped the German data because the in SOEP the financial satisfaction information was not available, and in the original ECHP sample has only waves. We dropped Sweden since the nature of our study requires longitudinal samples. With respect to the UK, we dropped the 3 waves from the original ECHP sample and worked with the 8 waves of BHPS cloned data.

How satisfied are you with your present financial situation?

1. *not at all satisfied*
2. *largely unsatisfied*
3. *mildly unsatisfied*
4. *mildly satisfied*
5. *largely satisfied*
6. *fully satisfied*

Note that respondents in the United Kingdom (who were participating in the British Household Panel Survey) answer a similar question with only 5 categories.[5]

To provide a sense of the data, and to motivate the subsequent analysis, we provide some descriptive statistics in Tables 1, 2 and 3. To keep the tables manageable, we focus on data from three countries: the Netherlands, the United Kingdom and Spain.

Table 1 documents, for each of these three countries, the responses to the above question among single men, single women, co-habiting men and co-habiting women. Again, the data are individuals living alone or with a just with a spouse or partner. The first panel, for the Netherlands, suggests that single men are more satisfied with their financial situation than single women. The same appears to be true in Spain (third panel) but less so in the U.K. (middle panel). In all three countries, cohabitation is associated with greater financial satisfaction for both men and women, but the differential appears to be larger for women.

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[5] The BHPS question is:
How well would you say you yourself are managing financially these days? Would you say you are . . .
Living comfortably.................................... 1
Doing alright 2
Just about getting by.................................. 3
Finding it quite difficult 4
Finding it very difficult? 5
Don't know 8
The scale is inverted to harmonize with the ECHP.

Table 2 exploits the household structure of the data. Here we cross-tabulate the financial satisfaction of the male and female partner. The numbers presented are row percentages. So, for example, the top row of the first panel can be read as the percentage of cohabitating Dutch women whose partner was very dissatisfied with their financial situation that gave each of the responses (from very dissatisfied to very satisfied). There is clearly a strong correlation between partners' responses to this question, but it is not a perfect correlation. In each country (panel) there are significant off-diagonal terms: partners differ in their reported financial satisfaction. One measure of agreement between two ratings is the Kappa statistic. This measure adjusts for the amount of agreement that would arise randomly. A value of 0 indicates the same agreement as would arise by chance. A value of 1 indicates complete agreement. The Kappa statistics reported at the bottom of Table 2 suggest that the degree of intra-household agreement is relatively similar in the three countries.

Table 3 exploits the longitudinal nature of the data. For each country and gender, the distribution of year-on-year changes in the (categorical) measure of financial satisfaction are reported, for four different subgroups: those that remained single from one year to the next, those that moved to into cohabitation, those who moved out of cohabitation, and those in remained in cohabitation for one year to the next. These numbers should be interpreted with considerable caution. In particular, if one assumes (as is often assumed, and as we shall assume below) that the categorical responses are related to a continuous underlying latent index, it is not necessarily the case that the difference in the categorical indicators is monotone in the difference in the latent variables.[6] Nevertheless, interesting patterns are apparent. For example, movements out of

[6] Note that out structural estimates (below) do not involve differencing the categorical data, and so do not suffer from this problem.

cohabitation appear to be particularly associated with decreases in financial satisfaction for women in The Netherlands and in Spain.

From this preliminary analysis of the data, we take four messages. First, partners view their finances differently, which is at least suggestive of unequal resource allocation within the household, and possible further evidence against the unitary model. This point has been made previously by Bonke and Browning (2003), based on their cross-sectional analysis of the Danish subset of the ECHP. Second, changes in financial satisfaction with changes cohabitation status are, on average, different for men and women. This is certainly a pattern that we would like to be able to interpret further. Third, the patterns in the data differ significantly across countries. This again suggests that further investigation may be fruitful. Finally, the patterns in the data are complicated. This suggests that a model is needed to interpret them.

3. Model

We now present our structural model. This is a collective household model, intended to capture both returns to scale in household consumption and unequal allocation within households. The model intentionally follows Browning, Chiappori, and Lewbel (2003) (BCL), although it is simpler than the model they develop in ways that will be indicated below.

Individual Utility

Individuals have (random) PIGLOG preferences. The indirect utility function for PIGLOG preferences is

$$V = \frac{1}{b(p)}(\log x - a(p)) = \alpha(p) + \beta(p)\log x\,. \qquad (1)$$

We do not have price data, but can allow that prices will differ across countries and through time by allowing preference parameters (α and β) to be time and country specific. We also allow $a(p)$ and hence $\alpha(p)$ to vary with observable individual characteristics (such as age and education), possibly a scalar unobservable characteristic (an individual specific, time - invariant effect), and an idiosyncratic time-varying error term. Thus for individual *i*, living in country *c* at time *t*,

$$V_{ict} = \alpha_{ct}(z_{ict}) + \beta_{ct} \ln x_{ict} + \mu_i + \varepsilon_{ict}, \qquad (2)$$

where V_{ict} is utility, z_{ict} is observable characteristics, x_{ict} is total private consumption and μ_i is an individual specific effect and ε_{ict} is the idiosyncratic time-varying error term.

There are two key assumptions here. First, preferences are egoistic. Although there may be sharing and other sources of returns to scale (or, alternatively, congestion), individuals care about their own consumption.[7]

Second, (1) depends only on individual consumption and prices, and not on living arrangements directly (though the relationship between household income and individual consumption will depend on living arrangements, as we discuss below). Effectively we are modeling economic (or material) wellbeing and assuming that, if positive or negative utility is derived directly from cohabitation, such effects are additively separable from the consumption of goods and services.

Household Income and Individual Consumption

We assume that single individuals consume their (real) income[8]:

$$x_{ict} = y_{ict}. \qquad (3)$$

[7] We could allow for specific kinds of caring; the key assumption is that intrahousehold allocation can be described by a sharing rule.
[8] Inter-temporal issues are certainly important, but we abstract from them in this analysis.

However, for couples, things are more complicated, in two ways. First, consumption of couples can exceed their combined income through sharing and other sources of returns to scale in households. Second, the total consumption of couples is divided between them according to a sharing rule. Thus for couples:

$$x_{ict} = \eta_{ict} F^{-1}(y_{ict}) . \qquad (4)$$

Where η_{ict} is the share of "Total expenditure" and the function $F^{-1}()$ captures the returns to scale in household consumption in a general way. With returns to scale, $F^{-1}(y) > y$, but congestion or other negative consumption externalities might give the opposite.) With respect to household returns to scale, we follow BCL in assuming a linear household consumption technology. However, because we will work with data on overall satisfaction (utility) and not with expenditure data (and relative prices), we are forced to assume a simpler version of the consumption technology. In particular, we can model "overall" returns to scale, but not substitutions induced by the price-like effects of different returns to scales in different goods.[9] Thus we have:

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[9] BCL specify:

$$z_{ict} = F(q^1_{ict} + q^2_{ict}) = \mathrm{A}(q^1_{ict} + q^2_{ict}) + a$$

where z_{ict} is a (observable) vector of household consumption quantities of *n* goods, q^1_{ict} and q^2_{ict} are (unobserved) *n*-vectors of private consumption, A is an *n* x *n* nonsingular matrix and α is an *n*-vector. The budget constraint is:

$$p' z_{ict} \le y_{ict} ,$$

where p is a vector of prices and y_{ict} is income (and observable scalar.) This structure nests familiar cases. For example, with A diagonal and $a = 0$, the setup is analogous to Barten scales though for a collective model (see BCL for further discussion). In their analysis, the elements of A and α are identified via the modeling of demands for the n goods, which is not feasible with our data. . Effectively, we assume that a = 0 and that A is a diagonal matrix with identical elements *A* along the diagonal. Thus, we have an Engel scale rather than a Barten Scale. Assuming the budget constraint holds with equality we have:

$$(x^1 + x^2) = \frac{y}{A}$$

Note that economies of scale imply that $0.5 < A \leq 1$. We can now restate (4) as:

$$x_{ict} = \eta_{ict} \frac{y_{ict}}{A}. \qquad (5)$$

We specify the sharing rule for the first (arbitrarily, female) partner as:

$$\eta^1 = \frac{e^{\gamma(p,y,w)}}{1 + e^{\gamma(p,y,w)}} \qquad (6)$$

(so that $\eta^2 = \frac{1}{1 + e^{\gamma(p,y,w)}}$). The sharing rule depends on prices and income, and on variables, w, that affect the intrahousehold allocation. These are *distribution factors* in the terminology of (for example) Browning, Chiappori and Lechene (2003). Again, we do not have data on relative prices, but we can allow $\gamma()$ to vary across countries and time. With respect to distribution factors, we focus on the (current) share of the first (female) partner's income in household income, which we denote w^1.[10] Thus we specify:

$$\gamma(p, y, w) = \gamma_{ct}^0 + \gamma_{ct}^1 \ln y + \gamma_{ct}^2 w^1 \qquad (7)$$

$$\begin{aligned} y &= p'z = (p'\mathrm{A}q^1 + p'\mathrm{A}q^2) \\ &= A(p'q^1 + p'q^2) \\ &= A(x^1 + x^2) \end{aligned}$$

Or

$$(x^1 + x^2) = \frac{y}{A}$$

[10] We are estimating a static model on dynamic data. One possible motivation is that partners can't commit. This would mean that the allocation at each point in time depends only the distribution factors at that point in time (see the discussion in Browning, Chiappori and Lechene, 2003)

4. Empirical Implementation and Econometric Issues

Combining equations (2), (3), (8) and (9) gives the indirect utility function, in terms of observables, for single men, single women and both members of a couple. For singles,

$$V_{ict} = \alpha_{ct}(z_{ict}) + \beta_{ct} \ln y_{ict} + \mu_i + \varepsilon_{ict}\,, \qquad (8)$$

For female members of couples

$$V_{ict} = \alpha_{ct}(z_{ict}) + \beta_{ct}\left\{\gamma^0_{ct} + \gamma^1_{ct}\ln y + \gamma^2_{ct} w^1 - \ln(1+e^{\gamma^0_{ct}+\gamma^1_{ct}\ln y + \gamma^2_{ct} w^1}) + \ln y_{ict} - \ln A\right\} + \mu_i + \varepsilon_{ict}\,,$$

(9)

and for male members of couples

$$V_{ict} = \alpha_{ct}(z_{ict}) + \beta_{ct}\left\{-\ln(1+e^{\gamma^0_{ct}+\gamma^1_{ct}\ln y + \gamma^2_{ct} w^1}) + \ln y_{ict} - \ln A\right\} + \mu_i + \varepsilon_{ict}\,, \qquad (10)$$

To ease estimation we take one further step, which is to linearize $\ln(1+e^{\gamma^0_{ct}+\gamma^1_{ct}\ln y_{ict} + \gamma^2_{ct} w^1_{ict}})$ around zero. We construct our data so that these variables have a mean of zero for each country (and the country specific means are subsumed in the constant). Thus the linearization is

$$\ln(1+e^{\gamma^0_{ct}+\gamma^1_{ct}\ln y_{ict} + \gamma^2_{ct} w^1_{ict}}) \approx \ln(1+e^{\gamma^0_{ct}}) + \frac{e^{\gamma^0_{ct}}}{1+e^{\gamma^0_{ct}}}(\gamma^1_{ct}\ln y_{ict} + \gamma^2_{ct} w^1_{ict}) \qquad (11)$$

Our final specifications for men and women living in couples are:

$$V_{ict} = \alpha_{ct}(z_{ict}) + \beta_{ct}\left\{\begin{array}{l}\gamma^0_{ct} + \gamma^1_{ct}\ln y_{ict} + \gamma^2_{ct} w^1_{ict} - \ln(1+e^{\gamma^0_{ct}}) \\ -\dfrac{e^{\gamma^0_{ct}}}{1+e^{\gamma^0_{ct}}}(\gamma^1_{ct}\ln y_{ict} + \gamma^2_{ct} w^1_{ict}) + \ln y_{ict} - \ln A\end{array}\right\} + \mu_i + \varepsilon_{ict}\,, \qquad (12)$$

for women, and

$$V_{ict} = \alpha_{ct}(z_{ict}) + \beta_{ct}\left\{\begin{array}{l}-\ln(1+e^{\gamma^0_{ct}}) - \dfrac{e^{\gamma^0_{ct}}}{1+e^{\gamma^0_{ct}}}(\gamma^1_{ct}\ln y_{ict} + \gamma^2_{ct} w^1_{ict}) \\ +\ln y_{ict} - \ln A\end{array}\right\} + \mu_i + \varepsilon_{ict}\,, \qquad (13)$$

for men.

Defining D^C_{ict} as a dummy indicating membership in a couple and D^F_{ic} as dummy for female gender, (8),(12) and (13) are trivially combined into a single, reduced-form, individual-level model:

$$\begin{aligned} V_{ict} &= \pi^0_{ct}(z_{ict}) + \pi^1_{ct}\ln y_{ict} + \pi^2_{ct}D^C_{ict} + \pi^3_{ct}D^C_{ict}D^F_{ic} \\ &\quad + \pi^4_{ct}D^C_{ict}D^F_{ic}\ln y_{ict} + \pi^5_{ct}D^C_{ict}D^F_{ic}w^1_{ict} \\ &\quad + \pi^6_{ct}D^C_{ict}(1-D^F_{ic})\ln y_{ict} + \pi^7_{ct}D^C_{ict}(1-D^F_{ic})w^1_{ict}\,, \\ &\quad + \mu_i + \varepsilon_{ict} \\ &= W_{ict}\Pi_{ct} + \mu_i + \varepsilon_{ict} \end{aligned} \tag{14}$$

$$\begin{aligned} \pi^0{}_{ct}(z_{ict}) &= \alpha_{ct}(z_{ict}) \\ \pi^1{}_{ct} &= \beta_{ct} \\ \pi^2{}_{ct} &= -\beta_{ct}(\ln A + \ln(1+e^{\gamma^0_{ct}})) \\ \pi^3{}_{ct} &= \beta_{ct}\gamma^0_{ct} \\ \pi^4{}_{ct} &= \beta_{ct}\left(1-\frac{e^{\gamma^0_{ct}}}{1+e^{\gamma^0_{ct}}}\right)\gamma^1_{ct} = \frac{\beta_{ct}\gamma^1_{ct}}{1+e^{\gamma^0_{ct}}} \\ \pi^5{}_{ct} &= \beta_{ct}\left(1-\frac{e^{\gamma^0_{ct}}}{1+e^{\gamma^0_{ct}}}\right)\gamma^2_{ct} = \frac{\beta_{ct}\gamma^2_{ct}}{1+e^{\gamma^0_{ct}}} \\ \pi^6{}_{ct} &= \beta_{ct}\left(-\frac{e^{\gamma^0_{ct}}}{1+e^{\gamma^0_{ct}}}\right)\gamma^1_{ct} = \frac{-\beta_{ct}e^{\gamma^0_{ct}}\gamma^1_{ct}}{1+e^{\gamma^0_{ct}}} \\ \pi^7{}_{ct} &= \beta_{ct}\left(-\frac{e^{\gamma^0_{ct}}}{1+e^{\gamma^0_{ct}}}\right)\gamma^2_{ct} = \frac{-\beta_{ct}e^{\gamma^0_{ct}}\gamma^2_{ct}}{1+e^{\gamma^0_{ct}}} \end{aligned}$$

Measuring Utility

This gives us an equation we could estimate if V_{ict} were observable. To proceed, we interpret responses to the “Financial Satisfaction” question as a measure of economic wellbeing or utility from the consumption of goods and services. Specifically, denoting financial satisfaction of individual i in country *c* at time *t* as FS_{ict}, we assume:

$$\begin{aligned}
FS_{ict} &= 6 \Leftrightarrow V_{ict} > k_5 \\
FS_{ict} &= 5 \Leftrightarrow k_5 > V_{ict} > k_4 \\
FS_{ict} &= 4 \Leftrightarrow k_4 > V_{ict} > k_3 \\
FS_{ict} &= 3 \Leftrightarrow k_3 > V_{ict} > k_2 \\
FS_{ict} &= 2 \Leftrightarrow k_2 > V_{ict} > k_1 \\
FS_{ict} &= 1 \Leftrightarrow k_1 > V_{ict}
\end{aligned} \tag{15}$$

Thus, in the absence of unobserved individual heterogeneity ($\mu_i = 0$), and assuming that the ε_{ict} are normally distributed, (14) could be estimated as an ordered probit model. Note that we are assuming that wellbeing is interpersonally ordinally comparable (for further discussion, see Ferrer-i-Carbonell and Frijters, 2004).

Unobserved Heterogeneity

It is desirable, however, to allow for time-invariant, unobserved individual heterogeneity ($\mu_i \neq 0$) in preferences, or perhaps in reporting behaviour.[11] Our set up suggests a random effects ordered probit model. However, that model assumes the independence of FS_{ict} given μ_i and W_{ict}; intuitively this says that the unobserved, time-varying determinants of utility (captured by ε_{ict} in Equation 14) cannot be serially correlated. It is easy to imagine that there is some persistence in unobserved, time-varying determinants of utility, and so we wish to avoid imposing this assumption. A pooled ordered probit provides consistent estimates of Π up to scale without imposing this assumption. That is, we can estimate $\pi_{ct}^k / (1+\sigma_\mu^2)^{1/2}$ under fairly mild conditions (normality of the disturbances, ε_{ict}; observables, W_{ict}, uncorrelated with the individual specific effect μ_i, and with the contemporaneous disturbance. See Wooldridge, 2002, section 15.8). Fortunately, estimating the π^k up to a scalar is sufficient to identify the structural

[11] So long as this unobserved heterogeneity is in preferences, we are continuing to assume interpersonal ordinal comparability.

parameters of interest (A and the elements of γ) - because these can be recovered from ratio's of the $\pi^k /(1+\sigma_\mu^2)^{1/2}$. For example: $\gamma_{ct}^o = \frac{\pi_{ct}^3}{\pi_{ct}^1} = \frac{\pi_{ct}^3 / \sqrt{1+\sigma_\mu^2}}{\pi_{ct}^1 / \sqrt{1+\sigma_\mu^2}}$.

For further details, see the appendix.[12]

This estimation strategy does not allow for correlation between the unobservable individual effect μ_i and observed covariates, W_{ict}. To relax this restriction, we follow the Mundlak (1978) version of the Chamberlain (1980) suggestion, and model the individual effects, μ_i, as a linear function of the individual specific means of a subset of the right-hand side variables, W_{ict}. Denoting a subset of W_{ict} (including the constant) by w_{ict} and the individual specific means of these variables by $\overline{w_{ic}}$, our assumption is:

$$\mu_{ic} = \overline{w_{ic}}\delta + \zeta_{ic} \tag{16}$$

with $\zeta_{ic} \mid w_{ic} \sim N(0, \sigma_\zeta^2)$, $w_{ic} = (w_{ic1} ...... w_{icT})'$. Thus our formulation of latent, indirect utility becomes:

$$V_{ict} = W_{ict}\Pi_{ct} + \overline{w_{ic}}\delta + \zeta_{ic} + \varepsilon_{ict} \tag{17}$$

Again we estimate by pooled ordered probit in order to avoid assuming serial independence of the disturbances. The key reduced form parameters continue to be identified up to scale. A test of $\delta = 0$ is a test of the assumption that the individual effects are uncorrelated with observables.

While the addition of the "Mundlak" terms, $\overline{w_{ic}}$, relaxes somewhat the assumption that the individual effects are uncorrelated with observables, it does impose an additional restriction. In particular, the Chamberlain/Mundlak procedure requires strict exogeneity of the observables,

[12] Parameters of the utility function (α and β), however, are identified only up to the scale factor.

W_{ict}. Strict exogeneity would imply, for example, that future cohabitation is uncorrelated with the unobserved determinants of current financial satisfaction.[13]

Recovering Structural Parameters

Given the reduced form estimates, the structural parameters can be recovered with a minimum distance step. This also provides a useful over-identification test. Because the dimension of this maximization is the number of parameters, it is very fast. (See also the appendix.)

5. Results

We now turn to estimating our simple collective model on ECHP data from ten European countries. The key structural parameters are the parameters of the sharing rule, and the parameter *A*, which captures household returns to scale. Before presenting our estimates, it is useful to consider some points of comparison.

Note that with equal sharing ($\eta = 0.5$) 2*A* is a traditional equivalence scale (divide household income by 2*A* to give the equivalent income for a single individual). The "original" OECD equivalence scale that gives a weight of 1 to the first adult and a weight of 0.7 to the second adult implies a implies a value for *A* of 1.7/2 = 0.85. The "modified" OECD equivalence scale (deVos and Zaidi, RIW, 1997) gives a weight of 0.5 to the second adult and so implies a value of 0.75 for *A*. The common "square-root of household size" equivalence scale implies that $A = \frac{\sqrt{2}}{2} \approx 0.7$.

[13] While we think the case for these two estimation strategies is good, we did experiment (unsuccessfully) with other panel estimators for ordered responses. These included random effects ordered probit, random effects ordered probits with Chamberlain/Mundlak terms, and a procedure for implementing a fixed effects ordered logit suggested by Andersen (1973) (see also Das and van Soest, 1997). These estimators also require strict exogeneity.

In a paper that shares some methodological aspects with our work, Schwarze (2003) uses financial satisfaction questions in the German Socio Economic Panel to estimate equivalence scales (but assumes equal allocation within households.) His estimates imply a value for A of 0.61 to 0.63, which suggests larger returns to scale than the OECD or "square-root" equivalence scales (the second adult gets a weight of approximately 0.25). It is a common finding that equivalence scales based on "subjective information" suggest larger returns to scales than equivalence scales based on demand system estimation or expert opinion.

As noted above, our simple collective household model is similar to (thought somewhat less rich than) the model that Browning, Chiappori and Lewbel (2003) develop and then estimate on Canadian data (using methods quite different from our own.) BCL posit Barten scales, so that the returns to scale explicitly differ across goods, but they can, and do, calculate an "overall" return to scale from their estimates. Their estimates imply a value of A of 0.79.

With respect to sharing rule parameters BCL find that a woman of the same age and personal income of her spouse, and median household income, enjoys a 65% share of potential household consumption. That the female share exceeds the male share reflects the fact couples' demands are more similar to those of single women than to those of single men. BCL find that the female share is larger in richer households, but find no effect of the age difference between the female and her spouse or of the income share of the female. The finding that income shares do not affect intra-household allocation contradicts earlier findings by Browning et al., (1994). The earlier findings are based on a different identification strategy (a strategy that assumes that particular goods are assignable).

We begin by estimating Equation (14) by pooled ordered probit. This is the base model, without any Chamberlain/Mundlak terms to account for correlation between observables and unobservable individual effects. We estimate separately by country, and include a full set of time dummies. However to keep things manageable, we do not allow other parameters to vary over the 8 years covered by the data. This amounts to assuming that relative prices do not differ substantially within countries over this period. Among the key variables in the specification captured by Equation (14) are household income and the female income share. Country specific means for these variables are reported in Table 4. Our observable utility shifters, z_{ict}, (which could also be interpreted as determinants of reporting behavior), are gender (a female dummy), education (captured by two dummy variables) and age and age-squared.

Reduced form parameters estimates are reported in Table 5a. Reassuringly, financial satisfaction is increasing in income in all countries. However, in most countries, the effect of income is different for individuals living alone or cohabiting. Among cohabiting individuals, the female income share is a significant determinant of financial satisfaction, for both men and women, in all countries. These reduced for parameters are difficult to interpret however, so it is natural to move to our structural parameter estimates. These are presented in Table 5b.

For most countries we get small but reasonable estimates of the returns to scale parameter A. Note that a small value of this parameter indicates substantial returns to scale. A value of 0.5 indicates that a couple's potential total consumption is double their income; this in turn implies that all consumption is public. A value of 1 indicates no returns to scale; all consumption is private. The estimates of A and associated confidences intervals are presented graphically in Figure 1. A traditional "equivalence scale" is obtained by multiplying the parameter A by two. This gives the value by which a couple's income should be divided to give the income that a single person would require to have the same per capita total consumption. In only two countries (Netherlands and Belgium) are the theoretical restrictions on this parameter ($0.5 \le A \le 1$) rejected by the data at conventional levels of statistical significance. The point estimates for Denmark and France also lie outside the theoretical range. Among the other countries, the estimates range from 0.526 (U.K) to 0.767 (Portugal); the implied equivalence scales range from 1.05 to 1.53. These estimates indicate substantial returns to scale. For example, for every country but Portugal the estimated returns to scale exceed those implied by the “modified OECD” equivalence scale.

Turning to the sharing rule parameters, we find that the female income share is a statistically significant determinant of consumption shares in seven of our ten countries. The exceptions are the Netherlands, Belgium and Ireland. In all ten countries the sign of the sharing rule coefficient on female income share is positive. This indicates that, holding income constant, an increased female income share raises the financial satisfaction of the female in a couple and lowers the financial satisfaction of her male partner. This seems to us to be strong evidence against the unitary model.

Household income is a statistically significant determinant of the female share only in Denmark and the United Kingdom (and in Spain at the 10% level).

To aid in the interpretation of these parameters, we calculate female consumption shares at mean household income and alternative assumptions about the female income share. In particular, we calculate the female’s share of total consumption if the couple has average income and the female share of income is 0.25, 0.5, 0.75 or the mean female income share for that country (from Table 4). These calculations are reported at the bottom of Table 5b. Female consumption shares rise steeply with female income shares in some countries, notable Denmark, France, Spain and Portugal. The same relationship is notably flat in the Netherlands and Ireland.

A striking feature of the results is that, in all countries, our estimates suggest the female share of total consumption is almost always greater than one half.

Finally, we also note that the over-identification tests reject the null in all countries. This is perhaps to be expected. The fairly tightly specified model we are using to interpret the data is not parameter rich.

These estimates, based on a pooled ordered probit, do not fully exploit the longitudinal nature of our data. In particular, all of the parameters are identified by both cross-sectional and longitudinal variation in the relevant variables. We therefore now turn to estimates based on the Chamberlain-Mundlak procedure described in the previous section (see especially Equations 16 and 17).

The "Mundlak" terms that we include are person-specific means of the couple dummy, and the couple-gender interaction. This means that the reduced-form coefficients on the couple dummy and the couple-gender interaction are identified only by within-person variation. In turn, this means that the returns to scale parameter A is identified only by within- person variation (as it is recovered from these reduced form parameters.) We do not include person-specific means of income or income share variables. In our short panel, within-person variation in these variables is dominated by transitory income shocks and measurement error, and we did not think it advisable to estimate parameters only with such variation. The consequence is that other structural parameters, notably the sharing rule parameters, continue to reflect both between- and within – person variation. We nevertheless feel that this specification represents the limit of what can reasonably be asked of the data.

The resulting reduced form estimates are presented in Table 6a and the corresponding estimates of the structural parameters are presented in Table 6b. Table 6a also reports (in the second to last row) tests of the joint statistical significance of the person-specific means ("Mundlak terms"). These are statistically significant at the 5% level in half of our countries

(Denmark, Belgium, France, Ireland, and United Kingdom) and at the 10% level in a further three (Denmark, Greece and Portugal). These results suggest that the individual effects are correlated with cohabitation status.

Turning to the resulting estimates of the structural parameters (Table 6b), we see that the returns to scale parameter, A, is now somewhat less precisely estimated. The estimates of A and associated confidences intervals are presented graphically in Figure 2. It is now the case that the theoretical restrictions on this parameter ($0.5 \leq A \leq 1$) are not rejected by the data for any country. However, the data do contain useful information about this parameter, as large parts of the theoretical range are excluded in many countries. The estimates again suggest quite large returns to scale. Only for Portugal, Ireland and Greece do the estimated returns to scale exceed those implied by the "modified OECD" equivalence scale.

Turning to the sharing rule parameters, we find that the female income share is statistically significant at the 5% level in five countries (Denmark, France, Italy, Greece, and Spain) and at the 10% level in a further two (Portugal and the U.K.). Again the sign in all countries is positive, indicating that, holding income constant, an increased female income share raises the financial satisfaction of the female in a couple and lowers the financial satisfaction of her male partner.

With these new estimates we repeat our calculations of female consumption shares at mean household income and alternative assumptions about the female income share. The results are presented in the bottom of Table 6b and also in Figure 3. In Figure 3, countries are arrayed along the horizontal axis. The female share of a couple's total consumption is measured on the vertical axis. This is calculated in three ways, all employing country-specific estimates of the sharing rule parameters. First, we assume that the female contributes 25 percentage of household income (plotted as a circle.); second, we assume that the female contributes 75 percent of household income (plotted as a triangle); and finally, we set the female contribution to household income equal to the country mean (plotted as a diamond). In all three cases, household income is set to the country specific means. Thus the diamonds give a sense of women's share of total consumption in an "average" couple in each country. The vertical distance between the circles and triangles give, for each country, a sense of the responsiveness of the sharing rule to the female income share (with a greater vertical distance indicating a more responsive sharing rule). The figure exhibits considerable variability across country in the share of an "average" couples' total consumption that is enjoyed by the female partner. There are also considerable differences in the responsiveness of that share to the fraction of household income that is contributed by the female partner. For example, the estimated female share of total consumption is lowest in Denmark, Spain and the U.K. However, only in the U.K is it less than one half (when evaluated at the means of the data.) Our estimates suggest that Denmark, Spain and France are countries where the sharing to rule is most sensitive to the fraction of household income that is contributed by the female partner.

To summarize, our preferred estimates are those that use the Chamberlain-Mundlak procedure to allow for some correlation between unobserved individual heterogeneity and observable characteristics (notably cohabitation status). Both these estimates and our base estimates suggest that cohabitating individuals enjoy returns to scale in consumption that are towards the larger end of the range of estimates reported in the literature (or equivalently that the implied equivalence scale is towards the smaller end of the range of plausible values). They also suggest that, in most of the countries we study, the share of household income provided by the female partner is a significant determinant of her share of household consumption.

6. Application to Inequality Measurement

One application of our estimates is the measurement of inequality. The typical approach is to calculate an inequality measure (for example the Gini index) at the individual level. Individuals are assigned the "equivalent income" of their household, which is just household income adjusted by an equivalence scale. Implicitly or explicitly, such analyses assume equal allocations within households. Inequality studies that account for intra-household inequality with direct evidence on individual consumption are very rare.[14]

In principal, knowledge of the returns to scale and sharing rule parameters allow for the calculation of individual consumption, and hence, an examination of individual inequality without the assumption of equal intra-household allocations. An early paper exploiting this idea is Phipps and Burton (1995), who explore the sensitivity of Canadian poverty statistics to alternative assumptions about sharing rule and returns to scale parameters. More recently, Lise and Seitz (2004) estimate a collective model on U.K. data, use the estimates to calculate individual consumptions and then study the evolution of individual consumption inequality in the U.K. They conclude that failure to account for unequal intra-household allocations leads one to overestimate the growth in inequality since the 1970s. One possible concern with this important paper is the assumptions they make in order to estimate parameters of the collective model. In particular, they assume that leisure is an assignable good. Our estimates allow the calculation (from equation (8)) of a private consumption measure that allows for both returns to scale in consumption and unequal intra-household allocation. Of course, our private consumption measure also depends on the (different) assumptions we make to identify sharing rule and returns to scale parameters.

[14] Haddad and Kanbar (1990) is one well-known study using Philippine data.

To illustrate, we calculated Gini coefficients for individual inequality among singles and couples in our ten countries for 2001. These are displayed in Figure 4. We calculate Gini-coefficients for three measures of individual resources. First is equivalised income, where we use the common $\sqrt{n}$ equivalence scale (so each single person is assumed to consume their net income, and each member of a couple is assumed to enjoy consumption of $1/\sqrt{2} \approx 70\%$ of household net income.) This quantity is measured on the horizontal axis in Figure 4. Next, we use our (country-specific) estimates of returns to scale and sharing rule parameters to calculate personal consumption for each person in our data (using Equations 3 and 5 for singles and couples respectively). In Figure 4, the Gini for personal consumption is plotted against the Gini for equivalized income with squares (so that former is read off the y-axis and the latter is read off the x-axis). The square for each country is labeled with the country's acronym. The difference between these two Ginis is the vertical distance of the relevant square from the 45 degree line. In Figure 4, all of the squares lie above the 45 degree line, indicating that, in every country, personal consumption is more inequitably distributed than equivalent income. In some cases the differences are very small (for example, the U.K and Greece) while in other cases they are larger (for example, Denmark and the Netherlands). Changing the measure of individual resources from equivalent income to personal consumption leads to only small changes in the rank ordering of countries. There are reversals in the relative positions of Denmark and the Netherlands, Italy and France, and Belgium and Greece; but there are no large changes in position.

Personal consumption, as we calculate it, differs from equivalent income both because we allow for inequitable allocation of consumption within couples and because we use country-specific estimated equivalent scales rather than $\sqrt{n}$. The choice of equivalence scale can have a significant impact on the amount of "between-group" inequality (between singles and couples.)

To decompose the effects of these two changes, we calculate, for each individual, an "intermediate case". To do this, we return to the minimum-distance step that recovers the structural parameters from our reduced forms and impose that $A = \frac{\sqrt{2}}{2} \approx 0.7$. We then use this value of A (in every country) and the corresponding (country-specific) restricted estimates of the sharing rule parameters to calculate the "intermediate case" resource measure for every individual in every country.[15]

Country-specific Gini coefficients for this "intermediate case" are also plotted against Gini coefficients for equivalent income in Figure 4, with this combination plotted as circles. Thus the figure can be read as follows: for each country, the vertical distance from the 45 degree line to the circle gives the increase in measured inequality that results from accounting for intra-household inequality but using a standard ($\sqrt{n}$) equivalence scale. The vertical distance from the circle to the square gives the additional increment in inequality that results from also using the country-specific estimate of the equivalence scale (i.e., the equivalence scale implied by the country-specific estimates of the returns to scale parameter, A). An examination of Figure 4 reveals that in most countries, the two changes contribute roughly equally to the increase in inequality (if any) as one moves from equivalised income to personal consumption. The exception is Belgium, where using the estimated equivalent scale has no effect, while accounting for intra-household inequality has a substantial impact.

[15] The restricted estimates of the structural parameters are those values of the structural parameters that minimize the relevant distance given the restriction on the parameter A.

7. Conclusion

In this paper we have used survey data on financial satisfaction to estimate a collective household model. The parameters of interest are the household consumption technology (returns to scale in consumption) and the parameters of a sharing rule that determines the allocation of resources within households.

Estimation of the model delivers plausible estimates of the returns to scale in household consumption. We also find significant effects of female income shares on the sharing rule, in the majority of countries. This is evidence against the unitary model and emphasizes the importance of modeling intra-household allocation.

Our results add to the existing literature on collective intra-household models at least two ways. First, our approach uses a different kind of data, and in particular different identifying assumptions. Thus, the range of evidence against the unitary model is expanded, as is the set of alternatives for researchers wishing to estimate collective models. Some of the assumptions made in the previous literature are quite strong (for example, that non-market time is private leisure), so that alternative identification strategies (even if they involve different strong assumptions) are very useful.

The second virtue of our approach is that is computationally very straight forward and the data requirements are quite modest. This opens up possibilities for the wide use of these kinds of estimates. We were able to generate estimates of sharing rule parameters and the returns to scale in household consumption for ten European countries. We illustrated how these estimates could be used to conduct international inequality comparisons that account intra-household allocation. A second line of possible research is to relate differences in sharing rule parameters across countries to institutions such as divorce law. This would build on research based on U.S. data by

Gray (1998) and Chiappori et al., (2002). Because our methodology can generate sharing rule estimates for many jurisdictions, it expands the range of institutional factors that can be studied.

A surprising feature of our results is that the sharing rule in most countries favours women, in the sense that at average household income and an average female share of income, the female share of a couple's total consumption is greater than one half. This is, in fact, consistent with earlier work, including BCL and Lise and Seitz (2004). In BCL's analysis this finding reflects the fact that couples' spending patterns more closely resemble the spending patterns of single women than the spending patterns of single men. In the case of Lise and Seitz it may reflect the fact that non-market time is interpreted as an assignable good (of which women enjoy more). In our analysis, the same finding reflects a third distinct data feature. In particular, it seems that, holding per capita income constant, both men and women experience greater financial satisfaction if cohabiting, but the increment for women is larger. Our structural model interprets the increment in financial satisfaction from cohabiting that is common to men and women as returns to scale to in consumption. It attributes the gender differential in this increment to the sharing rule. Since the increment is larger for women, the estimated sharing rule favours them in the sense described above.

It is surprising that collective models estimated in such different ways should all indicate that sharing rules favour women – most researchers' prior would probably be the opposite. Understanding these findings is an obvious priority for future research.

Another important avenue for future research – and one that may help resolve the puzzle just noted – is to incorporate subjective information on satisfaction with other life domains (such as time, stress and health.) Aggregating information on satisfaction in multiple domains poses additional, difficult, methodological problems, and we reserve this for future work.

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Appendix

In the paper, the following relation between the reduced form and structural parameters has been derived (country index suppressed, up to now we assume that that the structural parameters are time invariant (no variation in relative prices over time)):

$$
\begin{aligned}
\pi_1 &= \beta \\
\pi_2 &= -\beta(\ln(A) + \ln(1+e^{\gamma_0})) \\
\pi_3 &= \beta\gamma_0 \\
\pi_4 &= \frac{\beta\gamma_1}{1+e^{\gamma_0}} \\
\pi_5 &= \frac{\beta\gamma_2}{1+e^{\gamma_0}} \\
\pi_6 &= \frac{\beta\gamma_1 e^{\gamma_0}}{1+e^{\gamma_0}} \\
\pi_7 &= \frac{\beta\gamma_2 e^{\gamma_0}}{1+e^{\gamma_0}}
\end{aligned}
$$

The reduced parameter vector $\mathring{\pi} = (\mathring{\pi}_1, \ldots, \mathring{\pi}_7)'$ has been estimated by means of pooled ordered probit. Notice that $\mathring{\pi}_j = \pi_j / \sqrt{1+\sigma_\zeta^2}$) where σ_ζ^2 is the variance of the random effect, cf. equation (17). In other words, we have $\hat{\mathring{\pi}}$ and the estimated covariance matrix $\hat{V}(\hat{\mathring{\pi}})$. For estimation purposes, it is handy to rewrite the system above and obtain an alternative reduced form parameter vector $\pi^* = (\pi_1^*, \ldots, \pi_7^*)'$. This system can be rewritten as follows:

$$\pi_1^* = \overset{\circ}{\pi}_1 = \beta^* = \beta / \sqrt{1+\sigma_\zeta^2}$$

$$\pi_2^* = -\left(\frac{\overset{\circ}{\pi}_2}{\overset{\circ}{\pi}_1} + \ln\left(1 + e^{\frac{\overset{\circ}{\pi}_3}{\overset{\circ}{\pi}_1}}\right)\right) = \ln(A)$$

$$\pi_3^* = \frac{\overset{\circ}{\pi}_3}{\overset{\circ}{\pi}_1} = \gamma_0$$

$$\pi_4^* = \frac{\overset{\circ}{\pi}_4 \left(1 + e^{\frac{\overset{\circ}{\pi}_3}{\overset{\circ}{\pi}_1}}\right)}{\overset{\circ}{\pi}_1} = \gamma_1$$

$$\pi_5^* = \frac{\overset{\circ}{\pi}_5 \left(1 + e^{\frac{\overset{\circ}{\pi}_3}{\overset{\circ}{\pi}_1}}\right)}{\overset{\circ}{\pi}_1} = \gamma_2$$

$$\pi_6^* = -\frac{\overset{\circ}{\pi}_6 \left(1 + e^{\frac{\overset{\circ}{\pi}_3}{\overset{\circ}{\pi}_1}}\right)}{\overset{\circ}{\pi}_1 e^{\frac{\overset{\circ}{\pi}_3}{\overset{\circ}{\pi}_1}}} = \gamma_1$$

$$\pi_7^* = -\frac{\overset{\circ}{\pi}_7 \left(1 + e^{\frac{\overset{\circ}{\pi}_3}{\overset{\circ}{\pi}_1}}\right)}{\overset{\circ}{\pi}_1 e^{\frac{\overset{\circ}{\pi}_3}{\overset{\circ}{\pi}_1}}} = \gamma_2$$

From this set of equations it becomes clear that the parameter β cannot be identified. However, the parameter β^* $(= \beta / \sqrt{1+\sigma_\zeta^2}\,)$ can be estimated. Given $\hat{\overset{\circ}{\pi}}$, consistent estimates for π^* can be obtained in a trivial way. The variance covariance matrix of $\hat{\pi}^*$, $\hat{V}(\hat{\pi}^*)$ can be obtained in the following way:

$$\hat{V}(\hat{\pi}^*) = \hat{F}\,\hat{V}(\hat{\overset{\circ}{\pi}})\,\hat{F}'$$

where

$$F=\frac{\partial\pi^*}{\partial\mathring{\pi}'}=\begin{pmatrix}
1 & 0 & 0 & 0 & 0 & 0 & 0\\
\frac{1}{\mathring{\pi}_1^2}\left(\mathring{\pi}_2+\frac{\mathring{\pi}_3}{1+e^{\frac{\mathring{\pi}_3}{\mathring{\pi}_1}}}\right) & -\frac{1}{\mathring{\pi}_1} & -\frac{1}{\mathring{\pi}_1\left(1+e^{\frac{\mathring{\pi}_3}{\mathring{\pi}_1}}\right)} & 0 & 0 & 0 & 0\\
-\frac{\mathring{\pi}_3}{\mathring{\pi}_1^{2}} & 0 & \frac{1}{\mathring{\pi}_1} & 0 & 0 & 0 & 0\\
-\frac{\mathring{\pi}_3\mathring{\pi}_4e^{\frac{\mathring{\pi}_3}{\mathring{\pi}_1}}}{\mathring{\pi}_1^{3}}-\frac{\mathring{\pi}_4\left(1+e^{\frac{\mathring{\pi}_3}{\mathring{\pi}_1}}\right)}{\mathring{\pi}_1^{2}} & 0 & \frac{\mathring{\pi}_4e^{\frac{\mathring{\pi}_3}{\mathring{\pi}_1}}}{\mathring{\pi}_1^{2}} & \frac{1+e^{\frac{\mathring{\pi}_3}{\mathring{\pi}_1}}}{\mathring{\pi}_1} & 0 & 0 & 0\\
-\frac{\mathring{\pi}_3\mathring{\pi}_5e^{\frac{\mathring{\pi}_3}{\mathring{\pi}_1}}}{\mathring{\pi}_1^{3}}-\frac{\mathring{\pi}_5\left(1+e^{\frac{\mathring{\pi}_3}{\mathring{\pi}_1}}\right)}{\mathring{\pi}_1^{2}} & 0 & \frac{\mathring{\pi}_5e^{\frac{\mathring{\pi}_3}{\mathring{\pi}_1}}}{\mathring{\pi}_1^{2}} & 0 & \frac{1+e^{\frac{\mathring{\pi}_3}{\mathring{\pi}_1}}}{\mathring{\pi}_1} & 0 & 0\\
\frac{\mathring{\pi}_6\left(1-\frac{\mathring{\pi}_3}{\mathring{\pi}_1}+e^{\frac{\mathring{\pi}_3}{\mathring{\pi}_1}}\right)}{\mathring{\pi}_1^2e^{\frac{\mathring{\pi}_3}{\mathring{\pi}_1}}} & 0 & \frac{\mathring{\pi}_6}{\mathring{\pi}_1^2e^{\frac{\mathring{\pi}_3}{\mathring{\pi}_1}}} & 0 & 0 & -\frac{1+e^{\frac{\mathring{\pi}_3}{\mathring{\pi}_1}}}{\mathring{\pi}_1e^{\frac{\mathring{\pi}_3}{\mathring{\pi}_1}}} & 0\\
\frac{\mathring{\pi}_7\left(1-\frac{\mathring{\pi}_3}{\mathring{\pi}_1}+e^{\frac{\mathring{\pi}_3}{\mathring{\pi}_1}}\right)}{\mathring{\pi}_1^2e^{\frac{\mathring{\pi}_3}{\mathring{\pi}_1}}} & 0 & \frac{\mathring{\pi}_7}{\mathring{\pi}_1^2e^{\frac{\mathring{\pi}_3}{\mathring{\pi}_1}}} & 0 & 0 & 0 & -\frac{1+e^{\frac{\mathring{\pi}_3}{\mathring{\pi}_1}}}{\mathring{\pi}_1e^{\frac{\mathring{\pi}_3}{\mathring{\pi}_1}}}
\end{pmatrix}$$

Given $\hat{\pi}^*$ and $\hat{V}(\hat{\pi}^*)$ estimation of the structural parameter vector $\theta=(\beta^*,\ln(A),\gamma_0,\gamma_1,\gamma_2)'$ can be done by means of feasible GLS:

$$\hat{\theta}=\left(X'\hat{V}(\hat{\pi}^*)^{-1}X\right)^{-1}X'\hat{V}(\hat{\pi}^*)^{-1}\hat{\pi}^*$$

where

$$X = \begin{pmatrix} 1 & 0 & 0 & 0 & 0 \\ 0 & 1 & 0 & 0 & 0 \\ 0 & 0 & 1 & 0 & 0 \\ 0 & 0 & 0 & 1 & 0 \\ 0 & 0 & 0 & 0 & 1 \\ 0 & 0 & 0 & 1 & 0 \\ 0 & 0 & 0 & 0 & 1 \end{pmatrix}$$

Obviously, $\hat{V}(\hat{\theta}) = \left(X'\hat{V}(\hat{\pi}^{*})^{-1} X \right)^{-1}$

Tables

Table 1: Distribution of Financial Satisfaction, by Country, Gender and Cohabiting, Singles and Couples (no children), ECHP 1994-2001 (column %)

	Single Men	Single Women	Cohabiting Men	Cohabiting Women
		Netherlands		
1. very dissatisfied	3.7	5.8	1.2	1.0
2. dissatisfied	7.4	9.6	2.7	2.3
3. A bit dissatisfied	13.3	16.7	8.0	7.2
4. A bit satisfied	24.7	26.0	23.7	20.7
5. Satisfied	35.9	29.7	44.4	45.2
6. Very satisfied	14.9	12.2	19.9	23.7
no. obs	3,991	6,220	10,749	10,747
		United Kingdom		
1. finding it very difficult	3.0	3.0	1.2	1.1
2. finding it quite difficult	7.3	6.3	2.9	3.3
3. just about getting by	25.8	29.7	22.9	19.7
4. doing alright	31.9	31.3	32.3	34.9
5. living Comfortably	31.9	29.7	40.7	41.0
no. obs	3,777	6,071	9,308	9,318
		Spain		
1. very dissatisfied	10.4	14.7	9.1	9.8
2. dissatisfied	15.8	20.9	16.3	16.8
3. A bit dissatisfied	22.4	24.5	24.9	24.5
4. A bit satisfied	24.0	21.2	25.3	24.5
5. Satisfied	20.7	14.4	19.4	19.1
6. Very satisfied	6.8	4.3	5.0	5.4
no. obs	2,271	4,473	8,834	8,867

χ^2 **Tests of Independence**:
Single men versus single women rejects in all countries
Single men versus cohabiting men rejects in all countries
Cohabiting men versus cohabiting women does not reject in Spain

Table 2: Within-Household Patterns of Financial Satisfaction
Couples (no children), ECHP 1994-2001

Netherlands
(n=10,737, Row %)

	Female partner						
Male partner	1	2	3	4	5	6	**Total**
1. Very dissatisfied	41.6	22.4	17.6	10.4	6.4	1.6	100
2. dissatisfied	7.9	33.5	30.7	16.7	8.2	3.1	100
3. a bit dissatisfied	2.3	7.3	36.2	34.5	17.3	2.4	100
4. a bit satisfied	0.3	1.2	9.4	44.6	38.6	5.9	100
5. satisfied	0.1	0.4	2.0	13.9	66.4	17.4	100
6. Very satisfied	0.0	0.1	0.7	3.3	24.5		100

United Kingdom
(n=9,298) (Row %)

	Female partner					
Male partner	1	2	3	4	5	**Total**
1. finding it very difficult	30.6	31.5	27.0	10.8	0	100
2. finding it quite difficult	9.9	37.4	37.7	12.8	2.2	100
3. just about getting by	1.5	5.7	54.1	27.9	10.8	100
4. doing alright	0.1	1.4	13.3	56.1	29.2	100
5. living Comfortably	0.1	0.3	3.9	24.2	71.5	100

Spain
(n=8,782) (Row %)

	Female partner						
Male partner	1	2	3	4	5	6	**Total**
1. Very dissatisfied	60.4	21.2	10.8	4.3	2.3	1.1	100
2. dissatisfied	13.6	49.4	22.4	9.6	4.1	1.0	100
3. a bit dissatisfied	4.5	17.5	49.9	19.4	7.6	1.0	100
4. a bit satisfied	2.7	7.0	21.6	49.1	17.2	2.4	100
5. satisfied	1.1	3.4	9.0	24.7	54.6	7.2	100
6. Very satisfied	0.7	1.8	3.9	7.1	29.5	57.1	100

Kappa Statistics

	Expected Agreement, %	**Actual Agreement, %**	**Kappa (SE)**
Netherlands	30.4	58.6	0.41 (0.006)
United Kingdom	32.6	61.1	0.42 (0.007)
Spain	19.9	51.8	0.40 (0.005)

Table 3: Changes in Financial Satisfaction
Singles and Couples (no children), ECHP 1994-2001

Note: change in satisfaction >=2 means considerable improvement, <=-2 means considerable deterioration

Netherlands
(column %)

	Male				female			
Change in Satisfaction	single single	single cohabiting	cohabiting single	cohabiting cohabiting	single single	Single cohabiting	cohabiting single	cohabiting cohabiting
<=-2	5.1	7.9	8.7	4.0	5.8	3.3	16.9	3.6
-1	19.0	17.9	30.2	18.5	19.7	10.8	27.9	18.4
0	46.8	40.0	36.5	52.1	43.8	31.7	37.2	53.7
1	21.2	17.9	16.7	20.6	23.0	29.2	12.6	20.1
>=2	7.9	16.4	7.9	4.8	7.7	25.0	5.5	4.2
Obs.	2,891	140	126	7,971	4,704	120	183	7,966

United Kingdom
(column %)

	Male				female			
Change in Satisfaction	single single	single cohabiting	cohabiting single	cohabiting cohabiting	single single	Single cohabiting	cohabiting single	cohabiting cohabiting
<=-2	3.5	2.9	8.9	2.9	3.7	2.2	7.9	2.8
-1	16.2	14.3	23.1	15.5	16.7	16.4	28.3	15.0
0	57.5	43.6	42.0	60.2	54.1	36.6	42.9	61.2
1	18.4	30.0	18.9	17.8	20.3	30.6	17.3	17.4
>=2	4.4	9.3	7.1	3.7	5.21	14.2	3.7	3.6
Obs.	2,743	140	169	7,150	4,741	134	191	7,168

Spain
(column %)

	Male				female			
Change in Satisfaction	single single	single cohabiting	cohabiting single	cohabiting cohabiting	single single	Single cohabiting	cohabiting single	cohabiting cohabiting
<=-2	13.6	13.3	5.2	12.6	13.2	15.8	28.7	12.9
-1	20.0	22.2	26.0	21.3	21.3	15.8	24.3	21.3
0	31.7	20.0	39.0	30.2	30.0	29.0	19.9	30.1
1	20.0	26.7	11.7	21.8	21.5	23.7	13.2	20.8
>=2	14.8	17.8	18.2	14.1	14.0	15.8	14.0	14.9
Obs.	1,621	45	77	6,403	3,401	38	136	6,433

Test of Gender Equality
(p-values)

	S-P	**P-S**
Netherlands	0.020	0.246
United Kingdom	0.635	0.523
Spain	0.856	<0.001

Table 4: Selected Means, by Country
Singles and Couples (no children), ECHP 1994-2001

	ln real household income at PPP	Female income share (Couples only)
Denmark	9.74	0.41
Netherlands	9.81	0.29
Belgium	9.74	0.27
France	9.72	0.30
Ireland	9.46	0.27
Italy	9.49	0.29
Greece	8.98	0.24
Spain	9.31	0.20
Portugal	8.88	0.32
United Kingdom	9.74	0.37

Notes: Household income is the sum of personal incomes. Personal income is net, and is the sum all income components, over the year preceding the survey.

Table 5a: Reduced Form Parameter Estimates, Base Specification

	DK	NL	BE	FR	IE	IT	GR	ES	PT	UK
Ln(income$_{it}$) (π^1)	0.374***	0.374***	0.374***	0.374***	0.374***	0.374***	0.374***	0.374***	0.374***	0.374***
	(-0.044)	(-0.044)	(-0.044)	(-0.044)	(-0.044)	(-0.044)	(-0.044)	(-0.044)	(-0.044)	(-0.044)
Couple$_{it}$ (π^2)	-0.068	-0.068	-0.068	-0.068	-0.068	-0.068	-0.068	-0.068	-0.068	-0.068
	(-0.047)	(-0.047)	(-0.047)	(-0.047)	(-0.047)	(-0.047)	(-0.047)	(-0.047)	(-0.047)	(-0.047)
Couple$_{it}$*female$_{it}$ (π^3)	0.259***	0.259***	0.259***	0.259***	0.259***	0.259***	0.259***	0.259***	0.259***	0.259***
	(-0.063)	(-0.063)	(-0.063)	(-0.063)	(-0.063)	(-0.063)	(-0.063)	(-0.063)	(-0.063)	(-0.063)
Couple*female*ln(income) (π^4)	0.155**	0.155**	0.155**	0.155**	0.155**	0.155**	0.155**	0.155**	0.155**	0.155**
	(-0.063)	(-0.063)	(-0.063)	(-0.063)	(-0.063)	(-0.063)	(-0.063)	(-0.063)	(-0.063)	(-0.063)
Couple$_{it}$*female$_i$*income_share_female$_{it}$ (π^4)	-0.347**	-0.347**	-0.347**	-0.347**	-0.347**	-0.347**	-0.347**	-0.347**	-0.347**	-0.347**
	(-0.14)	(-0.14)	(-0.14)	(-0.14)	(-0.14)	(-0.14)	(-0.14)	(-0.14)	(-0.14)	(-0.14)
Couple$_{it}$*male$_i$*ln(income$_{it}$) (π^6)	0.285***	0.285***	0.285***	0.285***	0.285***	0.285***	0.285***	0.285***	0.285***	0.285***
	(-0.064)	(-0.064)	(-0.064)	(-0.064)	(-0.064)	(-0.064)	(-0.064)	(-0.064)	(-0.064)	(-0.064)
Couple$_{it}$*male$_i$*income_share_female$_{it}$ (π^7)	-0.772***	-0.772***	-0.772***	-0.772***	-0.772***	-0.772***	-0.772***	-0.772***	-0.772***	-0.772***
	(-0.14)	(-0.14)	(-0.14)	(-0.14)	(-0.14)	(-0.14)	(-0.14)	(-0.14)	(-0.14)	(-0.14)
Female$_i$	-0.117**	-0.117**	-0.117**	-0.117**	-0.117**	-0.117**	-0.117**	-0.117**	-0.117**	-0.117**
	(-0.048)	(-0.048)	(-0.048)	(-0.048)	(-0.048)	(-0.048)	(-0.048)	(-0.048)	(-0.048)	(-0.048)
(Upper) secondary education$_i$	0.0582	0.0582	0.0582	0.0582	0.0582	0.0582	0.0582	0.0582	0.0582	0.0582
	(-0.037)	(-0.037)	(-0.037)	(-0.037)	(-0.037)	(-0.037)	(-0.037)	(-0.037)	(-0.037)	(-0.037)
Post secondary education$_i$	0.00781	0.00781	0.00781	0.00781	0.00781	0.00781	0.00781	0.00781	0.00781	0.00781
	(-0.033)	(-0.033)	(-0.033)	(-0.033)	(-0.033)	(-0.033)	(-0.033)	(-0.033)	(-0.033)	(-0.033)
Age$_{it}$	-0.0116**	-0.0116**	-0.0116**	-0.0116**	-0.0116**	-0.0116**	-0.0116**	-0.0116**	-0.0116**	-0.0116**
	(-0.0046)	(-0.0046)	(-0.0046)	(-0.0046)	(-0.0046)	(-0.0046)	(-0.0046)	(-0.0046)	(-0.0046)	(-0.0046)
Age$^2_{it}$	0.000334***	0.000334***	0.000334***	0.000334***	0.000334***	0.000334***	0.000334***	0.000334***	0.000334***	0.000334***
	(-4.6E-05)	(-4.6E-05)	(-4.6E-05)	(-4.6E-05)	(-4.6E-05)	(-4.6E-05)	(-4.6E-05)	(-4.6E-05)	(-4.6E-05)	(-4.6E-05)
Observations	18751	18751	18751	18751	18751	18751	18751	18751	18751	18751

Notes: Specification also contains time dummies; Standard errors in parentheses; Standard errors account for clustering;
*** p<0.01, ** p<0.05, * p<0.1; Estimation method: Pooled Ordered Probit.

Table 5b: Structural Parameter Estimates, Base Specification

	DK	NL	BE	FR	IE	IT	GR	ES	PT	UK
Sharing Rule Parameters										
Intercept	0.340**	0.519***	0.627***	0.261**	0.329**	0.171*	0.146*	0.379***	0.242**	0.144
(γ^0)	(0.17)	(0.10)	(0.21)	(0.12)	(0.17)	(0.089)	(0.076)	(0.13)	(0.11)	(0.13)
Ln(income)	-0.410**	-0.0862	-0.0704	-0.0505	0.0398	-0.0769	-0.0689	-0.258*	-0.0422	-0.236**
(γ^1)	(0.17)	(0.10)	(0.19)	(0.11)	(0.16)	(0.070)	(0.045)	(0.14)	(0.071)	(0.12)
Female Income Share	1.693***	0.298	0.289	0.808***	0.0799	0.422**	0.405***	0.822***	1.008***	0.721**
(γ^2)	(0.56)	(0.22)	(0.39)	(0.29)	(0.34)	(0.17)	(0.12)	(0.27)	(0.22)	(0.33)
Household Consumption Technology (returns to scale parameter) *(Given equal allocation, the equivalence scale is 2A)*										
A	0.459***	0.346***	0.292***	0.456***	0.605***	0.551***	0.659***	0.539***	0.767***	0.526***
	(0.045)	(0.023)	(0.036)	(0.030)	(0.053)	(0.025)	(0.026)	(0.040)	(0.042)	(0.038)
Estimated Female Consumption Shares, Mean Household Income and Alternative Female Income Shares										
$\eta^1(\overline{\ln(y)}, 0.25)$	0.517***	0.624***	0.651***	0.556***	0.581***	0.538***	0.537***	0.603***	0.543***	0.515***
	(0.054)	(0.025)	(0.048)	(0.030)	(0.041)	(0.022)	(0.019)	(0.031)	(0.028)	(0.036)
$\eta^1(\overline{\ln(y)}, 0.50)$	0.621***	0.641***	0.667***	0.605***	0.586***	0.564***	0.562***	0.651***	0.605***	0.560***
	(0.036)	(0.024)	(0.048)	(0.027)	(0.039)	(0.023)	(0.019)	(0.031)	(0.023)	(0.029)
$\eta^1(\overline{\ln(y)}, 0.75)$	0.714***	0.658***	0.683***	0.652***	0.591***	0.590***	0.587***	0.696***	0.663***	0.603***
	(0.041)	(0.029)	(0.056)	(0.033)	(0.047)	(0.028)	(0.021)	(0.037)	(0.025)	(0.035)
$\eta^1(\overline{\ln(y)}, \overline{w})$	0.584***	0.627***	0.652***	0.565***	0.582***	0.543***	0.536***	0.594***	0.560***	0.536***
	(0.041)	(0.024)	(0.048)	(0.029)	(0.040)	(0.022)	(0.019)	(0.032)	(0.027)	(0.031)
Overidentification test -p value	0.000	0.000	0.133	0.000	0.003	0.043	0.000	0.000	0.000	0.000

Notes: *** p<0.01, ** p<0.05, * p<0.1

Table 6a: Reduced Form Parameter Estimates, Chamberlain/Mundlak Estimator

	DK	NL	BE	FR	IE	IT	GR	ES	PT	UK
$\text{Ln(income}_{it})$ (π^1)	0.376***	0.488***	0.370***	0.401***	0.519***	0.608***	0.747***	0.366***	0.573***	0.426***
	(-0.044)	(-0.032)	(-0.048)	(-0.028)	(-0.061)	(-0.035)	(-0.032)	(-0.033	(-0.035)	(-0.03)
Couple_{it} (π^2)	-0.0739	-0.135**	-0.0495	-0.199***	-0.650***	-0.163**	-0.293***	-0.154*	-0.359***	-0.192***
	(-0.067)	(-0.068)	(-0.087)	(-0.055)	(-0.13)	(-0.078)	(-0.088)	(-0.084	(-0.094)	(-0.059)
$\text{Couple}_{it}\text{*female}_{it}$ (π^3)	0.116	0.305***	0.119	0.192**	0.545***	0.173	-0.0053	0.195*	0.048	0.108
	(-0.09)	(-0.091)	(-0.13)	(-0.078)	(-0.16)	(-0.11)	(-0.11)	(-0.12	(-0.12)	(-0.084)
Couple*female*ln(income) (π^4)	0.151**	0.193***	-0.0273	0.136***	0.02	-0.0216	-0.0277	0.205***	-0.0595	0.275***
	(-0.063)	(-0.051)	(-0.066)	(-0.043)	(-0.083)	(-0.045)	(-0.038)	(-0.048	(-0.044)	(-0.045)
$\text{Couple}_{it}\text{*female}_{i}\text{*income\_share\_female}_{it}$ (π^4)	-0.337**	-0.244***	-0.15	-0.162**	-0.390***	-0.0261	-0.145**	-0.135*	-0.315***	-0.198**
	(-0.14)	(-0.083)	(-0.11)	(-0.081)	(-0.12)	(-0.075)	(-0.064)	(-0.072	(-0.084)	(-0.099)
$\text{Couple}_{it}\text{*male}_{i}\text{*ln(income}_{it})$ (π^6)	0.285***	0.168***	0.0105	0.123***	0.0281	0.0219	0.027	0.216***	-0.0104	0.297***
	(-0.064)	(-0.049)	(-0.067)	(-0.041)	(-0.084)	(-0.045)	(-0.037)	(-0.049	(-0.043)	(-0.045)
$\text{Couple}_{it}\text{*male}_{i}\text{*income\_share\_female}_{it}$ (π^7)	-0.772***	-0.219***	-0.157	-0.428***	-0.239*	-0.247***	-0.426***	-0.343***	-0.784***	-0.407***
	(-0.14)	(-0.08)	(-0.11)	(-0.079)	(-0.13)	(-0.071)	(-0.064)	(-0.071	(-0.082)	(-0.096)
Female_i	-0.129**	-0.213***	-0.135**	-0.101**	-0.127*	-0.114**	-0.133**	-0.236***	-0.229***	-0.0952**
	(-0.051)	(-0.044)	(-0.062)	(-0.041)	(-0.069)	(-0.049)	(-0.056)	(-0.047	(-0.057)	(-0.048)
(Upper) secondary education_i	0.0596	0.253***	0.369***	0.295***	0.343***	0.410***	0.343***	0.347***	0.252***	0.233***
	(-0.037)	(-0.034)	(-0.044)	(-0.03)	(-0.061)	(-0.049)	(-0.042)	(-0.034	(-0.064)	(-0.03)
Post secondary education_i	0.00821	0.144***	0.141***	0.0962***	0.209***	0.270***	0.340***	0.209***	0.144***	0.195***
	(-0.033)	(-0.028)	(-0.035)	(-0.025)	(-0.045)	(-0.028)	(-0.035)	(-0.033	(-0.053)	(-0.033)
Age_{it}	-0.0119**	-0.0243***	-0.00952*	-0.0242***	-0.0283***	0.00922**	0.00122	0.00438	-0.0112***	-0.0278***
	(-0.0046)	(-0.0042)	(-0.0054)	(-0.0035)	(-0.0064)	(-0.0041)	(-0.0041)	(-0.0038	(-0.0042)	(-0.0039)
Age^2_{it}	0.000337***	0.000305***	0.000248***	0.000324***	0.000423***	-5.9E-05	1.11E-05	3.99E-05	0.0000946**	0.000365***
	(-4.6E-05)	(-4.1E-05)	(-5.1E-05)	(-3.4E-05)	(-6.1E-05)	(-3.8E-05)	(-3.6E-05)	(-3.6E-05	(-3.9E-05)	(-3.8E-05)
$\overline{\text{Couple}}_i$	0.00562	0.160**	0.113	0.208***	0.433***	0.0584	0.0623	-0.0381	0.0705	0.139*
	(-0.078)	(-0.076)	(-0.1)	(-0.064)	(-0.14)	(-0.088)	(-0.1)	(-0.092	(-0.11)	(-0.073)
$\overline{\text{Couple*female}}_i$	0.163	0.0563	0.139	-0.0609	-0.276	-0.0857	0.108	0.0578	0.0944	0.0751
	(-0.11)	(-0.1)	(-0.15)	(-0.089)	(-0.18)	(-0.12)	(-0.13)	(-0.13	(-0.13)	(-0.1)
Joint Statistical significance, Mundlak terms - p value	0.077	0.001	0.039	0	0.003	0.761	0.088	0.894	0.1	0.002
Observations	18751	31346	15793	35717	9243	23564	21598	24434	22276	28011

Notes: Specification also contains time dummies; Robust standard errors in parentheses; Standard errors account for clustering;
*** $p<0.01$, ** $p<0.05$, * $p<0.1$;

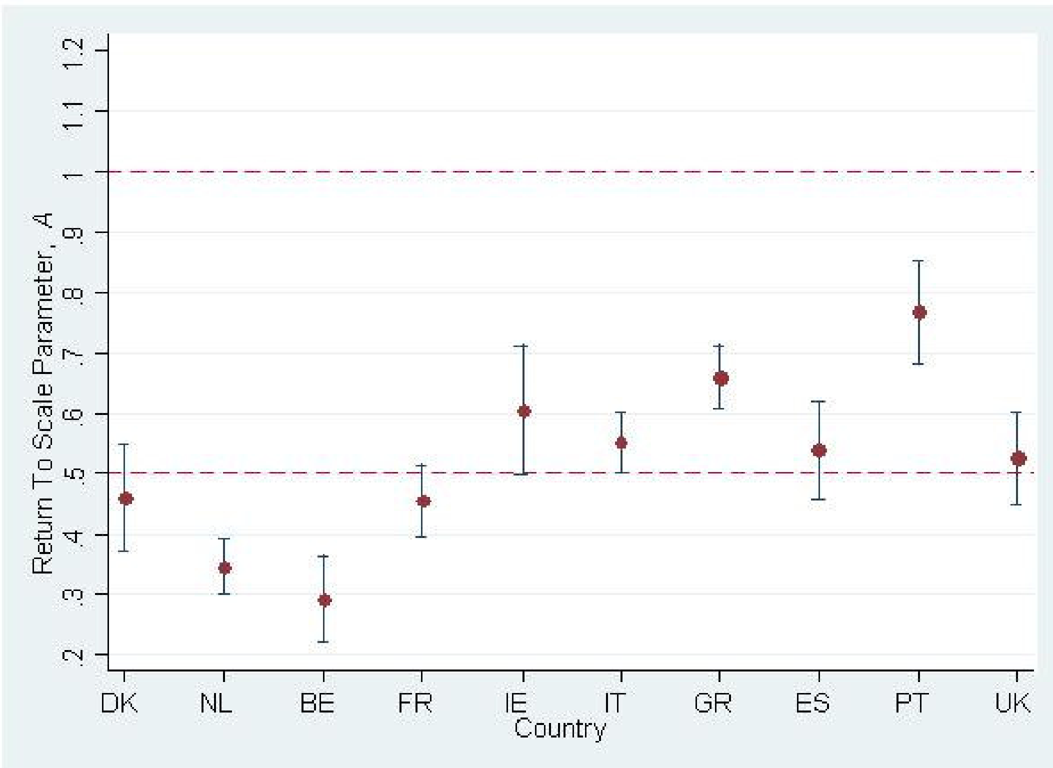
Table 6b: Structural Parameter Estimates, Chamberlain/Mundlak Estimator

	DK	NL	BE	FR	IE	IT	GR	ES	PT	UK
Sharing Rule Parameters										
Intercept	0.0833	0.240	0.213	0.219	0.472*	0.323*	0.192	0.0390	0.457**	-0.0406
(γ^0)	(0.24)	(0.17)	(0.36)	(0.19)	(0.26)	(0.17)	(0.15)	(0.28)	(0.20)	(0.19)
Ln(income)	-0.372*	-0.155	-0.0908	-0.115	0.00175	-0.0766	-0.0736*	-0.356*	-0.0754	-0.255
(γ^1)	(0.21)	(0.12)	(0.19)	(0.12)	(0.16)	(0.070)	(0.045)	(0.20)	(0.071)	(0.16)
Female Income Share	1.288**	0.375	0.217	1.008***	0.422	0.436**	0.289**	0.919***	0.542*	0.691*
(γ^2)	(0.64)	(0.23)	(0.43)	(0.30)	(0.33)	(0.17)	(0.13)	(0.31)	(0.28)	(0.36)
Household Consumption Technology (returns to scale parameter) *(Given equal allocation, the equivalence scale is 2A)*										
A	0.595***	0.477***	0.487***	0.632***	0.766***	0.566***	0.778***	0.565***	1.003***	0.689***
	(0.074)	(0.049)	(0.088)	(0.062)	(0.14)	(0.050)	(0.060)	(0.094)	(0.10)	(0.068)
Estimated Female Consumption Shares, Mean Household Income and Alternative Female Income Shares										
$\eta^1(\overline{\ln(y)}, 0.25)$	0.469***	0.556***	0.552***	0.543***	0.614***	0.575***	0.548***	0.521***	0.604***	0.470***
	(0.077)	(0.043)	(0.089)	(0.048)	(0.063)	(0.043)	(0.037)	(0.067)	(0.051)	(0.053)
$\eta^1(\overline{\ln(y)}, 0.50)$	0.550***	0.579***	0.566***	0.604***	0.638***	0.602***	0.566***	0.578***	0.636***	0.513***
	(0.053)	(0.039)	(0.083)	(0.040)	(0.056)	(0.041)	(0.033)	(0.059)	(0.039)	(0.044)
$\eta^1(\overline{\ln(y)}, 0.75)$	0.628***	0.602***	0.579***	0.663***	0.662***	0.627***	0.584***	0.633***	0.666***	0.556***
	(0.051)	(0.039)	(0.085)	(0.039)	(0.056)	(0.041)	(0.032)	(0.055)	(0.032)	(0.044)
$\eta^1(\overline{\ln(y)}, \overline{w})$	0.521***	0.560***	0.553***	0.555***	0.616***	0.580***	0.548***	0.510***	0.612***	0.490***
	(0.060)	(0.042)	(0.089)	(0.046)	(0.062)	(0.043)	(0.037)	(0.069)	(0.047)	(0.048)
Overidentification test p value	0.000	0.000	0.138	0.000	0.022	0.062	0.000	0.000	0.000	0.000

Notes: *** p<0.01, ** p<0.05, * p<0.1

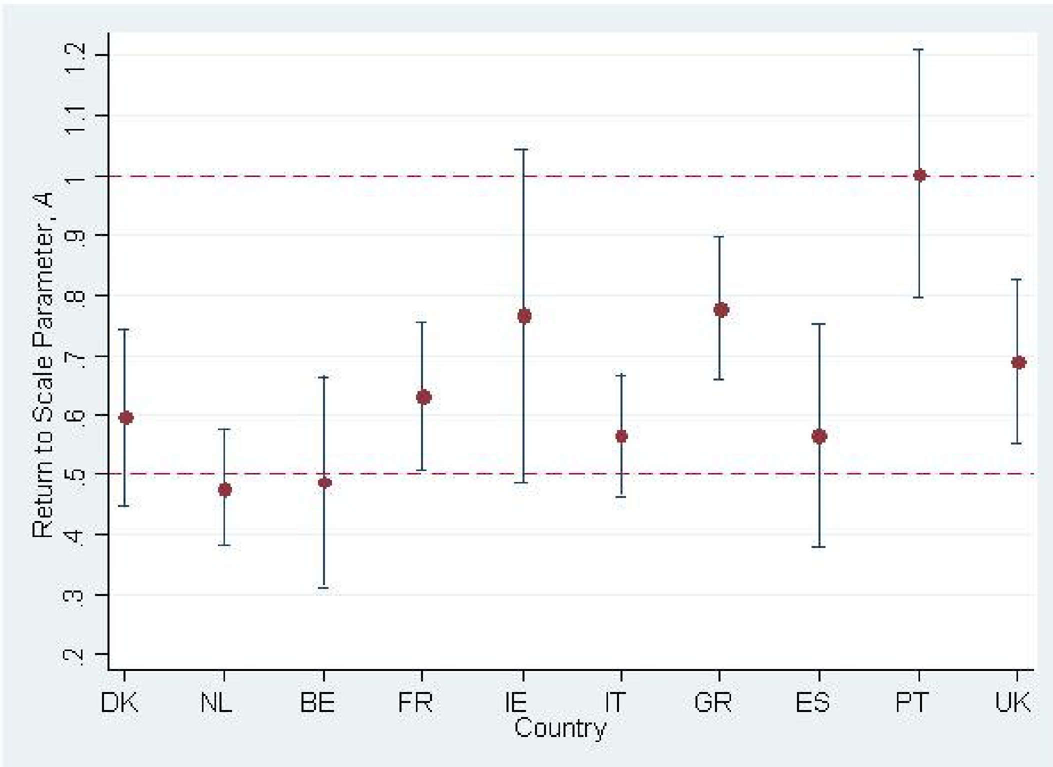
Figures

Figure 1: Estimates of the Returns to Scale in Household Consumption, Base Estimates, 10 European Countries, 1994-2001



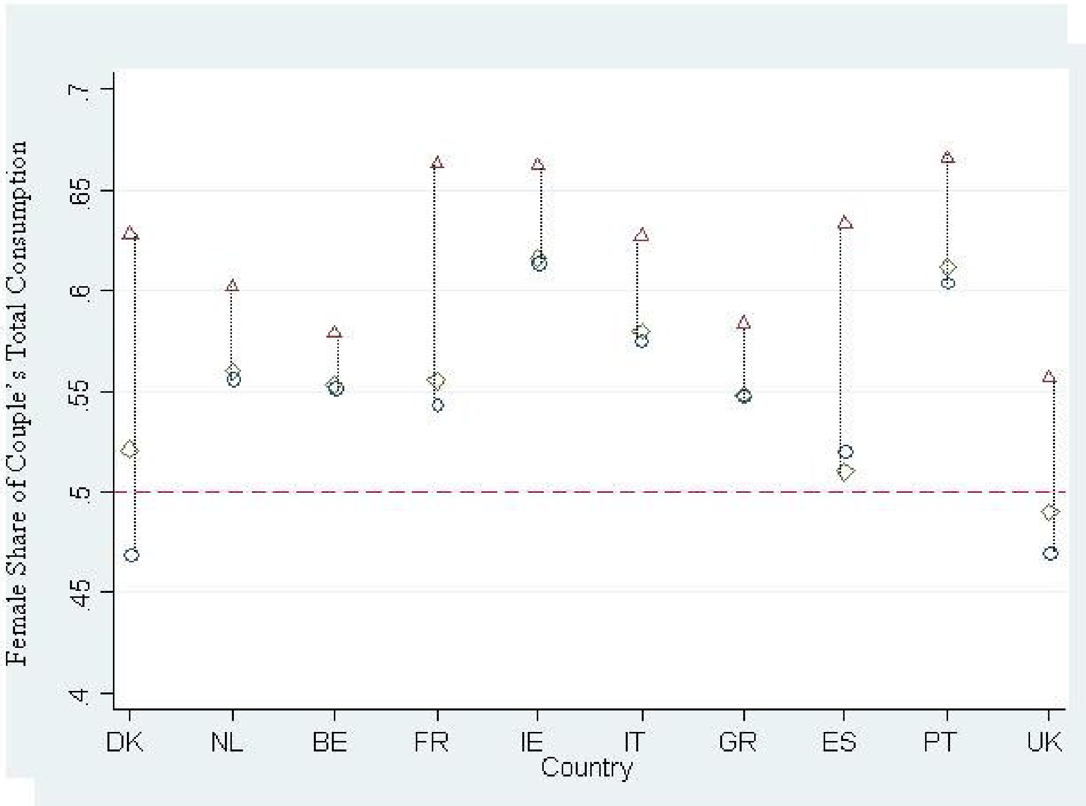
Explanation; Country-specific estimates of A and associated confidences intervals. A value of 0.5 indicates that a couple's potential total consumption is double their income; this in turn implies that all consumption is public. A value of 1 indicates no returns to scale; all consumption is private. A traditional "equivalence scale" is obtained by multiplying the parameter A by two. This gives the value by which a couple's income should be divided to give the income that a single person would require to have the same per capita total consumption. These estimates correspond to our base specification (Equation 14 in the text).

Figure 2: Estimates of the Returns to Scale in Household Consumption, Chamberlain/Mundlak Estimator, 10 European Countries, 1994-2001



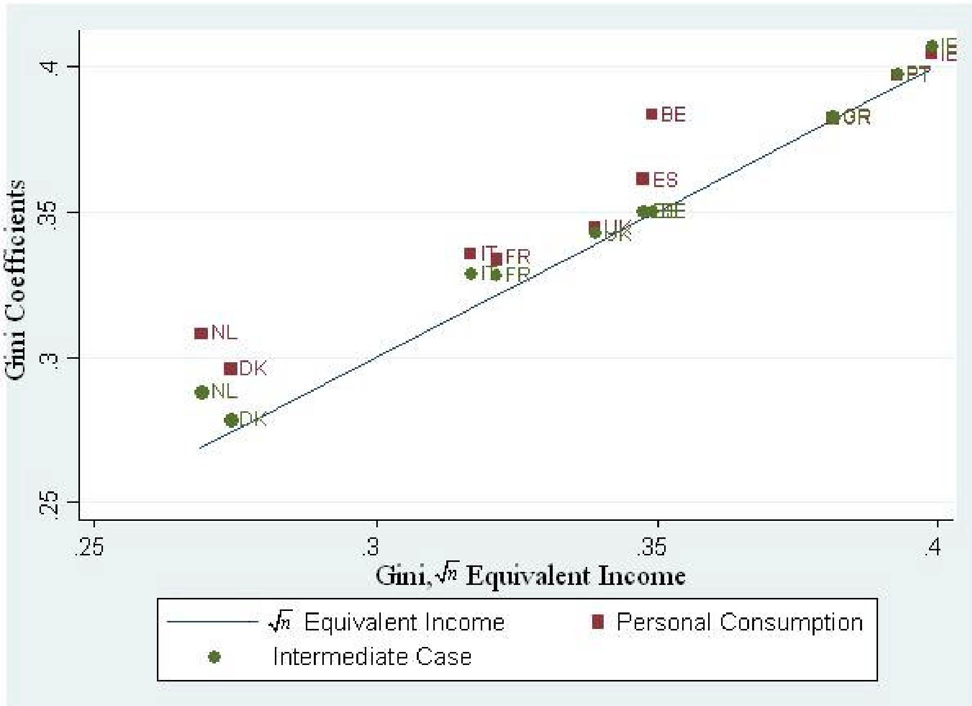
Explanation: Same as Figure 1 except that these estimates come from our second ("Chamberlain/Mundlak") specification. That specification includes person specific means of explanatory variables to control for potential correlation between those variables and time-invariant, unobserved individual effects. (See Equations 16 and 17 in the text.)

**Figure 3: Estimates of the Female Share of a Couple's Total Consumption,
10 European Countries, 1994-2001**



Explanation: Countries are arrayed along the horizontal axis. The female share of a couple's total consumption is measured on the vertical axis. This is calculated in three ways, all employing country-specific estimates of the sharing rule parameters. First, we assume that the female contributes 25 percentage of household income (plotted as a circle.); second, we assume that the female contributes 75 percent of household income (plotted as a triangle); and finally, we set the female contribution to household income equal to the country mean (plotted as a diamond). In all three cases, household income is set to the country specific means. Thus the diamonds give a sense of women's share of total consumption in an "average" couple in each country. The vertical distance between the circles and triangles give, for each country, a sense of the responsiveness of the sharing rule to the female income share (with a greater vertical distance indicating a more responsive sharing rule).

Figure 4: Inequality in Equivalent Income and Personal Consumption, 10 European Countries, 2001



Explanation: The Gini for personal consumption is plotted against the Gini for equiavlized income with squares (so that former is read off the y-axis and the latter is read off the x-axis). The square for each country is labeled with the country's acronym. The difference between these two Gini's is the vertical distance of the relevant square from the 45 degree line. Personal consumption, differs from equivalent income both because we allow for inequitable allocation of consumption within couples and because we use country-specific estimated equivalent scales rather than $\sqrt{n}$. To decompose the effects of these two changes, we calculate, for each individual, an "intermediate case" by imposing the $\sqrt{n}$ equivalence scale when we estimate our model (but allowing the data to determine sharing rule parameters, given this restriction.) Country-specific Gini coefficients for this "intermediate case" are also plotted against Gini coefficients for equivalent income in Figure 4, with this combination plotted as circles. Thus the figure can be read as follows: for each country, the vertical distance from the 45 degree line to the circle gives the increase in measured inequality that results from accounting for intra-household inequality but using a standard ($\sqrt{n}$) equivalence scale. The vertical distance from the circle to the square gives the additional increment in inequality that results from also using the country-specific estimate of the equivalence scale (ie., the equivalence scale implied by the country-specific estimates of the returns to scale parameter, A).