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The economies of scale of living together and how they are shared: estimates based on a collective household model

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Abstract How large are the economies of scale of living together? And how do partners share their resources? The first question is usually answered by equivalence scales which assume equal sharing of resources within the household. Recent evidence based on collective household models rejects this equal sharing assumption. This paper uses data on financial satisfaction to simultaneously estimate the sharing rule and the economies of scale in a collective household model. The estimates indicate substantial scale economies of living together. Furthermore, wives receive on average almost 50 % of household resources, but the estimated shares vary between 30 and 60 %. Female resource shares increase with the ratio of female to male wages. Consumption inequality is underestimated by 16 % if unequal sharing is ignored.

Keywords Collective household models · Sharing rule · Equivalence scale · Subjective data

JEL Codes D12 · C21 · D19

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1 Introduction

How large are the economies of scale associated with living together? And how are the resources allocated to household members? These questions are important for many economic topics such as designing welfare benefits, determining alimony and life insurance payments (Lewbel 2003), measuring inequality and poverty, and measuring the willingness to pay for public goods (Munro 2005). Traditionally, the first question is answered by equivalence scales which allow to compare well-being across households with different sizes. In this concept, within household distribution of resources is not an issue because they are implicitly assumed to be equally distributed. However, if there is unequal distribution within the household, traditional equivalence scales are biased and misleading in practice. In order to address this problem, a richer model of household behavior is necessary.

From a theoretical point of view, models of household behavior can be classified in two main groups. One strand is the general household utility framework, which is based on the assumptions that husband and wife have identical preferences.[1] Among those who challenge this simple model and recognize that each household member might have different preferences are Samuelson (1956), Apps and Rees (1988), Grossbard-Shechtman (1984), and Chiappori (1992). Moreover it is assumed in the general household utility framework that household utility is maximized subject to a single budget constraint. Accordingly, it is irrelevant who earns the money in the household. Redistribution of income within the household does not change household behavior. This so-called income pooling assumption has been frequently tested and rejected. Example include Attanasio and Lechene (2002), Bourguignon et al. (1993), Donni (2007), Fortin and Lacroix (1997), Lundberg et al. (1997), Phipps and Burton (1998), Thomas (1990), and Ward-Batts (2008). However, as Browning et al. (2006) point out, rejecting income pooling is not sufficient in order to reject the unitary model.

The second strand of household behavior is based on collective models, which allow family members to have distinct preferences. The standard collective approach assumes that the outcomes of decision making within the household are Pareto efficient (Chiappori 1988).[2] A standard result of welfare theory is that Pareto efficient decisions can be written as a constrained maximization of the weighted sum of individual utilities $\mu U^f(x^f) + U^m(x^m)$. The Pareto weight μ may depend on prices, total expenditures and on so-called distribution factors. These are defined as variables with no direct influence on preferences, technology or the budget constraint. From a bargaining perspective, the Pareto weight μ can be seen as a measure of the influence of household member f on the decision process. One difficulty with using μ as a measure of the share given to member f is that the magnitude of μ will depend on arbitrary cardinalizations of the utility functions U.

[1] Alternatively, there can also be an altruistic dictator who controls a larger share of the family income.

[2] Alternative models include cooperative household models as proposed by McElroy and Horney (1981) or Manser and Brown (1980), who describe household behavior as result of a Nash-bargaining game, where the bargaining power depends on the emerging expenditure patterns on the options outside marriage. Non-cooperative household models describe household behavior as a non-cooperative game with no binding and enforceable contracts between the household members and resulting inefficiencies.

Recently, Browning, Chiappori, and Lechene (2013, BCL hereafter) have shown that under specific assumptions there exists a unique Pareto weight corresponding to any sharing rule function η and vice versa. Household behavior can be described as allocating the fraction η to member f and the fraction $(1-\eta)$ to member m. The sharing rule η is invariant to cardinalizations of the utility function. This concept of a sharing rule is part of the standard collective household model. The BCL model is richer than the standard collective models due to the inclusion of a consumption technology function. That is, when two single individuals start cohabiting, their financial resources are altered due to two reasons: first, returns to scale in consumption let their joint consumption exceed the sum of what they could individually consume when living alone. Second, if there is unequal sharing of resources within the couple, one individuals consumption will rise by more than is implied by returns to scale, while the consumption of the other will rise by less. BCL derive the conditions necessary to estimate the consumption technology function, the sharing rule, and individual preferences and estimate their model using functional form assumptions for these functions.

The BCL model is hard to estimate, and consequently several simplifications have been proposed (Lewbel and Pendakur 2008; Cherchye et al. 2012). Lise and Seitz (2011) also follow the BCL approach, but focus on the demand for leisure and a composite consumption good. A different approach, also following the main ideas of BCL, has been suggested by Alessie, Crossley, and Hildebrand (2006, ACH hereafter). They use subjective panel data on financial satisfaction to estimate the parameters of the individual utility functions, the sharing rule and a consumption technology parameter. Already in the early 1970s van Praag and co-authors estimated individual utility functions based on subjective income evaluation data (e.g. Van Praag and Kapteyn 1973, for an early contribution, and Van Praag and Frijters 1999, for a survey). Nowadays, subjective data are increasingly used in the happiness literature (see Frey and Stutzer 2002, for an overview), but also in the collective household framework (e.g. Bonke and Browning 2009; Lührmann and Maurer 2007).

Our paper extends the ACH approach. By using a transformation of the financial satisfaction variable proposed by Van Praag and Ferrer-i-Carbonell (2004) we are able to directly estimate the structural parameters of the household consumption technology and resource allocation process by nonlinear least squares taking correlated unobserved heterogeneity into account. As ACL we monitor individuals over time and are therefore able to observe them moving in and out of cohabitation. Assuming stable (but possibly heterogeneous) individual preferences and satisfaction reporting, we can learn something about the household's economies of scale in consumption and sharing rule from the observed adjustments in the reported financial satisfaction. As is common with virtually all empirical analyzes of collective household models, we focus on households without children. The incorporation of children in these models is the major challenge for future research.

Using data from the Swiss Household Panel (SHP) we find substantial scale economies of living together. Total private consumption exceeds household income by almost 40 %. The consumption share of wives increases significantly with the wage ratio, which we define as the female hourly wage relative to the male hourly wage. On average, a wife has a share of 0.46 of total private consumption. About 30 % of the women have a consumption share that is significantly <0.5. Taking this unequal distribution of resources within households into account in the measurement of consumption inequality increases the Gini coefficient by roughly 15 % compared to the the classic approach, which assumes equal sharing.

The paper is structured as follows: Sect. 2 outlines the theoretical model and discusses indifference scales and the problems of measuring utility. Section 4 describes the data. The empirical model and the estimation results are discussed in Sects. 5, 6. Section 7 concludes.

2 A simple structural collective household model

We specify a collective household model which captures both returns to scale in household consumption and unequal allocation of resources within the household. As stated in the introduction, collective household models are based on individual preferences that are aggregated into household utility according to some rule. Hence, we first have to specify individual preferences. We assume that individual indirect utility can be described by a simple PIGLOG specification[3]

$$V_{it} = \mathbf{z}_{it}\alpha + \beta ln(x_{it}) + \mu_i + \varepsilon_{it}, \tag{1}$$

where V_{it} is utility of individual i in period t, z_{it} is a vector of characteristics, x_{it} are total consumption expenditure, μ_i is an individual fixed effect, and ε_{it} is the error term. Throughout this section we assume that V is observable. To simplify notation we drop the individual subscript i and the time subscript t unless it is necessary for clarity.

This specification implies that preferences are egoistic, that is people only care about their own consumption. Single individuals are assumed to consume their income in each period, i.e. $x = y^h$, where y^h denotes household income. The level of individual consumption in couple households, however, may depend on a sharing rule and returns to scale. Returns to scale exist if total private consumption of both household members f and m exceeds household income. This effect can be captured by writing

$$(x^m + x^f) = \tau y^h, \tag{2}$$

where x^i denotes consumption of person $i = f, m$. The scalar τ can be seen as representing a household consumption technology that transforms household income y_h into total household consumption $(x^m + x^f)$. The household consumption technology captures the fact that some goods are at least partly public, e.g. housing, household operation, transportation or newspapers. If $\tau = 1$ there are no returns to scale (all consumption is private). The logical upper bound for τ is 2 (all consumption is public). The specification adopted in this paper is a simple version of

[3] We test and cannot reject the linearity assumption in the empirical section.

the more complex consumption technology used by BCL, who estimate a collective household model using expenditure data with good-specific returns-to-scale.[4]

Given (2) we can write individual consumption of wife and husband as

$$x^f = \eta\tau y^h \quad and \quad x^m = (1-\eta)\tau y^h, \tag{3}$$

where η is the sharing rule that determines which share of τy^h is allocated to the wife.

The sharing rule depends on so-called household distribution factors $\mathbf{d^h}$. These distribution factors are the same for both household members. We specify a simple linear sharing rule given by

$$\eta = \gamma^0 + \mathbf{d}^h \boldsymbol{\gamma} d, \tag{4}$$

where $\boldsymbol{\gamma}^d$ is a vector of unknown parameters.

Using Eqs. (1), (3), and (4), we can write indirect utility of singles and both members of couples as follows. For singles, it is simply

$$V_{it} = \mathbf{z}_{it}\alpha + \beta lny_{it}^h + \mu_i + \varepsilon_{it}, \tag{5}$$

For females in couples we get

$$\begin{aligned} V_{it} &= \mathbf{z}_{it}\alpha + \beta\left\{\ln\left[(\gamma^0 + \gamma^d \mathbf{d}_{it}^h)y_{it}^h\tau\right]\right\} + \mu_i + \varepsilon_{it} \\ &= \mathbf{z}_{it}\alpha + \beta\ln(\gamma^0 + \gamma^d \mathbf{d}_{it}^h) + \beta\ln\tau + \beta\ln y_{it}^h + \mu_i + \varepsilon_{it} \end{aligned} \tag{6}$$

while for males we have

$$\begin{aligned} V_{it} &= \mathbf{z}_{it}\alpha + \beta\left\{\ln\left[\left(1 - \left(\gamma^0 + \gamma^d \mathbf{d}_{it}^h\right)\right)y_{it}^h\tau\right]\right\} + \mu_i + \varepsilon_{it} \\ &= \mathbf{z}_{it}\alpha + \beta\ln\left[1 - \left(\gamma^0 + \gamma^d \mathbf{d}_{it}^h\right)\right] + \beta\ln\tau + \beta\ln y_{it}^h + \mu_i + \varepsilon_{it} \end{aligned} \tag{7}$$

The term in square brackets is individual consumption determined by household income, the returns to scale and the sharing rule. These equations are essentially linear except that consumption is nonlinear in the parameters τ and $\boldsymbol{\gamma}$ for couples. The model is estimated by fixed effects nonlinear least squares using Eq. (5) for singles, Eq. (6) for women in couples, and Eq. (7) for men in couples (see Sect. 4 for more details). Using fixed effects has the advantage of obtaining unbiased estimates of the preference parameters in the presence of unobserved heterogeneity contained in μ_i. Identification of β is obtained by assuming that it is the same for singles and couples. This identification assumption has been made by Browning et al. (2013), Lise and Seitz (2011), Lewbel and Pendakur (2008), Cherchye et al.

[4] BCL specify $z_j = \mathbf{A}(q_j^f + q_j^m) + a$, where z_j is an observable vector of household j's consumption of n goods, q_j^f and q_j^m are unobservable vectors of private consumption of both spouses. $\mathbf{A}$ is an $n \times n$ nonsingular matrix and a is a n-vector. This linear consumption technology nests familiar cases, e.g. the well-known Barten scales if $\mathbf{A}$ is diagonal and $a = 0$. By contrast to BCL we can only identify an aggregate consumption technology, i.e. $\mathbf{A}$ is diagonal with identical elements A. Assuming the budget constraint holds with equality we have $y_j = p'z_j = (p'\mathbf{A}q_j^f + p'\mathbf{A}q_j^m)$. Given the restriction on $\mathbf{A}$ this simplifies to $A(x_j^f + x_j^m) = y_j$, and $\tau = A^{-1}$.

(2012), and Alessie et al. (2006). However, we do not require preferences to be the same for women and men.

2.1 Equivalence and indifference scales

Traditionally, an equivalence scale is defined as the ratio of the expenditures (or income) of two different household types with the same standard of living. Formally, this corresponds to the ratio of the cost functions of two household types evaluated at the same utility level. This requires the comparison of the utility levels of different households. But, as BCL state, the notion of household utility is flawed because individuals have utility, not households.

BCL introduce indifference scales as opposed to traditional equivalence scales. An indifference scale equates the utility of an individual living alone to the utility of the same person if she lived in a couple. In other words, it determines the expenditure change necessary to put the individual on the same indifference curve in both living situations. Taking the couple as point of reference, the indifference scale measures the proportion of household income an individual living in a couple would require to reach the same utility when living alone. This is the relevant statistic for issues like alimony, life insurance, and pensions payments. In our specification, the female indifference scale is obtained by setting $V(y^s) = V(\eta y^c \tau)$, where y^c denotes household income of the couple household and y^s is the income of the single household. Independent of any transformations of the utility function, the indifference scales are $\theta^f = \eta\tau$ for women and $\theta^m = (1-\eta)\tau$ for men. If $\eta = 0.5$ equivalence scales and indifference scales are formally identical in our approach. If $\eta \neq 0.5$, there is no unique equivalence scale for both spouses in a couple household.

3 Measurement of utility

In Sect. 2 we presented a structural individual-level model for indirect utility, assuming that utility is observable. Our estimation strategy is based on the assumption that individual satisfaction with own financial situation is a valid measure of own individual indirect utility. This can be criticized on several grounds. The most important are: (1) utility is an ordinal concept and cannot be measured on a cardinal scale; (2) utility can be measured but interpersonal comparisons are not valid. In our specific context, further objections are: (3) singles are less satisfied independent of their living situation, invalidating the identifying assumption; and (4) satisfaction with one's financial situation depends on some reference or comparison income, e.g. the partner's income.

Criticism (1) can be addressed by the fact that such measures have repeatedly been validated by psychologists to be a reasonable proxy for utility or well-being (see e.g. Kahneman et al. 1999). In economics, Van Praag and several co-authors challenged this criticism since the early 1970s. They claim that subjective income evaluation questions can be used to estimate utility functions, i.e. the answers to these questions are valid proxies of utility. Over the years, this so-called "Leyden-school" produced a series of credible evidence for the usefulness of the approach for empirical welfare analysis (see for a survey Van Praag and Frijters 1999). The happiness literature is also based on the assumption that answers to subjective satisfaction questions provide valid proxies for happiness (or utility). A recent overview is given in Frey and Stutzer (2002).

With respect to criticism (2) it can be argued that individuals are given a well-defined scale for their evaluation including verbal descriptions. Therefore, it is plausible that they reply in a comparable manner. At least, this approach seems to work well in a variety of settings (see, e.g. Diener and Suh 1997). Still, it is possible that individuals have different baseline satisfaction levels. For this reason, according to Ferrer-i Carbonell and Frijters (2004), controlling for unobserved heterogeneity is the most important methodological issue in empirical happiness research. The empirical analysis in this paper controls for observed and unobserved heterogeneity using a fixed effects model. Our results also differ depending on whether we control for unobserved heterogeneity.

Criticism (3) refers to a possible selection effect in the sense that singles are less satisfied independent of their living situation. This selection effect is documented by e.g. Stutzer and Frey (2006) using the German Socioeconomic Panel (GSOEP). They find that singles remaining singles over the entire observation period are less satisfied with life than singles who eventually marry. However, Zimmermann and Easterlin (2006), also analyzing the GSOEP, show that the selection effect becomes insignificant when cohabiting non-married persons are not classified as singles who will marry later, as was done by Stutzer and Frey (2006). We are able to address this issue by using the fact that a large fraction of the singles in fact have a partner, but do not live together. Using this subgroup allows to check the robustness of our results. Again, controlling for unobserved heterogeneity will also alleviate the selection effect. Note also that all research discussed above analyzes satisfaction with life, not satisfaction with one's financial situation. It may well be that potential selection effects are less pronounced when looking at satisfaction with one's financial situation.

Criticism (4) can be addressed by analyzing whether individual satisfaction with one's financial situation depends on the level of the partner's income. We estimate standard fixed effects reduced form models of individual well-being measured both by financial satisfaction and by satisfaction with life. We control for both household income and partner's income. In both regressions the partner's income has the expected negative effect on well-being (conditional on household income), but it is not significant. From this we conclude that the comparison effect is not strong and should not bias our results. It should also be mentioned that respondents are explicitly asked to evaluate both their personal income and their financial situation, so they should be aware that the financial satisfaction question is not related directly to personal income but to the individual's financial situation within the household.

It is important to note that when we would estimate our model with general life satisfaction it would not be possible to identify a couple's indifference scales. That is, if being married has a direct impact on satisfaction, then the estimated effect is not the indifference scale but the indifference scale plus an effect that arises from non-monetary benefits from marriage. When we estimate our model with either life satisfaction or satisfaction with personal income as proxies for the indirect utility the results for the indifference scale are implausible from an economic point of view. Using satisfaction with one's financial situation generates sensible results as we show below (and has been shown repeatedly in other empirical applications). We interpret this as informal evidence that satisfaction with one's financial situation is indeed a useful proxy for indirect utility.[5]

4 Data

The data source is the Living in Switzerland Survey conducted by the SHP.[6] This survey is based on a representative sample of the population of permanent Swiss residents. It has been conducted annually since 1999. There is a household and an individual questionnaire. The household questionnaire includes questions about housing, living standard, financial situation, the household structure and family organization, whereas the individual questionnaire covers topics such as household and family, health and life events, social origin, education, work, income, integration and social networks, politics and values, as well as leisure and internet use.

Our analysis is based on data from the years 2000 to 2008.[7] The main selection rules we impose are that all persons are aged between 20 and 60 and employed. The second selection rule is imposed to make the identification assumption more credible. In addition, we exclude divorced singles because we suspect that alimony payments and receipts are not fully covered in the data. Especially the disposable income of divorced men may be severely biased in this case. Individuals living in the same household and stating to be partners are defined as a couple. They do not have to be married. After eliminating missing observations, we have an unbalanced panel on 2,035 individuals. 965 of these are singles and 800 are cohabiting in all observed years. 270 individuals change their living situation in the observation period. Overall, we have person-year observations for 3,250 singles and 1,142 individuals living in couples. 730 person year observations relate to individuals that changed their living condition at least once. These observations are crucial for the identification of the consumption technology.

We use the answer to the following survey question to measure individual satisfaction with income:

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5 We are also able to distinguish between single living individuals with and without a partner. As we show in Sect. 4, single living individuals with partner are equally satisfied with their financial situation as single living individuals without partner. There is, however, a difference in the life satisfaction between these two group. This finding indicates that financial satisfaction is not necessarily altered by having a partner and therefore a better measure for indirect utility.

6 See also www.swisspanel.ch.

7 Since not all necessary variables are available for the first wave in 1999, this wave is excluded from this study.

> Overall, how satisfied are you with your financial situation, if 0 means "not at all satisfied" and 10 "completely satisfied"?

Figure 1 displays the distribution of reported financial satisfaction levels by sex and household type (singles and couples). By and large, the distributions are rather similar across these cases. For both household types, women more often report the highest satisfaction levels 9 and 10, whereas a larger fraction of men are reporting level 7.[8]

Table 1 reports descriptive statistics of the person- and household-specific characteristics included in the regression analysis. Net annual income is defined as labor as well as non-labor income net of taxes. Single women's net annual income is on average 17 % lower than the net annual income of single men. Furthermore, single women are on average less educated, but more satisfied with their financial situation than single men.

Men living in couples are older and on average better educated than their spouses. Regarding the distribution factors, we observe an average household income of roughly CHF 128,000 and an average wage ratio of 0.89. The wage ratio is defined as the female hourly wage divided by male hourly wage. The income ratio, i.e. women's total income divided by men's total income, is even smaller which is due to the fact that women work less hours than men.

The data also contain a variable indicating who is mainly responsible for household finances. This information is not used in the estimation, but we analyze *ex post* whether the estimated sharing rules are correlated with the way the household finances are managed. Table 2 presents a descriptive analysis of the financial responsibility variable. 42 % of the couples manage household income together. Furthermore, in 10 % of the couples the partners manage their income separately. In the remaining 47 % one partner is mainly responsible for household income. Interestingly, if women are responsible for the household finances, the average wage ratio is a bit close to one and the average household income is below average. Hence, these households are characterized by relatively low earning husbands.

As mentioned in Sect. 3, a potential problem with measuring the indirect utility by subjective financial satisfaction is that singles are on average less satisfied than married individuals. This might either reflect that individuals get happier after marriage or that happy people are more likely to get married. We address this possible selection effect by using the subgroup of individuals living in a single household but report having a partner (not singles in the narrower sense) to check the robustness of our results. Table 3 presents descriptive statistics for the satisfaction of these two types of individuals living in a single household. The material well-being does not differ between individuals with and without a partner. However, the satisfaction with life is lower for male and female individuals who do not have a partner. This indicates that having a partner affects life satisfaction positively, but it does not affect the financial satisfaction. For this reason we expect

[8] This finding is not new to the literature. Women mostly report higher values of general satisfaction (see, e.g. Stevenson and Wolfers 2009) and also financial satisfaction (see, e.g. Bonke and Browning 2009) than men.

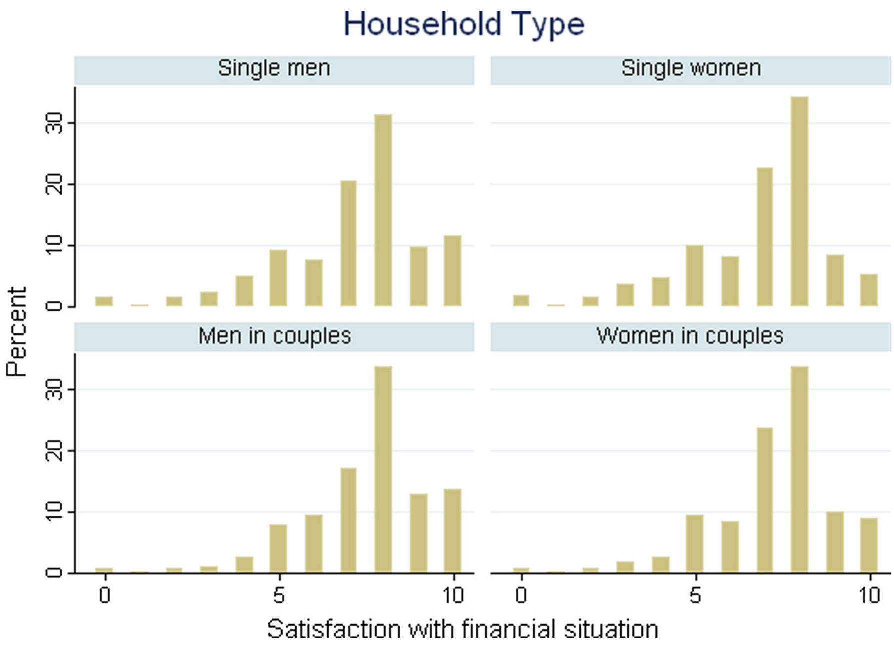

Fig. 1 Frequencies of reported levels of satisfaction with financial situation. *Data source*: SHP, 2000–2008

that estimation results based on financial satisfaction are similar for both definitions of singles.

5 Estimation strategy

Financial satisfaction, our measure for indirect utility is observed on an ordinal scale. A natural estimator in this case is an ordered response model, e.g. ordered probit or logit. As noted above, we observe individuals over time and therefore also their movement in and out of cohabitation. The panel structure of our data therefore allows us to control for unobserved heterogeneity. Estimating fixed effects ordered response models is a problem for which solutions have been proposed only recently, most prominently by Ferrer-i Carbonell and Frijters (2004). They extend the binary conditional logit estimator of Chamberlain (1980) to accommodate ordered response variables. The estimator is based on an individual specific binary recoding procedure using the individual specific information of the second derivative of the log likelihood function for the conditional logit estimator. Baetschmann et al. (2011) show that this estimator may be biased in small samples and propose another approach they call "blow up and cluster". This estimator is consistent.

However, due to the nonlinearity of the utility function with respect to the parameters τ and γ all the above approaches cannot be used to directly estimate the structural parameters. One solution is to follow ACH and estimate the reduced form

Table 1 Descriptive statistics

	Singles		Couples	
	Women	Men	Women	Men
Household income	61,227.7	69,315.4	128,112.6	
	(22,687.1)	(25,921.7)	(38,593.8)	
Age	37.70	37.99	37.00	39.36
	(9.295)	(8.787)	(9.937)	(9.858)
Low education	0.106	0.0683	0.0946	0.0394
High education	0.245	0.366	0.263	0.377
French speaking	0.261	0.240	0.244	0.242
Swiss	0.926	0.882	0.902	0.869
Financial satisfaction	7.171	6.890	7.539	7.293
	(2.028)	(1.989)	(1.790)	(1.739)
General satisfaction	7.831	7.544	8.183	8.148
	(1.351)	(1.406)	(1.251)	(1.109)
Wage ratio			0.891	
			(0.335)	
Income ratio			0.769	
			(0.399)	
Observations	1,668	1,582	1,142	

Own computations using SHP, 2000–2008

Sample means and standard deviations in parentheses

Table 2 Distribution of responsibility for household income

Responsible for household finances	Proportion of households	Wage ratio	Household income
Women	0.253	0.962	122,172.2
Men	0.219	0.816	129,158.8
Together	0.421	0.882	130,827.1
Separately	0.104	0.915	129,954.6
Observations	1,142		

Own computations using SHP, 2000–2008

parameters in a first step using for example the “blow up and cluster” estimator and to obtain the structural parameters in a second step.

An obvious alternative approach is to estimate the structural parameters directly applying a nonlinear regression model using the reported satisfaction levels as the dependent variable. The drawback is that this will attach cardinal values to the reported satisfaction levels, with equal distances between satisfaction levels and a restricted support of the dependent variable.

Recently, van Praag and Ferrer-i-Carbonell (2004) have proposed to replace the equidistant responses to satisfaction questions by suitable transformations that take

Table 3 Well-being of singles with and without partners

	Single women		Single men	
	With partner	Without partner	With partner	Without partner
Financial satisfaction	7.220	7.138	6.898	6.885
	(1.932)	(2.093)	(1.993)	(1.987)
General satisfaction	8.034	7.691	7.802	7.376
	(1.267)	(1.389)	(1.310)	(1.441)
Observations	681	987	625	957

Own computations using SHP, 2000–2008

Sample means and standard deviations in parentheses

account of the sample distribution of the reported satisfaction levels. The transformed variable can be used as the dependent variable in a linear regression. Van Praag and Ferrer-i-Carbonell call this procedure Probit-OLS (POLS) approach. A detailed discussion of POLS is given in their book.

The transformation of the original response $v \epsilon \{1, 2, \ldots, k\}$ into the new dependent variable involves three steps:

(a) compute the relative frequencies of discrete responses $p_i, i = 1, 2, \ldots, k$
(b) compute $z_i, i = 1, 2, \ldots, k-1$ such that $p_i = \Phi(z_i) - \Phi(z_{i-1})$, where $z_0 = -\infty$ and $z_k = \infty$
(c) compute $\overline{v} = E(V|z_{i-1} < V < z_i) = \frac{\phi(z_{i-1}) - \phi(z_i)}{\Phi(z_i) - \Phi(z_{i-1})}$

The transformed variable $\overline{v}$ is used in the regression analysis as left-hand side variable instead of the original v. Obviously, $\overline{v}$ is still an ordinal variable, but not equidistant. Rather, the distance depends on the sample probabilities of the satisfaction levels. The estimated coefficients can now be interpreted as shifting the thresholds that generate the sample distribution of responses, exactly as in the ordered probit model.

Van Praag and Ferrer-i-Carbonell (2004) do not formally prove that their transformation yields consistent estimates. However, both heuristically and in several applications they show that POLS is virtually identical to the traditional ordered probit analysis up to a scale factor. If we are mainly interested in marginal effects or in ratios of coefficients, POLS seems to give identical results compared to ordered probit. This has recently been confirmed in a careful Monte Carlo study by Geishecker and Riedl (2011). They evaluate among others the POLS and the Blow up and Cluster approach. They conclude that "if the researcher is more interested in the ratios of the parameter estimates, the linear fixed effects model ... essentially delivers the same results as the more elaborate binary recoding schemes and is much easier to compute" (page 15). Therefore, a natural way to proceed is to estimate a reduced-form model by linear fixed effects and to apply the ACH transformation to obtain the structural parameters, which are all ratios of reduced-form parameters.

It is also possible to estimate the structural parameters directly. As stated in Sect. 2, we want to control for correlated unobserved heterogeneity by estimating the model using a nonlinear fixed effects estimator. For example, unobserved heterogeneity variables are likely to influence the utility, the decision-making process, and also the individual hourly wage rate. These variables could also be correlated with the cohabitational status if the marriage market endogenously selects individuals with specific preferences. The nonlinear fixed effects estimation is possible because the estimation equations are linear in consumption, controls and the unobserved fixed effect. Only individual consumption within couples is a nonlinear function of the consumption technology and the sharing rule. For example, for women (both single and cohabiting) the differenced estimation equation can be written as

$$(V_{it} - \bar{V}_i) = \beta(\ln y_{it}^h - \overline{\ln y_i^h}) + \beta\Big[\Big(\ln(\gamma^0 + \gamma^d \mathbf{d}_{it}^h) - \overline{\ln(\gamma^0 + \gamma^d \mathbf{d}_i^h)}\Big) + (\ln \tau - \overline{\ln \tau})\Big] \\ + (\mathbf{z}_{it} - \bar{\mathbf{z}}_i)\alpha + (\varepsilon_{it} - \bar{\varepsilon}_i), \tag{8}$$

where the variables with overbars are the individual specific means within the observation period. For single women the restrictions $\gamma^0 = \tau = 1$ and $\gamma^d = 0$ apply, which allow to identify β. The individual fixed effect and all time-constant elements of $\mathbf{z}$ are eliminated by this differencing. This includes mother tongue, education, and nationality. The returns to scale parameter τ is identified by those individuals who change their cohabitation status. The corresponding equation for men is derived in the same manner.

As far as we know the performance of POLS in a nonlinear setting has not been analyzed yet. We conducted a small Monte Carlo simulation[9] and found that the estimates of the utility parameters α and β depend on the scaling of the dependent variable. However, in all simulations we obtained unbiased estimates of γ and τ. This makes intuitive sense because monotone transformations of the dependent variable will change the intercept and the slope parameters of the estimated utility function, but not the transformation of household income into individual consumption.[10] The question whether this is true for only modest monotone transformation is left to future research.

It is important to note that our estimation strategy does not depend on the assumption that the sharing rule is independent of income. We explicitly allow the sharing rule to differ with household income.

We tested the assumption that utility is linear in $\ln y$ by reduced form fixed effects models with higher degrees of polynomials of $\ln y$ for each group (sex and household type). In all cases, we could not reject the linear specification and conclude that the linear specification is adequate in our case.

[9] The data generation process was designed to mimic the theoretical model of Sect. 2. We generated a continuous latent utility which was split into 11 categories such that the empirical distribution of the ordinal responses was replicated.

[10] For this reason, also using the original reported satisfaction level as dependent variable hardly affects the estimates of γ and τ.

6 Results

In this section we discuss the estimation of the structural model. The dependent variable is the transformed financial satisfaction $\overline{v}$. The model is estimated by nonlinear least squares with standard errors adjusted for clustering due to the panel structure of the data. Table 4 displays the estimation results. As distribution factors we use the wage ratio, i.e. women's hourly wage divided by men's hourly wage, and the total household income. We also used the square of the wage ratio as a distribution factor, but the results indicate no significant nonlinearity. We also experimented with the duration of the relationship and the age difference between partners as distribution factors, but both turned out to be completely insignificant in the sharing rule. The distribution factors are both normalized to mean zero. Hence, γ^0 is an estimate of the share of total household consumption a wife with mean wage ratio and mean household income receives. The first set of results is obtained using all singles. In the Column II, only singles who are in a relationship but do not live together are included.

In Column I of Table 4, the estimated consumption technology parameter τ is 1.39 indicating that the sum of individual private consumption of both spouses exceeds household income by 39 %. The estimated parameters of the sharing rule indicate that at the mean of the distribution factors, women have a consumption share of 0.46. This estimate is not significantly smaller than 0.5. The female share increases significantly with the wage ratio. Increasing the wage ratio by 0.1, e.g. from 0.8 to 0.9, increases her consumption share by roughly 1.5 %-points. The household income on the other hand has no significant effect on the sharing rule.[11]

The estimation results in Column II of Table 4 include only singles who are in a relationship but do not live together with their partner. Compared to Column I the estimated parameters are almost identical. This reflects the finding in Sect. 4 that the financial satisfaction of singles is not affected by having a partner or not (as opposed to life satisfaction). For this reason we focus on the results displayed in Column I in the following discussion of the results.

6.1 The estimated sharing rule

Figure 2 displays the distribution of the predicted female consumption share in our sample. This distribution is clearly centered to the left of the equal share of 0.5. Although the empirical model does not impose this, all estimated shares are within [0,1] and therefore within the logical range. This finding underlines that the applied method produces reasonable results. All shares below 0.4 are significantly lower than 0.5, which covers about 30 % of all women. Overall, this picture suggests that consumption shares in Switzerland are unequally distributed even among couples in which both spouses are working.

[11] We also tested a more flexible specification where the coefficient β is allowed to vary by gender. The coefficient of the interaction term of β and the gender dummy is not significant on a 5 % level. Furthermore, allowing β tol vary by gender has no effect on the parameters of the sharing rule.

Table 4 Estimated parameters of consumption technology and sharing rule

	I	II
τ	1.39 (0.16)	1.38 (0.18)
γ^0	0.46 (0.05)	0.43 (0.06)
(wage ratio) γ^1	0.15 (0.06)	0.15 (0.06)
(household income) γ^2	0.01 (0.01)	0.01 (0.01)
β	0.69 (0.08)	0.70 (0.12)
Age and age squared	Yes	Yes
Year dummy	Yes	Yes
Observations	5534	3590

Nonlinear least squares fixed effects

Dependent variable: satisfaction with financial situation

I: All singles, II: singles with partners

Clustered standard errors in parentheses

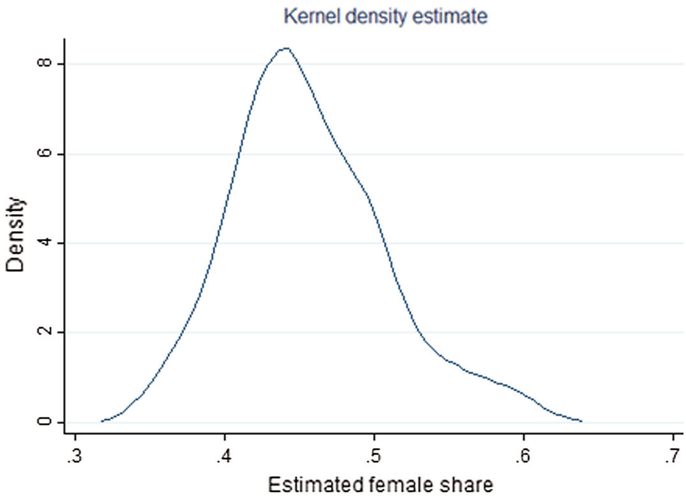

Fig. 2 Distribution of estimated female share estimates based on column I in Table 4

In Table 5 we analyze how the estimated shares correlate with the reported responsibility for the household's financial decision-making. To our knowledge, this is the first application linking reported responsibility for household finances with the estimated sharing rule. We regress the predicted female shares on the indicators for financial responsibilities, controlling for the age difference of the partners. None of these variables have been used to estimate the sharing rule. Recall that only one person per household responded to the financial responsibility question.

We find that if men are responsible for household finances they are able to shift about one percent of household consumption to their own use compared to joint financial responsibility. This result makes intuitive sense and gives some credibility to the estimated sharing rules. Women, however, who are responsible for the household finances are not able (or willing) to shift consumption to their own use.

6.2 Indifference scales

In our model the estimated indifference scales are proportional to the sharing rule ($\theta^f = \eta\tau$ for women and $\theta^m = (1-\eta)\tau$ for men) and always sum up to the return to scale factor τ. At the mean of the distribution factors, for example, the female indifference scale is 0.64 and the male indifference scale is 0.75. In other words, a woman would need 64 % of household income to reach the same utility when she lives alone, while a man would need 75 %. The difference seems rather large, but it is not significantly different from zero. When we evaluate the indifference scales at a wage ratio of 0.5, women's scale drops to 0.56. That is, she would need half of household income to be as well off when living alone. The husband, on the other hand, would need 83 % of the household income in that situation. This does not imply that wives are more modest and happy to receive the smaller share. The indifference scale is the consequence of the result that husbands who have much higher wages than their wives have substantial bargaining power and allocate more resources to their private consumption. Hence, in order to maintain their current well-being wives need a smaller fraction of household income if they were to live alone. These findings indicate a feedback effect between the labor market and the household domain. Unequal wages lead to less bargaining power for women which in turn makes it more difficult to enforce their preferences in the household. As a result the wage differences may persist. Breaking this circle will probably need some generations.[12]

We can use indifference scales to measure consumption inequality among couples. We compute the Theil index of inequality which is additively decomposable into within household inequality and between household inequality. In our sample of couples, assuming equal sharing of income within the household yields an estimate of 0.043 for the Theil index, which measures between household inequality. Using the indifference scales, which differ both within and across couples, depending on their specific value of η, the Theil index is 0.052. This is an estimate of the overall consumption inequality among individuals living in a couple. Within inequality is the difference between these two estimates and amounts to 16 % of overall inequality. This example illustrates that taking the distribution of resources within the household into account may affect inequality and poverty measurement substantially. Similar results are provided by Lise and Seitz (2011).Their decomposition of total inequality among couples in the UK in the 1990s attributes 25 % to within inequality. One reason for their higher share of

[12] As shown by Fernández et al. (2004), the growing presence of a new type of man-one raised by a working mother-contributes significantly to the increasing female labor supply over time. Hence, family attitudes and their intergenerational transmission played a quantitatively significant role in transforming women's role in the economy.

Table 5 Sharing rule and financial responsibility

Men decides	−1.09 (0.40)
Women decides	0.61 (0.39)
Separate decision	0.48 (0.53)
Age difference	−0.12 (0.03)
Constant	45.6 (0.23)
Observations	1,139

OLS estimates

Dep. variable: Estimated female share $\times$ 100

Joint decision making is reference group

Based on column I in Table 4

within inequality is that their sample includes non-working individuals who receive rather small shares of total consumption.

6.3 Cultural differences

Switzerland has four official languages which are spoken in different areas. Almost two third of the population is German speaking and 20 % is French speaking. The two other languages (Italian and Rhaeto-Romanic) are spoken by a smaller group of the population. As mentioned above, measuring the indirect utility by subjective financial satisfaction relies on the assumption that individuals are given a well-defined scale for their evaluation including verbal descriptions. Since our sample includes both French and German speaking individuals, who are interviewed in either French or German, the verbal scale might be understood differently in the two different language areas. Furthermore, the cultural gap between the two main language groups is rather large and is also reflected in a smaller female wage gap in the French speaking areas. In order to test whether these differences translate into different sharing rules we estimate the model for both groups separately.

The estimated parameters of the consumption technology and the sharing rule for the two language areas are reported in Table 6. It turns out that the point estimates of τ and γ are very similar, however, due to smaller sample sized not significant. As none of the estimates differs significantly between the two cultural groups, we do not find evidence for cultural differences with respect to resource sharing.

6.4 Comparison to previous research

In this section we briefly discuss the results from other papers based on the BCL approach. BCL use Canadian expenditure data from 1974 to 1992. They specify a richer consumption technology that differs across consumption goods, i.e. they estimate good-specific Barten scales. As distribution factors, BCL use the wife's contribution to household income, the age difference between the spouses, a home-ownership dummy, and total household expenditure. They compute an overall measure of the economies of scale in consumption that varies between 1.27 and 1.41

Table 6 Estimated parameters differentiated by language area

	III	IV
τ	1.40 (0.21)	1.45 (0.29)
γ^0	0.47 (0.07)	0.44 (0.07)
(wage ratio) γ^1	0.15 (0.08)	0.17 (0.10)
(household income) γ^2	−0.01 (0.01)	0.02 (0.02)
β	0.64 (0.09)	0.79 (0.22)
Age and age squared	Yes	Yes
Year dummy	Yes	Yes
Observations	3,957	1,371

Nonlinear least squares fixed effects

Dependent variable: satisfaction with financial situation

III: German, IV: French

Clustered standard errors in parentheses

(see Table 4 in BCL).[13] Our estimate is 1.39 (= τ). BCL's benchmark estimate of the sharing rule is 0.65, which appears to be quite large. BCL argue that this may be explained by the fact that household demand functions look more like women's demand functions than men's demand functions (p. 31). The estimated indifference scales vary between 0.58 and 0.74 for women, and 0.50 and 0.70 for men, depending on the restrictions imposed on the model.

Cherchye et al. (2012) analyze expenditure data for Dutch pensioners from 1978 to 2004. They compute the same overall measure of the economies of scale of living in couple as BCL. On average, this measure is 1.32. The average income share of the female is 0.49. They use real total expenditure and the education difference as distribution factors. The indifference scales for women in couples increase from 0.49 (evaluated in the bottom total real expenditure quartile) to 0.82 (evaluated in the top total real expenditure quartile). For men, the reverse pattern can be observed, i.e. the indifference scales drop from 0.81 to 0.50.

Lewbel and Pendakur (2008) estimate a simplified version of the BCL model also using Canadian expenditure data (1990 and 1992, no price variation). The simplification is achieved by imposing a shape invariance restriction on the Engel curves. They obtain estimates of the scale economy parameter of 0.70 for women and of 0.78 for men with average characteristics, but the estimates are very imprecise. Their benchmark estimate of the sharing rule is in the range of 0.36–0.46. As distribution factors, Lewbel and Pendakur (2008) use the female contribution to household income, and male and female age and education. Overall, their results are close to ours.

ACH is closest to our paper in term of methodology. They estimate a reduced form version of the model described in Sect. 2 for 10 European countries.[14]

[13] Their overall measure R is defined as (equivalent expenditures/actual expenditures) -1; hence R is in the range 0.27–0.41 in their Table 4.

[14] The structural parameters can be obtained from the reduced form parameters by a straightforward minimum distance step (c.f. their paper).

Compared to our results the estimates of τ in ACH are rather large, often above 1.65, in some cases even above 2.[15] On the other hand, their estimate of the sharing rule evaluated at the mean of the distribution factors is above 0.5 in almost all cases. The results which are most similar to ours refer to the UK with estimates of $\tau = 1.45$ and a mean sharing rule of 0.49 (Table 6 in ACH).

7 Conclusion

Based on a collective household model, this paper provides estimates of the returns to scale of living together and of the rule of sharing consumption among spouses taking correlated unobserved heterogeneity into account. Household income is transformed into individual consumption by a consumption technology (the returns of scale) and a rule that determines how much resources each member receives (the sharing rule). The sharing rule is a function of distribution factors that affect the individual bargaining power within the household. Assuming that preferences do not change by living together, it is possible to identify the returns to scale and the sharing function from data on singles and couples. This identification result is one of the major contributions of Browning et al. (2013). In this setup, it is possible to identify so-called indifference scales which allow to make welfare comparisons between different living conditions for the same person.

We use data on financial satisfaction as a measure of indirect utility received from individual consumption. We focus on single and couple households without children. The estimated consumption technology parameter in our preferred specification implies that scale economies increase the sum of individual consumption of both members to 139 % of household income.

The estimated sharing rule varies significantly with the distribution factor female wage ratio which is defined as the wife's hourly wage relative to the husband's hourly wage. The other distribution factor, total household income, has no significant effect. The estimated sharing rule is 0.46 at the mean of the distribution factors and 0.47 if women's and men's hourly wages are equal. The estimated female shares vary between 0.3 and 0.6. About 30 % of the women have a consumption share that is significantly <0.5.

At the mean of the estimated sharing rule the female indifference scale is 0.64, while the male indifference scale is 0.75. These numbers measure which proportion of the couple's total income each member would need to be equally well off when living alone. Using these indifference scales to analyze consumption inequality among individuals living in a couple allows to attribute 16 % of total inequality to within household inequality. Hence we conclude that unequal sharing among spouses has a non-negligible effect on inequality.

The analysis can be extended in several ways. The most important issue is to include children in the model. Collective household models so far mostly apply to households without children or treat children as public goods. The important question, however, is which share of household resources is transferred to the

[15] ACH denote the consumption technology $A = 1/\tau$ in their model

children and how this transfer affects the parents' shares (see Dunbar et al. 2013 who find, using data from Malawi, that children mostly reduce their mothers' share). This extension is necessary to make collective household models useful for inequality and poverty analysis. From an econometric point of view, more work needs to be done with respect to estimation of models with satisfaction data, especially nonlinear models with panel data. Another extension would be to consider more flexible specifications for individual utility. A very promising extension of this paper would be a combination of subjective satisfaction data with expenditure data. This would allow to compare the results from the BCL approach or an approach based on the BCL-model (see, e.g. Lewbel and Pendakur 2008; Cherchye et al. 2012; Lise and Seitz 2011) to the result using the approach with subjective financial satisfaction measures.

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