

Consumption Inequality and Intra-household Allocations

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The consumption literature uses adult equivalence scales to measure individual-level inequality. This practice imposes the assumption that there is no within-household inequality. In this paper, we show that ignoring consumption inequality within households produces misleading estimates of inequality along two dimensions. To illustrate this point, we use a collective model of household behaviour to estimate consumption inequality in the U.K. from 1968 to 2001. First, the use of adult equivalence scales underestimates the initial level of cross-sectional consumption inequality by 50%, as large differences in the earnings of husbands and wives translate into large differences in consumption allocations within households. Second, we estimate the rise in between-household inequality has been accompanied by an offsetting reduction in within-household inequality. Our findings also indicate that increases in marital sorting on wages and hours worked can simultaneously explain two-thirds of the decline in within-household inequality and between a quarter and one-half of the rise in between-household inequality for one and two adult households.

Keywords: Collective model, Consumption inequality, Marital sorting, Adult equivalence scales

JEL Codes: D12, D13, D31, J22

1. INTRODUCTION

A recent literature has documented a large rise in consumption inequality in several developed countries.¹ Underlying these measures of inequality is the use of adult equivalence scales, which are used to assign consumption levels to each member of a household. This has been necessary as there do not exist comprehensive measures of individual-level consumption for households with more than one member. The drawback of this approach is that it implicitly assumes that there is no inequality among adults within the household. In particular, the use of adult

1. See Blundell and Preston (1998) for the U.K., Pendakur (1998) for Canada, and Barrett, Crossley and Worswick (2000) for Australia. The evidence on consumption inequality in the U.S. is mixed. Slesnick (2001) and Krueger and Perri (2006) do not find a large rise in consumption inequality since the 1970's. Alternatively, Cutler and Katz (1992) and Johnson and Shipp (1997) report a rise in inequality. Attanasio, Battistin and Ichimura (2007) use two sources of data on consumption and also find that consumption inequality was rising over time. Also see the collection of papers to appear in a special issue of *Review of Economic Dynamics* available at <http://www.econ.umn.edu/~fperri/Cross.html>.

equivalence scales implies a very restrictive model of the household in which husbands and wives split consumption equally, regardless of preferences or the source of income.²

This equal division assumption is inappropriate in the study of consumption inequality as a large theoretical and empirical literature routinely rejects the assumption that the consumption allocation does not vary with the source of income in the household (Browning *et al.*, 1994; Lundberg, Pollak and Wales, 1997; Manser and Brown, 1980; McElroy and Horney, 1981; Chiappori, 1988, 1992; Browning and Chiappori, 1998; Chiappori, Fortin and Lacroix, 2002; Donni, 2007, among many others). Since there has been a sizable increase in women's wages and labour supply over the last half century, the share of household earnings provided by the wife has changed substantially over time. If consumption allocations depend on the source of income and the sources of income within households have changed over time, then adult equivalence scales will produce an inaccurate picture of the trends in consumption inequality.

The goal of this paper is to document the trends in consumption inequality once within-household inequality is taken into account. We construct and estimate a static model of intra-household allocations to examine how changes in the source of income in the household translate into changes in individual-level consumption allocations. The model is estimated on a sample of two-person households (couples without children) from the U.K. Family Expenditure Survey (FES) for the years from 1968 to 2001. We then use our estimates to construct a new measure of consumption inequality across individuals. We have two main findings. First, ignoring the potential for intra-household inequality may underestimate the initial level of individual-level consumption inequality by 50%, as differences in earnings across husbands and wives generate substantial within-household inequality. Second, the rise in consumption inequality measured between households has been largely offset by a corresponding decline in inequality within households. In 1970's, half of the inequality between individuals was attributable to dispersion within households; by 2000, this had fallen to 25%.

A recent literature examines several dimensions of the relationship between inequality and household behaviour. Kremer (1997) and Fernández and Rogerson (2001) consider the extent to which increases in marital sorting cause increases in inequality. Dahan and Gaviria (2001) and Fernández, Guner and Knowles (2005) present empirical evidence of a positive correlation between sorting and inequality.³ Gould and Paserman (2003) consider whether male wage inequality has a causal effect on marriage and find that increases in inequality lead to reductions in marriage rates. We contribute to this literature by providing evidence on the importance of several potential explanations for the rise in consumption inequality between households and the fall in inequality within households since 1970's in the U.K. Changes in the demographic composition of the population and the number of single versus married couples appear to play a limited role, accounting for at most 27% of the increase in between-household dispersion. On the other hand, an increase in marital sorting has profound effects on the trends in consumption inequality. In particular, the rise in marital sorting on wages and hours observed in the data accounts for between 20% and 53% of the rise in between-household dispersion and two-thirds of the decline within households.

The remainder of the paper is organized as follows. Section 2 describes in detail the stylized facts on consumption, wages, and labour supply that provide the motivation for our study. Section 3 outlines the theoretical framework and the identification strategy for estimating the

2. Recently, there has been a substantial departure from this literature. Browning, Chiappori and Lewbel (2010) relax the assumption that household members split consumption equally in the construction of adult equivalence scales. Hong and Ríos-Rull (2007) use information on the purchase of life insurance to estimate equivalence scales.

3. In this instance, the degree of marital sorting is measured by the correlation between characteristics such as education across spouses.

rule to allocate consumption to individuals within a household. Section 4 describes the data set and Section 5, the strategy for estimating the model. The estimation results and sensitivity analysis are presented in Section 6. Section 7 presents a decomposition of consumption inequality and considers the importance of several explanations for the trends in consumption inequality. We also consider the implications for inequality in a wider definition of consumption that includes the market value of leisure. Section 8 concludes.

2. TRENDS IN CONSUMPTION AND EARNINGS INEQUALITY IN THE U.K.

In this section, we outline the main stylized facts regarding consumption and income inequality in the U.K. between 1968 and 2001. The data we use to conduct our analysis comes from the U.K. FES. The FES contains information on household consumption expenditures and earnings over the period 1968 to the present. In the construction of the following stylized facts, we restrict the sample to individuals between the ages of 16 and 65 and eliminate students, retirees, and the self-employed. We are particularly interested in the following features of the data:

- 1 There has been a large rise in earnings inequality between households. Figure 1 documents the rise in the variance of log household earnings from 1968 to 2001; increasing from 0.35 to 0.58. This rise in earnings inequality in the U.K. has been well documented in the literature (e.g. Blundell and Preston, 1998; Blundell and Etheridge, 2010).
- 2 There has been a corresponding rise in consumption inequality. To account for economies of scale, we construct a standard measure of individual-level consumption by dividing total household consumption by the square root of household size. The trend in the variance of log consumption, measured both at the household level and using the adult equivalence scale is also presented in Figure 1. The level of inequality is everywhere higher and the increase in inequality is more pronounced after 1980's for earnings than for consumption. The relative stability of consumption inequality prior to the 1980's, followed by a sharp increase during the 1980's and then relative stability again in the 1990's is a

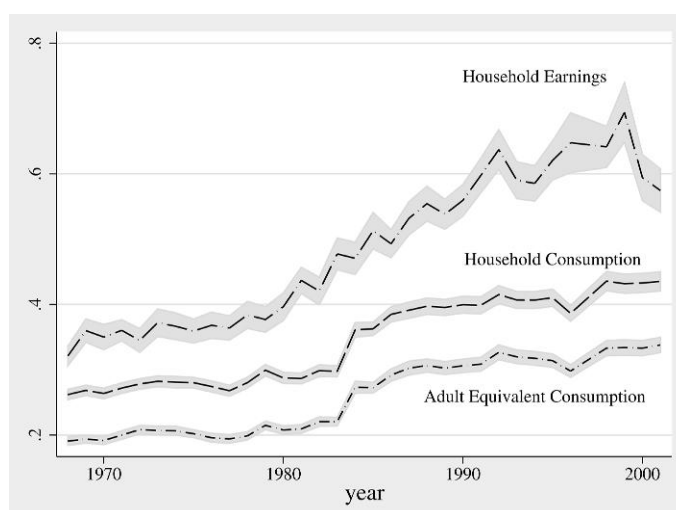


FIGURE 1
Trends in variance of log earnings and consumption

Note: Own calculations from the FES. Shaded area represents \pm two standard errors

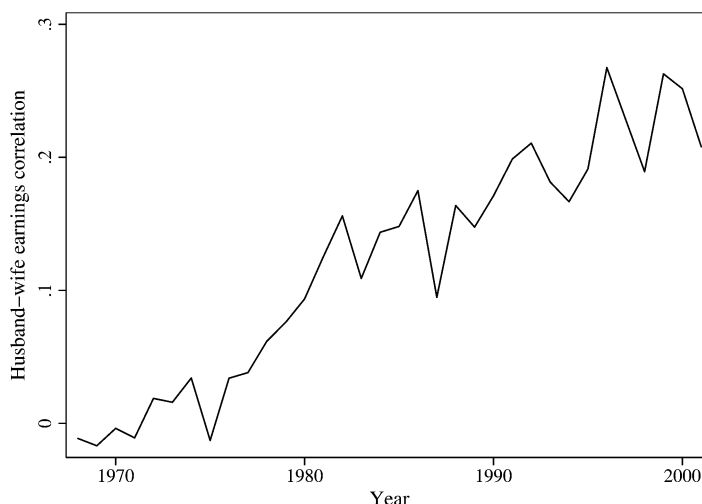


FIGURE 2
Correlation in earnings across husbands and wives

Note: Own calculations from the FES

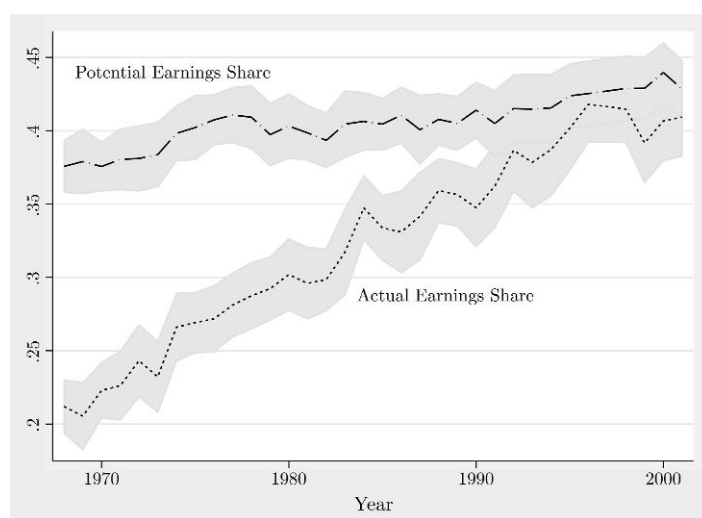


FIGURE 3
Fraction of actual and potential household earnings provided by wife

Note: Own calculations from the FES. Shaded area represents \pm two standard errors

well-documented feature of the U.K. data (see [Blundell and Preston, 1998](#); [Blundell and Etheridge, 2010](#)).

- 3 Turning our focus now to couples, we see from Figure 2 that the correlation between the earnings of husbands and wives increased dramatically over time. At the same time, the gender wage gap decreased and employment rates of married men and women converged. Figure 3 highlights the dramatic change in the gender wage gap and in women's contribution to household labour income between 1968 and the present. The dashed line

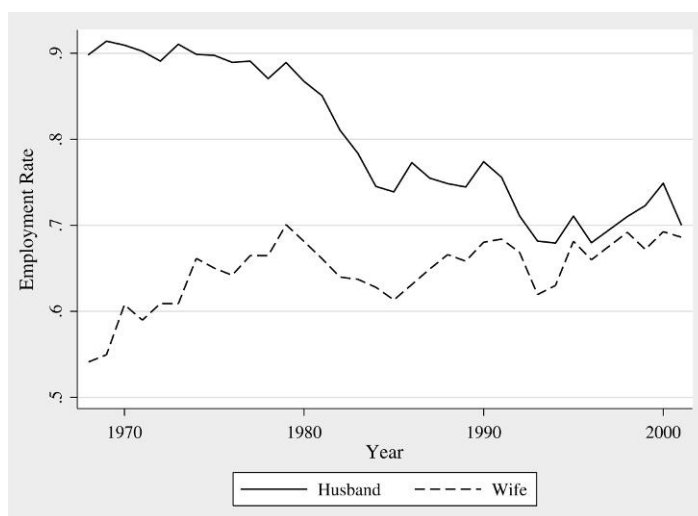


FIGURE 4

Employment rates of married men and women, 1968–2001

Note: Own calculations from the FES

represents the wife's share of potential earnings, defined as the share of labour earnings that would be contributed by the wife if both spouses worked full time.⁴ The dotted line represents the wife's share of actual household earnings. Overall, potential earnings of wives increased by 5 percentage points, and the share of earnings in the household increased by 20 percentage points over the sample period. The latter partly reflects the increase in women's wages relative to those of men, but also the large convergence in employment rates for married men and women in the U.K. since the late 1960's. In Figure 4, we can see the dramatic rise in the participation rates of married women, from 55% in 1968 to 70% in 2001. We also see the equally dramatic fall for married men, from 90% in 1968 to 70% in the year 2000. Employment rates were roughly equal for married men and women from the second half of the 1990's.

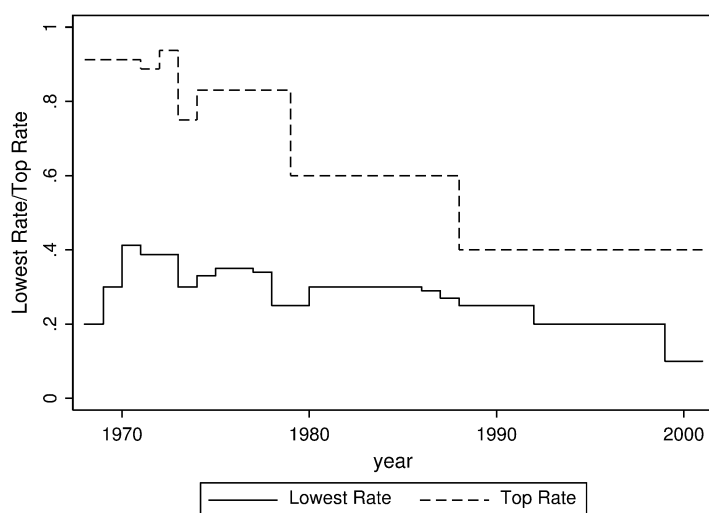
In summary, the evidence presented here highlights the fact that there has been a large rise in earnings and consumption inequality *between* households, while at the same time there has been a fall in inequality in the earnings distribution *within* households.

3. MODEL

The model we use to infer individual-level consumption is based on Chiappori's (1988; 1992) collective model of household decision making. This framework is less restrictive than the model of equal division underlying adult equivalence scales, as the only restriction on the intra-household allocation process is that households reach Pareto-efficient allocations.

We modify the basic collective model along three dimensions to address several issues of relevance to our analysis. First, we make a distinction between private and public consumption in the model. As public consumption is consumed by both members of the household, ignoring the presence of public consumption may overestimate the degree of inequality within households. We follow the strategy of Browning *et al.* (1994) and partition all expenditures into either public

4. See Section 4 for a discussion of the treatment of missing wages for non-participants.



National Statistics (1968–2001a,b)

FIGURE 5

Bottom and top marginal tax rates

National Statistics (1968–2001a,b)

or private expenditures. The goal of the estimation exercise is to determine the allocation of private expenditures across household members.

Second, our analysis spans the years 1968 through 2001, a time during which there have been major changes to the U.K. tax system, including dramatic changes in progressivity, and a move from taxation based on household earnings to taxation based on individual earnings. In Figure 5, we plot the time series for the top and bottom marginal tax rate over this period. Given these large changes that will affect both dispersion between and within households (the change from joint to individual taxation especially), it is important to account for taxes in our analysis. To this end, we assume that individuals choose from a discrete set of hours and base decisions on after-tax earnings.⁵

Third, the question we aim to address in this paper is whether measures of consumption inequality from the collective model differ from measures in the literature based on standard equivalence scales. To answer this question, we require an estimate of the full sharing rule to uncover the share of income allocated to each household member for consumption. In general, the sharing rule is identified only up to an additive constant (Chiappori, 1988, 1992). To identify the location of the sharing rule, we assume that when husbands and wives have equal wages, they share resources equally. This assumption has been imposed in related work, for example, Browning *et al.* (1994).

We start with a general description of the intra-household allocation decision of married couples and the model restrictions that allow for the identification of the share of private consumption allocated to each household member.⁶

5. See Blundell, MaCurdy and Meghir (2007b) for a discussion of the advantages of treating hours as discrete when the tax system induces a non-linear and possibly non-convex budget set in a unitary setting. With few exceptions, *e.g.* Donni (2003) and Vermeulen (2006), the literature on collective models ignores taxation.

6. Of course, there are aspects of the household's problem that we do not address, including savings and risk-sharing behaviour. See Mazzocco (2004, 2007) for important work on these issues.

3.1. Collective labour supply

Consider a two-member household, where each member has distinct preferences over own leisure (ℓ^i), own private consumption (c^i), and household public consumption (c^P). We assume that we can distinguish in the data between private and public consumption but not the allocation of private consumption between spouses. Thus, for each household, we observe $\{\ell^f, \ell^m, c, c^P\}$, where $c = c^f + c^m$. We assume that preferences over private consumption and leisure are separable from consumption of the public good. We test the sensitivity of our results to this assumption in Section 6.2. Since public consumption is consumed by both household members, it is only necessary to uncover the sub-utility for private consumption and leisure and the sharing rule to determine the share of total consumption allocated to each household member.⁷

Under the assumptions that preferences are egoistic and that allocations are Pareto efficient, the household's allocation can be represented as the solution to the problem:

$$\max_{c^f, c^m, c^P, \ell^f, \ell^m} \lambda(\pi, y, \mathbf{z}) U^f(u^f(c^f, \ell^f), c^P) + (1 - \lambda(\pi, y, \mathbf{z})) U^m(u^m(c^m, \ell^m), c^P), \quad (1)$$

subject to

$$c^f + c^m + pc^P + [\bar{\omega}^f - \omega^f(h^f, h^m)] + [\bar{\omega}^m - \omega^m(h^m, h^f)] = \bar{\omega}^f + \bar{\omega}^m + y^{nl}, \\ c^f + c^m = c, \quad h^g + \ell^g = T, \quad \text{and} \quad h^g \in H, \quad (2)$$

where the Pareto weight, $\lambda(\pi, y, \mathbf{z})$, represents the wife's relative bargaining power within the household and depends on prices (π), total resources (y), and distribution factors (\mathbf{z}).

The budget constraint is written with full income on the right-hand side, expenditures on consumption, and leisure on the left-hand side. Full income, y , comprises potential after-tax earnings and non-labour income, $\bar{\omega}^f$ and $\bar{\omega}^m$ denote potential after-tax earnings of females and males, respectively, when both partners work the maximum hours, and y^{nl} is non-labour income. The notation $\omega^f(h^f, h^m)$ accommodates the fact that the tax system in the U.K. is progressive and, prior to 1990's, was based on household as opposed to individual incomes with the wife explicitly defined as the secondary earner. As a result, after-tax earnings and non-labour income depend on the labour supply decisions of both spouses. With this notation, total expenditure on leisure for the wife is written as $[\bar{\omega}^f - \omega^f(h^f, h^m)]$ and similarly for the husband.

In the case without joint taxation, [Blundell, Chiappori and Meghir \(2005\)](#) show that the intra-household allocation problem with public goods can be decentralized by considering a two-stage process. In the first stage, the husband and wife decide on the level of public good expenditure (pc^P) and on how to divide the remaining full income $\bar{y} = \bar{\omega}^f + \bar{\omega}^m + \bar{y}^{nl}$, where $\bar{y}^{nl} = y^{nl} - pc^P$. The assumption that leisure and private consumption are separable from consumption of the public good is the key to allowing the allocation of public consumption to occur in the first stage.

The joint taxation of incomes complicates the decentralization of the household problem. To address this complication, we define the *conditional* earnings for individual i , $\omega^i(h^i, h_*^i)$, as the after-tax earnings of an individual working h^i hours, conditional on the equilibrium labour supply h_*^j of their spouse. The household's problem can then be decentralized as follows. In the first stage, spouses agree on a level of public goods expenditures, on conditional earnings functions for each household member $\omega^f(h^f, h_*^m)$ and $\omega^m(h^m, h_*^f)$, and on a sharing rule to divide the remaining full income. Define the sharing rule $\phi(\bar{\omega}^f, \bar{\omega}^m, y, \mathbf{z})$ as the non-labour income net of public goods expenditures transferred to the wife. This transfer may be positive or negative.

7. The implicit price of public consumption (*i.e.* the Lindahl price) is useful for welfare analysis, an issue we do not consider here. We do not attempt to estimate preferences for public consumption, rather we condition on total public consumption expenditures pc^P , where p is the relative price of public consumption.

The sharing rule is assumed to depend on potential earnings for the wife and husband, $\bar{\omega}^f$ and $\bar{\omega}^m$, full income, y , and distribution factors \mathbf{z} . The distribution factors are variables that influence the Pareto weight but do not directly affect either preferences or the budget set. In the second stage, each household member chooses labour supply and private consumption to maximize utility. The problem facing the wife in the second stage is

$$\begin{aligned} \max_{c^f, \ell^f} u^f(c^f, \ell^f) \\ \text{subject to } c^f + [\bar{\omega}^f - \omega^f(h^f, h_*^m)] &= \bar{\omega}^f + \phi(\bar{\omega}^f, \bar{\omega}^m, y, \mathbf{z}), \\ \ell^f + h^f &= T, \quad \text{and } h^f \in H, \end{aligned} \quad (3)$$

and the husband faces the problem

$$\begin{aligned} \max_{c^m, \ell^m} u^m(c^m, \ell^m) \\ \text{subject to } c^m + [\bar{\omega}^m - \omega^m(h^m, h_*^f)] &= \bar{\omega}^m + \bar{y}^{\text{nl}} - \phi(\bar{\omega}^f, \bar{\omega}^m, y, \mathbf{z}), \\ \ell^m + h^m &= T, \quad \text{and } h^m \in H. \end{aligned} \quad (4)$$

Optimal working hours for each spouse are obtained from a system of reaction functions, similar to a Nash equilibrium, that yield Pareto-efficient solutions. Empirically, estimation proceeds by searching for preference and sharing-rule parameters that maximize the probability that household member i chooses labour supply h_*^i given their spouse chooses h_*^j and the household public goods expenditure is pc_*^P .

3.2. Double indifference

In the case of a collective model in which the wife makes a continuous labour supply choice and the husband decides only whether to participate or not, [Blundell et al. \(2007a\)](#) show that efficiency requires *double indifference*: at the reservation wage of the husband, both the husband and the wife must be indifferent to whether he works or not. In order for the wife to be indifferent at the reservation wage, the husband must consume all the additional earnings associated with switching between not working and working, keeping the total consumption of the wife unchanged.

The same logic for double indifference holds in our model where all labour supply choices are discrete. If the husband is indifferent between working h_j and h_k hours, then efficiency requires the wife is indifferent as well. Again, the husband must consume all the difference in earnings resulting from working h_j versus h_k hours. It is important to note here that the sharing rule does not depend on the labour supply choice of either partner; the husband's share of household full income does not change with his labour supply choice.⁸ However, his share of household consumption does. When one member of the household increases labour supply, leading to an increase in own earnings, their share of household consumption increases but in a very particular way. Their consumption increases by exactly the amount that household consumption increased, leading to an increase in their share of total household consumption.

8. The exclusion of the labour supply choice from the sharing rule is a direct result of the maintained assumption that households reach efficient allocations. If the labour supply choice of an individual affected both the level of their earnings and their share of household full income then individuals would typically work more than the efficient number of hours, in contradiction of the maintained identifying assumption of efficiency.

3.3. Identification

As mentioned above, our primary interest is in allocating private consumption between the household members. This requires that we obtain an estimate of the sharing rule $\phi(\bar{\omega}^f, \bar{\omega}^m, y, \mathbf{z})$. Generically, this sharing rule is identified up to an additive constant (Chiappori, 1988, 1992; Chiappori, Fortin and Lacroix, 2002). The intuition behind identification is the following. Consider the wife's problem presented in equation (3). Her choice of labour supply depends on her own wage and the sharing rule ϕ . It depends on her husband's wage, household full income, and on the distribution factors \mathbf{z} only through their effect on the sharing rule. Thus, observed variation in the wife's labour supply choices in response to variation in either her husband's wage, full income, or the distribution factors are informative about how the sharing rule changes with changes in these variables. The same intuition applies to observed variation in the husband's choices in response to the wife's wage, household full income, and distribution factors. A formal presentation of how the sharing rule can be identified from the observed conditional choice probabilities for hours worked is presented in Appendix A.

In addition to identifying how the sharing rule changes with respect to changes in wages, non-labour income, and distribution factors, our exercise requires identification of the location of the sharing rule. As in Browning *et al.* (1994), among others, we impose a symmetry assumption that partners with equal earnings potential will share household full income equally. As a point of comparison, the adult equivalence scale approach assumes that partners share total consumption equally regardless of either differences in incomes or preferences. The implication of our identifying assumption imposes this equality only in the special case in which the husband and wife have equal earnings potential and, in addition, happen to make identical hours choices. In other words, we are assuming husbands and wives have the same consumption when they are identical in terms of resources and preferences. We do not impose any particular pattern of sharing when partners differ in terms of earnings potential or preferences.

3.3.1. Caring preferences and joint spousal leisure. Up to this point, we have assumed that individuals have egoistic preferences. This hardly seems reasonable within a family. It is not feasible to identify a model in which each household member has preferences over both own and spousal private consumption and leisure (paternalistic preferences). It is however feasible to identify a model in which household members have Beckerian caring preferences, where couples derive utility from their spouse's consumption only in so far as their spouse derives utility from it (Becker, 1991). Specifically, preferences of the form

$$\begin{aligned} V^f &= F^f(u^f(c^f, \ell^f), u^m(c^m, \ell^m)), \\ V^m &= F^m(u^m(c^m, \ell^m), u^f(c^f, \ell^f)) \end{aligned}$$

pose no obstacle to estimating the sharing rule. The only issue is one of interpretation. With caring preferences, our estimated sharing rule will be a mixture of caring and sharing.

The following simple example from Browning, Chiappori and Lechene (2006) illustrates the point. A special case of caring preferences is

$$\begin{aligned} V^f &= u^f(c^f, \ell^f) + \mu^f u^m(c^m, \ell^m), \\ V^m &= u^m(c^m, \ell^m) + \mu^m u^f(c^f, \ell^f), \end{aligned}$$

with μ^f and μ^m both between zero and one. Now, the overall household utility can be expressed as

$$\lambda(\pi, y, \mathbf{z})V^f + (1 - \lambda(\pi, y, \mathbf{z}))V^m = (\lambda + (1 - \lambda)\mu^m)u^f(c^f, \ell^f) + (\lambda\mu^f + (1 - \lambda))u^m(c^m, \ell^m),$$

which is simply a reweighted version of equation (1), where the Pareto weights still depend on (π, y, \mathbf{z}) . Our estimated sharing rule then will be consistent with either egoistic or caring preferences, where the estimated sharing rule absorbs the degree of caring.⁹

The model we present in Section 3.1 assumes that leisure is a private good. It is quite natural to think husbands and wives would have a preference for joint leisure time. Fong and Zhang (2001) present analysis of a collective model in which non-market time is divided into private and joint leisure. The formal implications are essentially unchanged. The same two-stage decision process can be applied as discussed above. With joint leisure, the husband and wife will decide in the first stage the amount of time to spend together, and then in the second stage individually choose their personal allocation of remaining time to leisure and market work. This process is exactly analogous to deciding total public goods expenditure in the first stage. In the first stage, the couple decides how much of full income to spend on public goods and joint leisure. In the second stage, they independently allocate their share of the remaining full income between private leisure and consumption.

The implications that public goods and joint spousal leisure have for the sharing rule are the same. The larger the share of full income devoted to public goods and joint leisure, the smaller is the scope for within-household inequalities in private consumption. In principle, it is possible to identify preferences for joint versus private leisure time. This of course requires data with information on time use for household members, including time spent together. This information is not available in the U.K. FES.

4. DATA

The data we use come from the U.K. Family Expenditure Survey (FES, Department of Employment, 1969–2001, National Statistics (1968–2001a,b)). These data are ideal for the study of consumption inequality for three reasons. First, the FES contains detailed information on private and public consumption expenditures for households, on wages and labour supply for individuals, and on demographics including age, education (from 1978 onwards), and region of residence. Second, the FES has fewer problems with measurement issues than the leading contenders in the U.S. and elsewhere.¹⁰ The FES uses a weekly diary to collect data on frequently purchased items and uses recall questions to collect data on large and infrequent expenditures. Finally, the FES contains information over the period 1968 to the present that allows the study of changes in consumption inequality over a long period of time.

Our sample includes single-person households and couples without children. We exclude households with children in this paper to abstract from the intra-household allocation of resources for children's consumption. This is obviously an important issue. To this end, our estimates of the sharing rule and the comparison of various inequality measures only apply to households without children.¹¹ Focussing our attention on single individuals and childless couples, we see in Figure 6 that the pattern of increase in inequality for this group is virtually identical to the wider sample but everywhere shifted up 0.05 points.

We restrict the age range in the sample to individuals between the ages of 22 and 65, who were born between 1910 and 1969, and eliminate students and the self-employed. For robust-

9. Indeed, we can define the new weight as $\tilde{\lambda}(\pi, y, \mathbf{z}) = (\lambda + (1 - \lambda)\mu^m) / (1 + \lambda\mu^f + (1 - \lambda)\mu^m)$. With $\lambda \in [0, 1]$, we have $\tilde{\lambda} \in [\mu^m / (1 + \mu^m), 1 / (1 + \mu^f)]$, which moves the weight towards one-half as the degree of caring increases.

10. Battistin (2003) documents reporting errors in the U.S. Consumer Expenditure Survey (CEX) due to survey design.

11. We leave to future work an analysis of consumption inequality for the entire sample of households.

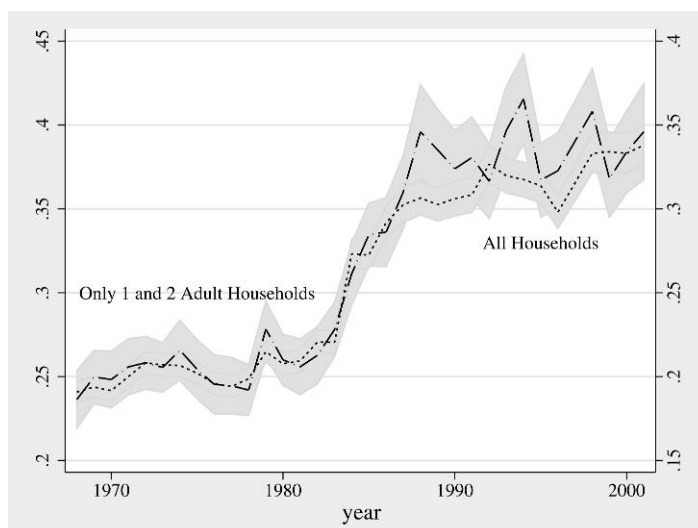


FIGURE 6

Trends in the variance of log consumption for all households (dotted, right scale) and one- and two-adult households (dashed, left scale)

Note: U.K. National Statistics (1968–2001). Shaded area represents \pm two standard errors. Consumption is by adult equivalent, using square root of household size

ness, households in which one of the individuals is in the top 1% of the wage distribution are also excluded. The resulting sample contains 81,949 individuals.¹² Descriptive statistics for our entire sample are presented in Table 1.

We define consumption and non-labour income measures as follows. Total consumption is defined as total household expenditures. Public consumption is defined as expenditures on housing, light and power, and household durable goods. Private household consumption is total expenditures net of public consumption. Non-labour income is defined as total household expenditures minus net labour income. We use this expenditure-based definition of non-labour income as it reduces measurement error and accounts for sources of wealth we do not observe including pension wealth (Blundell and Walker, 1986; Blundell *et al.*, 2007a).

To construct the level of consumption corresponding to each labour supply decision, including zero hours, we need to assign a wage to all individuals. For those who are working, we use the usual hourly wage, defined as weekly earnings divided by usual weekly hours. For non-participants, we use a predicted wage, computed based on a reduced-form selection-corrected wage equation. The log of the wage is estimated as a function of age, birth cohort, year, quarter, and regional dummies, plus the age at which full-time education was completed and its square. The selection equation is identified by the exclusion from the wage equation of household non-labour income, marital status, and the age, education, and labour income of the spouse.¹³

Labour supply is assumed to take a value in $H = \{0, 5, 10, \dots, 65\}$. After-tax earnings are subsequently computed for each person individually, at each possible hours choice, by converting weekly wage income to an annual base, deducting the appropriate personal allowance and then

12. The sample sizes by birth cohort, sex and marital status are presented in Table 1.

13. Results are available from the authors upon request. Using predicted wages in this way assumes that wage profiles are exogenous, ruling out human capital accumulation due to learning by doing.

TABLE 1

	Male			Female		
	Single	Married		Single	Married	
1968-2001	Mean	Standard deviation	Mean	Standard deviation	Mean	Standard deviation
Age (22-65), years	44.29	12.95	48.50	13.17	50.50	12.74
Zero hours dummy	0.28	0.45	0.19	0.39	0.42	0.48
Hourly wage	4.96	2.42	4.78	2.27	3.88	1.95
Non-labour income	13.03	90.92	25.82	129.83	40.25	76.47
Total expenditures	118.28	99.17	193.77	129.63	100.71	76.39
Public expenditures	39.44	41.80	59.85	62.23	39.48	38.73
Observations						
1910 birth cohort	807		4,963		2,263	
1920 birth cohort	1,612		7,203		2,821	
1930 birth cohort	1,809		6,086		2,360	
1940 birth cohort	2,069		5,506		1,737	
1950 birth cohort	2,135		3,814		1,424	
1960 birth cohort	1,825		2,279		1,235	
Total	10,257		29,851		11,840	
Observed wages	7.158		23.650		6.711	

applying the appropriate tax rate. Personal allowances and marginal tax rates are from the *Inland Revenue Statistics* (National Statistics, 1968–2001a). For the years prior to 1990, taxation was done at the household level. For these years, after-tax earnings for each hours choice are calculated conditional on the actual earnings of the spouse, $\omega^i(h^i, h_*^i)$. All monetary values are expressed in 1987 pounds. The resulting income measure is treated as known and is also used to construct the potential share of after-tax household labour income contributed by the wife, $\bar{\omega}^f/(\bar{\omega}^f + \bar{\omega}^m)$. Individuals may also be entitled to income related to earnings when working zero hours, for instance unemployment benefits, so we also predict unemployment and other benefits for those who are working based on the *Official Yearbook of the United Kingdom* (National Statistics, 1968–2001b).

5. ECONOMETRIC SPECIFICATION

We specify a CEX sub-utility function

$$u(c, \ell) = [\alpha c^\rho + (1 - \alpha)\ell^\rho]^{1/\rho} \quad (5)$$

allowing the share coefficient (α) to depend on observables and the substitution coefficient (ρ) to vary by sex. To ensure that during estimation neither spouse receives either negative full income or full income that exceeds the household resources, we define $\varphi \in [0, 1]$ to be the share of household full income transferred to the wife. This is simply a convenient transformation of the transfer ϕ : $\varphi \equiv \frac{\bar{\omega}^f + \phi}{\bar{y}}$ and $(1 - \varphi) \equiv \frac{\bar{\omega}^m + \bar{y}^{nl} - \phi}{\bar{y}}$. We parameterize this share as

$$\varphi(\mathbf{z}) = \frac{\exp(\mathbf{z}'\varphi)}{1 + \exp(\mathbf{z}'\varphi)}. \quad (6)$$

Here, \mathbf{z} contains the wife's potential share of household earnings, $\bar{\omega}^f/(\bar{\omega}^f + \bar{\omega}^m)$, the log of full household income ($\log y$), and the spousal age gap. Wages enter the sharing rule here as the wife's potential share of earnings to capture the notion that relative earnings power is likely to influence the intra-household allocation of resources. The dependence on full income allows the possibility for sharing to be more or less equal depending on the total household resources. We remove time effects from $\log y$ by projecting off a complete set of time dummies. Thus, the variation comes from dispersion in the cross section not from growth over time.

We make the following additions to the model before estimation. First, we introduce several sources of individual-specific heterogeneity. We allow the weighting of private consumption relative to leisure to differ across individuals according to a set of demographic characteristics including a quadratic in age, education, and dummy variables for decade of birth:

$$\alpha(\mathbf{x}_i) = \frac{1}{1 + \exp(\mathbf{x}_i'\alpha)},$$

where the coefficients α differ by sex. We also introduce additive unobserved individual heterogeneity in preferences η_{ij} , specific to individual i and labour supply choice j

$$\eta_{ij} = v_i h_{ij} + \varepsilon_{ij} \quad (7)$$

which has two components. The first component is a normally distributed, individual-specific component that allows unobserved tastes for work (possibly correlated with non-labour income as discussed below). The second component is i.i.d. type I extreme value preference heterogeneity.

Second, we treat non-labour income as endogenous and measured with error. Much of non-labour income depends on past labour income and thus may be endogenous to current labour

supply decisions. To control for this potential endogeneity, we instrument non-labour income using capital gains in the local housing market as instruments, following [Hurst and Lusardi \(2004\)](#). Capital gains in the housing market are measured by quarterly year-over-year changes in real house prices in each of 12 regions in the U.K., interacted with a home ownership indicator. The estimating equation for non-labour income for household i is

$$y_i^{\text{nl}} = \mathbf{w}_i' \beta + e_i, \quad (8)$$

where \mathbf{w}_i includes the capital gains, plus the demographics \mathbf{x}_i .¹⁴ We then control for endogenous non-labour income by conditioning the unobserved heterogeneity v_i on the residual, $\hat{e}_i = y_i^{\text{nl}} - \mathbf{w}_i' \hat{\beta}$ (see [Telser, 1964](#); [Heckman, 1979](#); [Smith and Blundell, 1986](#)).

Making use of these functional forms for utility (5) and the sharing rule (6), and substituting in the budget and time constraints (3) and (4), we can write the value to individual married woman i of labour supply choice $h_{ij}^f \in H$ as

$$\begin{aligned} V_{ij}^f(\mathbf{Z}_i, \theta) = & (\alpha^f(\mathbf{x}_i)[\omega^f(h_{ij}^f, h_{i*}^m) - \bar{\omega}_i^f + \varphi(\mathbf{z}_i)\bar{y}_i]^{\rho_f} \\ & + (1 - \alpha^f(\mathbf{x}_i))[T - h_{ij}^f]^{\rho_f})^{1/\rho_f} + v_i^f h_{ij}^f + \varepsilon_{ij}^f, \end{aligned}$$

and similarly for married man as

$$\begin{aligned} V_{ij}^m(\mathbf{Z}_i, \theta) = & (\alpha^m(\mathbf{x}_i)[\omega^m(h_{ij}^m, h_{i*}^f) - \bar{\omega}^m + (1 - \varphi(\mathbf{z}_i))\bar{y}_i]^{\rho_m} \\ & + (1 - \alpha^m(\mathbf{x}_i))[T - h_{ij}^m]^{\rho_m})^{1/\rho_m} + v_i^m h_{ij}^m + \varepsilon_{ij}^m, \end{aligned}$$

where \mathbf{Z}_i contains all relevant observables relating to both the preference heterogeneity, and the sharing rule, and θ contains all preference parameters and sharing-rule parameters to be estimated.

Estimation proceeds in the following steps. First, we estimate a selection-corrected wage equation and predict wages for individuals that are not working and predict after-tax labour earnings for each possible hours choice, as described in Section 4. Second, we estimate the reduced-form equation for non-labour income to obtain the residuals, \hat{e}_i . Third, we obtain estimates of the preference parameters α^g , ρ^g , σ_v^g , and $\sigma_{v_e}^g$ and the sharing rule parameters φ by estimating a mixed Logit model for the discrete labour supply choice using our sample of married men and women.¹⁵ The reported standard errors are bootstrapped.¹⁶

14. We would like to thank a referee for this suggestion. The F-statistic for the instruments not included in labour supply and the sharing rule (regional change in house price interacted with tenure) is 5.64.

15. The contribution of an individual i to the likelihood function is the probability of observing individual i making labour force decision h_{ij} and has the mixed Logit form ([McFadden and Train, 2000](#)):

$$\mathcal{L}(h_{ij}, \mathbf{Z}_i, \theta) = \int \frac{1}{\sum_{k \in H} \exp(V_{ik}^g(\mathbf{Z}_i, \theta) - V_{ij}^g(\mathbf{Z}_i, \theta))} dF(v | \hat{e}_i),$$

where we integrate out over unobserved tastes for work, conditional on the unpredictable component of non-labour income. To implement the dependence of v on \hat{e} , we assume that v , conditional on \hat{e} , is normally distributed, standardize the residual to have variance one, and make use of the Choleski factorization $\begin{pmatrix} \hat{e} \\ v \end{pmatrix} = \begin{pmatrix} 1 & 0 \\ s_{21} & s_{22} \end{pmatrix} \begin{pmatrix} \hat{e} \\ \eta \end{pmatrix}$, with $\eta \sim N(0, 1)$. We then replace $v = (s_{21}\hat{e} + s_{22}\eta)$ in equation (7) and integrate over η . The standard deviation of v and the correlation with \hat{e} are estimated as $\sigma_v = \sqrt{s_{21}^2 + s_{22}^2}$ and $\text{corr}(v, \hat{e}) = (s_{21}/s_{22})$.

16. This is a non-parametric bootstrap in which we resample with replacement from the original data, clustering by birth cohort. The reported standard errors are calculated as the standard deviation of 99 bootstrapped estimates. As a robustness exercise, replications of size 999 were carried out for the 1930's birth cohort. Both produced very similar estimates of the standard errors.

6. ESTIMATION RESULTS

We begin with estimates of the preference parameters, presented in Table 2. The CEX specification guarantees the utility function is increasing and quasi-concave for all the data. The elasticity of substitution ($\frac{1}{1-\rho}$) indicates that the degree of substitutability between private consumption and leisure is similar across gender, 1.82 for women and 1.53 for men. The share parameters suggest a slightly lower weight on consumption for women (0.146) than for men (0.212) at the mean of the data, although the share parameter for males tends to vary across observables to a greater extent than for females. Males also display slightly less variation in unobserved tastes for work, as indicated by the variance of v . Finally, the results suggest that the unobserved tastes for work are positively correlated with the unobserved component of non-labour income for men, while the correlation is small and negative for women.

We now turn to the estimates of the sharing-rule parameters, the parameters that allow us to infer the share of consumption attributed to each adult in the household. As discussed in

TABLE 2
Estimates of the preference parameters

	Base estimates		Separability test	
	Women	Men	Women	Men
ρ	0.453 (0.006)	0.348 (0.007)	0.432 (0.016)	0.315 (0.013)
$\alpha(\mathbf{x})$ at mean of data	0.146 (0.003)	0.212 (0.004)	0.088 (0.008)	0.137 (0.015)
α :				
Age	-0.319 (0.070)	-0.846 (0.068)	-0.259 (0.057)	-0.716 (0.150)
Age ²	0.782 (0.071)	1.342 (0.070)	0.844 (0.070)	1.268 (0.112)
Education	0.023 (0.013)	0.107 (0.010)	0.094 (0.033)	0.138 (0.029)
1910 cohort	-0.097 (0.038)	-0.681 (0.033)	-0.295 (0.072)	-0.864 (0.135)
1920 cohort	-0.085 (0.033)	-0.299 (0.028)	-0.203 (0.085)	-0.410 (0.121)
1940 cohort	0.083 (0.024)	0.231 (0.032)	0.173 (0.080)	0.304 (0.064)
1950 cohort	0.131 (0.040)	0.401 (0.040)	0.250 (0.081)	0.485 (0.064)
1960 cohort	0.242 (0.066)	0.564 (0.038)	0.529 (0.100)	0.828 (0.081)
α_0	1.766 (0.028)	1.314 (0.023)	2.345 (0.088)	1.845 (0.171)
σ_v	0.484 (0.020)	0.427 (0.006)	0.477 (0.014)	0.410 (0.041)
$\text{corr}(v, e)$	-0.038 (0.018)	0.144 (0.018)	-0.026 (0.017)	0.155 (0.025)
Public goods expenditure			0.0060 (0.0002)	0.0057 (0.0004)

Note: $\alpha_i = \frac{1}{1+\exp(\mathbf{x}_i'\alpha)}$. Bootstrapped standard errors in parentheses.

TABLE 3
Sharing rule estimates

Estimates	Pooled	Separability Test	Birth cohort					
			1910	1920	1930	1940	1950	1960
φ (mean of data)	0.396 (0.049)	0.386 (0.041)	0.345 (0.109)	0.374 (0.161)	0.394 (0.066)	0.405 (0.079)	0.424 (0.043)	0.442 (0.043)
Marginal effect of								
Wife's potential share of earnings	1.021 (0.414)	1.114 (0.394)	1.169 (0.416)	1.039 (0.466)	1.016 (0.288)	1.025 (0.360)	1.044 (0.284)	0.895 (0.319)
log of full income	-0.137 (0.036)	-0.161 (0.063)	-0.199 (0.142)	-0.173 (0.304)	-0.156 (0.106)	-0.140 (0.219)	-0.129 (0.163)	-0.063 (0.122)
Husband–wife age gap	-0.001 (0.002)	-0.001 (0.002)	-0.004 (0.098)	-0.0003 (0.021)	-0.003 (0.104)	-0.002 (0.027)	0.001 (0.104)	0.002 (0.066)
Test null of separability			Test null of equality of estimates					
<i>p</i> -value	0.878		0.921					

Note: Bootstrapped standard errors in parentheses. The sharing rule is parameterized as $\varphi(\mathbf{z}_i) = \frac{\exp(\mathbf{z}_i' \boldsymbol{\varphi})}{(1 + \exp(\mathbf{z}_i' \boldsymbol{\varphi}))}$, and the marginal effect is evaluated at the mean of \mathbf{z} . The estimated φ coefficients corresponding to the first column are $\hat{\varphi} = [4.2676, -0.5711, -0.0041]'$, with $\mathbf{z} \equiv \left[\frac{\hat{\omega}_f}{\hat{\omega}_f + \hat{\omega}_m} - \frac{1}{2}, \log y - \log \bar{y}_t, (\text{age}_f - \text{age}_m) - (\text{age}_f - \text{age}_m) \right]'$. The variables entering the vector \mathbf{z} are transformed such that the sharing rule is normalized to equal one-half when husband and wife have equal potential wages, evaluated at the (year-by-year) mean of log full income and the mean husband–wife age gap. We remove any trend from log of full income by projecting it off of a full set of time dummies. The test for the null of separability is a test that the first two columns are equal.

Section 3, we impose the identifying assumption that equal sharing of full income occurs at equal potential earnings (evaluated at mean log full income and mean spousal age difference).¹⁷ In the first column of Table 3, we present estimates of the sharing rule. At the mean of the data, the point estimates imply that the wife receives 39.6% of full household income, net of public goods expenditure.

The marginal effect of the share of potential earnings due to the wife is 1.021; A one percentage point increase (from the mean) in the wife's share of potential earnings translates into slightly more than a one percentage point increase in her share of full income. The negative coefficient on log full income (−0.137) indicates that as full income increases, a smaller share of this increase is allocated to the wife than to the husband (although the level of full income she receives may increase, as illustrated below). The effect of the spousal age gap on sharing is effectively zero, implying that identification of the sharing rule requires the choice probabilities to be twice differentiable in potential earnings and non-labour income (see Appendix A).¹⁸

17. In an earlier version of this paper, we identified the location of the sharing rule by assuming that the sub-utility over private consumption and leisure did not depend on marital status (see Vermeulen, 2006; Browning, Chiappori and Lewbel, 2010, for similar identification strategy). We then used data on singles to identify preferences and data on married couples to identify the full sharing rule. Estimates from that version of the model implied that equal sharing occurred at equal wages.

18. This requirement will be satisfied by any empirical specification that requires probabilities to be between zero and one, including the mixed logit specification used here.

To be concrete, consider a household in which, initially, the husband and wife have equal potential earnings, $\bar{\omega}^f = \bar{\omega}^m = \bar{\omega}$, and non-labour income and public goods expenditures are zero $\bar{y}^{nl} = 0$. Full income is $\bar{y} = 2\bar{\omega}$. Assume that $\bar{\omega}$ is such that this household has mean log full income and has the average age gap. Then full income is divided equally between the husband and wife.

Now, consider an increase in non-labour income that results in a 10% increase in household full income. The new value of the sharing rule becomes $\varphi' = \varphi - 0.137 \times \log 1.1 = 0.487$. The wife now receives a smaller share of full income. But, household full income has also increased, leading to an increase in the wife's full income from $\bar{y}^f = 0.5\bar{y}$ to $\bar{y}_{\text{new}}^f = 0.487 \times 1.1\bar{y} = 0.536\bar{y}$, an increase of 7.12%. The corresponding increase to the husband's full income is 12.88%.

Second, consider a 20% rise in the wife's potential earnings, holding the husband's potential earnings constant, resulting in an increase to household full income of 10%. The resulting new value of the sharing rule is $\varphi' = \varphi + 1.021 \times 0.045 - 0.137 \times \log 1.1 = 0.5328$. The wife is now receiving a larger share of full income (the level of which has also increased), increasing her individual full income by 17.23%. At the same time, she shares some of her wage increase with her husband, his full income rises by 2.78%.

Finally, consider the case in which the wife's potential earnings increases by 10% and the husband's decreases by 10%, leaving full income unchanged. Here, the change in full income going to each spouse is nearly exactly equal to the change in their potential earnings. The new value of the sharing rule is $\varphi' = \varphi + 1.021 \times 0.05 = 0.551$.

6.1. *Adult equivalence scales revisited*

One of the main goals of this paper is to determine whether measures of consumption inequality using standard adult equivalence scales provide an accurate estimate of consumption inequality across individuals. Recall, adult equivalence scales assume that husbands and wives share household consumption equally, independently of the source of income, or distribution factors. We have made an identifying assumption of equal sharing of full income at equal wages (with the implication of equal consumption when the partners also have identical preferences). Equal sharing at *all* wages would require that the effect of the wife's potential share in earnings were zero. A zero effect of the wife's potential share in earnings is rejected.

6.2. *Sensitivity analysis*

We consider the sensitivity of our estimation results to several modifications. The first check we consider is whether the results are sensitive to our assumption that private consumption and leisure are separable from public consumption expenditure. Under the null of separability, public goods expenditure should only affect allocations through the effect on the budget constraint: $\bar{y} = \bar{\omega}^f + \bar{\omega}^m + y^{nl} - pc^P$. We can test the separability assumption by testing whether pc^P , appropriately controlling for the fact that it is endogenous, affects the estimated parameters beyond the direct effect through the budget constraint we already include (see [Browning and Meghir, 1991](#)). We proceed by estimating a reduced-form equation for pc^P and calculate the residual.¹⁹ We then allow public expenditures and the residual to influence the consumption–leisure choice by interacting it with the hours term in the equation for unobserved tastes for

19. The reduced form regresses public expenditures on total expenditures plus age, age squared, and education of each spouse, interacted with year dummies. Total expenditures are instrumented by full income and capital gains in the local housing market.

work equation (7).²⁰ While this is not the specification we would naturally choose if we were interested in estimating preferences for public consumption, it is flexible enough to test the separability assumption since, under the null of separability, adding pc^P should not affect the estimates of our parameters of interest (the sharing rule). This is a Hausman test where under the null, both the estimates are consistent but excluding pc^P is efficient. Under the alternative, excluding pc^P produces biased estimates. The last two columns of Table 2 present the estimates of the preference parameters allowing for non-separability. Although public expenditures do have a significant coefficient (the last row), and the other preference parameters do move some, we see in Column 2 of Table 3 that the effect on the sharing rule estimates is not significant. The null hypothesis that the sharing rule parameters are unaffected by the separability assumption is not rejected, with a p -value of 0.878. While we may not want to take this as definitive evidence for separability, it is clear that our sharing-rule estimates, which are the parameters of interest, are not sensitive to this assumption.

The next specification we estimate allows for differences in the sharing-rule parameters for each birth cohort in our pooled sample. The sample covers a long time period and a wide age range in every year; we thus estimate separate sharing rules for each ten-year birth cohort in the data. The parameter estimates are presented in the last 6 columns of Table 3. There appears to be some slight variability in the point estimates across cohorts and a slight trend over time, however, the parameters are not statistically different from each other (p -value of 0.921).

7. CONSUMPTION INEQUALITY

In this section, we compare the inequality measure implied by our model to a conventional measure of consumption inequality. We use the estimated sharing rule to divide full income between the husband and wife in each household. We subsequently construct private consumption based on the individual's share of full income and his or her personal net labour earnings based on each individual's private budget constraint. Our *sharing rule* measure of individual consumption, for married individuals, is then equal to individual private consumption plus household public consumption expenditure:

$$\begin{aligned}\hat{c}_i^f &= \omega^f(h_*^f, h_*^m) - \bar{\omega}^f + \hat{\phi}(\mathbf{z}_i)\bar{y}_i + pc_i^P, \\ \hat{c}_i^m &= \omega^m(h_*^m, h_*^f) - \bar{\omega}^f + (1 - \hat{\phi}(\mathbf{z}_i))\bar{y}_i + pc_i^P.\end{aligned}$$

Single individuals consume their entire labour and non-labour income.

Given our estimates of individual consumption, we construct an estimated time series for the variance of log consumption from 1968 to 2001. Indexing individuals by i and households by k , the total variance across individuals can be decomposed into the between household and the within-household variance, respectively,

$$\text{var}(\log \hat{c}_{it}) = \text{var}(\mathbb{E}(\log \hat{c}_{it} \mid i \in k)) + \mathbb{E}(\text{var}(\log \hat{c}_{it} \mid i \in k)).$$

It is useful to recall that the use of adult equivalence scales to obtain a measure of individual-level consumption implies that this last term is zero.

20. As discussed in footnote 15, we model the dependence of v on the residuals using a Choleski factorization $\begin{pmatrix} \hat{e} \\ \sigma \hat{u} \\ v \end{pmatrix} = \begin{pmatrix} 1 & 0 & 0 \\ s_{21} & s_{22} & 0 \\ s_{31} & s_{32} & s_{33} \end{pmatrix} \begin{pmatrix} \hat{e} \\ \hat{u} \\ \eta \end{pmatrix}$, where the variance of the residuals is normalized to one and $\eta \sim N(0, 1)$. We replace $v = (s_{31}\hat{e} + s_{32}\hat{u} + s_{33}\eta)$ in equation (7) and integrate over η . The standard deviation of v is estimated as $\sigma_v = \sqrt{s_{31}^2 + s_{32}^2 + s_{33}^2}$ and $\text{corr}(v, \hat{e}) = (s_{31}/s_{33})$. Since our test does not imply a structural interpretation for the coefficient on the public expenditures residual, we report s_{32} rather than $\text{corr}(v, \hat{u})$ in the last row of Table 2.

7.1. *Between and within-household consumption inequality*

The time-series pattern of total, between-, and within-household inequality, measured by the variance of log consumption, for the years 1968 to 2001 is presented in Figure 7. Several findings are of interest. First, looking at the series for the between-household variance established using adult equivalence scales, we see it aligns with the well-documented pattern of stability prior to the 1980's, a rise in variance during the 1980's, and a levelling off again in the 1990's (Figure 6). This is not surprising as inequality measures based on equivalence scales measure between-household inequality.

Second, turning to the estimated series for within-household inequality, we see that at the beginning of the sample, the contribution of within-household inequality was equal to that of between-household inequality. This pattern continued until the early 1980's, at which point the contributions diverged: between-household inequality began to rise, while within household inequality fell. By the end of our sample period, the within-household variance is estimated to be responsible for only 25% of the total variance, one-half of its contribution in 1968. Finally, while there appears to be some variability in the series for total inequality, there is no clear trend. The offsetting movements in between and within inequality result in total inequality being only slightly higher in 2001 than it was in 1968.

The stylized facts presented in Section 2 point to two main reasons for the decline in within-household inequality: the fall in the gender wage gap and the rise in female labour supply. As women's wages rose and as married women increased their labour supply, the share of income contributed to the household by the wife increased. The share of consumption allocated to wives increased accordingly. The distinction between the effect of the convergence in wages and convergence in hours can be seen by comparing Figure 8 (a) and (b). In Figure 8, we plot the distribution of the estimated sharing rule, $\hat{\phi}(z_{it})$, for the years 1970 and 2000. The means of the distributions are 0.38 and 0.43, respectively. This increase in the mean of the sharing rule in large part reflects the convergence in potential earnings of husbands and wives, presented in

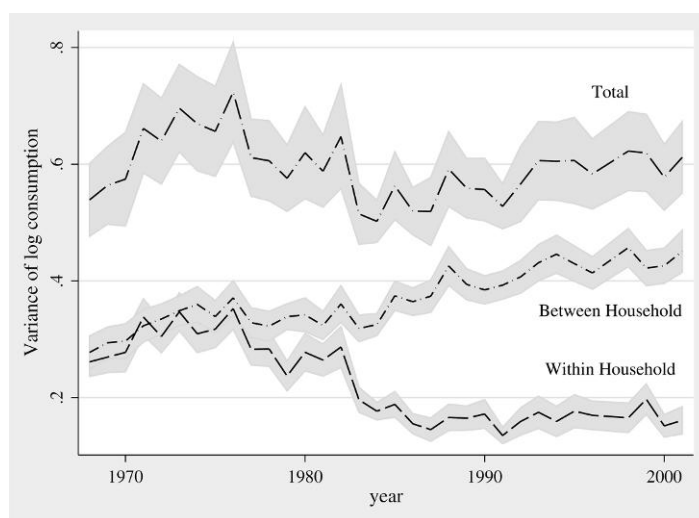
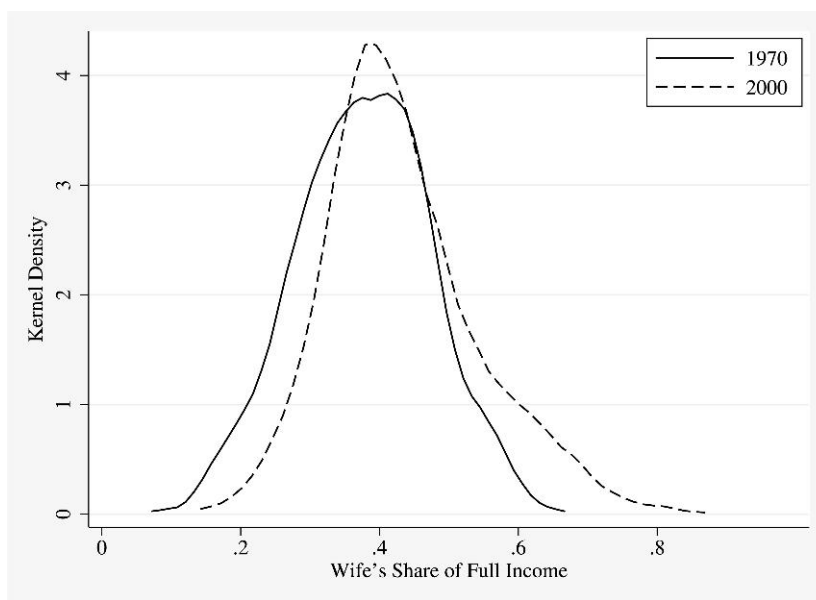


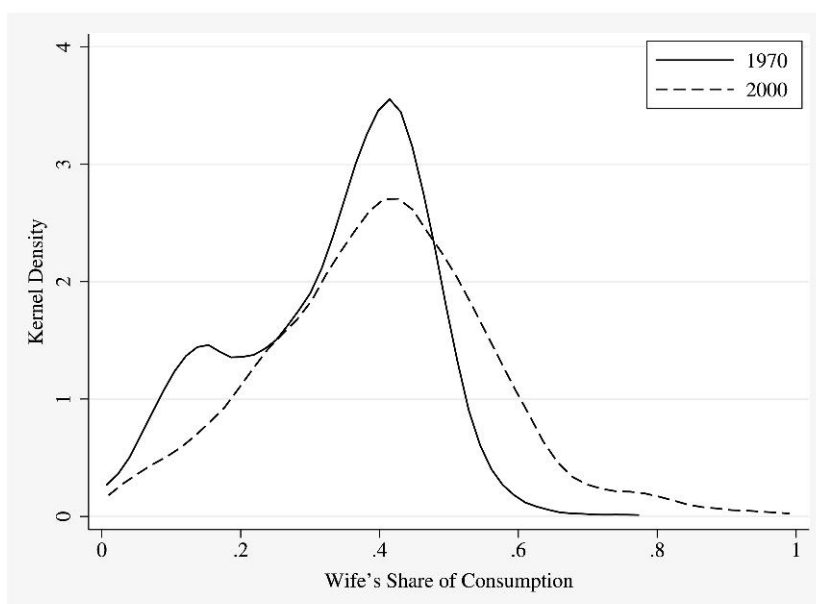
FIGURE 7

Total, between-, and within-household decomposition of trends in the variance of log consumption

Note: Own calculations from the FES. Shaded area represents \pm two standard errors



(a) The mean share is 0.38 in 1970 and 0.43 in 2000.



(b) The mean share is 0.33 in 1970 and 0.40 in 2000.

FIGURE 8

Distribution of wife's share of full income and consumption

Note: Based on sharing rule estimates for the pooled sample. (A) The mean of wife's share of full income is 0.38 in 1970 and 0.43 in 2000. (B) The mean of wife's share of consumption is 0.33 in 1970 and 0.40 in 2000.

Figure 3. Consider next the distribution of consumption inferred from our model for the years 1970 and 2000 in Figure 8. The share of consumption reflects both the changes in the sharing

rule of full income and changes in actual earnings of husbands and wives. The means of the consumption share for the wife are 0.33 in 1970 and 0.40 in 2000. The impact of actual earnings, as opposed to just potential earnings, is evident in the bimodal distribution in 1970. As we saw in Figure 4, only 60% of the married women in our sample were participating in paid work, compared with 90% of married men. This shows up in the estimated consumption distribution as a mass of non-working women with a low consumption share. By 2000, the participation rate of both married men and women are similar, as are the consumption shares.

7.2. Accounting for the trends in consumption inequality

In this section, we examine several explanations for the change in between and within-household consumption inequality observed over the period from 1968 to 2001. To reduce noise, and give a clean picture of the change from the beginning to the end of the period, we pool three years of data centred around 1970 and 2000. Results are presented in Table 4. The first column shows that the between-household component of the variance of log consumption increased by 0.139 between 1970 and 2000, with an offsetting decrease in the within-household component of -0.14 .

In order to decompose the change in inequality by changes in the distribution of the observables, we reweight the year 2000's data to have the same distribution of observables as the 1970 data. The weights used in the exercise are obtained as follows. Define a binary variable that takes a value of 1 if the year is 2000 and 0 if the year is 1970. We estimate the probability that the individual is observed in 2000 conditional on the observables of interest, \mathbf{x}_i , and call this $\hat{\pi}(\mathbf{x}_i)$. We then define a weight for each observation in 2000 as

$$w(\mathbf{x}_i) = \frac{1 - \hat{\pi}(\mathbf{x}_i)}{\hat{\pi}(\mathbf{x}_i)}.$$

Applying these weights to the 2000 data create counterfactual 2000 data, where the distribution of observables \mathbf{x}_i is the same as in the year 1970. We estimate the probability $\hat{\pi}(\mathbf{x}_i)$ using

TABLE 4
Decomposing changes in consumption inequality 1970–2000

Effect of Change in:									
	Total change	Household composition	Age dist.	Public share	Individual hours	Household hours	Individual wages	Household wages	unexplained
Left to right									
Between household	0.139	0.018 (13.1)	0.019 (13.6)	−0.016 (−11.6)	0.044 (32.1)	0.024 (17.0)	−0.020 (−14.7)	0.004 (3.0)	0.066 (47.5)
Within household	−0.140	−0.043 (30.8)	0.044 (−31.2)	−0.020 (14.4)	0.001 (−1.0)	−0.008 (6.4)	−0.009 (6.8)	−0.085 (60.9)	−0.018 (12.9)
Right to left									
Between Household	0.139	0.000 (0.0)	−0.017 (−12.1)	0.001 (0.5)	0.014 (10.3)	0.054 (38.8)	0.001 (0.6)	0.020 (14.5)	0.066 (47.5)
Within household	−0.140	0.000 (0.0)	0.003 (−2.0)	−0.54 (38.5)	0.029 (−20.4)	0.017 (−12.4)	−0.007 (4.9)	−0.110 (78.5)	−0.018 (12.9)

Note: Percentage of total change in parenthesis. The decomposition is sequential and therefore depends on the order in which the variables are introduced. The top panel presents the results when the variables are sequentially introduced starting with household composition and ending with household wages, while the second panel reverses this order.

a Probit. The set of observed characteristics we use to reweight the data include household composition (single or married), age, the share of household expenditures devoted to public goods, hours worked by individuals, joint hours worked by members of the household, wages of individuals, and the joint wages of household members.

To assess the contribution of each observable separately, we introduce them sequentially. This has the advantage of ensuring the total contributions (including the unexplained component) sum to one. However, this procedure is not invariant to the order in which we introduce the variables. To assess the importance of the order, we also present the results with the order reversed. Reweighting based on the change in observed characteristics accounts for 52.5% of the change in between-household inequality and 87.1% of the change in within-household inequality.

7.2.1. Household composition and demographics. Consider the large change in household composition that occurred alongside the rise in inequality. In particular, with delays in marriage and a rise in divorce rates, the fraction of households with one-adult increased relative to the fraction of two-adult households. Although single-person households have no within-household inequality by definition, it is still the case that there may exist substantial inequality between single-adult households. We present the results of reweighting on household composition in Table 4, Column 2. The change in the share of single versus married households accounts for between 0% and 13.1% of the increase in the between-household variance and between 0 and 30.8% of the fall in within-household variance. Thus, at most 30% of the measured fall in within-household inequality can be accounted for by the increased share of single-person households.

We also consider the hypothesis that the rise in consumption inequality is capturing cohort effects due to the changes in the age structure of the population. To this end, we add age to the set of observables and reweight the 2000 data to have the age structure of 1970. The results of this exercise, presented in the third column of Table 4, suggest that changes in the age distribution explain -12.1% to 13.6% of the increase between households. The effect on within-household inequality moves completely in the wrong direction; we see that the effect is between -31.2% and -2.0%. Taking the effect of household composition and age together, the net effect of demographics explains -12.1% to 26.7% of the between-household increase and -2.0% to 0.4% of the change within households. Thus, controlling for demographics (age and marital status) alone does not help us to understand the fall in within-household dispersion.

7.2.2. Public goods. The share of public goods expenditure in the household has a direct effect on our measurement of within-household inequality. As the share of public goods in the household increases, the inequality created by differences in private consumption between spouses gets down weighted. Indeed, if all goods were public, there would be no within-household dispersion. We present the results of reweighting the data so that the distribution of the share of public goods in 2000 matches the 1970 distribution in Column 4. Changes in the share of public goods accounts for between 14.4% and 38.5% of the fall in the within-household variance. At the same time, this does not explain the change in between-household variance, accounting for between -11.6% and 0.5% of the increase.

7.2.3. Labour supply and wages. According to the stylized facts presented in Section 2, two of the most salient trends over time are the closing of the gender gap in wages and the convergence of participation rates of husbands and wives. To what extent do changes in the distribution of wages and hours account for the rise in consumption inequality?

We start by examining changes in hours worked by individuals, presented in Column 5. This column accounts for changes in labour supply across men and women but not changes in the joint distribution of husbands' and wives' hours within households. The change in the distribution of hours worked explains between 10.3% and 32.1% of the increase in between-household dispersion but predicts a counterfactual change in within-household inequality. Next, we consider the effect of increased marital sorting or the fact that in 2000, husbands and wives work more similar hours to each other than in 1970's. Reweighting the data based on the *joint* hours decision of couples, we see in Column 6 that this explains 17.0% to 38.8% of the increase in the between-household variance. As with individual hours, joint hours do little to explain the decline in the within-household variance, accounting for between -12.4% and 6.4% of the decline.

Finally, we turn to the effect of wages. Wages enter the determination of individual consumption in our model both directly through the effect on earnings and indirectly through the effect on the sharing rule. Our first experiment reweights wages in 2000 to have the same sex composition as in 1970's. This experiment captures the fall in the gender wage gap plus the increase in wage dispersion. We take out the effect of growth, so the wage distributions have the same mean in both years. The decomposition then accounts for the increase in the variance of wages and the change in the sex composition of the distribution; between 1970 and 2000, the probability that a woman is found in the top quartile of wages has increased substantially. This explains -14.7% to 0.6% of the increase in between-household variance and 4.9% to 6.8% of the fall within households, as illustrated in Column 7. In the second experiment, we focus on changes in the joint wage distributions of husbands and wives. We reweight using the joint spousal distribution of wages to capture the effect of sorting. The results in Column 8 indicate that the change in the joint distribution of spousal wages explains 60.9% to 78.5% of the decline in within-household variance and 3.0% to 14.5% of the rise between households.

What is interesting about the results on hours and wages is that, taken together, sorting can simultaneously explain both the rise in consumption inequality across households and the fall in consumption inequality within households: sorting on wages and hours alone can explain 20% to 53.3% of the rise in between-household dispersion and 66.1% to 67.3% of the decline within households. It is also interesting to note that it is wage sorting within households in particular that is important, not simply changes in the relative wages of women and men in general, as this explains little of the changes in inequality.

Over time, the distribution of marriages shifted from specialized families, with one income or a mix of low and high earners within the household, to a distribution with a larger proportion of high earnings couples and low earnings couples. The exercises conducted above illuminate the dramatic role of sorting in determining the distribution of consumption across individuals. The evidence presented here is consistent with the existing literature on marital sorting and inequality that finds a positive relationship between marital sorting and inequality (Kremer, 1997; Dahan and Gaviria, 2001; Fernández and Rogerson, 2001; Fernández, Guner and Knowles, 2005).

7.3. Full Consumption

While it is standard in the literature to focus on inequality in consumption, more precisely expenditures, this is not necessarily the object we might be interested in. Indeed, in our model, individuals make choices to allocate their full income between consumption of private goods and leisure. Since individuals with different tastes for leisure will necessarily make different consumption-leisure choices when faced with the same full income, it is not clear that we should consider only differences in consumption when calculating inequality. Following Becker (1965), we can define full consumption as the market value of the individuals' consumption of private

goods, private leisure, and public goods, which is simply equal to each individual's share of full income plus public goods expenditure

$$\begin{aligned}\hat{c}_{\text{full}}^f &= \hat{c}^f + [\bar{\omega}^f - \omega^f(h_*^f, h_*^m)] + pc^P = \hat{\phi} \bar{y} + pc^P, \\ \hat{c}_{\text{full}}^m &= \hat{c}^m + [\bar{\omega}^m - \omega^m(h_*^m, h_*^f)] + pc^P = (1 - \hat{\phi}) \bar{y} + pc^P.\end{aligned}$$

Inequality within the household in full consumption is then directly related to resources allocated to each spouse via the sharing rule and avoids including differences in the leisure–consumption choices of spouses in the measure of dispersion.

In Figure 9, we plot the trends for total, between-, and within-household inequality in full consumption. Several points are of interest. First, total inequality is much lower for full consumption than for goods consumption. This is a direct result of counting the market value of leisure. Also, within-household inequality accounts for a much smaller fraction of the total.

In Figure 10, we plot the share of full consumption allocated to the wife for the years 1970 and 2000. The differences in the estimated share of full consumption are much smaller than the estimated shares of goods consumption presented in Figure 8. Indeed, while the estimates for share of goods consumption range between 0 and 1, the range for full consumption is narrower, between 0.2 and 0.8.

8. CONCLUSION

The literature on consumption inequality has focused on the question of how changes in income inequality translate into changes in consumption inequality at the level of the individual. In this paper, we show that current measures of consumption inequality only reflect inequality at the household level, as it is assumed that there is no inequality within households. We provide evidence that this assumption produces a very inaccurate picture of both the level and the trend in consumption inequality for two reasons. First, the large dispersion in incomes within the household are highly inconsistent with equal division of consumption. When equal division is relaxed, we estimate that there is substantial inequality within the household. This suggests

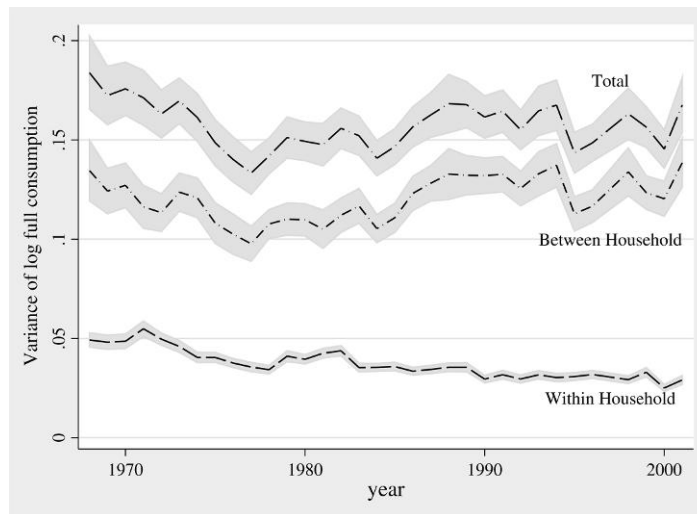


FIGURE 9

Total, between, and within-household decomposition of trends in the variance of log *full* consumption

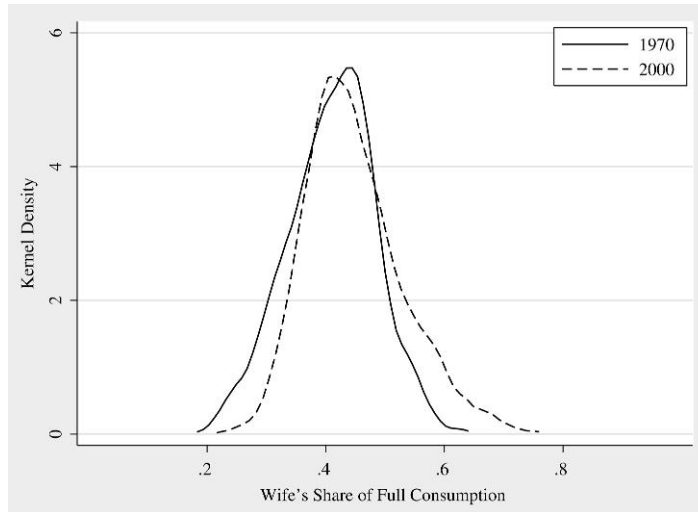


FIGURE 10
Distribution of wife's share of *full* consumption

previous work underestimates the level of individual consumption inequality by between 25% and 50%. Second, the earnings gap within households has closed over the past 30 years, resulting in less inequality within the household over time. As a result, we estimate that the rise in inequality between households documented in the literature may be fully offset by declines within the household. Our results show that the conventional approach for measuring inequality is only appropriate for measuring inequality for single individuals and for couples with the same earnings.

Our work highlights the fact that looking inside the household is as important as looking between households for the study of consumption inequality across individuals. What happens in the household not only changes our estimates of the level and trend in inequality but also changes our understanding of the forces behind the trends. We find that increased marital sorting on earnings can explain most of both the fall in within-household inequality and the rise in between-household inequality over time. These results are complementary to those of Fernández and Rogerson (2001), Fernández, Guner and Knowles (2005), and Dahan and Gaviria (2001), among others, on sorting and income inequality and suggest an important avenue for further research.

APPENDIX A. IDENTIFICATION OF THE SHARING RULE

Identification is straightforward in the presence of distribution factors and follows directly from Chiappori, Fortin and Lacroix (2002). Define the observable (reduced form) conditional choice probability $\hat{P}_j^f(\omega_j^f, \bar{\omega}^f, \bar{\omega}^m, y, \mathbf{z}, \mathbf{x}^f) \equiv \Pr(h^f = h_j | \omega_j^f, \bar{\omega}^f, \bar{\omega}^m, y, \mathbf{z}, \mathbf{x}^f)$ as the probability of observing the wife making hours choice h_j conditional on earnings from that choice, the potential earnings of both spouses, household non-labour income, distribution factors, and preference shifters. The structural (collective) conditional choice probability depends on the same variables but requires potential earnings, non-labour income, and distribution factors to enter only via the sharing rule: $\hat{P}_j^f(\omega_j^f, \phi^f(\bar{\omega}^f, \bar{\omega}^m, y, \mathbf{z}), \mathbf{x}^f)$. We have similar expressions for the husband's observed conditional choice probability, $\hat{P}_j^m(\omega_j^m, \bar{\omega}^f, \bar{\omega}^m, y^{nl}, \mathbf{z}, \mathbf{x}^m)$, and structural conditional choice probability, $\hat{P}_j^m(\omega_j^m, \phi^m(\bar{\omega}^f, \bar{\omega}^m, y, \mathbf{z}), \mathbf{x}^m)$. Recall that full household income net of public goods expenditure is $\bar{y} = \bar{\omega}^f + \bar{\omega}^m + \bar{y}^{nl}$, where $\bar{y}^{nl} = y^{nl} - pc^P$. We can divide the household full income between the wife and the husband using the sharing rule ϕ . Now, the full income of the wife is $\bar{y}^f = \bar{\omega}^f + \phi$ and the full income

for the husband is $\bar{y}^m = \bar{\omega}^m + \bar{y}^{nl} - \phi$, with $\bar{y}^f + \bar{y}^m = \bar{y}$. Assuming that the choice probabilities and the sharing rule are sufficiently smooth, we can relate the observable changes to structural changes as

$$\begin{aligned}\frac{\partial \hat{P}_j^f}{\partial \bar{\omega}^m} &= \frac{\partial \tilde{P}_j^f}{\partial \phi} \phi_{\bar{\omega}^m}, & \frac{\partial \hat{P}_j^f}{\partial \bar{y}^{nl}} &= \frac{\partial \tilde{P}_j^f}{\partial \phi} \phi_{\bar{y}^{nl}}, & \frac{\partial \hat{P}_j^f}{\partial z_k} &= \frac{\partial \tilde{P}_j^f}{\partial \phi} \phi_{z_k}, \\ \frac{\partial \hat{P}_j^m}{\partial \bar{\omega}^f} &= -\frac{\partial \tilde{P}_j^m}{\partial \phi} \phi_{\bar{\omega}^f}, & \frac{\partial \hat{P}_j^m}{\partial \bar{y}^{nl}} &= \frac{\partial \tilde{P}_j^m}{\partial \phi} (1 - \phi_{\bar{y}^{nl}}), & \frac{\partial \hat{P}_j^m}{\partial z_k} &= -\frac{\partial \tilde{P}_j^m}{\partial \phi} \phi_{z_k},\end{aligned}$$

where the notation X_y refers to $\partial X / \partial y$. Now, define

$$\begin{aligned}A &= \frac{\partial \hat{P}_j^f / \partial \bar{\omega}^m}{\partial \hat{P}_j^f / \partial \bar{y}^{nl}} = \frac{\phi_{\bar{\omega}^m}}{\phi_{\bar{y}^{nl}}}, & B &= \frac{\partial \hat{P}_j^m / \partial \bar{\omega}^f}{\partial \hat{P}_j^m / \partial \bar{y}^{nl}} = -\frac{\phi_{\bar{\omega}^f}}{1 - \phi_{\bar{y}^{nl}}}, \\ C &= \frac{\partial \hat{P}_j^f / \partial z_k}{\partial \hat{P}_j^f / \partial \bar{y}^{nl}} = \frac{\phi_{z_k}}{\phi_{\bar{y}^{nl}}}, & D &= \frac{\partial \hat{P}_j^m / \partial z_k}{\partial \hat{P}_j^m / \partial \bar{y}^{nl}} = -\frac{\phi_{z_k}}{1 - \phi_{\bar{y}^{nl}}},\end{aligned}$$

and note that the left-hand side is observable. Solving this system of equations gives us all the derivatives of the sharing rule in terms of observables:

$$\phi_{\bar{\omega}^f} = \frac{BC}{D-C}, \quad \phi_{\bar{\omega}^m} = \frac{AD}{D-C}, \quad \phi_{\bar{y}^{nl}} = \frac{D}{D-C}, \quad \phi_{z_k} = \frac{CD}{D-C}.$$

Identification requires that $A, B, C, D \neq 0$, and $C \neq D$, for at least one distribution factor z_k , which simply states that the distribution factor must have an observable effect on the conditional choice probabilities. The sharing rule is identified up to an additive constant.

In the absence of distribution factors (or when the distribution factors are not valid in the sense of having no effect on the conditional choice probabilities), identification still obtains but relies on higher derivatives of the choice probabilities (Chiappori, 1988, 1992). Making use of the notation $X_{yz} = \partial^2 X / \partial y \partial z$, we have

$$\begin{aligned}A_{\bar{y}^{nl}} &= \frac{\phi_{\bar{\omega}^m \bar{y}^{nl}} \phi_{\bar{y}^{nl}} - \phi_{\bar{\omega}^m} \phi_{\bar{y}^{nl} \bar{y}^{nl}}}{(\phi_{\bar{y}^{nl}})^2}, & A_{\bar{\omega}^f} &= \frac{\phi_{\bar{\omega}^f \bar{\omega}^m} \phi_{\bar{y}^{nl}} - \phi_{\bar{\omega}^m} \phi_{\bar{\omega}^f \bar{y}^{nl}}}{(\phi_{\bar{y}^{nl}})^2}, \\ B_{\bar{y}^{nl}} &= -\frac{\phi_{\bar{\omega}^f \bar{y}^{nl}} (1 - \phi_{\bar{y}^{nl}}) + \phi_{\bar{\omega}^f} \phi_{\bar{y}^{nl} \bar{y}^{nl}}}{(1 - \phi_{\bar{y}^{nl}})^2}, & B_{\bar{\omega}^m} &= -\frac{\phi_{\bar{\omega}^f \bar{\omega}^m} (1 - \phi_{\bar{y}^{nl}}) + \phi_{\bar{\omega}^f} \phi_{\bar{\omega}^m \bar{y}^{nl}}}{(1 - \phi_{\bar{y}^{nl}})^2}.\end{aligned}$$

Finally, defining

$$\alpha = \left(1 - \frac{BA_{\bar{y}^{nl}} - A_{\bar{\omega}^f}}{AB_{\bar{y}^{nl}} - B_{\bar{\omega}^m}} \right)^{-1},$$

it follows directly that

$$\phi_{\bar{y}^{nl}} = \alpha, \quad \phi_{\bar{\omega}^f} = B(\alpha - 1), \quad \phi_{\bar{\omega}^m} = A\alpha.$$

Identification in the absence of distribution factors then further requires that $BA_{\bar{y}^{nl}} - A_{\bar{\omega}^f} \neq AB_{\bar{y}^{nl}} - B_{\bar{\omega}^m}$.

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