ECONtribute
MARKETS & PUBLIC POLICY

ECONtribute
Discussion Paper No. 409

Beyond Vows: Family Structure and Consumption Inequality

Zainab Iftikhar Theresa Linhard Hanna Schwank

May 2026 www.econtribute.de

UNIVERSITÄT BONN UNIVERSITÄT ZU KÖLN

Funding by the Deutsche Forschungsgemeinschaft (DFG, German Research Foundation) under Germany's Excellence Strategy – EXC 2126/2 – 390838866 – is gratefully acknowledged.

Beyond Vows: Family Structure and Consumption Inequality *

Zainab Iftikhar[†] Theresa Linhard [‡] Hanna Schwank [§]

Abstract

This paper studies how family structure shapes consumption inequality and poverty in the USA. Using PSID data and a collective household model, we estimate sharing rules for married and cohabitating couples and recover individual-level consumption. In the full sample, cohabitating couples appear more egalitarian on average, with women receiving a share of household resources 9% higher than married women. These differences reflect systematic differences in characteristics across union types and largely disappear when comparing otherwise similar couples. Half of the economy-wide consumption inequality is explained by inequality between and within married households. 7% comes from cohabitation, 23% from between singles while the rest is explained by inequality between these three groups. Quantitatively, distinguishing cohabitation increases the role of between-group inequality and changes the assessment of poverty.

Keywords: consumption inequality, marriage, cohabitation, sharing rule, bargaining

JEL codes: D12, D13, D31, J12, J22

*We thank Alexander Bick, Chiara Lacava, Georg Dürnecker, Jeremy Lise, Nicolo Russo, Marc Chan, Victoria Baranov for helpful comments. We also thank the participants at the 9th International PhD Meeting in Economics (Thessaloniki), the Melbourne PhD Brown Bag Seminar, and the Bonn Internal Seminar. Further thanks go to audiences at the Mountain Seminar and various workshops for their helpful feedback. The first and third authors would like to thank the Deutsche Forschungsgemeinschaft (DFG, German Research Foundation), under Germany's Excellence Strategy – EXC 2126/2 – 390838866 and CRC TR 224 for financial support. The first and second author would like to thank DFG for funding through Research Unit FOR 5399 (project number 462655750).

[†]University of Bonn, CEPR, email: iftikhar@uni-Bonn.de

[‡]University of Bonn, email: theresa.linhard@uni-bonn.de

[§]University of Bonn, email: hschwank@uni-bonn.de

1 Introduction

A large body of literature employs adult equivalence scales to examine consumption inequality (see, for example, Blundell and Preston, 1998, Attanasio et al., 2005, Aguiar and Bils, 2015). This approach abstracts from within-household consumption inequality by implicitly assuming that resources are shared equally among household members. Evidence from the literature using the collective household framework, however, documents the presence of intra-household consumption inequality (Bobonis, 2009; Cherchye et al., 2012; Dunbar et al., 2013; Lise and Seitz, 2011), suggesting that the allocation of resources within households may affect population-wide consumption inequality. Most of this literature focuses on singles and married couples, largely overlooking the distinction between marriage and cohabitation (non-marital consensual unions).

A related literature argues that cohabitation is institutionally and economically distinct from marriage (Adamopoulou et al., 2025; Blasutto, 2024; Calvo, 2025; Smock, 2000; Smock and Manning, 1997). In contrast to marriage, cohabitating unions typically lack legal commitments, such as access to marital property or spousal support obligations. On the one hand, cohabitation may resemble marriage, in sharing of public good, companionship and joint decision-making. On the other hand, in the absence of formal legal and institutional protection, cohabitating partners may retain greater financial independence, implying weaker intra-household insurance. These institutional differences shape the incentives and constraints faced by married and cohabitating households, and may therefore influence how family arrangements contribute to population-wide consumption inequality.

To examine this relationship, this paper analyzes consumption inequality among childless individuals in the United States, focusing on three primary family arrangements: marriage, singlehood, and cohabitation. For this purpose, we first estimate sharing rules for couples.[1] Using the estimated sharing rules, we construct individual consumption measures and decompose population-wide consumption inequality into within-group and between-group components across these arrangements. Additionally, a poverty analysis is conducted to identify the incidence of poverty among households by type of family arrangement.

[1] The sharing rule is an empirical concept and can be expressed in monetary terms, for example, as the individual income share in household income. A higher individual income share implies a higher bargaining power of individual household members over resource allocation (Browning et al. (2013)).

The empirical analysis builds on the collective household framework of Lise and Seitz (2011). We extend their model in two ways. First, time spent on housework is treated as an input in the production of household public goods, allowing us to distinguish between leisure and non-market work. Second, marital status is included as a distribution factor in the empirical estimation.[2] The model is estimated using Panel Study of Income Dynamics (PSID) data for the period 1999–2017 and focuses on childless households. Estimating sharing rules allows us to recover individual consumption for all individuals in the sample.

The estimated sharing rule is used to construct individual-level consumption for married and cohabitating couples, allowing for a direct comparison of resource allocation across union types. We document three main findings. First, we identify similarities and differences between cohabitating and married couples in terms of resource sharing. In the full sample, cohabitating couples appear more egalitarian on average, with women receiving a 9% higher share of household resources than their married counterparts. These differences reflect that married and cohabitating couples differ systematically in their characteristics. When comparing otherwise similar couples in a matching approach, the average differences largely disappear, indicating that cohabitating couples exhibit similar patterns of resource sharing as married couples once observable differences are accounted for. In contrast, a key difference remains: even among observably comparable couples, cohabitating households exhibit greater dispersion and thicker tails in the sharing-rule distribution, pointing to higher heterogeneity in intra-household allocation. Second, family structure is a quantitatively important dimension of consumption inequality. Inequality among married individuals accounts for about 50% of total inequality, singles account for 23%, and cohabitants for approximately 7%, while differences between these groups themselves account for about one-fifth of overall inequality. Third, accounting for cohabitation and intra-household allocation jointly alters the assessment of poverty across family arrangements. In particular, cohabitating individuals are substantially more exposed (19% higher than married individuals) to low individual consumption, and gender differences within couples become more pronounced, indicating that standard approaches that ignore both cohabitation and unequal sharing misrepresent the distribution of economic well-being.

These findings highlight the role of family arrangements in shaping measured con-

[2] Distribution factors are exogenous variables that affect intra-household resource allocation without affecting preferences or the household budget constraint (Chiappori, 1988).

sumption inequality. Married and cohabitating couples differ systematically in both characteristics and patterns of resource allocation, leading to distinct contributions to population-wide inequality. Distinguishing between marriage, cohabitation, and singlehood therefore provides a more informative picture of the distribution of consumption and poverty across individuals. Ignoring cohabitation, or imposing equal sharing within households, may bias inequality measurement and obscure systematic differences in individual consumption within couples, including gender gaps that vary across union types. Moreover, the greater heterogeneity in intra-household allocation among cohabitating couples implies that the management of household-level transfers can differ across union types. For example, lump-sum transfers to households or transfers targeted to women need not translate one-to-one into individual consumption, and the disposition of resource sharing suggests that such transfers may be more strongly redistributed within married couples than among cohabitants.

This paper contributes to four related literateurs. First, it contributes to the literature on cohabitation and its implications for household behavior and economic outcomes. Cohabitation has become increasingly common in the United States: Manning (2013) estimates that the fraction of women who ever cohabited rose from 33% to 60% between 1987 and 2010. Several studies suggest that some of the gains traditionally associated with marriage can also arise in co-residential partnerships, including the sharing of public goods, division of labor, and risk pooling (Avellar and Smock, 2005; Brien et al., 2006; Lundberg et al., 2016).[3] At the same time, a broad literature emphasizes that family structure has important consequences for household members (Becker, 1973; Bumpass and Lu, 2000; Gemici and Laufer, 2011; Graefe and Lichter, 1999; Lundberg et al., 2016; Smock, 2000). While this literature documents behavioral differences across union types, little is known about whether these differences translate into different patterns of intra-household resource allocation. Our paper contributes to this literature by estimating sharing rules separately for married and cohabitating couples.

Second, the paper contributes to the literature on the measurement of individual welfare and consumption inequality. Early economic research focused primarily on income as a measure of welfare, but a growing body of work argues that consumption provides a more informative indicator of long-run well-being because it reflects households' ability

[3] Gains from marriage include sharing of public goods, division of labor and risk pooling (Becker, 1973, 1974; Chiappori et al., 2002; Lam, 1988).

to smooth income fluctuations over time (Attanasio and Pistaferri, 2016; Blundell and Preston, 1998; Cutler and Katz, 1992). Earlier approaches to studying consumption inequality were typically based on the unitary household model, which assumes a single utility function, pooled resources, and equal sharing within the household (Becker, 1965, 1973, 1974). By recovering individual consumption through estimated sharing rules, this paper contributes to the measurement of consumption inequality by explicitly accounting for intra-household resource allocation.

Third, the paper contributes to the literature on intra-household allocation in collective household models. Non-cooperative models characterize household outcomes as Nash equilibria in which each partner maximizes their own utility given the choices of the other (Bergstrom et al., 1986; Del Boca and Flinn, 2012; Lundberg and Pollak, 1994). Collective models, introduced by Chiappori (1988, 1992), instead assume Pareto-efficient allocations and recover individual consumption through a sharing rule reflecting bargaining power within the household. A large empirical literature has used this framework to study intra-household allocation and consumption inequality (Bobonis, 2009; Browning et al., 2013; Dunbar et al., 2013; Lise and Seitz, 2011; Oreffice, 2011), while related work combines collective models with revealed preference methods to identify bounds on the sharing rule (Cherchye et al., 2018, 2015, 2011). Existing research shows that intra-household inequality can account for a substantial share of aggregate consumption inequality (Lise and Seitz, 2011) and that sharing rules may adjust in response to shocks (Lise and Yamada, 2019). Our paper extends this literature by allowing sharing rules to differ across union types.

Finally, the paper relates to research on how family structure shapes the distribution of resources in the population. Household formation, assortative mating, and marital sorting influence the distribution of income and consumption across households (Chiappori et al., 2017; Eika et al., 2019; Fernandez et al., 2005; Fernández and Rogerson, 2001; Greenwood et al., 2014). For example, Johnson and Shipp (1997) attribute part of the rise in U.S. consumption inequality during the 1980s to shifts in population composition across family types. Related links between family structure and consumption inequality have also been documented in developing-country contexts such as Senegal, China, Ethiopia, and Argentina (Belete et al., 2022; De Vreyer and Lambert, 2021; Echeverría et al., 2019; Xue et al., 2026). A key dimension of family structure that has received little attention in this literature is the distinction between married and cohabitating couples. By distinguishing these union types in the estimation of sharing rules and in the measurement of individual consumption, this paper highlights how family arrangements contribute to population-wide consumption inequality and poverty.

The remainder of the paper is organized as follows. Section 2 describes the data, institutional background, and presents stylized facts highlighting differences between marriage and cohabitation. Section 3 outlines the model used to estimate the sharing rule, and Section 4 presents the econometric specification. Section 5 discusses the main results and robustness checks. Section 6 analyzes inequality and poverty using the estimated sharing rules. Section 7 concludes.

2 Data and Institutional Background

2.1 Data

We use the Panel Study of Income Dynamics (PSID) for the analysis. PSID is a representative sample of U.S. population that began in 1968 with annual waves up to 1997, from 1997 until present the dataset is biennially released. It contains data on relationship histories, wages, non-labor income, labor market and housework hours of the partners and information on household consumption. PSID records information on the relationship of the household members with the household head which allows us to differentiate between married and cohabitating households. The household expenditure data is collected from 1999 onwards. Therefore, we use the waves 1999-2017 for our analysis.

We impose certain restrictions on the sample. First, we restrict the analysis to heterosexual couples, with at least one partner between 24 and 65 years old, and at least one employed partner. Furthermore, we only include married and cohabitating couples who do not have any children present in the family unit at the time of the interview. We do this to abstract from decisions regarding fertility and investments in children.

The sample has a total of 3,810 couples giving us 10,359 year-couple observations on household heads and their partners. We identify 87.86% as married individuals while the rest are cohabitating. We have in total 30,406 observations: 20,718 from coupled individuals and 9,688 singles (47.31% of singles are female). Table 1 presents the sample profile by gender and relationship status. On average, married individuals are older and earn a higher hourly wage than their cohabitating and single counterparts. Single individuals work the longest hours per week, regardless of gender. A clear gender gap in housework hours exists among cohabitating and married couples: women spend considerably more hours on housework per week in both relationships, whereas no such gap exists among singles. Married respondents make up the largest group for both men and women.

2.2 Institutional Background

The empirical analysis focuses on heterosexual, childless couples observed in the PSID between 1999 and 2017. The incentives and constraints faced by partners may depend on the institutional environment governing different union types. We therefore briefly summarize key institutional differences between marriage and cohabitation.

While cohabitation may generate some of the same gains as marriage—such as joint consumption of public goods, division of labor, and risk pooling (Avellar and Smock, 2005; Brien et al., 2006; Lundberg et al., 2016)—it remains institutionally and behaviorally distinct. cohabitating unions are typically characterized by lower levels of commitment and match quality, which may limit cooperation and resource sharing within the household (Bumpass and Lu, 2000; Graefe and Lichter, 1999; Smock, 2000). Accordingly, cohabitation should be viewed as a distinct form of union rather than an informal version of marriage (Gemici and Laufer, 2011).

Institutional differences further shape intra-household resource allocation. Marriage provides legal protections to the lower-earning partner, including access to marital property and potential alimony, and entails divorce-related costs that are largely absent in cohabitating unions (Adamopoulou et al., 2025; Blasutto and Kozlov, 2025; Calvo, 2025; Matouschek and Rasul, 2008; Wydick, 2007). These features affect partners' rights and obligations, and may influence incentives to pool resources, specialize in market and non-market activities, and cooperate in the production of household public goods. In contrast, cohabitating partners generally do not have legally guaranteed access to such protections, which may lead to more individualized resource management within the household.

Behavioral differences between married and cohabitating couples have been documented even in contexts where cohabitation is institutionalized similarly to marriage. Evidence from Sweden, Canada, and the Netherlands demonstrates that married and cohabitating households differ in income organization, labor supply, fertility decisions, and

Table 1: Descriptive statistics of the study sample from the PSID 1999–2017, by gender and relationship type

Variable	Male						Female					
	Single		Cohabiting		Married		Single		Cohabiting		Married	
	Mean	SD	Mean	SD	Mean	SD	Mean	SD	Mean	SD	Mean	SD
Age, years	37.98	11.10	36.26	11.74	49.02	13.07	41.90	12.84	34.31	11.66	47.06	12.72
Years education	13.65	2.27	13.37	2.32	13.89	2.46	14.24	2.17	13.83	2.24	14.00	2.33
Hourly wage (reported / estimated)	18.43	17.98	18.21	17.02	25.88	35.05	18.01	16.62	15.92	20.12	18.16	19.60
Work hours (week)	41.01	14.66	36.80	17.30	37.45	17.23	37.80	12.80	31.40	16.67	29.74	17.14
Housework hours (week)	7.34	6.61	6.84	6.59	6.96	6.92	7.68	6.27	10.72	8.50	13.89	9.79
Total expenditure, week	462.96	345.35	680.65	428.18	822.12	688.52	466.06	333.27	680.65	428.18	822.12	688.52
Public expenditure, week	187.63	175.70	263.73	186.77	330.13	474.12	204.46	224.49	263.73	186.77	330.13	474.12
Observations	5104		1258		9101		4584		1258		9101	

Notes: The sample is drawn from the PSID 1999–2017 and restricted to heterosexual couples with at least one partner aged 24–65 and at least one employed partner. All singles and couples are restricted to those without children present in the household. Hourly wage is the self-reported wage where available; if missing, it is imputed using a Heckman selection model.

relationship stability (Chen and van Ours, 2020; Goussé and Leturcq, 2022; Heimdal and Houseknecht, 2003). These findings suggest a potential role for selection into different family arrangements.

A related literature views cohabitation as a transitional stage preceding marriage (Axinn and Thornton, 1992; DeMaris and Leslie, 1984; Hall and Zhao, 1995; Rao and Trussell, 1989; Ridley et al., 1978; Rosenfeld and Roesler, 2019; Trost, 1975). Even if cohabitation serves as a transition to marriage for some couples, multiple family arrangements coexist within the population at any point in time. Differences in institutional settings and in the characteristics of couples across single, married, and cohabitating households may therefore give rise to systematic differences in intra-household resource allocation. The stylized facts presented below document these differences in our data and provide insights into how and why resource sharing and individual consumption vary across family arrangements.

2.3 Stylized facts

The differences between married and single households are clear and have been discussed in the literature in several contexts in detail. By definition, single households do not involve intra-household resource allocation, whereas married and cohabitating households may exhibit within-household inequality as partners bargain over the allocation of household resources (Chiappori, 1988, 1992). However, less is known about the differences between the two unions and the implications of these differences for consumption inequality. In the following section, we discuss several stylized facts that highlight critical differences between married and cohabitating couples. These differences generate mechanisms that can shape the consumption inequality between and within households.

1. ***Match quality is lower for cohabitation.*** Panel 1a of Figure 1 shows that married unions last substantially longer than cohabitating relationships. While some cohabitating unions transition into marriage, a large fraction dissolves. In our sample, about 40% of first-time cohabitation unions end within three years, whereas only 11.7% of first marriages dissolve within the same period. One interpretation of this pattern is that cohabitating relationships involve lower match quality on average. Consistent with this view, Blasutto and Kozlov (2025) find that cohabitating couples experience lower values of companionship or emotional gains

from partnership than married couples. Lower match quality may reduce the degree of cooperation between partners and prevent them from fully realizing the potential gains from union, such as the sharing of public goods, division of labor, and risk pooling (Avellar and Smock, 2005; Brien et al., 2006; Lundberg et al., 2016).

2. ***cohabitating couples are younger, less educated, and have lower labor earnings than married couples.*** Figure 1 (panels 1b, 1c, and 1d) shows that cohabitating couples are on average significantly younger, somewhat less educated, and have lower after-tax labor earnings than married couples. Age and education may influence household consumption patterns (Ando and Modigliani, 1963; Fair and Dominguez, 1991; Fernandez-Villaverde and Krueger, 2007; Hougaard Jensen et al., 2022; Modigliani and Brumberg, 1954). If older or more educated couples allocate a larger share of resources to public good, within-household consumption inequality may differ systematically across union types. Consistent with the patterns in our data, previous research documents that cohabitation is more prevalent among younger and less educated individuals, while marriage is more common among individuals with higher education and is correlated with the economic position of men (Avellar and Smock, 2005; Blasutto, 2024; Smock and Manning, 1997). These differences in observable characteristics suggest that selection into union type may affect both intra-household resource allocation and each group's contribution to overall consumption inequality.

3. ***Sorting by education differs by marital status.*** We observe systematic differences in educational gaps between partners across union types. Among cohabitating couples, women are, on average, more educated than their male partners, with a gap approaching 0.5 years, whereas the gap among married couples is negligible (Fig.1e). The average age gap of approximately two years is similar across union types. These patterns suggest that educational assortative matching varies by marital status, while age-based matching does not. Consistent with these observations, Gemici and Laufer (2011); Schoen and Weinick (1993) document differences in assortative matching between married and cohabitating couples. Prior research shows that educational disparities affect intra-household resource allocation (Belloc et al., 2022; Cherchye et al., 2012; Gobbi et al., 2018; Oreffice, 2011, 2014; Vermeulen,

2005; Vermeulen et al., 2006). Gemici and Laufer (2011) finds that couples in which women are more educated than men are more likely to cohabit, suggesting weaker incentives for specialization. The subsequent stylized fact provides further support for this interpretation.

4. ***cohabitating women work more in the market than their married counterparts.*** Figure 1 (panels 1f and 1g) shows that cohabitating women supply about 1.5 more weekly hours of market work and spend over 3 fewer hours in home production than married women. Accounting for differences in their male partners' labor supply, the gaps in market and home hours are larger among married couples than among cohabitants, suggesting weaker specialization in cohabitating households. As a result, cohabitating women have about 1.5 more hours of weekly leisure than married women. These patterns indicate that time allocation within households varies systematically by marital status, which may in turn affect intra-household consumption inequality.

5. ***Female potential and actual share of household labor earnings are higher among cohabitating couples.*** Figure 1 (panel 1h) shows that the average female share of both potential and actual household labor earnings is higher among cohabitating couples than among married couples.[4] These findings align with previously documented patterns: cohabitating women are generally more educated than their partners, which contributes to a higher share of potential earnings. By allocating more hours to market work and fewer to home production, cohabitating women also account for a larger proportion of household realized labor income. The female share of potential full income is an important distributional factor and may influence the allocation of intra-household consumption (Browning et al., 1994; Chiappori et al., 2002). In contrast, the difference in the female share of realized earnings reflects variation in household specialization by union type and is also associated with differences in leisure consumption.

6. ***cohabitating women own a larger share of the household's non-labor income.*** Figure 1, panel 1i shows that women in cohabitation also own a larger share of household's non-labor income. Browning et al. (1994) and Cherchye et al.

[4]Potential earnings refer to the labor income if both partners worked 65 hours per week.

(2012) argue that the share of non-labor income is an important determinant of the distribution of power within a household, meaning that cohabitating women may have more influence on intra-household resource allocation decisions.

7. ***cohabitating couples face higher effective tax rates than married couples with similar income.*** Panel 1j of Figure 1 shows that cohabitating couples pay higher average tax rates than married couples with similar labor earnings.[5] The comparison is based on a matched sub-sample of cohabitating and married couples with similar observable characteristics, including labor income. One explanation for this difference is that married couples in the United States can file jointly for tax returns and may benefit from lower effective tax rates on household income, while cohabitating couples typically file separately and therefore do not qualify for these benefits.[6] Joint taxation can alter the incentives and bargaining positions of primary and secondary earners within married couples and may therefore contribute to variation in intra-household resource allocation across union types.

8. ***cohabitating couples exhibit lower between-household consumption inequality.*** Consumption inequality between households is lower among cohabitating couples than among married couples, as measured by the variance of log household consumption (Figure 1, panel 1k). One possible explanation is that marriage attracts a broader range of couples, leading to greater heterogeneity in observable characteristics among married households. Alternatively, cohabitating couples may be more similar to each other along observable dimensions, resulting in lower between-household variation in consumption.

9. ***The prevalence of cohabitation has increased over time.***

   We conclude by noting that, among couples in the PSID, the proportion who are married has declined substantially since the 1980s (Fig. 1l). Consequently, an increasing share of couples are in union types that differ in resource allocation, labor supply, and income organization, as previously documented. Therefore, dis-

[5]We model taxes $T(Y)$ as a function of total income Y, following the specification in Borella et al. (2023): $T(Y) = Y - (1-\lambda)Y^{1-\tau}$, where λ denotes the average tax rate when $Y = 1$, and τ captures the degree of tax progressivity. Using the parameter estimates for λ and τ derived by Borella et al. (2023) we compute after-tax income $Y - T(Y)$.

[6]In some states that recognize Common Law Marriage, joint taxation is possible. However, this is rarely observed in the data.

tinguishing between marriage and cohabitation has become increasingly important for understanding patterns of consumption and inequality among couples.

3 Model

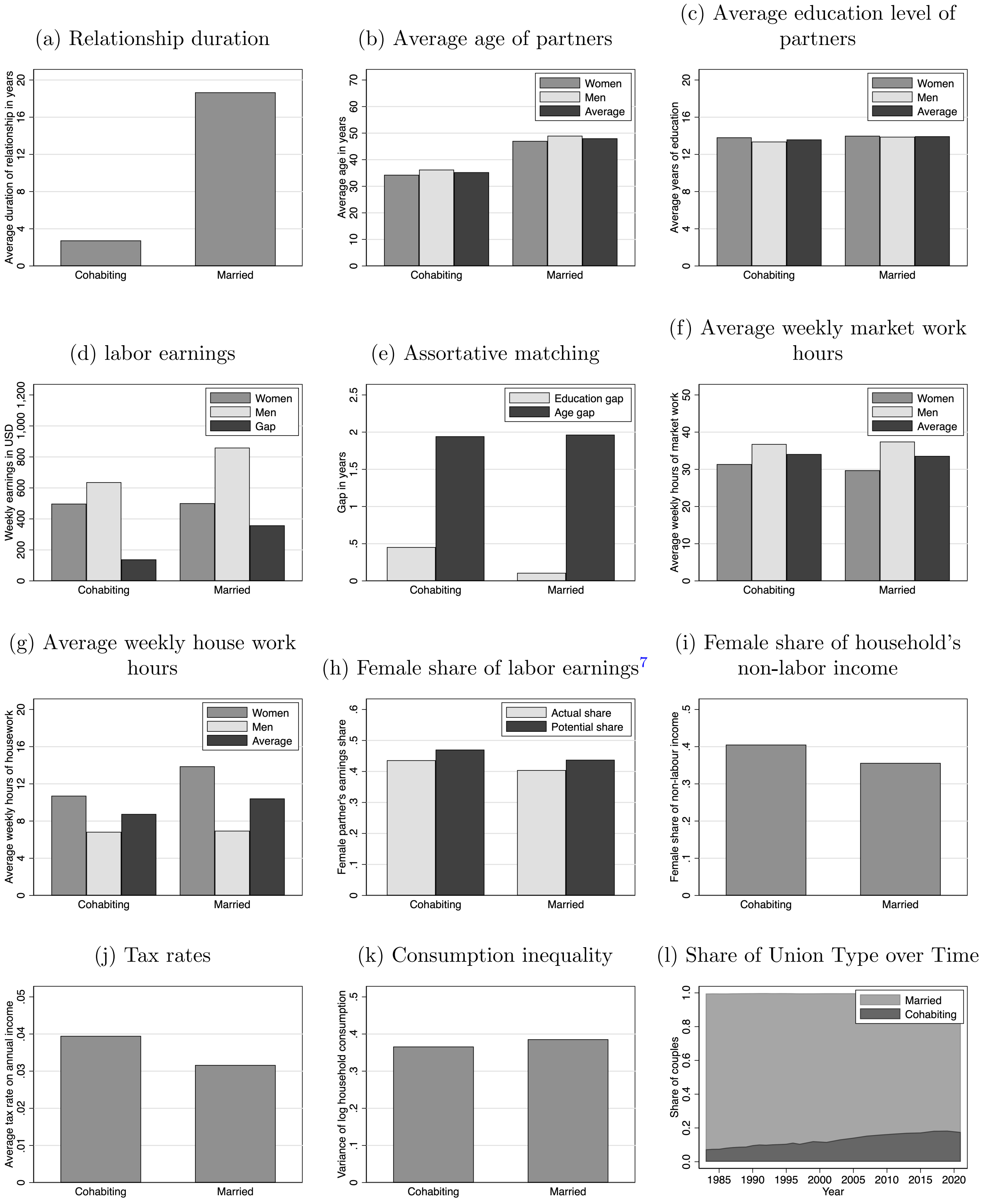
We build on the approach developed by Lise and Seitz (2011) to estimate individual-level consumption inequality. Our analysis introduces two main innovations. First, we distinguish between marriage and cohabitation, whereas Lise and Seitz (2011) focus exclusively on married couples. Second, we incorporate time spent on housework as an input in home production. In Lise and Seitz (2011), time devoted to home production is treated as leisure and monetary expenditures are assumed to be the only input in the production of household public goods. However, the data indicate that housework is an important component of household production and that its allocation differs systematically across partners and union types. As shown in Figure 1, panel 1g, women spend more time on housework than their male partners, and this difference also varies by marital status. Treating housework as leisure would therefore overstate women's leisure and bias the estimated sharing rule, leading to an underestimation of intra-household consumption inequality.

Consider an economy populated with heterogeneous households. Households consist of one member (singles) or two members (unions). The two-member households are married or cohabitating couples. The single households do not face a resource-sharing problem and entirely consume the total household resources. Therefore, these households only contribute to between household consumption inequality. The two-member household faces the problem of sharing time and monetary resources between partners. Therefore, they contribute to both within- and between-household consumption inequality. To identify the household sharing rule for resource allocation between partners, we describe a collective household model below, which provides details of the intra-household decision-making process and introduces the sharing rule representation that will be used in the following sections.

Each two-member household is defined by a set of four exogenous variables: marital

[7] Actual share refers to the female share of labor earnings at the current hours choice of both partners. Potential share refers to the female share of labor earnings if both partners worked full time (65 hours per week).

Figure 1: A comparison of married and cohabitating couples



Notes: The figures are based on authors' calculations using PSID. Sample period: 1999 – 2017. N = 20,718 (10,359 couples). 12.1% are cohabitating, the rest is married.

status, wage rate of male partner, wage rate of female partner, and non-labor income of the household $\{K, w^m, w^f, y^{nl}\}$ respectively, where $K \in \{M, C\}$ denotes the marital status, marriage, and cohabitation, respectively. Marital status is important for determining the tax regime a household faces and whether they have legal access to income from marital property. Married couples can file taxes jointly, whereas cohabitating couples are taxed separately on their incomes. However, joint taxation complicates the analysis therefore following Lise and Seitz (2011), we assume that couples make their decisions based on after-tax earnings. We use the tax functions estimated by Borella et al. (2023) to generate net labor earnings for each partner's discrete hours choices.

In the following discussion, we suppress the notation for the union type for ease of exposition. Each individual member (i) has distinct preferences over the household public good (Q), own private consumption (c^i), and own leisure (ℓ^i). We observe private consumption only at the household level, i.e. $c = c^f + c^m$. We assume that preferences over private consumption and leisure are separable from the consumption of the household public good. The household public good is produced by spending time (q) and monetary inputs (g) on housework using the following technology

$$Q(q^f, q^m, g) \equiv Q = (\kappa(q^f)^\gamma + (1-\kappa)(q^m)^\gamma)^{\frac{\delta}{\gamma}} g^{1-\delta} \tag{1}$$

where q^f, q^m denote the time of the female and male partner, respectively, spent on home production, κ_K is the union-specific productivity of the female in home production, γ captures the elasticity of substitution between male and female time inputs, and δ captures the degree of substitutability between time and financial inputs.[8] Assuming egoistic preferences and Pareto efficiency, the household's allocation can be represented

[8]We assume a union-specific κ_K because the female productivity may be affected by the degree of cooperation within a union. Since cohabitation is characterized by low match quality, we expect κ_c to be low in cohabitation. The share of female input in home production in PSID confirms our conjecture. The value of $\kappa_M = 0.62$ in marriage vs. $\kappa_C = 0.59$ in cohabitation.

as a solution to the problem:[9]

$$\max_{c^f,c^m,Q,\ell^f,\ell^m} \lambda(\pi,y,\mathbf{z})U^f(u^f(c^f,\ell^f),Q) + \big(1-\lambda(\pi,y,\mathbf{z})\big)U^m(u^m(c^m,\ell^m),Q) \tag{2}$$

$$\begin{aligned} \text{s.t.} \quad & c^f + c^m + g + \big[\bar{\omega}^f - \omega^f(h^f,h^m) - \omega^f(q^f,q^m)\big] \\ & \quad + \big[\bar{\omega}^m - \omega^m(h^m,h^f) - \omega^m(q^m,q^f)\big] = \bar{\omega}^f + \bar{\omega}^m + y^{nl} \qquad (3) \\ & c^f + c^m = c \qquad (4) \\ & Q = (\kappa(q^f)^\gamma + (1-\kappa)(q^m)^\gamma)^{\frac{\delta}{\gamma}} g^{1-\delta} \qquad (5) \\ & h^i + \ell^i + q^i = T \qquad (6) \end{aligned}$$

where the Pareto weight, $\lambda(\pi, y, \mathbf{z})$, represents the female partner's relative bargaining power within the household and depends on prices (π), total resources (y), and distribution factors ($\mathbf{z}$). The budget constraint (3) represents total household expenditure on the left-hand side and full household income on the right-hand side. Full household income y is the sum of potential after-tax earnings of men ($\bar{\omega}^m$) and women ($\bar{\omega^f}$) at wage rates w^m and w^f respectively, and the household non-labor income ($y^{n\ell}$). In the case of full-time work, both q^i and ℓ^i are zero. $\omega^f(h^f, h^m)$ denote women's after-tax earnings when working h^f hours which depends on choices of both partners. Let $\omega^f(q^f, q^m)$ denote the opportunity cost of women's time spent on housework for the given wage rate. Taking these together, $[\bar{\omega^f} - \omega^f(h^f, h^m) - \omega^f(q^f, q^m)]$ denotes women's expenditure on leisure.

The household decision problem is decentralized into two stages. In the first stage, expenditure and the hours allocated to home production (Q) are mutually agreed upon by the two partners. This is done by the following cost minimization:

$$\min \omega^f(q^f, q^m) + \omega^m(q^f, q^m) + g \quad \text{s.t.} \quad 1 \tag{7}$$

Additionally, the household also determines the conditional earnings functions for each household member and the sharing rule for dividing the remaining full income ($\bar{y}$) in the first stage. Where $\bar{y} = \bar{\omega}^f + \bar{\omega}^m - \omega^f(q^f, q^m) - \omega^m(q^f, q^m) + \bar{y}^{nl}$ with $\bar{y}^{nl} = y^{nl} - g$. The conditional earnings for individual i, $\omega(h^i, h^{j*})$, are the after-tax earnings of an individual working h^i hours, conditional on the equilibrium labor supply h^{j*} of their spouse.

[9] For discussion on implications of caring preferences and treatment of leisure as a public good, see Lise and Seitz (2011).

The sharing rule $\phi(\bar{\omega}^f, \bar{\omega}^m, y, \mathbf{z})$ is defined as the sum transferred to the female partner out of the household's non-labor income net of public goods expenditures. The sharing rule depends on the potential earnings of female and male partners ($\bar{\omega}^G$) and on distribution factors ($\mathbf{z}$).

In the second stage, each partner chooses their labor supply and private consumption to maximize their utility, taking the variables from the first stage as given. The female partner's problem in the second stage is

$$\begin{aligned} &\max_{c^f, \ell^f} u^f(c^f, \ell^f) \\ &\text{s.t.} \quad c^f + [\bar{\omega}^f - \omega^f(h^f, h^m_*) - \omega^f(q^f, q^m)] = \bar{\omega}^f - \omega^f(q^f, q^m) + \phi(\bar{\omega}^f, \bar{\omega}^m, y, \mathbf{z}) \\ &\qquad\qquad \ell^f + h^f + q^f = T, \end{aligned} \tag{8}$$

and the male partner's problem in the second stage is

$$\begin{aligned} &\max_{c^m, \ell^m} u^m(c^m, \ell^m) \\ &\text{s.t.} \quad c^m + [\bar{\omega}^m - \omega^m(h^m, h^f_*) - \omega^m(q^m, q^f)] = \bar{\omega}^m - \omega^m(q^m, q^f) + \bar{y}^{nl} - \phi(\bar{\omega}^f, \bar{\omega}^m, y, \mathbf{z}) \\ &\qquad\qquad \ell^m + h^m + q^m = T. \end{aligned} \tag{9}$$

We use the above model to estimate the sharing rule. An individual's choice of labor supply depends on their own wages and the sharing rule. It is influenced by their partner's wage, the household's full income and distribution factors, but only through their impact on the sharing rule. Therefore, observing how an individual's labor supply choice varies in response to a variation in their partner's wage, household full income or distribution factors provides insight into how the sharing rule changes with these variables. Hence, we can identify how the sharing rule changes with respect to changes in wages, non-labor income, and distribution factors. To estimate the location of the sharing rule, we impose the same identifying assumption following Lise and Seitz (2011) and Browning et al. (1994), that married partners share their resources equally when they have equal potential wages.[10] More specifically, symmetric sharing arises only under several conditions, including equal preferences, equal earning potential, identical labor supply, and symmetric access to non-labor income. While these conditions may be more plausi-

[10] see Lise and Seitz (2011) for details on identification.

ble for married couples due to institutional arrangements governing joint resources, they need not hold for cohabitating couples. In particular, even with identical preferences and earnings potential, asymmetric access to non-labor income can lead to unequal resource sharing in cohabitation. Since such asymmetries are not directly observed in the data, they constitute an important source of unobserved heterogeneity in the sharing rule. This implies that differences in the level of the sharing rule across union types may reflect both observable characteristics and unobserved asymmetries in access to non-labor income, while differences in dispersion are more directly informative about heterogeneity in intra-household allocation.

4 Empirical Strategy

In this section, we briefly discuss how we measure public expenditure and non-labor income in PSID, then we describe the empirical implementation of the model.

4.1 Public expenditure and non-labor income

We categorize all available expenditure data as either public or private expenditures. Expenditures on clothing, computers and electronics, education, food, health care, legal bills, transportation, and vacations and entertainment are classified as private expenditures. Expenditures on utilities, such as electricity, gas, heating, and water, as well as other housing expenditures, including rent, mortgage, and property tax, are categorized as monetary expenditures on the household public good (g). Additionally, we use the time spent by partners on housework and their hourly wages to monetize the value of the time spent on the household public good.

Non-labor income is defined as the difference between total household expenditures and net labor earnings. The expenditure-based definition of non-labor income reduces measurement error and accounts for wealth that is not observed in the data (Blundell et al., 2007, 1998; Blundell and Walker, 1986; MaCurdy, 1982). To account for inflation, all financial data are deflated to 2016 US dollars using the Consumer Price Index (World Bank, 2026).

4.2 Estimation

We assume a CES sub-utility function

$$u(c, \ell) = [\alpha c^{\rho} + (1 - \alpha)\ell^{\rho}]^{\frac{1}{\rho}} \tag{10}$$

where α determines the relative weight placed on private consumption c and leisure l, and ρ governs the elasticity of substitution between the two goods. Both α and ρ vary by sex but not by marital status, and in addition α depends on observables. In particular, we assume $\alpha(\mathbf{x_i}) = \frac{1}{(1+exp(\mathbf{x_i'}\alpha))}$, where $\mathbf{x_i}$ includes age, quadratic in age, education and birth cohort to allow for individual-level heterogeneity. Furthermore, we introduce additional unobserved individual heterogeneity in preferences η_{ij} for individual i and labor supply choice j with two components.

$$\eta_{ij} = \nu_i h_{ij} + \varepsilon_{ij} \tag{11}$$

The second component is i.i.d. type I extreme value preference heterogeneity. The first component allows for unobserved tastes for work and is correlated with non-labor income. Current non-labor income depends on previous labor income. To account for this endogeneity we instrument non-labor income using capital gains in the state housing market as proposed by Hurst and Lusardi (2004). Capital gains in the housing market are measured by the house price change indicator (HPI) for states in the US in our sample period. The data is collected by Federal Housing Finance Agency (2025). The measure is based on quarterly all-transactions HPI data (not seasonally adjusted), averaged over four quarters to obtain annual changes. These capital gains are interacted with home ownership. We estimate non-labor income as follows

$$y_i^{nl} = \mathbf{w}_i'\beta + e_i \tag{12}$$

To control for the endogeneity of non-labor income, we model unobserved heterogeneity, ν_i, as a function of the residual $\hat{e}_i = y_i^{nl} - \mathbf{w}_i'\hat{\beta}$.

Next, we transform ϕ to $\varphi \equiv \frac{\bar{\omega}^f - \omega^f(q^f, q^m) + \phi}{\bar{y}}$ and $1 - \varphi \equiv \frac{\bar{\omega}^m - \omega^m(q^m, q^f) + \bar{y}^{nl} - \phi}{\bar{y}}$, and define $\varphi \in [0, 1]$ to be the share of a household's remaining full income (net of expenditure on public good) transferred to the female partner, parameterized as

$$\varphi(\mathbf{z}) = \frac{\exp(\mathbf{z}'\varphi)}{1+\exp(\mathbf{z}'\varphi)} \tag{13}$$

Here, $\mathbf{z}$ is a vector of distribution factors, which contains the female partner's share of potential household earnings $(\bar{\omega}^f/(\bar{\omega}^f+\bar{\omega}^m))$, the log of full household income $(\log y)$ and the age gap between partners. In addition, we introduce a fourth distribution factor to capture the deviation in the sharing rule by union type. For this, the log of full household income $(\log y)$ is interacted with the indicator function K for union type with $K=0$ for marriage and $K=1$ for cohabitation. As discussed earlier, the legal access of partners to non-labor income varies by union type, influencing the bargaining positions of the spouses. Second, a higher match quality in marriage is associated with increased cooperation between partners in marriage, which affects income organization behavior, which in turn has implications for sharing rules (Eickmeyer et al., 2023; Heimdal and Houseknecht, 2003; Hiekel et al., 2014b). We capture these differences with our fourth distribution factor.

Using the functional form for the utility (10) and the sharing rule (13), and substituting in the budget and time constraints from (8) and (9), we get the value of labor supply choice $h^f_{ij} \in H$ for women i as

$$\begin{aligned} V^f_{ij}(\mathbf{Z}_i,\theta) = \Big(\alpha^f(\mathbf{x_i})\left[\omega^f(h^f_{ij},h^m_{i*}) - \bar{\omega}^f_i + \omega^f(q^f,q^m) + \varphi(\mathbf{z_i})\bar{y}_i\right]^{\rho_f} \\ + \left(1-\alpha^f(\mathbf{x_i})\right)\left[T - h^f_{ij} - q^f_{ij}\right]^{\rho_f}\Big)^{\frac{1}{\rho_f}} + \nu^f_i h^f_{ij} + \varepsilon^f_{ij}, \end{aligned} \tag{14}$$

and similarly for men

$$\begin{aligned} V^m_{ij}(\mathbf{Z}_i,\theta) = \Big(\alpha^m(\mathbf{x_i})\left[\omega^m(h^m_{ij},h^f_{i*}) - \bar{\omega}^m_i + \omega^m(q^m,q^f) + (1-\varphi(\mathbf{z_i}))\,\bar{y}_i\right]^{\rho_m} \\ + \left(1-\alpha^m(\mathbf{x_i})\right)\left[T - h^m_{ij} - q^m_{ij}\right]^{\rho_m}\Big)^{\frac{1}{\rho_m}} + \nu^m_i h^m_{ij} + \varepsilon^m_{ij} \end{aligned} \tag{15}$$

The vector $\mathbf{Z}_i$ captures all observable characteristics linked to both preference heterogeneity and the sharing rule. The parameter vector θ represents the preference and sharing-rule parameters that are subject to estimation. These value functions differ from those in Lise and Seitz (2011) in two key respects. First, time devoted to home production (the public good) enters explicitly into the value function. Second, marital status influences the value function through the distribution factors in the sharing rule.

The estimation proceeds in four steps using the dataset described in section 2. In the first step, we estimate a selection-corrected wage equation to predict potential wages for non-working individuals. In the second step, we calculate after-tax labor earnings for each discrete hours choice, using the tax parameters estimated by Borella et al. (2023).[11] In the third step, we estimate the reduced-form equation for non-labor income to derive the residuals $\hat{e}_i$. In the final step, we estimate the preference parameters $(\alpha^G, \rho^G, \sigma^G_\nu, \sigma^G_{\nu e})$ for men and women, and the sharing-rule parameters φ for cohabitating and married couples, using a mixed logit model for discrete labor supply choices. The distribution factors are transformed in such a way that our identifying assumption of equal sharing at equal potential earnings shares holds with $\mathbf{z} \equiv [\frac{\bar{\omega}^f}{\bar{\omega}^f + \bar{\omega}^m} - 0.5, \log y - \overline{\log y_t}, (age_f - age_m) - \overline{age_f - age_m}, (\log y - \overline{\log y_t}) \cdot K]$.

5 Results

5.1 Preference parameters

Table 2 shows the results for preference parameters computed at the mean of data and the bootstrapped standard errors. The value of ρ has a similar value across genders, implying a 1.12 elasticity of substitution between leisure and consumption for men and 1.16 for women. At the mean of data, men put a slightly lower weight on consumption compared to women. The variation in unobserved taste for work is higher for men than women as reflected in the variance of ν. The correlation between unobserved taste for work and non-labor income is negative for both men and women but stronger for men.

5.2 Sharing rule parameters

Table 3 reports the estimates of the sharing rule parameters for full sample. The value of φ implies that, at the mean of the data, women receive 43.8% of household full income net of public good expenditure. Consistent with existing evidence, the female share increases strongly with her potential earnings share, with a marginal effect of 1.046, and declines with household income, indicating that women capture a smaller fraction of additional resources as total household income rises. The key novel result emerges when allowing

[11] Following Lise and Seitz (2011), we discretize the hours choice by creating bins of five hours, from 0 to 65.

Table 2: Estimates of the preference parameters

	Base Estimates		**Separability Test**	
	Women	Men	Women	Men
ρ	0.1371 (0.0091)	0.1054 (0.0090)	0.1240 (0.0120)	0.0939 (0.0097)
$\alpha(\mathbf{x})$ at mean of data	0.0251 (0.0028)	0.0223 (0.0036)	0.0127 (0.0019)	0.0108 (0.0016)
α:				
Age	0.1362 (0.0539)	0.2104 (0.0587)	0.1702 (0.0769)	0.2946 (0.0856)
Age2	0.2327 (0.0678)	0.0711 (0.0387)	0.3085 (0.1076)	0.1563 (0.0687)
Education	0.0530 (0.0246)	0.0453 (0.0212)	0.1349 (0.0518)	0.1210 (0.0366)
1930 cohort	-0.0828 (0.0430)	0.0804 (0.0393)	-0.1058 (0.0570)	0.0878 (0.0531)
1940 cohort	-0.0537 (0.0308)	-0.1681 (0.0572)	-0.0406 (0.0261)	-0.1897 (0.0815)
1960 cohort	0.0926 (0.0425)	0.2198 (0.0922)	0.1107 (0.0526)	0.2532 (0.1086)
1970 cohort	0.3353 (0.0931)	0.1553 (0.0594)	0.3040 (0.0975)	0.1467 (0.0581)
1980 cohort	0.3913 (0.1038)	0.3516 (0.1004)	0.3835 (0.1233)	0.3415 (0.1331)
1990 cohort	0.5165 (0.1619)	0.5356 (0.1696)	0.5761 (0.1786)	0.5430 (0.1876)
α_0	3.6610 (0.1007)	3.7796 (0.1295)	4.3504 (0.1476)	4.5197 (0.1435)
σ_ν	0.8736 (0.0200)	0.9336 (0.0224)	0.6182 (0.0176)	0.6662 (0.0219)
corr(ν,e)	-0.0020 (0.0125)	-0.0085 (0.0126)	0.1456 (0.0358)	0.0584 (0.0191)
Public goods expenditure			0.0007 (0.0000)	0.0010 (0.0001)

Notes: This table presents the estimates of the preference parameters for women and men. The sample consists of married and cohabitating couples from the PSID ($N = 20,718$), sample period 1999 – 2017. The first two columns display the preference parameters for the base model, while the final two columns report results for the separability test, which includes public goods expenditure. Standard errors are bootstrapped with 1,000 replications and clustered at the couple level (reported in parentheses). The elasticity of substitution is parameterized as $\frac{1}{1-\rho}$ and the consumption share parameter as $\alpha_i = \frac{1}{1+\exp(\mathbf{x_i}'\alpha)}$.

this income gradient to differ by union type. The interaction between log income and an indicator for cohabitation is positive at 0.0223 and precisely estimated, implying that cohabitating women retain a significantly larger share of marginal increases in household income than married women. This result indicates that the income organization differs systematically across union types: while higher income is associated with more unequal sharing in marriage, this pattern is substantially attenuated in cohabitation.

To illustrate the implications of the estimated sharing rule, we replicate the counterfactual scenarios in Lise and Seitz (2011), allowing for a direct comparison of outcomes between married and cohabitating couples. Consider a couple with equal potential wages and zero expenditure on the public good. Each partner has potential labor incomes $\bar{\omega}$ such that the full log income at the mean of the data is $\bar{y} = 2\bar{\omega}$ and the partners have an average age-gap. This couple will share the resources equally irrespective of the union type because the effects are measured for deviations from the mean of data.

Now, let's redistribute earnings between partners by raising the female partner's wage by 10% and lowering the male partner's wage by 10%, so that the full household income remains the same. Therefore, the sharing rule adjusts exclusively in response to the redistribution of earnings between partners. The new sharing rule is identical for married and cohabitating couples and is given by $\varphi'_K = \varphi_K + 1.0455 \times 0.05 = 0.55$.

Next, consider a 10% increase in household income due to an increase in the female partner's share of potential wage by 20%. In the case of marriage, the wife's new share becomes $\varphi'_M = \varphi_M + 1.0455 \times 0.045 - 0.0705 \times log 1.1 = 0.544$. For cohabitating women the new sharing rule will be $\varphi'_C = \varphi_C + 1.0455 \times 0.045 - 0.0705 \times log 1.1 + 0.0223 \times log 1.1 = 0.545$. In both cases, the rise in female partner's share of labor earnings increase their share of household resources. However, women in both unions share a part of increase in their wage with their male partner, implied by the negative coefficient of full income. The positive coefficient of the interaction term captures, among other things, the lower degree of income sharing in cohabitation: cohabitating women end up sharing a slightly smaller share of their wage rise with their male partners compared to married women resulting in a somewhat larger sharing rule in cohabitation. Our results are in line with the literature that associates marriage with higher degree of resource sharing and increased cooperation due to better match quality and higher sense of commitment compared to cohabitation (Eickmeyer et al., 2023; Hamplová et al., 2014; Hiekel et al., 2014a; Kapelle et al., 2025;

Lyngstad et al., 2010), while still indicating that substantial resource sharing also takes place within cohabitating unions.

Finally, consider a 10% increase in full income due to a rise in non-labor income for both married and cohabitating households. The new sharing rule for marriage is as follows $\varphi'_M = \varphi_M - 0.0705 \times log1.1 = 0.497$ and for cohabitation $\varphi'_C = \varphi_C - 0.0705 \times log1.1 + 0.0223 \times log1.1 = 0.498$. In both unions, the share of household resources allocated to women reduces, but the level of full income she receives increase from $0.5 * \bar{y}$ to $0.497 * 1.1\bar{y} = 0.547\bar{y}$ in marriage and to $0.498 * 1.1\bar{y} = 0.548\bar{y}$ in cohabitation. cohabitating women observe a smaller decline in the sharing rule and a larger increase in full income they receive. The full income of women in cohabitation increases by 9.6% while in marriage it only increases by 9.4%. This suggests that both married and cohabitating couples share a substantial fraction of non-labor income, consistent with sharing rather than independent "roommate"-type behavior, while married couples exhibit a slightly higher degree of sharing.

Table 3: Sharing rule estimates

Estimates	Pooled	Separability Test	Birth Cohort 1930	1940	1950	1960	1970	1980	1990	Matched Sample
φ (mean of data)	0.4384 (0.0232)	0.4432 (0.0055)	0.4382 (0.0037)	0.4114 (0.0056)	0.4111 (0.0072)	0.4372 (0.0040)	0.4506 (0.0030)	0.4700 (0.0018)	0.4547 (0.0026)	0.4644 (0.0017)
Marginal effect of										
Female partner's potential earnings share	1.0455 (0.3371)	0.9662 (0.0858)	1.0325 (0.0345)	1.2336 (0.0445)	1.2043 (0.0565)	1.0384 (0.0366)	1.0947 (0.0368)	1.0531 (0.0332)	1.0849 (0.0344)	1.1133 (0.0538)
Log of full income	−0.0705 (0.0463)	−0.0931 (0.0059)	−0.0824 (0.0060)	−0.2581 (0.0168)	−0.2136 (0.0184)	−0.1483 (0.0157)	−0.0653 (0.0135)	−0.0995 (0.0175)	−0.0392 (0.0082)	−0.0698 (0.0118)
Log income × cohabitation	0.0223 (0.0088)	0.0418 (0.0051)	0.0237 (0.0020)	0.0204 (0.0023)	0.0222 (0.0080)	0.0154 (0.0056)	−0.0056 (0.0083)	0.0035 (0.0131)	0.0499 (0.0046)	0.0211 (0.0165)
Partner's age gap	−0.0315 (0.0228)	−0.0237 (0.0009)	−0.0280 (0.0055)	−0.0937 (0.0152)	0.0876 (0.0222)	0.3193 (0.0179)	−0.1822 (0.0128)	0.2054 (0.0204)	0.0800 (0.0064)	−0.0324 (0.0181)
Observations	20,718	20,718	467	3,717	6,492	3,373	3,252	3,002	415	4,952
	Test null of separability:					Test null of equality of estimates:				
p-value	0.6069					0.0000				

Notes: The sample consists of married and cohabitating couples from the PSID ($N = 20,718$), sample period 1999 – 2017. Bootstrapped standard errors clustered at couple level with 1,000 replications are in parentheses. The sharing rule is parameterized as $\varphi(\mathbf{z}_i) = \frac{\exp(\mathbf{z}_i'\varphi)}{1+\exp(\mathbf{z}_i'\varphi)}$, and the marginal effect is evaluated at the mean of $\mathbf{z} \equiv [\frac{\varpi^f}{\varpi^f+\varpi^m} - \frac{1}{2}, \log(y) - \overline{\log(y_t)}, (age_f - age_m) - \overline{(age_f - age_m)}, K \times (\log(y) - \overline{\log(y_t)})]'$. The variables entering the vector $\mathbf{z}$ are transformed such that the sharing rule is normalized to equal one-half when both partners have equal potential wages, evaluated at the mean of log full income and the mean age gap. The test for the null of separability is a test that the first two columns are equal. The test for the null of equality of estimates tests whether the estimated sharing rule coefficients are equal across all birth cohorts. The matched sample is constructed using propensity score matching which is based on female partner's potential earnings share, age difference between spouses, log full household income, age and age squared of both partners, years of education, and race.

5.3 The distribution of the sharing rule by union type

We use the results in Table 3 to compute the sharing rule separately for the cohabitating and married sub-samples. Figure 2a shows the distribution of the sharing rule by marital status. It shows that cohabitating couples are, on average, more egalitarian than married ones, as the average $\varphi_C > \varphi_M$, and therefore closer to 0.5. A larger mass of women in cohabitation shows up to the right of the distribution, receiving a higher share of full household income than married women. Figure 2b shows that the distribution of the sharing rule closely aligns with, though does not precisely mirror, the pattern observed for the female share of potential earnings. This share is among the key distribution factors that shape intra-household resource allocation. Similar to the average sharing rule, the average share of potential earnings is higher in cohabitation. Interestingly, the variance of female potential earnings share is larger for marriage. Particularly striking is the lower end of the distribution: The mass of women with very low shares of potential earnings is substantially higher in marriages than in cohabitating relationships. Nonetheless, despite greater variance in the female potential earnings share, the somewhat smaller variance in resource sharing among married households suggests that institutional settings that provide legal protection and social norms that may assign gender roles to married couples may limit the extent to which income differences translate into resource allocation disparities.

To further examine the household resource allocation by marital status, we report the share of household income women receive in each union type, based on their relative wage contributions. Table 4 presents the average sharing rule by tertiles of total potential household income and by groups defined according to the female share of potential earnings. The sample is divided into three income groups based on tertiles. Within each tertile, three sub-groups are defined by the female partner's contribution to household labor income: less than 40% (low), between 40% and 60% (mid), and at least 60% (high). In all specifications, the sharing rule increases substantially with the female partner's potential earning share, suggesting that relative earning capacity is a primary determinant of intra-household allocation. Differences between married and cohabitating couples are generally minor across most groups, but become more pronounced at higher income levels. Specifically, among couples where women have a higher share of potential earnings, cohabitating women receive substantially larger shares than their married counterparts, particularly in the upper income tertile. This finding indicates that the primary distinction between union types lies not in average allocation but in greater dispersion and a stronger association between bargaining power and realized shares among cohabitating couples.

To assess the implications of misclassifying housework as leisure, we re-estimate the sharing rule under an alternative specification that treats housework as part of leisure. The results, reported in Appendix A, Table 8, indicate a slightly higher sharing rule, implying that women receive a larger share of household resources than in the benchmark case. As the sample consists of childless couples, this overestimation is likely to be even more pronounced among couples with children, for whom housework plays a larger role. These findings underscore the importance of correctly classifying housework in such analyses.

5.4 Robustness

5.4.1 Testing separability and controlling for birth cohorts

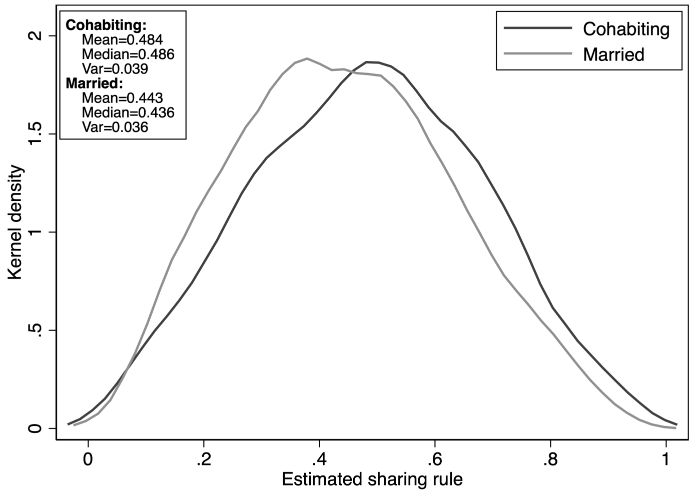
To assess the robustness of the model, we examine how sensitive the results are to the assumption that private consumption and leisure are separable from public good consumption. The null hypothesis is that public expenditure influences household behavior only through its impact on the budget constraint, $\overline{y} = \overline{\omega}^f + \overline{\omega}^m + y^{nl} - g - \omega^f(q^f, q^m) - \omega^m(q^m, q^f)$, without directly altering preferences. To test this, we investigate whether the level of public good spending, $g + \omega^f(q^f, q^m) + \omega^m(q^m, q^f)$, has any additional explanatory power for the estimated preference parameters beyond the budget effect.

In practice, we first regress public good spending on observable characteristics (age, age squared, and education of both partners, interacted with year fixed effects) and retain the residual, which captures the component of public good consumption not explained by these observables.[12] We then include both public good spending and its residual, interacted with hours of work, in the equation for unobserved tastes for work. Under separability, once we control for observable characteristics, variation in public good consumption should not systematically affect the allocation of private consumption and leisure. We therefore use this specification to assess whether public good consumption

[12] Total expenditures and leisure are instrumented by full earnings of both partners, full household income, and the interaction of home ownership with house price changes.

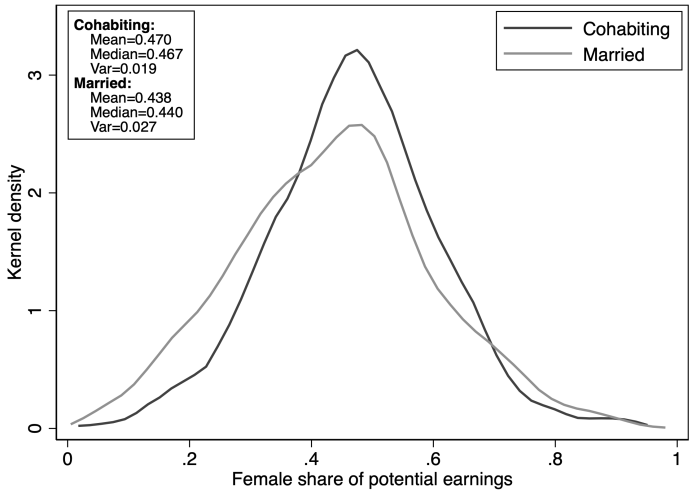
Figure 2: Distribution of sharing rule and female partner's potential wage by marital status.

(a) Estimated sharing rule



Notes: The sharing rule measures share of full household income, net of public goods expenditure, transferred to female partner.

(b) Female share of potential earnings



Notes: Potential earnings reflect the labor income that could be earned if the agents worked 65 hours per week.

Notes: The figures are based on authors' calculations using PSID. Sample period: 1999 – 2017. N = 20,718 (10,359 couples). 12.1% are cohabitating, the rest is married.

Table 4: Mean sharing rule by potential earnings tertiles and female share of potential household income groups (%)

Total potential earnings / female share	Total (%)	cohabitating (%)	Married (%)
Tertile 1 / Low	33.82 (11.1)	34.24 (12.2)	33.76 (11.0)
Tertile 1 / Mid	51.42 (17.1)	51.19 (22.8)	51.47 (16.3)
Tertile 1 / High	69.17 (5.1)	68.93 (6.8)	69.21 (4.9)
Tertile 2 / Low	32.33 (11.6)	33.05 (8.0)	32.26 (12.1)
Tertile 2 / Mid	50.16 (17.0)	51.77 (20.7)	49.87 (16.5)
Tertile 2 / High	63.54 (4.7)	66.64 (4.5)	63.13 (4.8)
Tertile 3 / Low	26.31 (16.6)	27.67 (9.0)	26.21 (17.7)
Tertile 3 / Mid	48.94 (11.1)	49.30 (11.0)	48.89 (11.1)
Tertile 3 / High	63.18 (5.6)	69.13 (4.8)	62.48 (5.7)
Total	44.74	48.24	44.25

Notes: The sample consists of married and cohabitating couples from the PSID ($N = 20,718$), sample period 1999 − 2017. Total potential earnings is calculated as the sum of potential full time earnings of both partners ($\overline{\omega}^f + \overline{\omega}^m$). Households are categorized into three income tertiles based on this total. Within each tertile, couples are classified by the female share of potential earnings: low (under 40%), mid (between 40% and 60%), and high (at least 60%). Main entries show the mean sharing rule. The percentage of couples who fall into each group are reported in parentheses.

affects the estimated sharing rule. The results are reported in table 3, column 2. A Wald test does not reject the equality of estimates between the base model and the separability model.

Lastly, to capture potential generational differences, we estimate a specification that allows the sharing-rule parameters to differ across cohorts. We estimate cohort-specific sharing rules for ten-year birth groups. The results reported in Table 3, columns 3–9 suggest that, over time, the sharing rule shifted in favor of women, consistent with broader economic and social changes that improved their bargaining position within the household.

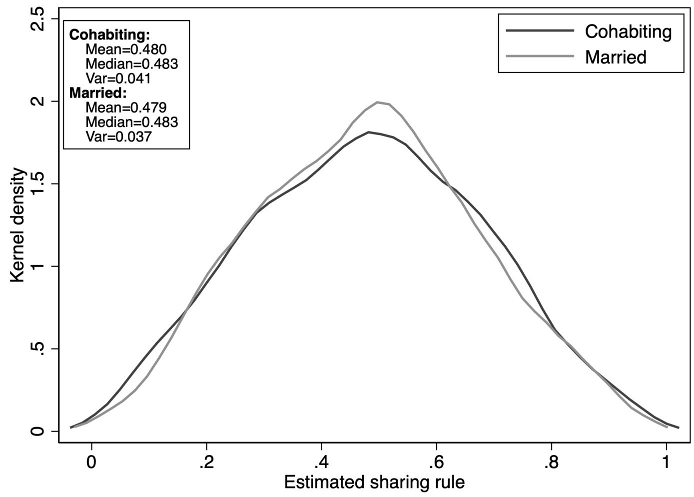
5.4.2 Propensity score matching

Given the substantial differences in observable characteristics between married and cohabitating couples—such as age, education, and income— we employ propensity score matching to improve the comparability of these groups. Specifically, we construct a subsample of married and cohabitating couples with similar observable characteristics, including the female partner's potential earnings share, age differences, full household income, age and age squared of both partners, years of education, and race. This procedure allows us to compare couples that are more similar along key dimensions, thereby reducing selection on observables and bringing the comparison closer to a causal interpretation of union type. At the same time, unobserved differences, such as preferences or access to non-labor income, may still vary across union types.

Figure 3a shows the distribution of the estimated sharing rule for married and cohabitating couples in the matched sample. In contrast to the full sample (Figure 2a), the mean and median sharing rule are now nearly identical across union types, indicating that differences in average allocation are largely driven by observable characteristics. However, the variance in the sharing rule is again larger for cohabitating unions, and there are differences in mass along the distribution. In particular, the cohabitation distribution remains slightly more dispersed in the tails, with more mass both at relatively low and relatively high values of the sharing rule. Consistent with this, the marginal effects reported in the last column of Table 3 show that the interaction between log income and cohabitation is virtually unchanged relative to the full sample, although estimated less precisely due to the smaller sample size. Thus, even among otherwise comparable couples, cohabitation is associated with substantially greater heterogeneity in intra-household allocation. This result highlights that the key distinction between union types lies not in average sharing, but in the distribution of bargaining outcomes within households. The detailed results for the matched sample are reported in Appendix A.

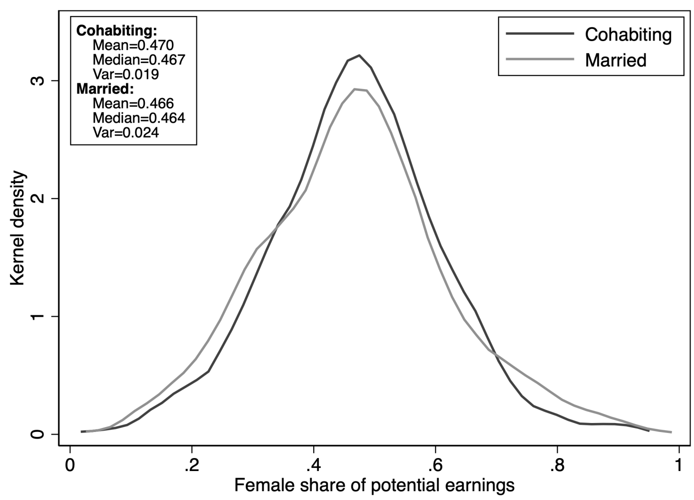
Figure 3: Distribution of sharing rule and female partner's potential wage by marital status in propensity score matched sample.

(a) Estimated sharing rule



Notes: The sharing rule measures share of full household income, net of public goods expenditure, transferred to female partner.

(b) Female share of potential earnings



Notes: Potential earnings reflect the labor income that could be earned if the agents worked 65 hours per week.

Notes: The figures are based on authors' calculations using PSID. Sample period: 1999 – 2017. N = 4,952. 50% are cohabitating, the rest is married. Potential earnings reflect the labor income that could be earned if the agents worked 65 hours per week.

6 Consumption inequality

In this section we discuss the implications of family arrangements for consumption inequality. We compare the distribution of the female share of consumption within a couple by union type, then show how much each family arrangement contributes to overall population consumption inequality, and finally end with a poverty analysis.

6.1 Distribution of Female Consumption Shares by Union Type

We first show the consumption inequality by union type predicted by our model for the US economy. Our model gives the following specification for individual consumption of each partner in both types of unions;

$$\hat{c}_i^f = \hat{\varphi}\overline{y} - \overline{\omega}^f + \omega^f(q^f, q^m) + \omega^f(h^f, h^m) + g + \omega^f(q^f, q^m) + \omega^m(q^m, q^f) \tag{16}$$

$$\hat{c}_i^m = (1-\hat{\varphi})\overline{y} - \overline{\omega}^m + \omega^m(q^m, q^f) + \omega^m(h^m, h^f) + g + \omega^f(q^f, q^m) + \omega^m(q^m, q^f) \tag{17}$$

It is the sum of expenditure on private consumption (first four terms) and household public good expenditure (last three terms). Using this specification and the estimated sharing rule for the full sample, Figure 4a plots the distribution of female consumption shares within households by union type. Two main patterns emerge. First, women's average consumption share is higher in cohabitating unions, while the distribution is more compressed among married couples. Second, consistent with the distribution of potential earnings shares, the actual earnings distribution exhibits thicker tails in marriage (Figure 4b), with a substantially larger fraction of married women contributing less than 20% to household labor income. Despite these pronounced differences in earnings, the distribution of female consumption shares is much more similar across union types, particularly in the tails. This suggests that married couples exhibit a higher degree of specialization combined with stronger income sharing, such that large disparities in earnings translate less directly into disparities in consumption.

Next, we estimate the full consumption of each partner. The full consumption of an individual is the sum of expenditure on private and public consumption and leisure. It is given as follows;

$$
\begin{aligned}
\hat{c}^f_{\text{full}} &= \hat{c}^f + [\bar{\omega}^f - \omega^f(h^f, h^m) - \omega^f(q^f, q^m)] + g + \omega^f(q^f, q^m) + \omega^m(q^m, q^f) \\
&= \hat{\varphi}\bar{y} + g + \omega^f(q^f, q^m) + \omega^m(q^m, q^f)
\end{aligned} \tag{18}
$$

$$
\begin{aligned}
\hat{c}^m_{full} &= \hat{c}^m + [\bar{\omega}^m - \omega^m(h^m, h^f) - \omega^m(q^m, q^f)] + g + \omega^f(q^f, q^m) + \omega^m(q^m, q^f) \\
&= (1 - \hat{\varphi})\bar{y} + g + \omega^f(q^f, q^m) + \omega^m(q^m, q^f)
\end{aligned} \tag{19}
$$

The distribution of full consumption is shown in Figure 4c. Compared to consumption excluding leisure, the distribution of full consumption shares is more compressed. At the same time, the relative patterns across union types are preserved: women in cohabitating unions receive, on average, a higher share of full consumption, and the distribution exhibits similar features to individual consumption, including differences in dispersion across union types, indicating that the inclusion of leisure does not eliminate the underlying differences in how resources are translated into individual consumption.

6.2 A Decomposition of Population Consumption Inequality by Marital Status

The preceding analysis highlights differences in intra-household resource allocation between marriage and cohabitation. On average, cohabitating partners exhibit a more egalitarian distribution of resources. However, these differences in female consumption shares do not by themselves reveal how much each family arrangement contributes to overall population consumption inequality. To address this, we use the estimated individual-level consumption measure to compute the variance of log consumption and decompose population inequality into between-household and within-household components across marital-status groups using the following specification (e.g. Denderski and Obermeier, 2025):

$$
\begin{aligned}
\text{Var}(\log \hat{c}_{it}) = p_M \text{Var}(\log \hat{c}_{it}|M) + p_C \text{Var}(\log \hat{c}_{it}|C) + p_S \text{Var}(\log \hat{c}_{it}|S) \\
+ \text{Var}(\log \hat{c}_{it}|\text{between M, C, S})
\end{aligned} \tag{20}
$$

Where p_M, p_C, p_S is the share of married, cohabitating and single individuals in the population respectively.[13] $\hat{c}_{it}$ is the estimated consumption of individual i at time t. Further for $K \in \{M, C, S\}$ we have

$$\text{Var}(\log \hat{c}_{it}|K) = \text{Var}(\mathbb{E}(\log \hat{c}_{it}|i \in k) + \mathbb{E}(\text{Var}(\log \hat{c}_{it}|i \in k)) \tag{21}$$

with k denoting the household in each marital group. Finally, for inequality between the three groups we have,

$$\text{Var}(\log c_{it}|\text{between M, C, S}) = \sum_K p_K[\mathbb{E}(\log \hat{c}_{it}|K) - \mathbb{E}(\log \hat{c}_{it})]^2 \tag{22}$$

Eq(20) together with eq(21) and eq(22) gives us six components of consumption inequality: 1) within married inequality, 2) between married inequality, 3) within cohabitation inequality, 4) between cohabitation inequality, 5) between singles inequality, 6) between three marital groups inequality.

Table 5 presents the results of the decomposition for consumption inequality in our full sample. The between-married component accounts for the largest share of overall consumption inequality (39.58%). The within-marriage component accounts for an additional 10% of consumption inequality, representing approximately 20% of total consumption inequality among married households. These findings differ somewhat from Lise and Seitz (2011) and Denderski and Obermeier (2025), who report that within-marriage inequality accounts for 25% and 27% of total consumption inequality among married individuals in the UK and the US, respectively. The estimated between-singles component (23.19%) is also lower than the 30% reported by Denderski and Obermeier (2025). A key difference relative to the existing literature is that we explicitly account for cohabitating couples. Cohabitors represent 8.3% of individuals in our sample and account for approximately 7% of total consumption inequality. While this contribution broadly reflects their population share, it is entirely absent in existing decompositions that do not distinguish cohabitation. Note that, in the decomposition, each group is weighted by its share in the population. Including cohabitation increases the share of between-group inequality in population consumption inequality to 20%, compared to 7% in Denderski and Obermeier (2025). The estimated population variance is 0.5533, which lies between the value of 0.45

[13] $p_M + p_C + p_S = 1$.

estimated by Aguiar and Bils (2015) for the US in 2005, and the value of 0.6 for the UK in 2000 reported by Lise and Seitz (2011), both using the variance of log consumption. Neither study considers the contribution of cohabitation to the population variance of log consumption. Our results show that distinguishing cohabitation from marriage substantially changes the contribution of between-group differences to overall inequality and provides a more complete picture of the distribution of consumption across individuals.

Table 5: Variance decomposition of individual log-consumption

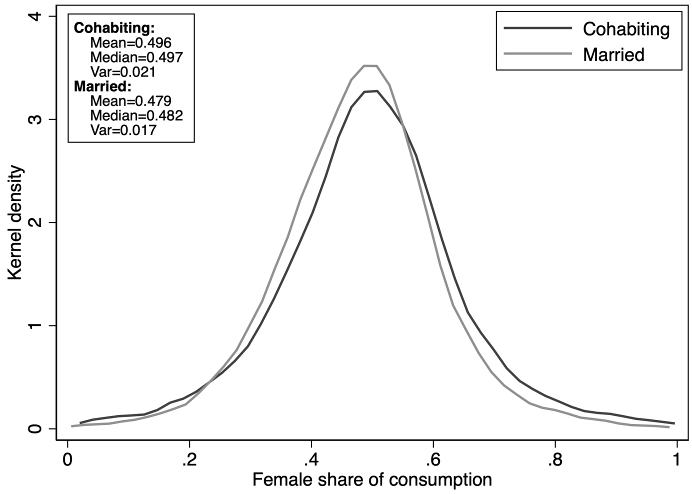
	Value	Fraction (%)
Between Married	0.2190	39.58
Within Married	0.0561	10.15
Between Cohabiting	0.0258	4.67
Within Cohabiting	0.0110	1.98
Between Singles	0.1283	23.19
Between Groups	0.1131	20.44
Population Variance	0.5533	100.00

Notes: The sample consists of single-person households, married and cohabitating couples from the PSID ($N = 30,406$), sample period 1999 − 2017. The table reports the decomposition of population variance of individual log-consumption into components by marital status group (married, cohabitating, single). Between households captures variance in mean consumption across households within each group. Within households captures variance in consumption between members of the same household. Singles are excluded from the within-household decomposition as they have no within-household variance by definition. Between Groups captures variance due to differences in mean consumption across the three marital status groups. All components sum to the population variance of 0.5533.

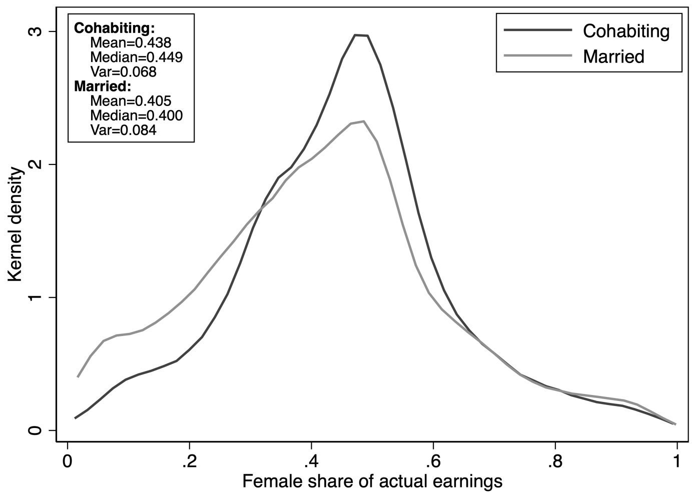
Table 11 in Appendix A presents the results based on full consumption. Population consumption inequality decreases substantially. The ranking of each component's contribution to total population variance remains consistent. These results suggest that a consumption inequality analysis that fails to account for union type may obscure significant heterogeneity. This is particularly relevant for policies that operate at the household level. For example, lump-sum transfers to households or tax credits targeted at families need not translate one-to-one into individual consumption, as the extent of intra-household redistribution depends on how resources are shared within couples. The greater heterogeneity in sharing among cohabitating couples implies that the outcomes of such policies may differ systematically across union types, even for households with similar observable characteristics.

Figure 4: Distribution of female partner's consumption, actual earnings and full consumption by marital status.

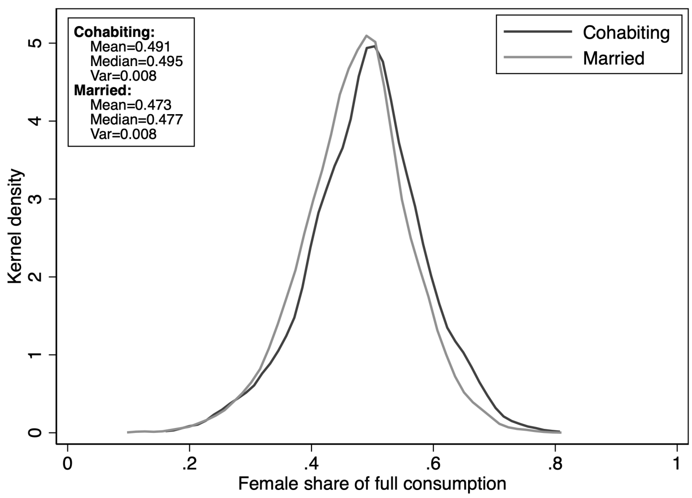
(a) Female share of consumption excluding leisure.



(b) Female share of actual earnings.

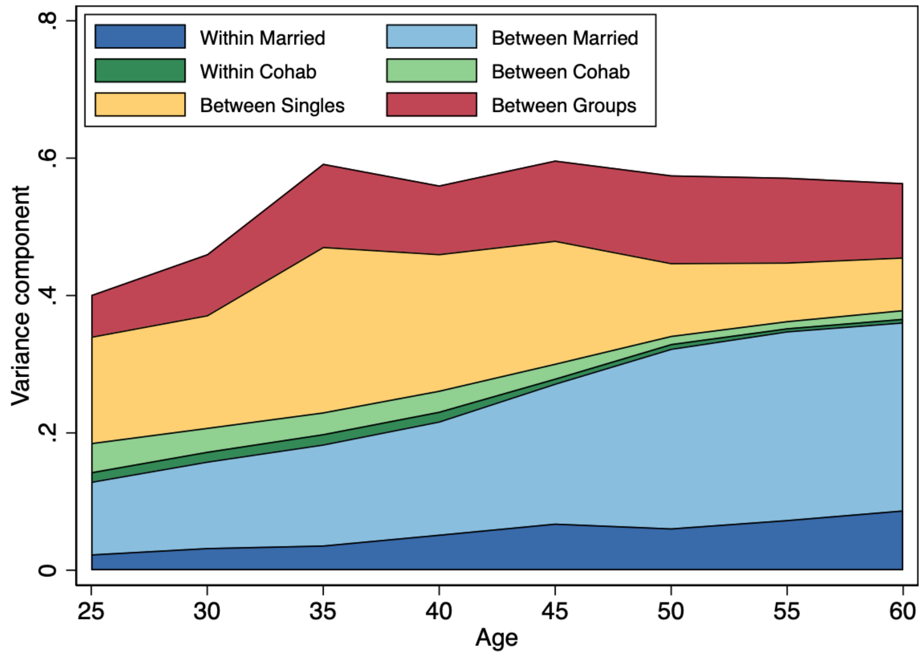


(c) Female share of full consumption leisure included.



Notes: The figures are based on authors' calculations using PSID. Sample period: 1999 – 2017. N = 20,718 (10,359 couples). 12.1% are cohabitating, the rest is married.

Figure 5: Consumption inequality by age



Notes: The sample consists of single-person households, married and cohabitating couples from the PSID ($N = 30,406$), sample period 1999 – 2017. The figure shows the decomposition of the variance of individual log consumption into components by marital status group and age. For married and cohabitating couples, age refers to the average age of both household members. Observations are grouped into 5-year age categories to ensure sufficient group sizes. Between households captures variance in mean consumption across households within each group. Within households captures the variance in consumption between members of the same household. Singles are excluded from the within-household decomposition as they have no within-household variance by definition. Between groups captures the variance due to differences in mean consumption across the three marital status groups.

Figure 5 presents the decomposition of inequality across age groups and reveals several patterns. First, overall consumption inequality increases with age, consistent with evidence that both income and consumption dispersion rise over the life cycle (Deaton and Paxson, 1994, 1997). Second, the contribution of inequality among married individuals increases with age, reflecting both rising within-group dispersion and the growing share of married households in older cohorts (see Fig. 7 in appendix). Third, the contribution of singles declines with age, consistent with their decreasing population share. In contrast, the contribution of cohabitating couples is concentrated in younger age groups and becomes negligible at older ages, primarily due to their declining prevalence. Despite these compositional shifts, the contribution of between-group inequality to overall inequality remains relatively stable across age groups. These patterns indicate that differences in family structure—and in the dispersion of intra-household allocation associated with them—matter primarily for inequality at younger ages, when cohabitation is more prevalent.

6.3 Poverty analysis by marital status

The inequality measures considered above summarize dispersion in consumption but do not capture how resources are distributed at the lower end of the distribution. To assess the extent to which individuals in different family arrangements are exposed to low consumption levels, we examine poverty incidence across marital-statuses. Estimates of the sharing rule allow us to conduct this analysis at the individual level. To determine poverty rates, we follow Cherchye et al. (2015). First, each household's full income and individual shares are computed using the sharing rule estimates from Section 5. The poverty line is defined as 60% of the median of individual-level income, where household income is initially split equally between partners to construct the reference distribution. We then compare poverty rates under two approaches: equal splitting of income between partners (Column 1 of Table 6) and allocation based on the estimated sharing rule (Columns 2–4), which accounts for intra-household inequality.

Comparing the two approaches yields three main findings. First, equal splitting understates individual-level poverty: 19.2% of individuals are classified as poor under equal splitting (Column 1), compared to 20.61% when using the sharing rule (Column 2). This is consistent with existing evidence that ignoring intra-household allocation leads to underestimation of poverty (Calvi, 2020; Cherchye et al., 2015; Dunbar et al., 2013). Second, equal splitting masks systematic within-household differences. When accounting for the sharing rule, women in both marriage and cohabitation exhibit higher poverty rates than their male partners (Columns 3 and 4). The gender gap is somewhat larger in marriage than in cohabitation, consistent with stronger specialization among married couples. Third, poverty differs markedly across union types: cohabitating individuals face higher poverty rates than married individuals under both approaches (Columns 1 and 2), and singles have the highest poverty rates overall, with single men being substantially poorer than single women.

Figure 6 complements these insights by showing substantial heterogeneity in poverty by age and family arrangement. While singles have the highest poverty rates in every age group, the ranking between married and cohabitating individuals varies over the life cycle, indicating that the aggregate poverty advantage of marriage does not hold

Table 6: Poverty rates

	Equal Sharing	Sharing according to sharing rule		
	All individuals (%)	All individuals (%)	Men (%)	Women (%)
All individuals	19.20	20.61	20.83	20.39
Single	34.62	34.62	37.28	31.65
Cohabiting	13.43	16.30	15.26	17.33
Married	11.79	13.76	12.37	15.14
Diff. (Cohab - Married)	1.64	2.54	2.89	2.19

Notes: The sample consists of single-person households, married and cohabitating couples from the PSID ($N = 30,406$), sample period 1999 – 2017. The table shows poverty rates in percent. The poverty line is defined as 60% of the median individual full income, calculated under the assumption of equal splitting within all households. Column 1 reports poverty rates assuming income is divided equally between partners. Columns 2–4 report poverty rates where income is allocated according to the estimated sharing rule from Section 5, thereby accounting for potential intra-household inequality. For single-person households both approaches yield identical results.

uniformly within age groups.

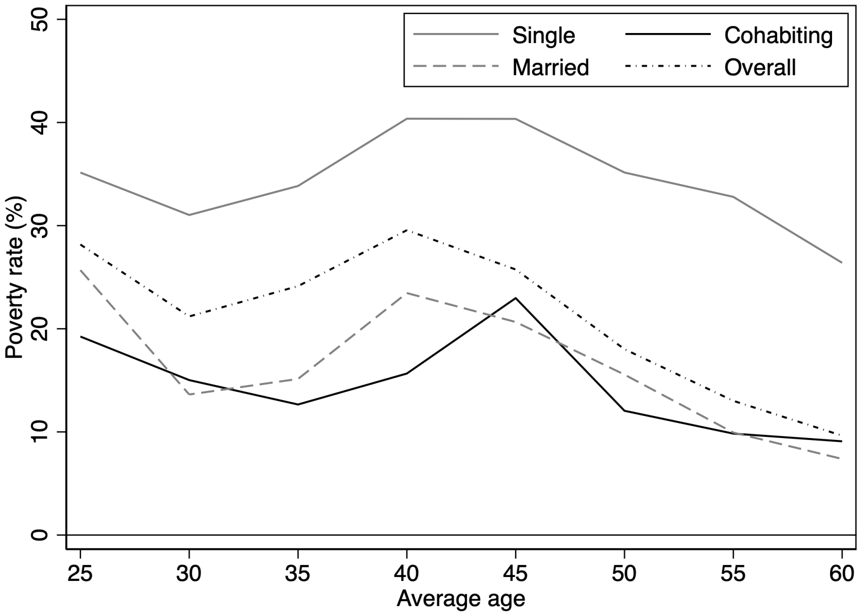
Overall, the poverty analysis highlights the importance of accounting for intra-household resource allocation. Equal splitting understates poverty and masks substantial differences within couples, in particular the higher incidence of poverty among women. While singles are consistently the most exposed group, differences between married and cohabitating individuals are heterogeneous and vary across the life cycle, suggesting that aggregate gaps partly reflect differences in group composition. Taken together, these results show that both selection into family arrangements and the distribution of resources within households are central for understanding who is at risk of poverty.

7 Conclusion

This paper studies how family structure shapes the measurement of individual consumption inequality in the United States. We extend the collective household framework by Lise and Seitz (2011) to distinguish between marriage and cohabitation and to incorporate housework into home production. Using PSID data, the analysis recovers sharing rules for married and cohabitating couples, constructs individual-level consumption measures, and examines how marriage, cohabitation, and singlehood contribute to inequality and poverty.

A central question is whether cohabitating couples should be expected to mirror mar-

Figure 6: Poverty rates by age and marital status



Notes: The sample consists of single-person households, married and cohabitating couples from the PSID ($N = 30,406$), sample period $1999 - 2017$. The figure depicts poverty rates by marital status in 5-year average age bins. For couples, age refers to the average age of both household members. The poverty line is defined as 60% of median individual full consumption, calculated under the assumption of equal splitting within all households. Full consumption is allocated according to the estimated sharing rule. Singles are assigned their individual consumption directly.

ried couples in their patterns of resource sharing. Without the institutional protections afforded to marriage, cohabitation might be expected to involve limited resource sharing and greater individual autonomy. Contrary to this expectation, the findings show that, after controlling for observable characteristics, cohabitating couples share resources at levels comparable to married couples. This suggests that substantial resource sharing occurs even outside of marriage. However, the distribution of sharing among cohabitating couples is more dispersed and exhibits thicker tails, indicating greater heterogeneity and less structured intra-household allocation.

These distinctions have important quantitative implications for the measurement of consumption inequality and poverty rates in the macroeconomy. Family structure is a key determinant of consumption inequality, with a large share of inequality arising both within and between different union types. Differentiating cohabitation from marriage increases the observed contribution of between-group differences and uncovers patterns of inequality that are not captured when cohabitating couples are subsumed under marriage. Moreover, accounting for intra-household allocation changes the assessment of poverty, uncovering higher poverty rates overall and systematic differences within couples that are obscured under equal-split approaches.

More broadly, the paper shows that both household composition and intra-household allocation are critical for understanding how resources are distributed across individuals. Treating cohabitation as equivalent to marriage, or assuming equal sharing within households, can lead to incomplete or misleading assessments of inequality and poverty. As cohabitation becomes increasingly prevalent, distinguishing between family arrangements is necessary for accurately measuring individual well-being and for evaluating the effectiveness of policies that operate at the household level.

References

Adamopoulou, E., A. Hannusch, K. A. Kopecky, and T. Obermeier (2025). Cohabitation, child development, and college costs. *Child Development, and College Costs. IZA Discussion Paper* (18237).

Aguiar, M. and M. Bils (2015). Has consumption inequality mirrored income inequality? *American Economic Review 105*(9), 2725–2756.

Ando, A. and F. Modigliani (1963). The "life cycle" hypothesis of saving: Aggregate implications and tests. *American Economic Review 53*(1), 55–84.

Attanasio, O., H. Low, and V. Sánchez-Marcos (2005). Female labor supply as insurance against idiosyncratic risk. *Journal of the European Economic Association 3*(2-3), 755–764.

Attanasio, O. P. and L. Pistaferri (2016). Consumption inequality. *Journal of Economic Perspectives 30*(2), 3–28.

Avellar, S. and P. J. Smock (2005). The economic consequences of the dissolution of cohabiting unions. *Journal of Marriage and Family 67*(2), 315–327.

Axinn, W. G. and A. Thornton (1992). The relationship between cohabitation and divorce: Selectivity or causal influence? *Demography 29*(3), 357–374.

Becker, G. S. (1965). A theory of the allocation of time. *The Economic Journal 75*(299), 493–517.

Becker, G. S. (1973). A theory of marriage: Part i. *Journal of Political Economy 81*(4), 813–846.

Becker, G. S. (1974). A theory of marriage: Part ii. *Journal of Political Economy 82*(2), S11–S26. Part 2.

Belete, G. Y., M. Menon, and F. Perali (2022). Children's resources and poverty: A collective consumption evidence from Ethiopia. SITES Working Paper 1.

Belloc, I., J. A. Molina, and J. Velilla (2022). How does intrahousehold bargaining power impact labor supply? European cross-country evidence (2004–2019). GLO Discussion Paper 1132.

Bergstrom, T., L. Blume, and H. Varian (1986). On the private provision of public goods. *Journal of Public Economics 29*(1), 25–49.

Blasutto, F. (2024). Cohabitation vs marriage: Mating strategies by education in the USA. *Journal of the European Economic Association 22*(4), 2899–2924.

Blasutto, F. and E. Kozlov (2025). (Changing) marriage and cohabitation patterns in the US: Do divorce laws matter? ECARES Working Paper 2025-02.

Blundell, R., P.-A. Chiappori, T. Magnac, and C. Meghir (2007). Collective labour supply: Heterogeneity and non-participation. *The Review of Economic Studies 74*(2), 417–445.

Blundell, R., A. Duncan, and C. Meghir (1998). Estimating labor supply responses using tax reforms. *Econometrica*, 827–861.

Blundell, R. and I. Preston (1998). Consumption inequality and income uncertainty. *The Quarterly Journal of Economics 113*(2), 603–640.

Blundell, R. and I. Walker (1986). A life-cycle consistent empirical model of family labour supply using cross-section data. *The Review of Economic Studies 53*(4), 539–558.

Bobonis, G. J. (2009). Is the allocation of resources within the household efficient? New evidence from a randomized experiment. *Journal of Political Economy 117*(3), 453–503.

Borella, M., M. De Nardi, M. Pak, N. Russo, and F. Yang (2023, 09). FBBVA Lecture 2023. The Importance of Modeling Income Taxes over Time: U.S. Reforms and Outcomes. *Journal of the European Economic Association 21*(6), 2237–2286.

Brien, M. J., L. A. Lillard, and S. Stern (2006). Cohabitation, marriage, and divorce in a model of match quality. *International Economic Review 47*(2), 451–494.

Browning, M., F. Bourguignon, P.-A. Chiappori, and V. Lechene (1994). Income and outcomes: A structural model of intrahousehold allocation. *Journal of Political Economy 102*(6), 1067–1096.

Browning, M., P.-A. Chiappori, and A. Lewbel (2013). Estimating consumption economies of scale, adult equivalence scales, and household bargaining power. *Review of Economic Studies 80*(4), 1267–1303.

Bumpass, L. and H.-H. Lu (2000). Trends in cohabitation and implications for children's family contexts in the United States. *Population Studies 54*(1), 29–41.

Calvi, R. (2020). Why are older women missing in India? The age profile of bargaining power and poverty. *Journal of Political Economy 128*(7), 2453–2501.

Calvo, P. (2025). The effects of institutional gaps between cohabitation and marriage. Working paper, Arizona State University.

Chen, S.-E. and J. C. van Ours (2020). Symbolism matters: The effect of same-sex marriage legalization on partnership stability. *Journal of Economic Behavior & Organization 178*, 44–58.

Cherchye, L., S. Cosaert, B. De Rock, P. J. Kerstens, and F. Vermeulen (2018). Individual welfare analysis for collective households. *Journal of Public Economics 166*, 98–114.

Cherchye, L., B. De Rock, A. Lewbel, and F. Vermeulen (2015). Sharing rule identification for general collective consumption models. *Econometrica 83*(5), 2001–2041.

Cherchye, L., B. De Rock, and F. Vermeulen (2011). The revealed preference approach to collective consumption behaviour: Testing and sharing rule recovery. *Review of Economic Studies 78*(1), 176–198.

Cherchye, L., B. De Rock, and F. Vermeulen (2012). Married with children: A collective labor supply model with detailed time use and intrahousehold expenditure information. *American Economic Review 102*(7), 3377–3405.

Chiappori, P.-A. (1988). Rational household labor supply. *Econometrica 56*(1), 63–90.

Chiappori, P.-A. (1992). Collective labor supply and welfare. *Journal of Political Economy 100*(3), 437–467.

Chiappori, P.-A., B. Fortin, and G. Lacroix (2002). Marriage market, divorce legislation, and household labor supply. *Journal of Political Economy 110*(1), 37–72.

Chiappori, P.-A., B. Salanié, and Y. Weiss (2017). Partner choice, investment in children, and the marital college premium. *American Economic Review 107*(8), 2109–2167.

Cutler, D. M. and L. F. Katz (1992). Rising inequality? Changes in the distribution of income and consumption in the 1980s. NBER Working Paper 3964.

De Vreyer, P. and S. Lambert (2021). Inequality, poverty, and the intra-household allocation of consumption in Senegal. *The World Bank Economic Review 35*(2), 414–435.

Deaton, A. and C. Paxson (1994). Intertemporal choice and inequality. *Journal of Political Economy 102*(3).

Deaton, A. and C. Paxson (1997). The effects of economic and population growth on national saving and inequality. *Demography 34*(1), 97–114.

Del Boca, D. and C. Flinn (2012). Endogenous household interaction. *Journal of Econometrics 166*(1), 49–65.

DeMaris, A. and G. R. Leslie (1984). Cohabitation with the future spouse: Its influence upon marital satisfaction and communication. *Journal of Marriage and the Family*, 77–84.

Denderski, P. and T. Obermeier (2025). Marriage, intra-household inequality, and wage risk. IFS Working Paper W25/47, Institute for Fiscal Studies.

Dunbar, G. R., A. Lewbel, and K. Pendakur (2013). Children's resources in collective households: Identification, estimation, and an application to child poverty in Malawi. *American Economic Review 103*(1), 438–471.

Echeverría, L., M. Menon, F. Perali, and M. Berges (2019). Intra-household inequality and child welfare in Argentina. CEDLAS Working Paper 241.

Eickmeyer, K. J., W. D. Manning, M. A. Longmore, and P. C. Giordano (2023). Exploring the married-cohabiting income pooling gap among young adults. *Journal of family and economic issues 44*(4), 990–1006.

Eika, L., M. Mogstad, and B. Zafar (2019). Educational assortative mating and household income inequality. *Journal of Political Economy 127*(6), 2795–2835.

Fair, R. C. and K. M. Dominguez (1991). Effects of the changing U.S. age distribution on macroeconomic equations. *American Economic Review 81*(5), 1276–1294.

Federal Housing Finance Agency (2025). House price index (hpi) datasets. Retrieved from Federal Housing Finance Agency. https://www.fhfa.gov/data/hpi/datasets?tab=quarterly-data.

Fernandez, R., N. Guner, and J. Knowles (2005). Love and money: A theoretical and empirical analysis of household sorting and inequality. *The Quarterly Journal of Economics 120*(1), 273–344.

Fernández, R. and R. Rogerson (2001). Sorting and long-run inequality. *The Quarterly Journal of Economics 116*(4), 1305–1341.

Fernandez-Villaverde, J. and D. Krueger (2007). Consumption over the life cycle: Facts from consumer expenditure survey data. *The Review of Economics and Statistics 89*(3), 552–565.

Gemici, A. and S. Laufer (2011). Marriage and cohabitation. Mimeo, New York University.

Gobbi, P. E., J. Parys, and G. Schwerhoff (2018). Intra-household allocation of parental leave. *Canadian Journal of Economics 51*(1), 236–274.

Goussé, M. and M. Leturcq (2022). More or less unmarried. legal setting of cohabitation and labor market outcomes. *European Economic Review 149*, 104259.

Graefe, D. R. and D. T. Lichter (1999). Life course transitions of American children: Parental cohabitation, marriage, and single motherhood. *Demography 36*(2), 205–217.

Greenwood, J., N. Guner, G. Kocharkov, and C. Santos (2014). Marry your like: Assortative mating and income inequality. *American Economic Review 104*(5), 348–353.

Hall, D. R. and J. Z. Zhao (1995). Cohabitation and divorce in Canada: Testing the selectivity hypothesis. *Journal of Marriage and the Family*, 421–427.

Hamplová, D., C. Le Bourdais, and É. Lapierre-Adamcyk (2014). Is the cohabitation–marriage gap in money pooling universal? *Journal of Marriage and Family 76*(5), 983–997.

Heimdal, K. R. and S. K. Houseknecht (2003). Cohabiting and married couples' income organization: Approaches in Sweden and the United States. *Journal of Marriage and Family 65*(3), 525–538.

Hiekel, N., A. C. Liefbroer, and A.-R. Poortman (2014a). Income pooling strategies among cohabiting and married couples: A comparative perspective. *Demographic Research 30*(55), 1527–1560.

Hiekel, N., A. C. Liefbroer, and A.-R. Poortman (2014b). Understanding diversity in the meaning of cohabitation across Europe. *European Journal of Population 30*(4), 391–410.

Hougaard Jensen, S. E., S. P. Olafsson, T. S. Sveinsson, and G. Zoega (2022). Mapping educational disparities in life-cycle consumption. CESifo Working Paper 9855.

Hurst, E. and A. Lusardi (2004). Liquidity constraints, household wealth, and entrepreneurship. *Journal of Political Economy 112*(2), 319–347.

Johnson, D. and S. Shipp (1997). Trends in inequality using consumption-expenditures: The US from 1960 to 1993. *Review of Income and Wealth 43*(2), 133–152.

Kapelle, N., N. Frémeaux, P. M. Lersch, and M. Leturcq (2025, 03). A cohabitation wealth premium for women and men: Considering the regulatory framework and normative acceptance in France and Germany. *Socio-Economic Review 23*(2), 591–620.

Lam, D. (1988). Marriage markets and assortative mating with household public goods: Theoretical results and empirical implications. *Journal of Human Resources*, 462–487.

Lise, J. and S. Seitz (2011). Consumption inequality and intra-household allocations. *Review of Economic Studies 78*(1), 328–355.

Lise, J. and K. Yamada (2019). Household sharing and commitment: Evidence from panel data on individual expenditures and time use. *The Review of Economic Studies 86*(5), 2184–2219.

Lundberg, S. and R. A. Pollak (1994). Noncooperative bargaining models of marriage. *The American Economic Review 84*(2), 132–137.

Lundberg, S., R. A. Pollak, and J. Stearns (2016). Family inequality: Diverging patterns in marriage, cohabitation, and childbearing. *Journal of Economic Perspectives 30*(2), 79–102.

Lyngstad, T. H., T. Noack, and P. A. Tufte (2010). Pooling of economic resources: A comparison of Norwegian married and cohabiting couples. *European Sociological Review 27*(5), 624–635.

MaCurdy, T. E. (1982). The use of time series processes to model the error structure of earnings in a longitudinal data analysis. *Journal of Econometrics 18*(1), 83–114.

Manning, W. (2013). Trends in cohabitation: Over twenty years of change, 1987-20. *Studies 54*, 29–41.

Matouschek, N. and I. Rasul (2008). The economics of the marriage contract: Theories and evidence. *Journal of Law and Economics 51*(1), 59–110.

Modigliani, F. and R. H. Brumberg (1954). Utility analysis and the consumption function: An interpretation of cross-section data. In K. K. Kurihara (Ed.), *Post-Keynesian Economics*, pp. 388–436. Rutgers University Press.

Oreffice, S. (2011). Sexual orientation and household decision making: Same-sex couples' balance of power and labor supply choices. *Labour Economics 18*(2), 145–158.

Oreffice, S. (2014). Culture and household decision making: Balance of power and labor supply choices of U.S.-born and foreign-born couples. *Journal of Labor Research 35*(2), 162–184.

Rao, K. V. and J. Trussell (1989). Premarital cohabitation and marital stability: A reassessment of the Canadian evidence: Feedback. *Journal of Marriage and the Family 51*, 535–540.

Ridley, C. A., D. J. Peterman, and A. W. Avery (1978). Cohabitation: Does it make for a better marriage? *Family Coordinator*, 129–136.

Rosenfeld, M. J. and K. Roesler (2019). Cohabitation experience and cohabitation's association with marital dissolution. *Journal of Marriage and Family 81*(1), 42–58.

Schoen, R. and R. M. Weinick (1993). Partner choice in marriages and cohabitations. *Journal of Marriage and Family 55*(2), 408.

Smock, P. J. (2000). Cohabitation in the United States: An appraisal of research themes, findings, and implications. *Annual Review of Sociology 26*(1), 1–20.

Smock, P. J. and W. D. Manning (1997). Cohabiting partners' economic circumstances and marriage. *Demography 34*(3), 331–341.

Trost, J. (1975). Married and unmarried cohabitation: The case of Sweden, with some comparisons. *Journal of Marriage and the Family*, 677–682.

Vermeulen, F. (2005). And the winner is... an empirical evaluation of unitary and collective labour supply models. *Empirical Economics 30*(3), 711–734.

Vermeulen, F. M. P., O. Bargain, M. Beblo, D. Beninger, R. Blundell, R. Carrasco, M.-C. Chiuri, F. Laisney, V. Lechene, N. Moreau, M. Myck, and J. Ruiz-Castillo (2006). Collective models of labor supply with nonconvex budget sets and nonparticipation: A calibration approach. *Review of Economics of the Household 4*(2), 113–127.

World Bank (2026). Inflation, consumer prices for the United States [FPCPITOTLZGUSA]. FRED, Federal Reserve Bank of St. Louis. Retrieved February 17, 2026.

Wydick, B. (2007). Grandma was right: Why cohabitation undermines relational satisfaction, but is increasing anyway. *Kyklos 60*(4), 617–645.

Xue, X., Y. Hao, and L. Han (2026). Effect of household structure changes on consumption inequality: An empirical study from household data in China. *Sage Open 16*(1), 21582440251411354.

A Appendix A

Table 7: Estimates of preference parameters without accounting for housework time

	Base Estimates	
	Women	Men
ρ	0.2405 (0.0043)	0.2522 (0.0041)
$\alpha(\mathbf{x})$ at mean of data	0.0647 (0.0028)	0.0971 (0.0046)
α:		
Age	-0.3757 (0.1969)	-0.6526 (0.1889)
Age2	0.6236 (0.2009)	1.0087 (0.1767)
Education	0.0126 (0.0155)	-0.0187 (0.0184)
1930 cohort	-0.0405 (0.1069)	-0.0034 (0.0753)
1940 cohort	-0.0571 (0.0684)	-0.0661 (0.0552)
1960 cohort	0.2525 (0.0672)	0.2079 (0.0782)
1970 cohort	0.1597 (0.0780)	0.3259 (0.1060)
1980 cohort	0.1315 (0.0934)	0.3808 (0.1173)
1990 cohort	0.2015 (0.1290)	0.3798 (0.1727)
α_0	2.6707 (0.0457)	2.2299 (0.0527)
σ_ν	0.6575 (0.0124)	0.7062 (0.0150)
corr(ν,e)	0.0274 (0.0195)	0.0704 (0.0187)

Notes: This table presents the estimates of the preference parameters for women and men. The sample consists of married and cohabitating couples from the PSID ($N = 20,718$), sample period 1999 – 2017. Time spent on housework is treated as leisure. Standard errors are bootstrapped with 1,000 replications and clustered at the couple level (reported in parentheses). The elasticity of substitution is parameterized as $\frac{1}{1-\rho}$ and the consumption share parameter as $\alpha_i = \frac{1}{1+\exp(\mathbf{x_i}'\alpha)}$.

Table 8: Sharing rule estimates without accounting for housework time

Estimates	Base Model
φ (at mean of data)	0.4461 (0.0008)
Marginal effect of:	
Female partner's potential earnings share	0.9154 (0.0129)
Log of full income	-0.0403 (0.0119)
Log income $\times$ cohabitation	0.0127 (0.0344)
Partner's age gap	-0.0005 (0.0019)

Notes: The sample consists of married and cohabitating couples from the PSID ($N = 20,718$), sample period 1999 – 2017. Time spent on housework is treated as leisure. Bootstrapped standard errors clustered at couple level with 1,000 replications are in parentheses. The sharing rule is parameterized as $\varphi(\mathbf{z}_i) = \frac{\exp(\mathbf{z}_i'\varphi)}{1+\exp(\mathbf{z}_i'\varphi)}$, and the marginal effect is evaluated at the mean of $\mathbf{z} \equiv [\frac{\overline{\omega}^f}{\overline{\omega}^f+\overline{\omega}^m} - \frac{1}{2}, \log(y) - \overline{\log(y_t)}, (age_f - age_m) - \overline{(age_f - age_m)}, K \times (\log(y) - \overline{\log(y_t)})]'$. The variables entering the vector $\mathbf{z}$ are transformed such that the sharing rule is normalized to equal one-half when both partners have equal potential wages, evaluated at the mean of log full income and the mean age gap. The test for the null of separability is a test that the first two columns are equal. The test for the null of equality of estimates tests whether the estimated sharing rule coefficients are equal across all birth cohorts.

Table 9: Estimates of preference parameters for propensity score matched sample

	Base Estimates		**Separability Test**	
	Women	Men	Women	Men
ρ	0.1762	0.1645	0.1890	0.1625
	(0.0133)	(0.0104)	(0.0167)	(0.0142)
$\alpha(\mathbf{x})$ at mean of data	0.0744	0.0445	0.0384	0.0306
	(0.0146)	(0.0074)	(0.0076)	(0.0049)
α:				
Age	-0.0149	0.0534	-0.0000	0.1369
	(0.1592)	(0.1207)	(0.0000)	(0.0812)
Age2	0.5826	0.0717	0.8091	0.1225
	(0.1924)	(0.1269)	(0.1739)	(0.0762)
Education	0.2471	0.1570	0.3909	0.2734
	(0.1086)	(0.0815)	(0.1019)	(0.0798)
1930 cohort	0.5077	-0.0233	-0.3528	-0.0131
	(0.5902)	(0.1534)	(0.2727)	(0.0112)
1940 cohort	-0.3888	0.2037	-0.4569	0.1068
	(0.3244)	(0.2263)	(0.3608)	(0.0923)
1960 cohort	0.1997	0.0387	0.1539	0.0617
	(0.1971)	(0.1273)	(0.1128)	(0.0497)
1970 cohort	0.5458	0.0229	0.2755	0.0999
	(0.3226)	(0.1472)	(0.1874)	(0.0816)
1980 cohort	0.9946	-0.0147	0.9321	0.0309
	(0.3435)	(0.1975)	(0.2530)	(0.0255)
1990 cohort	1.0844	-0.0198	1.1148	-0.0119
	(0.4201)	(0.1683)	(0.3816)	(0.0100)
α_0	2.5208	3.0661	3.2195	3.4548
	(0.2626)	(0.1475)	(0.2203)	(0.1918)
σ_ν	0.8071	0.8960	0.7299	0.6318
	(0.0350)	(0.0373)	(0.0429)	(0.0412)
corr(ν,e)	-0.0071	-0.0603	0.1586	0.2223
	(0.0326)	(0.0364)	(0.0793)	(0.0564)
Public goods expenditure			0.0011	0.0009
			(0.0001)	(0.0002)

Notes: This table presents the estimates of the preference parameters for women and men for the propensity score matched sample. The sample consists of married and cohabitating couples from the PSID ($N = 4,952$), sample period 1999 – 2017. The matched sample is constructed using propensity score matching based on female partner's potential earnings share, age difference between spouses, log full household income, age and age squared of both partners, years of education, and race. The first two columns display the preference parameters for the base model, while the final two columns report results for the separability test, which includes public goods expenditure. Standard errors are bootstrapped with 1,000 replications and clustered at the couple level (reported in parentheses). The elasticity of substitution is parameterized as $\frac{1}{1-\rho}$ and the consumption share parameter as $\alpha_i = \frac{1}{1+\exp(\mathbf{x_i}'\alpha)}$.

Table 10: Sharing rule estimates for propensity score matched sample

Estimates	Pooled	Separability Test	Birth Cohort 1930	1940	1950	1960	1970	1980	1990
φ (mean of data)	0.4644	0.4608	0.4169	0.4413	0.4420	0.4668	0.4575	0.4785	0.4623
	(0.0017)	(0.0028)	(0.0129)	(0.0071)	(0.0050)	(0.0021)	(0.0023)	(0.0020)	(0.0037)
Marginal effect of									
Female partner's potential earnings share	1.1133	1.2241	1.2769	1.0939	1.1874	1.1384	1.1381	1.1120	1.0435
	(0.0538)	(0.0874)	(0.1909)	(0.1297)	(0.0999)	(0.0726)	(0.0600)	(0.1045)	(0.1015)
Log of full income	−0.0698	−0.0601	−0.0656	−0.1230	−0.1487	−0.0549	−0.0512	−0.0713	−0.0459
	(0.0118)	(0.0100)	(0.0128)	(0.0366)	(0.0483)	(0.0362)	(0.0817)	(0.0833)	(0.0667)
Log income × cohabitation	0.0211	0.0186	0.0187	−0.0355	0.0156	0.0636	−0.0106	0.0407	0.0375
	(0.0165)	(0.0083)	(0.0035)	(0.0129)	(0.0453)	(0.0194)	(0.0394)	(0.0790)	(0.0839)
Partner's age gap	−0.0324	−0.0279	−0.0031	−0.0234	0.0613	0.0763	−0.0406	0.1560	0.0105
	(0.0181)	(0.0017)	(0.0207)	(0.0921)	(0.0935)	(0.0798)	(0.1232)	(0.0684)	(0.0680)
	Test null of separability:				Test null of equality of estimates:				
p-value	0.7982				0				

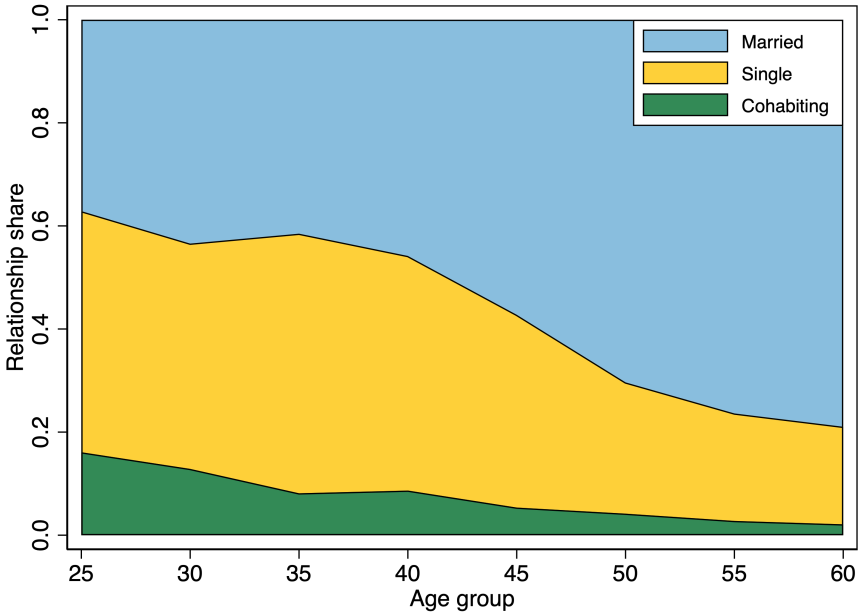
Notes: This table presents the estimates of the preference parameters for women and men for the propensity score matched sample. The sample consists of married and cohabitating couples from the PSID ($N = 4,952$), sample period 1999 − 2017. The matched sample is constructed using propensity score matching based on female partner's potential earnings share, age difference between spouses, log full household income, age and age squared of both partners, years of education, and race. Standard errors are bootstrapped with 1,000 replications and clustered at the couple level (reported in parentheses). The sharing rule is parameterized as $\varphi(\mathbf{z}_i) = \frac{\exp(\mathbf{z}_i'\varphi)}{1+\exp(\mathbf{z}_i'\varphi)}$, and the marginal effect is evaluated at the mean of $\varphi(\mathbf{z}_i) = \frac{\exp(\mathbf{z}_i'\varphi)}{1+\exp(\mathbf{z}_i'\varphi)}$, and the marginal effect is evaluated at the mean of $\mathbf{z} \equiv [\frac{\overline{\omega}^f}{\overline{\omega}^f+\overline{\omega}^m} - \frac{1}{2}, \log(y) - \overline{\log(y_t)}, (age_f - age_m) - \overline{(age_f - age_m)}, K \times (\log(y) - \overline{\log(y_t)})]'$. The variables entering the vector $\mathbf{z}$ are transformed such that the sharing rule is normalized to equal one-half when both partners have equal potential wages, evaluated at the mean of log full income and the mean age gap. The test for the null of separability is a test that the first two columns are equal. The test for the null of equality of estimates tests whether the estimated sharing rule coefficients are equal across all birth cohorts.

Table 11: Variance decomposition of full individual log-consumption

	Value	Fraction (%)
Between Married	0.1832	48.72
Within Married	0.0219	5.82
Between Cohabiting	0.0195	5.19
Within Cohabiting	0.0031	0.82
Between Singles	0.1114	29.61
Between Groups	0.0370	9.83
Population Variance	0.3762	100.00

Notes: The sample consists of single-person households, married and cohabitating couples from the PSID ($N = 30,406$), sample period 1999 − 2017. The table reports the decomposition of population variance of individual log-full consumption into components by marital status group (married, cohabitating, single). Between households captures variance in mean consumption across households within each group. Within households captures variance in consumption between members of the same household. Singles are excluded from the within-household decomposition as they have no within-household variance by definition. Between Groups captures variance due to differences in mean consumption across the three marital status groups. All components sum to the population variance of 0.3762.

Figure 7: Relationship shares by age



Notes: The sample consists of single-person households, married and cohabitating couples from the PSID ($N = 30,406$), sample period 1999 – 2017. The figure reports relationship shares by age. For married and cohabitating couples, age refers to the average age of both household members. Observations are grouped into 5-year age categories to ensure sufficient group sizes.