Children Costs in a One-Adult Household: Empirical Evidence from the UK.

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Abstract

This paper addresses two central questions in family and economic policy. First, to what extent are estimates of the cost of children derived from two-parent households applicable to single-parent families? Second, is the recently introduced *two-child limit* policy in the UK appropriate given the diversity of family structures? To address these questions, I propose a collective consumption model for one-adult households, apply it to three datasets—the Family Expenditure Survey, the Expenditures and Food Survey, and the Living Costs and Food Survey in the UK—and present two key findings. First, child cost estimates derived from two-parent households remain externally valid for single-parent families, at least for single mothers. Second, in low-income families, household size plays a crucial role in determining the proportion of resources allocated to children, a factor less relevant for higher-income families. This suggests that the *two-child limit* policy would likely exacerbate inequalities and increase child poverty within low-income families.

Keywords: Collective Model, Shadow price, Economies of scale, Identification, Resource sharing
JEL Classification: C30, D11, D12, D36, D63, I31, J12, J13

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1 Introduction

As debates over the equity and efficiency of social benefits intensify, a surprisingly underexplored question emerges: do the measurement tools and targeting instruments of public policy genuinely account for the specificity of single-parent families, or do they, by default, perpetuate standards designed for two-parent households? The economic literature on intra-household resource allocation has largely relied on data and models primarily developed for—and validated on—couples.[1] See, e.g., Bourguignon [1999], Browning et al. [2013], Cherchye et al. [2012], Lise and Seitz [2011], Dunbar et al. [2013, 2019], Bargain et al. [2022a], among others. This tradition is rooted in the assumption that structural differences—such as goods sharing, production complementarities, or economies of scale—are, in principle, adequately identified by collective models, thereby making results directly transferable to all families. However, recent research increasingly suggests that this hypothesis of transferability masks important disparities [DeLeire and Levy, 2004, Nieuwenhuis and Maldonado, 2018].[2] Family structure, the absence of a second adult, and the unique economic vulnerability of single parents introduce specific conditions that collective models might not fully capture. This observation calls into question the external validity of estimates of child costs derived from couple-based models, especially in the context of single-parent families.[3]

Alongside these empirical concerns, rising attention to child poverty and the implementation of structural reforms such as the *two-child limit* - which caps family benefits to the first two children in the United Kingdom (UK) - have fundamentally reshaped the issues of targeting and fairness in social policy.[4] While existing research consistently shows that the *two-child limit* has raised child poverty by lowering household income, it does not address the underlying mechanisms through which the policy may exacerbate inequality and poverty [González and Trommlerová, 2023, Chzhen and Bradshaw, 2025, Stewart et al., 2025a,b].[5]

Against this backdrop, this paper questions the validity of applying results from collective models estimated on couples to single-parent families. It also examines how family size (number of children in the household) and the parent's position in the income distribution influence children's well-being. Two main questions guide the analysis: (i) To what extent are estimates of the cost of children derived from couples relevant for single parents? (ii) Is the recently introduced *two-child limit* policy in the UK suitable for the diversity of family structures?

To address these questions, the paper extends collective models to the analysis of single-parent households in the UK, using expenditure surveys spanning four decades. The empirical validation is conducted on a collective consumption model inspired by Bargain, Donni, and Hentati [2022a,

[1] A few notable exceptions examining resource allocation in single-parent households include Bargain et al. [2014], Echeverría et. (2019), Bose-Duker et al. [2021], and Lechene et al. [2022].

[2] Nieuwenhuis and Maldonado [2018] describe the "triple bind" of single-parent families: limited resources, barriers to stable and well-paid employment, and policy frameworks that often fail to offset these disadvantages. Although single mothers frequently achieve relatively high employment rates [Meyer and Sullivan, 2004], they remain overrepresented in precarious, low-wage jobs, and the lack of a second parent combined with atypical working hours exacerbates the challenge of balancing work and family responsibilities.

[3] In this paper, single parents are defined as follows: a) widowed, divorced, or separated parents with dependent children living in their household; b) a child is defined as any biological child of the household head aged 0 to 15 years. The expression "living in the household" denotes children who regularly spend at least four nights per week at the parent's residence. This definition ensures the inclusion of households where children reside primarily, which is critical for analyzing costs.

[4] The *two-child limit*, introduced in the United Kingdom in April 2017, restricts means-tested child-related benefits (Child Tax Credit and the child element of Universal Credit) to the first two children. For any additional child born after that date, no supplement is paid, implying a loss of about £244 per month (roughly £3,000 per year) per child [Chzhen and Bradshaw, 2025, Reader et al., 2025]. Adopted in a context of fiscal austerity, the reform was intended both to reduce social spending and to encourage families to finance the cost of a third child without public support. By 2023, around 1.5 million children were already affected (57% in working families and 50% in single-parent households), a figure expected to rise to nearly 2.8 million once the policy is fully phased in (HMRC statistics).

[5] Recent poverty statistics reveal that the number of children living in relative poverty (after housing costs) rose from 3.6 million in 2010 to 4.2 million by 2022. The report is available on the Child Poverty Action Group website.

hereafter BDH]. The identification strategy relies on the allocation of resources to children based on assignable goods, under the assumption of preference stability. The empirical approach also incorporates a set of exclusion restrictions to identify the intra-household sharing rule.

This study develops and estimates a static model of intra-household allocation to analyze how parental and child characteristics affect the distribution of resources within the household. The theoretical framework rests on three main components: (1) an additive utility function, positing parental altruism towards children; (2) a consumption technology describing how households transform purchased goods into individual consumption; and (3) a sharing rule determining the allocation of resources among household members, defined as the share of total household resources assigned to each individual.

To identify the sharing rule, I use clothing as the assignable good—that is, a good for which individual consumption is directly observable by the researcher within the household. This choice enables the analysis of how adult demand for such good varies with the presence and number of children in the household. Identification further relies on a parametric characterization of preferences: adults with similar observable characteristics are assumed to share similar preferences for assignable goods, independent of parental status. In other words, becoming a parent does not substantially alter individual preferences for clothing; rather, it is primarily the overall budget constraint and the resource allocation—affected by the arrival of children and associated expenditures—that may change. This assumption, widely discussed in the literature, offers a relevant framework for identifying Engel curves. See, e.g., Browning, Chiappori, and Lewbel [2013, hereafter BCL] and BDH. Moreover, it is also adopted by the benchmark studies —BDH— against which my results are compared to assess their generalizability to single parents. Finally, the empirical strategy relies on an exclusion restriction. Certain explanatory variables influence demand for assignable good but do not directly affect either the shadow price or the intra-household sharing rule. The model is estimated using a sample of one-adult households, with and without children, from the UK Family Expenditure Survey (FES) over the period 1978–2022.[6]

The results show that the collective model predicts individual resource shares satisfactorily, while revealing significant heterogeneity by parent gender and household size. Notably, the analysis indicates male-headed single-parent households allocate a higher share of resources to children than female-headed ones. Ceteris paribus, single fathers with one child allocate, on average, 21.6% of their total resources to the child, compared to 14.8% for single mothers. For comparison, BDH estimate that in two-parent households with one child, the share allocated to children is 14.2%. Regardless of the identification strategy, the collective approach accurately captures the effects of family size and child age on the child's share. Furthermore, family size exerts a significant influence on resource allocation among households in the lower vigintile of the income distribution, an effect that fades in higher-income households. This suggests that the *two-child limit* policy would likely worsen disparities and increase child poverty among low-income families.

This work makes four main contributions to the literature. First, the paper provides one of the few empirical estimates of the cost of children in single-parent households and, to my knowledge, the first based on data from a high-income country. By contrast, the limited existing literature—such as Bargain et al. [2014] for Côte d'Ivoire, Echeverria et al. [2019] for Argentina, and Bose-Duker et al. [2021] for Jamaica—focuses mainly on low- and middle-income countries with substantially different economic and social contexts. Moreover, the last two studies are based on the methodology developed by Dunbar, Lewbel, and Pendakur [2013, hereafter DLP], which does not require data on childless single adults and uses less restrictive homogeneity assumptions on preferences (Similarity Across

[6] Family Expenditure Survey (FES) has been replaced by Expenditure and Food Survey (EFS) in 2001, then Living Costs and Food Survey from 2008 onwards. For convenience, all three are referred to as FES in this paper.

Type—SAT—and Similarity Across People—SAP).[7] While the former —SAT— assumes that, for a given person type (man, woman, child), the slope of individual Engel curves does not depend on the number of children, the latter —SAP— posits that for a given household type, individual Engel curves have the same slope. The present study follows the approach of Bargain et al. [2014], who identify the intra-household sharing rule using information from adults without children. However, it departs significantly by exploiting longitudinal data over multiple years, whereas Bargain et al. [2014] rely solely on cross-sectional data, which compels them to assume "independence of base"—that is, that economies of scale are independent of household resource level. Finally, this study provides new empirical evidence on the applicability of the collective approach for estimating the sharing rule and suggests that, under certain conditions, child cost estimates from couple-based models remain transferable to single-parent households.

Second, it proposes several identification tests for the sharing rule and economies of scale. Unlike BDH, I exploit price variability in the sharing rule and show that the estimates are robust to alternative functional forms for both the sharing rule and economies of scale.

Third, it sheds new light on the heterogeneity of resource shares allocated to children by parental gender. The results indicate that, all else equal, single fathers allocate a greater share of their resources to their children than single mothers, regardless of the number of children. These findings contrast with the prevailing consensus in the literature, which emphasizes women's stronger child-oriented preferences. For example, Duflo [2003], Duflo and Udry [2004] and Caiumi and Perali [2015] find that women allocate a larger proportion of their resources to child-related goods (food, health, education), whereas men tend to spend more on adult-oriented goods, such as tobacco and alcohol. Similarly, Bose-Duker et al. [2021] observe, in the Jamaican context, that children in female-headed households receive a larger share of household resources than those in male-headed households. Complementary evidence from BDH indicates that, within couples, mothers allocate a larger share to children than fathers. This reversal may be driven by disparities in external support. Single mothers receive more frequent transfers, including child benefits and alimony, from public programs or nonresident parents. As a result, their personal spending on children represents a smaller share of total resources.

Finally, this paper advances the literature by uncovering the intra-household allocation mechanisms through which the *two-child limit* contributes to individual poverty and offers new evidence on its equity and effectiveness [González and Trommlerová, 2023, Chzhen and Bradshaw, 2025, Reader et al., 2025, Stewart et al., 2025a,b]. The analysis shows that, rather than a relief, this reform imposes a burden on large, low-income families. It further emphasizes the existence of a minimum expenditure threshold below which government intervention is crucial to ensure children's well-being, regardless of family size.

The remainder of the paper is organized as follows. Section 2 presents the theoretical framework and identification assumptions. Section 3 details the empirical strategy. Section 4 describes the data. Section 5 discusses the main results and their implications. Section 6 concludes.

2 Theoretical Framework

A central objective of this study is to evaluate whether sharing rule estimated from a collective model—originally developed for couples with or without children—remains valid when applied to single-parent households. To address this, I develop a static model of household consumption inspired

[7] A growing body of research applies and extends these approaches in various contexts. See, e.g., Tommasi and Wolf [2018], Tommasi [2019], Calvi [2020], Brown et al. [2021], Dunbar et al. [2019], Penglase [2021], Bargain et al. [2022b], Lechene et al. [2022], Lewbel and Pendakur [2022], Calvi et al. [2023], Lewbel and Pendakur [2024].

by BDH, enabling direct comparison with analyses of couples in the same UK data. The theoretical analysis proceeds in two steps: first, by characterizing the consumption behavior of a single adult (without children), which serves as the baseline; and second, by extending the model to single adults living with children.

2.1 The Consumption Behavior of a Single Adult without Children

Consider a household consisting of a childless single adult (either male or female), who maximizes a well-behaved utility function $u(x^a, x^c)$, where x^a denotes consumption of an assignable good (e.g., adult clothing) and x^c denotes a composite good.[8] The utility function is assumed to be twice continuously differentiable, strictly increasing, and strictly concave.[9]

Let p denote the price of the assignable good, and let the price of the composite good be normalized to unity. The individual faces the following budget constraint:

$$x^a p + x^c = y \tag{1}$$

where y is total household expenditure.[10]

The household's optimization problem is:

$$\max_{x^a, x^c} \quad u(x^a, x^c) \text{ subject to (1)} \tag{2}$$

The first-order conditions yield the demand for the assignable good as a function of prices and total expenditure:

$$\omega = g(p, y) \tag{3}$$

where $\omega = px^a/y$ is the budget share allocated to the assignable good. Since the utility function is strictly increasing, the budget constraint binds at the optimum.

2.2 The Consumption Behavior of a Single-Adult with Children

Now consider a single adult living with $n \geq 1$ children under 16.[11] In line with established convention, the parent is assumed to exhibit Beckerian altruism, such that utility is a function of both the parent's own consumption and the utility of the children.

The utility function is specified as:

$$W = u(x^a, x^c) + \delta(\mathbf{n})u_k(x^k) \tag{4}$$

where x^a and x^c are as before, x^k denotes the composite good consumed by a representative child (intra-sibling allocation is not identified), and $\delta(\mathbf{n})$ captures the weight placed by the parent

[8] Assignable good and exclusive good are used interchangeably. See BCL for more details about exclusive and assignable goods.

[9] Individual utility is further influenced by preference-driven factors, which are incorporated into the budget share function in the empirical section. Preferences are assumed to be stable, enabling predictions about household behavior.

[10] As is standard in the literature, only non-durable goods are considered. When the proportion of purchased goods that remains unconsumed is negligible, purchased quantities can be taken as a proxy for actual consumption.

[11] In contrast to Penglase [2021], the model treats foster and non-foster children indiscriminately. Penglase [2021] explicitly separates the two groups of children, focusing on whether there is differential treatment in the allocation of resources for the consumption of foster and non-foster children. At this point, no distinction is made regarding the characteristics of the children. My assumption is limited to the child residing with either the father or the mother and younger than 16 years old.

on child consumption, as a function of $\mathbf{n}$.[12] This formulation is consistent with Bargain and Donni [2012b].

For simplicity, household income is treated as exogenous and entirely allocated to consumption; neither labor supply nor home production is modeled.[13] Total expenditure y is allocated across purchased quantities q^a (exclusive good) and Q (composite good), subject to:

$$q^a p + Q = y \tag{5}$$

Here, q^a and Q denote, respectively, the purchased quantities of the household's exclusive goods and household composite goods.

Several clarifications are warranted. First, children's consumption is included in Q. Second, two types of goods are considered: an adult-exclusive good (x^a), and composite goods (x^c, x^k), the latter of which may comprise both private non-assignable and public components. Third, household survey data typically do not record individual consumption within household. As a result, information on composite goods provides limited insights into the allocation of resources to children. Nevertheless, observing adult-exclusive goods can reveal relevant aspects of household behavior.

For a purely private assignable good:

$$q^a = x^a \tag{6}$$

However, in a household with at least two members—such as an adult and a child—certain goods exhibit public or shared properties, such that their effective consumption cannot be directly inferred from purchased quantities alone. To address the resulting economies of scale, two main modeling approaches are used.

The Independence of Base (IB) approach posits that the cost savings from joint consumption of public goods are independent of both market prices and the total level of household expenditure. Under this assumption, the mapping from purchased quantities to effective consumption depends solely on household composition (e.g., number and type of members), but is invariant to economic variables. This restriction is appropriate in studies based on cross-sectional data with little or no price variation [Lewbel and Pendakur, 2008, Bargain and Donni, 2012a, Dunbar et al., 2013].

The Barten scales approach, in contrast, relaxes the IB restriction by allowing the transformation from purchases to effective consumption to depend on household composition as well as economic factors such as prices and total expenditure. Specifically, the vector of market prices is rescaled by a transformation function, which captures how economies of scale vary with total expenditure, prices, and household composition [BCL]. This approach, as in BCL and BDH, is particularly suited to analyses with multi-year data and substantial price variation.

Assumption 1 *(Barten prices). For any adult in a household with $n \geq 1$ children, there exists a differentiable scalar function $\pi(y, p, \mathbf{z}^\pi, \mathbf{n})$ such that composite good purchases satisfy:*

$$Q = \pi(y, p, \mathbf{z}^\pi, \mathbf{n})x^c + x^k \tag{7}$$

where $\pi(y, p, \mathbf{z}^\pi, \mathbf{n})$ is a shadow price (Barten price) associated with the parent's consumption of the composite good, and $\mathbf{z}^\pi$ is a vector of sociodemographic variables included in the Barten price. For identification, the shadow price for children is normalized to unity. The shadow price serves as a deflator that measures the gains from joint consumption experienced by an adult due to household

[12] Strictly speaking, the argument $\mathbf{n}$ in $\delta(\mathbf{n})$ is a vector capturing all relevant characteristics of the children (e.g., number of children, average age of children, gender composition, sibling structure, etc.).

[13] A broader perspective on this topic is addressed by Apps and Rees [2001] and Cherchye et al. [2012].

economies of scale. The parent effectively receives a greater quantity of composite goods than the amount actually purchased, due to sharing within the household (e.g., heating). When $\pi(.) = 1$, goods are purely private and one member's consumption does not affect the other's. When $\pi(.) < 1$, there are positive economies of scale, since each unit of x_c is effectively less costly. When $\pi(.) > 1$, this indicates diseconomies of scale or stronger preferences for private consumption.

Substituting (6) and (7) into (5) gives:

$$x^a p + \pi(y, p, \mathbf{z}^\pi, \mathbf{n}) x^c + x^k = y \tag{8}$$

The household chooses (x_a, x_c, x_k) to maximize utility subject to the new budget constraint (8). By construction, parental decisions are Pareto efficient, since the single parent makes all consumption decisions.[14]

The allocation problem reduces to dividing resources between the parent and the child. The resulting optimization program is:

$$\begin{aligned} &\max_{x^a, x^c, x^k} u(x^a, x^c) + \delta(\mathbf{n}) u_k(x^k) \\ &\text{s.t.} \quad x^a p + \pi(y, p, \mathbf{z}^\pi, \mathbf{n}) x^c + x^k = y \end{aligned} \tag{9}$$

Adopting an additive utility function simplifies the transition to a decentralized approach. The allocation can be characterized in two steps:

1. Let φ and φ_k denote the shares of resources allocated to the parent and to the child, respectively, with $\varphi + \varphi_k = 1$. The resource allocation is obtained by solving:

$$\max_{\varphi, \varphi_k} \quad \nu\left(\frac{p}{\pi}, y\frac{\varphi}{\pi}\right) + \delta(\mathbf{n}) \nu_k(y\varphi_k) \quad \text{s.t.} \quad \varphi + \varphi_k = 1 \tag{10}$$

   where ν and ν_k are the indirect utility functions for the parent and the child, respectively.

2. Given $\varphi(y, p, \mathbf{z}^\varphi, \mathbf{n}) \leq 1$ for $n \geq 1$, the parent solves:

$$\max_{x^a, x^c} \quad u(x^a, x^c) \quad \text{s.t.} \quad x^a p + \pi(y, p, \mathbf{z}^\pi, \mathbf{n}) x^c = y \cdot \varphi(y, p, \mathbf{z}^\varphi, \mathbf{n}) \tag{11}$$

   where $\mathbf{z}^\varphi$ represents the vector of sociodemographic variables entering the sharing rule, and $\varphi(y, p, \mathbf{z}^\varphi, \mathbf{n})$ denotes the share of the budget retained by the parent. Accordingly, the remaining share, $1 - \varphi(y, p, \mathbf{z}^\varphi, \mathbf{n})$, is assigned to the child. In the absence of children, $\varphi = 1$, meaning the parent keeps the entire budget, as in a childless single-adult household.

The parent's demand for the assignable good is thus:

$$\frac{\omega}{\varphi(y, p, \mathbf{z}^\varphi, \mathbf{n})} = g\left(\frac{p}{\pi(y, p, \mathbf{z}^\pi, \mathbf{n})}, y\frac{\varphi(y, p, \mathbf{z}^\varphi, \mathbf{n})}{\pi(y, p, \mathbf{z}^\pi, \mathbf{n})}\right) \tag{12}$$

[14] In the context of single-parent households, the Pareto efficiency assumption is automatically satisfied, as the parent acts as a dictator, making all consumption decisions on behalf of the household. More generally, the identification of intrahousehold resource allocations traditionally relies on the assumption of collective efficiency [Chiappori, 1988] and group separability—assumptions that are widely regarded as reasonable for everyday consumption choices [Baland and Ziparo, 2018]. While the literature has documented instances of inefficiency, particularly for strategic or production decisions—especially in low-income countries [Udry, 1996]—these assumptions remain plausible for repeated and non-strategic consumption choices, as also emphasized by Baland and Ziparo [2018] and Lewbel and Pendakur [2022].

where $\omega = px_a/y$ is the budget share devoted to the assignable good. Notably, explicit specification of the child's utility function is not required for estimation as it does not dictate the outcome of the model.

2.3 Identification

Estimating the cost of children would be straightforward if individual-level consumption data were available. In practice, however, most household surveys provide information only at the aggregate household level, and many goods are jointly consumed within the household. A few exceptions exist, such as the Dutch LISS panel exploited by Cherchye et al. [2012] and the Bangladeshi survey data used by Bargain et al. [2022b], which contain information on individual consumption. Yet these datasets remain rare and limited in scope. Consequently, the identification of both resource shares within the household and economies of scale in consumption requires the introduction of additional structure.

The identification strategy relies on three pillars. First, I assume preference stability across childless single adults and single parents, following BCL. This assumption should be interpreted as the invariance of preferences for assignable goods when moving from a single adult without children to a single parent with comparable socio-demographic characteristics. Thus, observed differences in assignable consumption can be attributed to household composition, not preference shifts. This assumption enables effective use of information from childless adults to estimate the demand function for single parents. This is especially valuable in the case of single fathers, given small sample size.

Second, identification requires the observation of at least one assignable good. While Chiappori and Ekeland [2009] show that three goods are typically needed for identification, Bourguignon [1999] and Bourguignon et al. [2009] demonstrate that under specific conditions, a single assignable good suffices. In line with recent contributions (e.g., BCL; Bargain and Donni, 2012a; DLP; BDH), I focus on clothing expenditures, which are observable at the individual level.[15]

Third, I exploit the nonlinearity of Engel curves. Following Prais and Houthakker [1971], nonlinear Engel curves, augmented with socio-demographic controls, provide sufficient flexibility to recover heterogeneous demand responses. In my empirical application, this requirement is satisfied since budget share equations are nonlinear in log expenditures.[16]

Finally, two normalization conditions are imposed. In households without children, both the shadow price of the composite good and the resource share allocated to the adult are normalized to one.

Conditional on the previous conditions, Proposition 1 states the additional conditions under which the sharing rule and Barten scales are generically identified.

Proposition 1 *Let the demand function for an assignable good be, respectively, for childless single adults and single parents:*

$$\omega = g(\mathbf{z}^{\omega}, p, y),$$
$$\omega = g(\mathbf{z}^{\omega}, \pi(y, p, \mathbf{z}^{\pi}, \mathbf{n}) \cdot p, \varphi(y, p, \mathbf{z}^{\varphi}, \mathbf{n}) \cdot y),$$

[15] Lechene et al. [2022] argue that food is a preferable assignable good relative to clothing, due to its larger budget share and consistently downward-sloping Engel curves. In this paper, however, I follow BDH in using clothing in order to ensure comparability with their results. This choice is further supported by Bargain et al. [2022b], who show that the collective model predicts individual resources reasonably well when using clothing.

[16] See also Banks, Blundell, and Lewbel [1997], who show that budget share equations are typically nonlinear.

where $\pi(y, p, \mathbf{z}^{\pi}, \mathbf{n})$ represents the price transformation à la Barten, and $\varphi(y, p, \mathbf{z}^{\varphi}, \mathbf{n})$ denotes the sharing rule. Here, p is the price of the assignable good, y represents total expenditures, $\mathbf{n}$ is a vector of children characteristics, and $\mathbf{z}^{\omega}, \mathbf{z}^{\pi}, \mathbf{z}^{\varphi}$ are vectors of sociodemographic variables associated with ω, π, and φ, respectively.

The functions $\pi(\cdot)$ and $\varphi(\cdot)$ are generically identifiable if at least one of the following conditions holds:

1. *At least one variable in $\mathbf{z}^{\omega}$ is excluded from both $\mathbf{z}^{\pi}$ and $\mathbf{z}^{\varphi}$.*
2. *π and φ are independent of total expenditure y.*
3. *π and φ are independent of price p.*
4. *$\pi(\cdot)$ and $\varphi(\cdot)$ are known up to some parameters (semi-parametric identification).*
5. *$\pi(y, p, \mathbf{z}^{\pi}, \mathbf{n}) = \pi_1(y, p, \mathbf{z}^{\pi}, \tilde{\mathbf{n}}) \cdot \pi_2(n)$ with $\pi_2(1) = 1$, and $\varphi(y, p, \mathbf{z}^{\varphi}, \mathbf{n}) = \varphi(y, p, \mathbf{z}^{\varphi}, \tilde{\mathbf{n}}) \cdot \varphi_2(n)$ with $\varphi_2(1) = 1$.*

The proof is given in Appendix A. Here, the economic intuition is as follows:

1. $\mathbf{z}^{\omega}$ denotes a vector of variables that affect the demand for the assignable good ω but do not necessarily influence the shadow price π or the sharing rule φ. These variables may include socio-demographic characteristics such as age, education, or region of residence. The presence of a variable in $\mathbf{z}^{\omega}$—for example, decade dummies—that impacts only the demand for the assignable good while leaving π and φ unaffected provides an exclusion restriction. Exploiting variation in such a variable allows one to identify its direct effect on demand, thereby isolating the estimation of the model's structural parameters while holding the shadow price and sharing rule constant. For instance, decade dummies capture changes in fashion trends, relative prices of clothing, or shifts in consumption norms that affect adult clothing expenditures, without necessarily altering how resources are shared between parents and children or the degree of household economies of scale. Hence, variation across decades provides a valid exclusion restriction for identifying the demand equation.[17]

2. In this case, the shadow price π and sharing rule φ do not change with total expenditures y. That is, whether the household's total expenditure is £1000 or £2000, π and φ remain unchanged. If empirical evidence shows that the allocation of resources between the parent and the child does not vary with changes in total expenditure, then it is unnecessary to model π and φ as functions of y. This invariance greatly facilitates identification.[18]

[17] In the empirical analysis, this exclusion restriction is also employed to identify the sharing rule.

[18] One of the alternative to identify the sharing rule is the IB assumption made by DLP. Essentially, DLP demonstrate that the sharing rule within a household can be recovered using distribution factors, provided it does not depend on total expenditure. In other words, the share an individual receives within the household is independent of the household's overall consumption. However, they allow it to depend on closely related factors such as wealth. The crucial assumption here is that resource shares do not depend on total expenditure. This assumption, while essential for identification, cannot be tested within their framework. Cherchye, Rock, Lewbel, and Vermeulen [2015] propose a method that allows for testing this assumption and arrive at a surprising conclusion: resource shares do not depend on the household's full income. Full income is defined as the sum of both spouses' maximum potential labor income and non-labor income (excluding savings and spending on durables), with leisure factored in. Although DLP do not directly test their IB assumption, they illustrate their methodology by examining the impact of access to credit on within-household consumption allocation in Malawi. Their findings reveal that the effect of credit on resource shares varies significantly based on the type of credit—microcredit and agricultural credit tend to reduce resources allocated to children—and on the recipient of the credit. Specifically, credit received by women tends to shift resources away from men and towards children and women themselves.

3. Suppose the price of the assignable good increases, but this does not affect how the parent allocates their budget between themselves and their child, as captured by the sharing rule (φ). In such a case, both π and φ remain unchanged, allowing for their separate estimation, independent of the effect of price changes on demand.

4. If the functional forms of π and φ are known up to some parameters, identification reduces to estimating these unknown parameters rather than specifying the entire functional relationship. For example, if it is known that φ depends on the number of children, but the exact magnitude of this dependence is unknown, one might assume a specific parametric structure, such as $\varphi = \psi(n)$, where ψ could be a simple linear or logistic function. The empirical task then becomes estimating the parameters that govern how φ varies as the number of children increases.

5. Assuming that π and φ can be factorized into two multiplicative components significantly simplifies their structure. This specification allows the effects of household composition (e.g., the number of children) to be separated from those of prices and sociodemographic characteristics. For example, one might express π as $\pi_1(y, p, \mathbf{z}^\pi, \tilde{\mathbf{n}}) \cdot \pi_2(n)$, where π_1 captures the influence of prices and demographics, and $\pi_2(n)$ isolates the effect of the number of children. Such a formulation enables the estimation of the respective impacts independently.

3 Empirical Implementation

3.1 Econometric Specification

The empirical specification examines a demand system which is quadratic in logarithmic expenditure, as used in studies such as Browning et al. [1994] and BDH. This quadratic parametrization relaxes the restrictive linearity assumption under which marginal budget shares are invariant to the level of expenditure. The empirical model is specified as follows:

$$\omega_i = \alpha_i \mathbf{z}_i^\omega + \beta_i \ln p_i + \gamma_i \ln y_i + \eta_i (\ln y_i)^2 + \epsilon_i \tag{13}$$

for $i = f, m$, where $\alpha_i,\ \beta_i,\ \gamma_i,$ and η_i are the parameters to be estimated. The vector $\mathbf{z}_i^\omega$ is a linear function of a set of covariates, including a constant, education level, adult age and its square, a set of dummies for labor force participation, home ownership, decades, and region of residence, and ϵ_i, an error term capturing unobserved heterogeneity.[19] Notably the equations are gender-specific, with separate estimations for men and women. For each adult, the budget share devoted to clothing is defined as the ratio of weekly clothing expenditures to total weekly expenditures on non-durable goods.

For childless single adults, the equation simplifies from (12) to (3). To distinguish between single parents and childless adults, I introduce a dummy variable, $\mathbb{1}_i$, which equals 1 if the adult is a parent and 0 otherwise. The stochastic structure of the budget share equations for single adults and single parents is then expressed as follows:

$$\text{If} \quad \mathbb{1}_i = 0, \quad \text{then} \quad \epsilon_i = \omega_i - \alpha_i \mathbf{z}_i^\omega - \beta_i \ln p_i - \gamma_i \ln y_i - \eta_i (\ln y_i)^2 \tag{14}$$

$$\text{If} \quad \mathbb{1}_i = 1, \quad \text{then} \quad \epsilon_i = \frac{\omega_i}{\varphi_i} - \alpha_i \mathbf{z}_i^\omega - \beta_i \ln\left(\frac{p_i}{\pi_i}\right) - \gamma_i \ln\left(\frac{\varphi_i y_i}{\pi_i}\right) - \eta_i \left[\ln\left(\frac{\varphi_i y_i}{\pi_i}\right)\right]^2 \tag{15}$$

[19] Three decade dummies are included, corresponding to the periods 1991–2000, 2001–2010, and 2011–2022. Regional dummy variables are defined according to the classification detailed in the sample selection subsection 4.1.

The sharing rule is modeled as a deterministic logistic function of observable parental characteristics ($\mathbf{z}^{\varphi}$), child characteristics ($\mathbf{n}$), total expenditures, and prices.[20] Unlike some earlier contributions (Bourguignon, 1999 and BDH), the specification does not rely on a Taylor expansion for linearization:

$$\begin{aligned} \varphi_i(y,p,\mathbf{z}^{\varphi},\mathbf{n}) &= \frac{e^{\psi_i(y,p,\mathbf{z}^{\varphi},\mathbf{n})}}{1+e^{\psi_i(y,p,\mathbf{z}^{\varphi},\mathbf{n})}}, \\ \text{with} \quad \psi_i(y,p,\mathbf{z}^{\varphi},\mathbf{n}) &= \mathbf{z}_i^{\varphi\prime}\theta_z + \theta_y \ln y_i + \theta_p \ln p_i + \mathbf{n}_i{}'\theta_n \end{aligned} \tag{16}$$

where θ_z, θ_y, θ_p, and θ_n are vectors of parameters. Identification is achieved using the first condition of Proposition 1, with decade dummies serving as exclusion restrictions in $\mathbf{z}^{\varphi}$. The parental covariates $\mathbf{z}^{\varphi}$ include a constant, education, age, a set of dummies for labor force status, home ownership, and region of residence. The vector of child-related variables $\mathbf{n}$ encompass the number of children and its square, the average age of the children, the proportion of boys, and a dummy for the presence of same-gender siblings, with the latter three interacted with the number of children.

It is assumed that children exert no bargaining power in the household decision process. Resource allocation is determined entirely by the parent. Nonetheless, the sharing rule links parental and child characteristics to the intrahousehold distribution of resources. For example, older children are expected to receive a larger share of household expenditures due to higher consumption needs.

To test Point V of Proposition 1, an alternative specification is also considered in which the vector of child characteristics excludes the number of children. In this case, the number of children affects the sharing rule solely through a multiplicative scale factor $n^{-\rho}$.

$$\varphi_i(y,p,\mathbf{z}^{\varphi},\mathbf{n}) = \frac{e^{\psi_i(y,p,\mathbf{z}^{\varphi},\tilde{\mathbf{n}})}}{1+e^{\psi_i(y,p,\mathbf{z}^{\varphi},\tilde{\mathbf{n}})}} n^{-\rho} \tag{17}$$

Following BDH, shadow prices are specified as functions of both total expenditures and the number of children, capturing potential economies of scale:

$$\pi_i(p,y,\mathbf{n}) = \kappa_i \ln y_i \sqrt{n_i} \tag{18}$$

where κ_i is a parameter to be estimated. This specification effectively combines points (1), (3), and (5) of Proposition 1. For simplicity, the intercept is assumed to be zero.[21]

Relevance of decade dummies as excluded shifters. To justify the use of decade dummies as excluded shifters in the reference specification, I verify their relevance, namely, that these time indicators predict clothing demand once the influence of other covariates has been partialled out. Using the Frisch–Waugh–Lovell decomposition, the residualized outcome $M_Z\omega$ is regressed on the residualized decade dummies $M_Z D$, with inference based on a joint Wald test and standard errors clustered by year, the unit of variation of the shifters. The results, reported in Table 1, reject $H_0 : \delta = 0$ at conventional levels for both genders ($F(3,44) = 4.23$, p-value $= 0.010$ for men; $F(3,44) = 3.80$, p-value $= 0.017$ for women), implying that decade-level variation explains a significant share of residualized demand movements.

As shown in Appendix B, adjusted contrasts relative to the 2011–2022 baseline point to a sizeable dip in 1991–2000 for men (≈ -0.40 percentage point) and a marked peak in 2001–2010 for

[20] Recent work by Dunbar et al. [2019] develops a framework for identifying random resource shares, defined as varying across observationally identical households.

[21] As a robustness check, I also estimate the Barten price function including an intercept.

women ($\approx +0.52$ percentage point). The decade-to-decade differences are not merely random noise. Men in 2001–2010 exhibit about 0.67 percentage point higher residualized clothing demand than in 1991–2000 (p-value = 0.019), while the same comparison for women yields a gap of 0.43 percentage point (p-value = 0.032). The shape of the annual means of $M_Z\omega$ mirrors these findings, with a pronounced trough for men in the 1990s and a clear peak for women in the early 2000s, visually corroborating the statistical evidence and reinforcing the interpretation that decade variation is an economically meaningful and empirically strong source of identifying variation.

Table 1: Joint tests by gender

Sex	F	(df$_1$, df$_2$)	*p-value*	Partial R^2 (%)
Women	3.80	(3, 44)	0.0166	20.57
Men	4.23	(3, 44)	0.0104	22.37

Notes: This table reports the results of joint significance tests for residualized decade indicators in the Frisch–Waugh–Lovell regression of the residualized clothing budget share $M_Z\omega$ on $M_Z D$, where D includes the three decade dummies (1978–1990, 1991–2000, 2001–2010). Standard errors are clustered at the year level, which corresponds to the unit of variation of the excluded shifters. The 3-df F-test assesses the conditional relevance of the temporal shifters, and the partial R^2 measures the share of residual variance explained by the decade block. Results indicate that, for both women, decade variation explains a substantial portion of residual demand movements conditional on controls, thereby satisfying the relevance condition in the identification strategy.

3.2 Estimation Strategy and Instruments

The demand equations are estimated separately for men and women, reflecting the focus on unpartnered adults. The structural model, which jointly specifies preferences, the sharing rule, and the shadow price, is estimated in a single step for both childless adults and single parents. Identification is then achieved through the preference stability assumption.

The model addresses two main sources of endogeneity that may bias the estimated budget shares. The first arises from measurement error in total expenditures, which may reflect the infrequency and potential under-reporting of clothing purchases, as well as recall errors in household survey. Such measurement error may induce correlation between observed total expenditure and the error term in the budget share equation, leading to biased estimates. Following DLP, total income is used as an instrumental variable for total expenditures. Total income is uncorrelated with the consumption allocation error within a given time period, but correlated with total expenditures, making it a valid instrument. Total income also serves to address endogeneity stemming from recall errors, as long as income measurement errors are orthogonal to consumption recall errors, and income remains correlated with total expenditures.

Second, the literature documents that marital expectations may influence fertility choices, creating endogeneity in the number of children [Nakamura and Nakamura, 1992, Apps and Rees, 2001].[22]

[22] Single women can easily have children without resorting to adoption or assisted reproductive technologies. Consequently, single parents and childless individuals may have fundamentally different preferences regarding children, which undermines the assumption of stable preferences across marital statuses. This selection issue presents a challenge to inferring single parents' preferences from those of observably similar childless singles.

To mitigate this bias, the econometric model includes a decent set of sociodemographic controls. Moreover, recent empirical evidence has shown that predicted individual resource shares—especially those derived from assignable goods such as clothing—are reliable under the preference stability assumption [Bargain et al., 2022b].

To set the instruments suitably, I write the budget share equations (14) and (15) as a unique budget share equation. To do this, multiply equation (14) by $(1 - \mathbb{1}_i)$ if childless single adult and equation (15) by $\mathbb{1}_i$ if single parent to obtain:

$$\epsilon_i = (1-\mathbb{1}_i)\Big[\omega_i - \alpha_i \mathbf{z}_i^{\omega} - \beta_i \ln p_i - \gamma_i \ln y_i - \eta_i (\ln y_i)^2\Big] + \mathbb{1}_i \Bigg[\frac{\omega_i}{\varphi_i} - \alpha_i \mathbf{z}_i^{\omega} - \beta_i \ln\left(\frac{p_i}{\pi_i}\right) - \gamma_i \ln\left(\frac{\varphi_i y_i}{\pi_i}\right) - \eta_i \left(\ln\left(\frac{\varphi_i y_i}{\pi_i}\right)\right)^2\Bigg]$$

Rearranging the right-hand side gives:

$$\omega_i = \alpha_i \mathbf{z}_i^{\omega} + \beta_i \ln p_i + \gamma_i \ln y_i + \eta_i (\ln y_i)^2 + \mathbb{1}_i A_i + \epsilon_i \tag{19}$$

with

$$A_i = \beta_i \ln\left(\frac{1}{\pi_i}\right) + \ln\left(\frac{\varphi_i}{\pi_i}\right)\left[\gamma_i + \eta_i \ln\left(\frac{y_i^2 \varphi_i}{\pi_i}\right)\right] - \omega_i \frac{1-\varphi_i}{\varphi_i}.$$

Then, I estimate the budget share equations by setting the iterated Two Stage Least Square Method. The nonlinear estimators are iterated until the estimated parameters and error/orthogonality condition covariance matrices settle.

I use all the exogenous variables as instruments, except total expenditures which are instrumented using total income. In addition, the set of instruments includes the product of $\mathbb{1}_i$ and a second-order polynomial in all the exogenous variables included in A_i. This yields a total of 38 instruments per equation.

To obtain suitable starting values, the demand equations are first estimated on the sub-sample of childless adults. These initial estimates are then used as starting points for the estimation on the full sample (with and without children). Furthermore, the cross-equation covariance matrix from the first-stage estimation is used as the initial weighting matrix for the iterative optimization algorithm. This approach improves convergence and robustness by explicitly accounting for the correlation structure of the residuals across equations.

4 Data

4.1 Sample Selection

The analysis draws on harmonized data from the United Kingdom's Family Expenditure Survey (FES) for the period 1978–2022, including its successors, the Expenditure and Food Survey (EFS, from 2001) and the Living Costs and Food Survey (LCF, from 2008).[23] These surveys provide detailed socio-economic information on households, including income, expenditure structure, and

[23] These surveys have been previously used by Lise and Seitz [2011] and BDH.

regional location.[24]

The initial sample includes 136,960 households. The analysis is restricted to single adults, both with and without children, aged 18 to 55, excluding individuals with outlier values, negative total expenditures, or missing information. The final sample consists of 40,403 households: 14,070 single men, 10,911 single women, 1,579 single fathers, and 13,843 single mothers. Notably, single fathers account for only 10% of single-parent households. Among parents, 58% of single fathers and 51% of single mothers have only one child.

Table 2: Descriptive statistics from the FES 1978-2022: Single adults and single parents

		Single		Single Mother			Single Father		
		Women	Men	Children			Children		
				1	2	3	1	2	3
Expenditure data									
Female clothing	Weekly expenditure (in £)	9.51	-	7.58	6.20	5.26	-	-	-
		(18.47)	-	(15.19)	(13.44)	(11.42)	-	-	-
	Percentage of zeros	0.43	-	0.44	0.47	0.48	-	-	-
		(0.49)	-	(0.50)	(0.50)	(0.50)	-	-	-
Male clothing	Weekly expenditure (in £)	-	5.36	-	-	-	4.42	3.67	1.61
		-	(15.73)	-	-	-	(11.99)	(10.63)	(4.85)
	Percentage of zeros	-	0.72	-	-	-	0.70	0.71	0.82
		-	(0.45)	-	-	-	(0.46)	(0.45)	(0.39)
Total weekly expenditure		117.90	125.19	137.19	144.18	144.12	160.43	162.72	156.66
		(84.51)	(94.94)	(95.60)	(96.69)	(93.19)	(105.18)	(106.00)	(87.86)
Individual and household characteristics									
Women's labor participation		0.72	-	0.50	0.43	0.29	-	-	-
		(0.45)	-	(0.50)	(0.50)	(0.45)	-	-	-
Men's labor participation		-	0.65	-	-	-	0.55	0.51	0.41
		-	(0.48)	-	-	-	(0.50)	(0.50)	(0.49)
Women's education (in years)		12.51	-	11.78	11.62	11.31	-	-	-
		(3.42)	-	(2.45)	(2.26)	(2.06)	-	-	-
Men's education (in years)		-	12.35	-	-	-	11.36	11.48	11.56
		-	(3.44)	-	-	-	(2.20)	(2.25)	(2.33)
Women's age		39.23	-	35.00	34.03	33.44	-	-	-
		(11.13)	-	(9.20)	(7.06)	(5.92)	-	-	-
Men's age		-	38.43	-	-	-	38.66	37.54	35.99
		-	(10.21)	-	-	-	(9.12)	(7.86)	(7.02)
House owner		0.52	0.49	0.28	0.28	0.19	0.46	0.45	0.31
		(0.50)	(0.50)	(0.45)	(0.45)	(0.39)	(0.50)	(0.50)	(0.46)
Average age of children		-	-	7.84	7.88	7.86	9.03	8.26	8.18
		-	-	(4.83)	(3.73)	(3.08)	(5.18)	(3.99)	(3.24)
Proportion of boys		-	-	0.52	0.50	0.51	0.58	0.53	0.56
		-	-	(0.50)	(0.35)	(0.30)	(0.49)	(0.34)	(0.30)
Number of observations		10974	14096	7135	4669	1579	915	486	133

Notes: Expenditures are in 1987 pounds. Standard deviations are in parentheses.

4.2 Descriptive Statistics

The demand system includes two assignable goods (men's and women's clothings) alongside a composite good, which represents all other goods to ensure total budget shares sum to one. Prices are measured annually at the national level.

Control variables include years of completed education, age, employment status, home ownership, region of residence (12 regions), number of children (and its square), average age of children, proportion of boys, a dummy for the presence of same-gender siblings, four decade time dummies, annual prices, and weekly total expenditures.

The descriptive statistics in Table 2 reveal a substantial decline in adult clothing expenditures in the presence of children: from £9.6 to £7.6 for women, and from £5.3 to £4.4 for men, with further

[24] The twelve regions of Great Britain are: Northern, Northern Ireland, York and Humberside, East Midlands, West Midlands, East Anglia, Greater London, South-East, North Western, South Western, Wales, and Scotland.

reductions as the number of children increases (down to £2.1 for fathers and £5.2 for mothers). These results echo the Rothbarth's view, which posits that household size reduces parental well-being as measured by individual consumption. The high incidence of zero clothing expenditures supports the use of total income as an instrument for total expenditure, given the infrequency of clothing purchases [Keen, 1986].

5 Estimation Results and Discussion

5.1 Resource Share Equations

A key focus of this study is the effect of children on parental resource shares. As discussed, φ_i represents the share of total expenditures retained by the parent, while $\varphi_k = 1 - \varphi_i$ corresponds to the share allocated to the children. A negative coefficient in the sharing function indicates an increase in child-related resource allocation. Table 3 presents the results corresponding to the various identification restrictions discussed in Proposition 1. For each sample group, column (I) reports the estimates from the benchmark model, where φ_i is identified using decade dummies as exclusion restrictions. Columns (I) and (II) test the first point in Proposition 1, under the assumption that decade dummies do not directly enter the sharing rule. Column (I) uses the baseline sharing rule 16, while Column (II) adopts the alternative specification 17. Column (III) jointly evaluates restrictions II and V of Proposition 1, using exogenous variation in prices. Columns (IV) and (V) implement two distinct strategies to jointly test identification restrictions II and V as outlined in Proposition 1. In column (V), I augment the demand specification by including a dummy variable—entered in both $\mathbf{z}_i^{\omega}$ and φ_i—that equals one for households whose total expenditure exceeds the sample median. Finally, column (VI) examines the SAT hypothesis of DLP in conjunction with restriction V of Proposition 1.

The results presented in column (I) indicate that the presence of children has an augmenting effect on parental resources, for both fathers and mothers. Specifically, the negative sign associated with the number of children implies that the marginal cost of children rises significantly with family size, while the resources allocated per child decline as family size increases. This pattern is robust across all model specifications. The estimated coefficient ρ is positive and remains stable across specifications, indicating that the per-child cost declines with additional children. These findings are consistent with previous research, including Bargain and Donni [2012a], DLP, Penglase [2021], and BDH. In addition, the results suggest that older children are associated with higher parental costs. While most coefficients related to children's characteristics for fathers are not statistically significant, the signs are generally consistent with those observed for mothers.

As expected, the parameter associated with same-gender siblings is positive. Specifically, the coefficient indicates that mothers retain a larger share of total expenditures when the household includes siblings of the same gender, suggesting the presence of economies of scale. A similar pattern is observed for fathers, although this coefficient is not statistically significant. For example, same-gender siblings who are close in age often share clothing, thereby reducing per-child expenditures. This variable therefore captures both the effects of family size and the influence of sibling gender composition.

Table 3: Estimated parameters of the individual resource shares – φ_i – and individual prices – π_i

	Women						Men					
Parameters	I	II	III	IV	V	VI	I	II	III	IV	V	VI
Parent variables												
$\theta_{intercept}$	2.265***	1.669**	4.943**	−3.160	8.595**	0.341	1.238	−0.958	−6.374	6.773	−128.882	−5.346
	(0.697)	(0.726)	(2.078)	(4.634)	(4.290)	(1.532)	(2.007)	(2.711)	(17.245)	(9.246)	(132.200)	(7.069)
$\theta_{education}$	−0.044	−0.063*	−0.090*	−0.130**	−0.061	0.074	−0.027	−0.022	4.936*	−0.073	−3.356	0.101
	(0.028)	(0.034)	(0.050)	(0.052)	(0.050)	(0.072)	(0.138)	(0.185)	(2.668)	(0.146)	(3.523)	(0.486)
θ_{age}	0.018	0.023	0.021	0.017	0.025	0.076	0.058	0.090*	0.699	0.071	0.735	0.221
	(0.014)	(0.016)	(0.017)	(0.015)	(0.018)	(0.059)	(0.036)	(0.052)	(0.428)	(0.048)	(0.921)	(0.160)
$\theta_{labor\,participation}$	−0.066	0.000	−0.061	0.259	−0.109	0.486	0.460	0.440	0.476	0.378	0.444	0.672
	(0.176)	(0.214)	(0.242)	(0.221)	(0.242)	(0.446)	(0.519)	(0.540)	(6.056)	(0.421)	(1.852)	(0.766)
$\theta_{house\,owner}$	−1.340**	−1.435*	−1.710**	−0.923	−1.621***	–	−0.863	0.065	−43.234*	0.632	−3.647	–
	(0.674)	(0.736)	(0.743)	(0.669)	(0.594)	–	(1.646)	(2.339)	(22.490)	(2.209)	(6.740)	–
Financial variables												
θ_{lnp}	−1.010***	−1.275***	–	−11.698	5.490	−2.388*	−1.010	−0.765	–	5.724	−474.166	−2.044
	(0.289)	(0.338)	–	(7.456)	(3.470)	(1.326)	(0.807)	(0.950)	–	(10.365)	(481.400)	(2.716)
θ_{lny}	0.568	0.801	1.304*	0.646	–	1.465	2.168*	1.762	66.749**	1.805	–	2.595
	(0.721)	(0.772)	(0.786)	(0.775)	–	(1.125)	(1.182)	(1.430)	(33.791)	(1.417)	–	(1.898)
$\theta_{dummy\,for\,y}$	–	–	–	–	1.934*	–	–	–	–	–	20.966	–
	–	–	–	–	(0.992)	–	–	–	–	–	(22.186)	–
Child variables												
θ_{n}	−0.906***	–	–	–	–	–	−1.277*	–	–	–	–	–
	(0.193)	–	–	–	–	–	(0.713)	–	–	–	–	–
θ_{n^2}	0.104***	–	–	–	–	–	0.146	–	–	–	–	–
	(0.032)	–	–	–	–	–	(0.127)	–	–	–	–	–
$\theta_{kid\,age\times n}$	−0.018*	−0.058**	−0.067**	−0.065***	−0.071**	−0.136	−0.002	−0.063	2.032	−0.055	−0.080	−0.138
	(0.011)	(0.023)	(0.026)	(0.024)	(0.029)	(0.089)	(0.024)	(0.059)	(1.261)	(0.052)	(0.395)	(0.168)
$\theta_{proportion\,of\,boys\times n}$	−0.094	−0.253*	−0.243*	−0.333**	−0.245*	−0.520	−0.112	0.280	2.809	0.019	1.856	0.285
	(0.059)	(0.131)	(0.143)	(0.163)	(0.144)	(0.363)	(0.217)	(0.531)	(6.504)	(0.522)	(2.992)	(1.026)
$\theta_{kid\,gender\times n}$	0.077*	0.333*	0.373*	0.318	0.303	0.713	0.109	0.508	−12.173	0.398	2.648	−0.302
	(0.045)	(0.184)	(0.216)	(0.197)	(0.199)	(0.663)	(0.159)	(0.638)	(9.613)	(0.526)	(2.725)	(1.229)
ρ	–	0.341***	0.347***	0.335***	0.265***	0.309***	–	0.400***	0.498***	0.395***	0.335*	0.287**
	–	(0.042)	(0.047)	(0.042)	(0.046)	(0.030)	–	(0.126)	(0.115)	(0.134)	(0.172)	(0.118)
Shadow prices												
κ	−0.449**	−0.336*	−0.429**	−0.158	−0.615*	0.035	0.832*	0.722*	0.659	1.108	−0.371*	0.288
	(0.215)	(0.173)	(0.206)	(0.100)	(0.335)	(0.117)	(0.464)	(0.421)	(0.425)	(0.900)	(0.204)	(0.292)
Number of observations	24754	24754	24754	24754	24754	24754	15649	15649	15649	15649	15649	15649
(Number of free parameters, Instruments)	(47,59)	(46,58)	(48,58)	(49,58)	(50,58)	(35,58)	(47,59)	(46,58)	(48,58)	(49,58)	(50,58)	(35,58)

Notes: * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$. Standard errors are in parentheses. The reported coefficients correspond to the estimated parameters from the internal specification of φ_i, as defined within the budget share equation. The dependent variable is the budget share devoted to adult clothing expenditures. Results are derived from simultaneous estimation of demand functions for individuals with (15) and without (14) children using the Two-Stage Least Squares (2SLS) method, under the assumption of preference stability. Demand systems are estimated separately by gender. For brevity, parameter estimates for regional dummy variables are omitted from the table. In each group, column (I) reports results for the benchmark model. Columns (I) test point I of Proposition I, using decade dummies as exclusion restrictions. Columns (II)–(VI) provide robustness checks, all imposing restriction V of Proposition I. Specifically, columns (II) and (III) test, respectively, restrictions I and II. Columns (IV) and (V) provide two alternative approaches to jointly assessing identification restrictions II and V from Proposition 1. In column (V), the demand function is estimated with the inclusion of an additional dummy variable in both $\mathbf{z}_i^{\omega}$ and φ_i. This dummy takes the value one for households with total expenditure above the sample median. Column (VI) examines the SAT assumption.

5.2 The *Two-Child Limit*: Blessing or Burden?

Given the variability in family expenditures, Figure 1 displays predicted per-child resource shares across the distribution of household expenditures, partitioned into twenty vigintiles. Focusing on the first panel, resource shares per child differ substantially at the bottom of the expenditure distribution. In the lowest vigintile, families with one child allocate approximately 25% of household resources, compared to around 20% and 18% for families with two and three children, respectively. This pattern indicates that, among low-expenditure households, having fewer children is associated with a higher per-child allocation of resources. However, as total household expenditures increase, per-child resource shares for all family sizes decline and converge. In the top vigintile, regardless of the number of children, the resource share per child approaches approximately 8%. This convergence indicates that, in higher-expenditure households, family size has little impact on the per-child allocation of resources, resulting in a more uniform distribution of resources per child across households of different sizes.

This figure conveys several important insights. First, it reveals a strong association between family size and per-child resource shares among low-expenditure households, with children in larger families receive systematically lower allocations. This pattern points to the existence of a critical region in the expenditure distribution, where intra-household trade-offs become particularly binding. The observed non-linear relationship suggests that, below a certain level of resources, budget constraints prevent larger families from allocating as much per child as smaller families with comparable means. In other words, at low levels of household expenditure, sibship size becomes a key determinant of intra-household allocation. For instance, consider a single mother earning £1,500 per month, with fixed subsistence needs of £1,400 and no access to family benefits. With one child, the remaining £100 can be fully allocated to that child. With two or more children, however, that same £100 must be divided among them, as the parent's consumption cannot adjust further. These findings underscore the potential importance of targeted policy interventions for households operating near this critical zone, where the burden of resource dilution is most pronounced.

Second, the figure shows that among higher-expenditure households, children receive nearly equal resource shares regardless of family size. In these families, the marginal impact of additional siblings on each child's material well-being appears negligible. This convergence in resource shares is consistent with the potential existence of a minimum level of household total expenditures that does not vary with family size.

In this regard, the findings suggest that policies such as the *two-child limit* may fail to achieve the objectives of equity and child welfare. Rather than imposing a blanket limit, a more effective policy would involve means-tested support that takes both family size and household resources into account.

5.3 Intra-household Resource Allocation

Table 4 reports point estimates of children's resource shares evaluated at the sample-mean covariates. Column (I) presents results under the baseline sharing rule specification, while Column (II) reports estimates from the alternative specification. Specifically, the model implies that a representative single mother with one child allocates 14.8 percent of resources to the child, whereas a representative single father allocates 21.6 percent.[25] For comparison, BDH estimate that a representative two-parent household with one child devote 14.2 percent of resources to that child. Given the

[25] Throughout, a representative single parent refers to a prediction obtained with all covariates set to their sample means (prediction-at-the-mean). Results are similar when averaging household-specific predictions over the sample distribution of covariates (average-of-predictions).

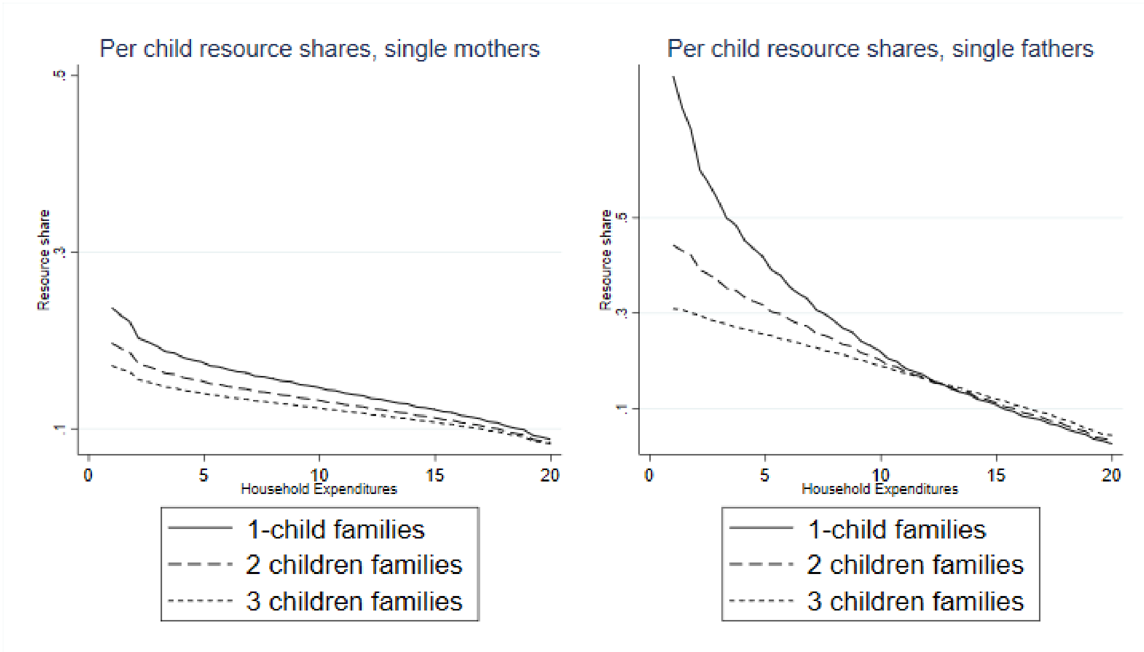

Figure 1: Children resource share by total expenditures

Note: This figure plots the predicted per-child resource shares allocated by parents across different points of the household expenditure distribution. The x-axis shows the distribution of total household expenditures divided into 20 vigintiles, ranging from the 1st to the 20th. The y-axis displays the per-child resource shares for mothers (left panel) and fathers (right panel). The solid line represents households with one child, the dashed line indicates households with two children, and the densely dashed line corresponds to households with three children. Both panels show that family size is a key determinant of children shares in low-income families, but not in high-income families.

overrepresentation of single mothers in the sample, comparisons with the couple estimates should primarily be made for this group. In this respect, the results suggest that, at least for mothers, couple-based estimates provide an externally valid reference point for single-mother households.

In contrast, single fathers appear to devote a substantially larger share of resources to children than single mothers—a pattern that reverses the gender gap observed in couple households, where mothers typically allocate more to children [BDH]. Figure 2 further shows that, for every family size considered, single fathers spend, on average, more on children than single mothers.[26] One plausible explanation for the higher child-related resource share observed among single fathers lies in differences in external support. Single mothers are considerably more likely to receive external transfers, such as child benefits or alimony, from governments or nonresident parents. In the data, 38% of fathers report receiving no child benefits at all, compared with only 3% of mothers, indicating a markedly lower coverage of transfers among fathers. Moreover, conditional on receiving alimony, single mothers obtain on average £80 more per week than single fathers. Such transfers reduce reliance on in-household resources for child-related needs, resulting in a lower measured share of mothers' resources devoted to children. Correspondingly, these differences in both the incidence and the level of external transfers are consistent with single fathers relying more heavily on their own resources to finance child-related expenditures. Another possible explanation is that single

[26] Cherchye et al. [2012] argue that empowering fathers may yield larger benefits for children than empowering mothers.

Table 4: Children resource share estimates – ϕ_k

Number of children	Single Mothers		Single Fathers	
	I	II	I	II
1	0.148**	0.152***	0.216**	0.218***
	(0.062)	(0.056)	(0.087)	(0.071)
2	0.257***	0.313***	0.385***	0.382***
	(0.076)	(0.045)	(0.079)	(0.058)
3	0.355***	0.398***	0.532***	0.496***
	(0.084)	(0.043)	(0.097)	(0.068)
Number of observations	13305		1534	

Notes: * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$. Standard errors are in parentheses. This table presents estimated average expenditures on children, calculated at the average point of the sample. Column (I) reports results based on the baseline sharing rule specification, while Column (II) provides estimates using an alternative specification.

mothers might compensate for lower monetary expenditures by investing more time in child care [García-Mainar et al., 2011, Campaña et al., 2023].

Figure 2 also highlights the nonlinear relationship between family size and the cost of children. For both genders, the marginal increase in child-related expenditures declines with family size. For example, the total cost attributed to children for a father with three children is less than twice that for a father with only one child.

5.4 Comparison with OECD-modified Equivalence Scale

The OECD-modified equivalence scale is widely used in income comparisons to adjust for differences in household size and composition. Under this scale, the first adult is assigned a weight of 1.0, additional adults a weight of 0.5, and children under 14 a weight of 0.3. Resource shares implied by the OECD scale can be directly compared to the child resource shares estimated in this paper. For example, in a single-parent household with one child under 14, the OECD scale implies that the child's share of household resources would be $0.3/1.3 = 23\%$, which exceeds my empirical estimates for both single mothers (14.7%) and single fathers (21.6%).

The OECD equivalence scale is a normative construct that imposes fixed resource shares based solely on household size and composition, without accounting for heterogeneity in actual spending patterns related to gender, or parental preferences. In contrast, this study provides empirically grounded estimates of resource allocation, capturing variation across households that reflects consumption choices and demographic characteristics.

Therefore, relying exclusively on the OECD equivalence scale may lead to substantial mismeasurement of the resources available to children, especially in contexts where allocation patterns deviate systematically from the scale's assumptions. This is particularly relevant for single-parent households, where observed spending behavior may diverge from the proportionality imposed by the OECD scale.

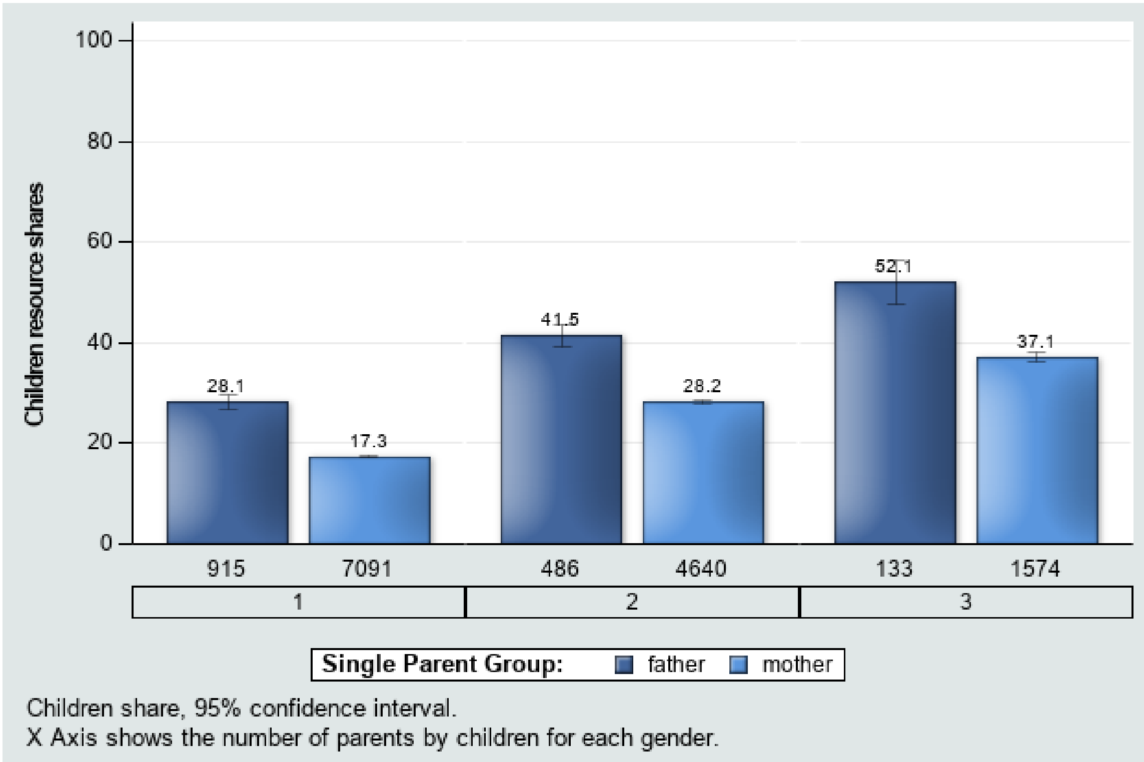

Figure 2: Children resource share by total expenditures

Note: This figure plots the average estimated parental expenditures on children, conditional on family size. The x-axis represents the number of parents, categorized by the number of children..

Table 5: Robustness tests

Models		Female				Male			
		Sargan statistics	LR-type statistics	Degrees of freedom	p-value	Sargan statistics	LR-type statistics	Degrees of freedom	p-value
Reference model		27.05		13		11.00		13	
Model with	region fixed effects excluded in φ	37.59	10.54	9	0.31	8.17	2.83	9	0.97
	prices excluded in φ	17.50	9.55	3	0.02	4.34	6.66	3	0.08
	total expenditures excluded in φ	19.97	7.08	3	0.07	4.73	6.27	3	0.10
	$\pi = s_0 \log y + s_1 \text{dkid}$	22.35	4.69	2	0.10	10.76	0.23	2	0.89
	$\pi = s_0 \log y \sqrt{n} + s_1 p$	22.00	5.05	2	0.08	10.07	0.92	2	0.63
	cubic term in Engel curves	19.98	7.06	4	0.13	3.01	7.98	4	0.09
Model without	economies of scale	26.52	0.52	2	0.77	12.10	1.10	2	0.58

Notes: This table reports Sargan and LR-type statistics for various specification of the model. The first column in each panel for both females and males shows the Sargan statistics, which are the objective function value times the number of observations. The LR-type statistics in the second column in each panel are computed as the absolute value of the difference between the Sargan statistics of the baseline model and those of the respective alternative model. It is worth noting that the objective function calculation for the alternative models is conducted using the identical baseline model weighting matrix.

5.5 Sensitivity Analysis

I assess identification stability using likelihood-ratio (LR)-type tests that compare the baseline model to alternatives exploiting different sources of variation and richer demand structures. In each case, I recompute the alternative model's objective function (J) using the baseline weighting matrix, so the difference in J is χ^2_q under the null.

Across both genders, most alternatives are not statistically different from the baseline. In particular, replacing time with regional variation in φ yields p-value far from rejection, indicating that the sharing rule parameters are not tied to a single dimension of variation. Excluding total expenditures from φ is at most marginally rejected. The only systematic tension arises when prices are excluded from φ: the p-value rejects for women ($p \approx 0.02$) and is borderline for men ($p \approx 0.08$). This pattern is theory-consistent—price movements are informative for the mapping from observables to needs.

I also probe robustness of the Barten scale and Engel curves. Adding a preschool-child term, allowing the scale to depend on prices, enriching Engel curves with cubic terms, and even removing economies of scale all leave the J-statistic essentially unchanged ($p\text{-}value > 0.1$). Hence, the baseline results do not hinge on a particular scaling or curvature assumption, and identification does not rely on functional-form knife-edges.

6 Conclusion

Several models have attempted to assess the cost of children for parents, but most focus exclusively on two-parent households, overlooking the increasing prevalence of single-parent families in OECD countries [Skinner et al., 2008, Nieuwenhuis and Maldonado, 2018]. This paper extends the collective consumption approach to explicitly account for the decision-making processes of single-parent households. By leveraging a model that exploits the stability of preferences and the adult-exclusive goods, I estimate the sharing rule and the associated cost of children in one-adult households.

Drawing on a unique, harmonized sample of single adults—with and without children—from three waves of UK consumption survey data, covering the years 1978–2022, I demonstrate that resource share estimates commonly applied to two-parent households remain externally valid for single-parent families, at least for single mothers. The analysis further reveals that family size exerts a pronounced effect on the intra-household allocation of resources among low-income single-parent families, whereas its impact diminishes markedly at higher income levels. Typically, this implies that in low-income households, the presence of additional siblings significantly reduces per-child allocations. Conversely, in high-income families, the number of children has a negligible impact on per child resources, suggesting that resource dilution effects are effectively mitigated by greater parental resources.

A comprehensive set of LR-type tests shows that my estimates are stable to the source of identifying variation and to rich perturbations of the Barten scale and Engel-curve specification. Switching from temporal to regional variation produces statistically equivalent fit. Also, allowing for additional curvature or alternative scaling leaves the objective function unchanged.

These findings have important policy implications. Most notably, they underscore the existence of a minimum expenditure threshold below which public intervention becomes necessary to ensure children's needs are met. Policy measures such as family allowances or targeted transfers should therefore be calibrated not only to household size but also to resource constraints, with particular emphasis on supporting large, low-income families most at risk of inadequate child expenditures.

A potential limitation of this paper concerns the endogeneity of fertility choices, which may be shaped by expectations about marriage and thus challenge the assumption of preference stability across marital statuses. While this issue is mitigated by the inclusion of rich sociodemographic controls, it is not entirely resolved. Future research could explore revealed preference approaches—such as those proposed by Cherchye et al. [2007, 2015]—to estimate bounds on intrahousehold sharing without relying on preference stability.

7 Statements and Declarations

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Appendix

A Identification Proof

Proof.

1. Let's write $\mathbf{z}^{\omega} = (z^{\omega 0}, z^{\omega 1})$ where $z^{\omega 1} \notin \mathbf{z}^{\pi}$ and $z^{\omega 1} \notin \mathbf{z}^{\varphi}$. Then consider two values of $z^{\omega 1}$, say $z_1^{\omega 1}$ and $z_2^{\omega 1}$. This provides a system of two equations with two unknowns:

$$\omega(\bar{p}, \mathbf{z}^{\omega}, \bar{\mathbf{z}}^{\pi}, \bar{\mathbf{z}}^{\varphi}, \bar{y}, \bar{\mathbf{n}}) = g(\bar{p}, z_1^{\omega 1}, \pi(\bar{p}, \bar{\mathbf{z}}^{\pi}, \bar{y}, \bar{\mathbf{n}}) \cdot \bar{p}, \varphi(\bar{p}, \bar{\mathbf{z}}^{\varphi}, \bar{y}, \bar{\mathbf{n}}) \cdot \bar{y})$$
$$\omega(\bar{p}, \mathbf{z}^{\omega}, \bar{\mathbf{z}}^{\pi}, \bar{\mathbf{z}}^{\varphi}, \bar{y}, \bar{\mathbf{n}}) = g(\bar{p}, z_2^{\omega 1}, \pi(\bar{p}, \bar{\mathbf{z}}^{\pi}, \bar{y}, \bar{\mathbf{n}}) \cdot \bar{p}, \varphi(\bar{p}, \bar{\mathbf{z}}^{\varphi}, \bar{y}, \bar{\mathbf{n}}) \cdot \bar{y})$$

Under some regularity conditions, this system of two equations generally has a unique solution for $\pi(\bar{p}, \bar{\mathbf{z}}^{\pi}, \bar{y}, \bar{\mathbf{n}})$ and $\varphi(\bar{p}, \bar{\mathbf{z}}^{\varphi}, \bar{y}, \bar{\mathbf{n}})$, and another for each choice of $(\bar{p}, \bar{\mathbf{z}}^{\pi}, \bar{\mathbf{z}}^{\varphi}, \bar{y}, \bar{\mathbf{n}})$.

2-3. Combine (2) and (3), the proof of this statement is similar to the previous one.[27]

4. Let's consider choosing a parametric specification for the sharing rule function, specifically a linear form that depends on k parameters. There are k degrees of freedom, representing the k identifiable parameters. The idea is that we need k equations to determine the unknown parameters.

5. Let
$$\omega = g(\mathbf{z}^{\omega}, \pi_1(p, \mathbf{z}^{\pi}, y) \cdot \pi_2(n) \cdot p, \varphi_1(p, \mathbf{z}^{\varphi}, y) \cdot \varphi_2(n) \cdot y)$$

By varying the values of y and n, we might obtain the following equations:

$$\omega = g(\mathbf{z}^{\omega}, \pi_1(\bar{p}, \bar{\mathbf{z}}^{\pi}, y_1, \tilde{\mathbf{n}}) \cdot \pi_2(n_1) \cdot \bar{p}, \varphi_1(\bar{p}, \bar{\mathbf{z}}^{\varphi}, y_1, \tilde{\mathbf{n}}) \cdot \phi_2(n_1) \cdot y_1)$$
$$\omega = g(\mathbf{z}^{\omega}, \pi_1(\bar{p}, \bar{\mathbf{z}}^{\pi}, y_1, \tilde{\mathbf{n}}) \cdot \pi_2(n_2) \cdot \bar{p}, \varphi_1(\bar{p}, \bar{\mathbf{z}}^{\varphi}, y_1, \tilde{\mathbf{n}}) \cdot \varphi_2(n_2) \cdot y_1)$$
$$\omega = g(\mathbf{z}^{\omega}, \pi_1(\bar{p}, \bar{\mathbf{z}}^{\pi}, y_2, \tilde{\mathbf{n}}) \cdot \pi_2(n_1) \cdot \bar{p}, \varphi_1(\bar{p}, \bar{\mathbf{z}}^{\varphi}, y_2, \tilde{\mathbf{n}}) \cdot \varphi_2(n_1) \cdot y_2)$$
$$\omega = g(\mathbf{z}^{\omega}, \pi_1(\bar{p}, \bar{\mathbf{z}}^{\pi}, y_2, \tilde{\mathbf{n}}) \cdot \pi_2(n_2) \cdot \bar{p}, \varphi_1(\bar{p}, \bar{\mathbf{z}}^{\varphi}, y_2, \tilde{\mathbf{n}}) \cdot \varphi_2(n_2) \cdot y_2)$$

The above example shows a set of 4 equations with 4 unknowns. Then one can identify the sharing function as well as the economies of scales. This completes the proof.

■

B Relevance Tests for Time Variation in Identification

I assess the relevance of the time variation used for identification by testing whether decade dummies significantly shift residualized demand once other factors are partialled out. Using the Frisch-Waugh-Lovell theorem, I regress the residualized outcome $r_{\omega} = M_Z \omega$ on the residualized decade indicators (r_{D1}, r_{D2}, r_{D3}), clustering standard errors at the year level—the unit of variation of the excluded shifters. A joint Wald test of $H_0 : \delta = 0$ provides the block-level significance, and the corresponding partial R^2 is computed from the F-statistic. As seen in Table 1 in the main text, in both the female

[27] The complete proof for the statement 2 is given by [Dunbar et al., 2013, online appendix].

and male subsamples, the null is rejected: decade variation explains a substantial share of residual demand movements conditional on controls, satisfying the relevance condition.

Table 6: Differences vs base 2011-2022, clustered year by gender

	Women			Men		
Decade dummies	Est. value	Std err.	p-value	Est. value	Std err.	p-value
1978-1990	0.003	0.002	0.165	0.002	0.003	0.539
1991-2000	0.001	0.002	0.638	-0.005	0.003	0.077
2001-2010	0.005	0.002	0.025	0.002	0.004	0.620

Notes: This table reports the adjusted means of the residual $M_Z\omega$ (residualized clothing budget share) by decade, controlling for the full set of covariates Z and clustering standard errors at the year level. The 2011–2022 decade serves as the reference category. Coefficients represent the average deviation (in budget share points) relative to this baseline. Among women, the 2001–2010 decade shows a higher residual demand level (+0.005, $p \approx 0.025$), while other decades are statistically equivalent from the baseline. Among men, the 1991–2000 decade displays a significantly lower level (-0.005, $p = 0.077$), with other decades showing no significant difference.

To quantify the magnitude of these shifts (in budget share points) and identify which decades deviate from the baseline, I estimate decade-specific adjusted means of $M_Z\omega$. For women, the level is highest in the early to mid-2000s, with the 2001–2010 period being approximately +0.0052 (p = 0.0246) relative to the 2011–2022 baseline and +0.0043 (p = 0.0317) higher than 1991–2000. For men, the largest gap occurs between 2001–2010 and 1991–2000, at about +0.0067 (p = 0.0185), while 1991–2000 is roughly −0.0069 (p = 0.0048) below 1978–1990 and −0.00495 (p = 0.077) below the 2011–2022 baseline. These estimates indicate that for women, the early 2000s stand out as a period of high residual demand, whereas for men the 1990s exhibit markedly lower values compared to both earlier and later decades.

Table 7: Pairwise decade contrasts, clustered year by gender

	Women			Men		
Parwise decade contrast	Est. value	Std err.	p-value	Est. value	Std err.	p-value
(2001-2010 vs 1991-2000)	0.004	0.002	0.032	0.007	0.003	0.019
(2001-2010 vs 1978-1990)	0.003	0.002	0.235	-0.000	0.003	0.943
(1991-2000 vs 1978-1990)	-0.002	0.002	0.252	-0.007	0.002	0.005

Notes: This table reports estimated mean differences in $M_Z\omega$ (residualized clothing budget share) between decades, controlling for the full set of covariates Z and clustering standard errors at the year level. Coefficients represent deviations in budget share points between the indicated decades. The results highlight gender-specific temporal patterns, with a peak in the early 2000s for women and a pronounced dip in the 1990s for men.

Finally, I include a visual diagnostic by plotting the annual mean of $M_Z\omega$ with 95% confidence intervals, shading the background by decade blocks. This provides a direct visual check that the annual trends support the decade-level breaks detected by the statistical tests. The figures confirm the patterns: a pronounced peak for women in the early-to-mid 2000s and a deep trough for men in the mid-1990s.

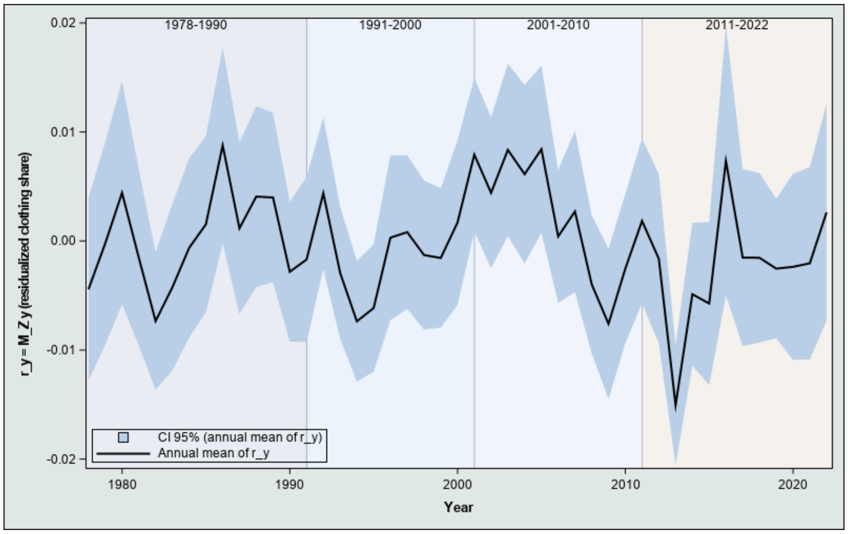

Figure 3: Annual Means of $M_Z\omega$ with Confidence Intervals and Decade Segments— Women

Note: This figure plots the annual mean of the residual $M_Z\omega$ (residualized clothing budget share) together with its 95% confidence interval, overlaid with shaded backgrounds denoting decade blocks. The graph provides a visual check on whether the annual trends align with the temporal breaks detected in the formal tests (Tables 6-7) and highlights potential peaks or troughs in the series.

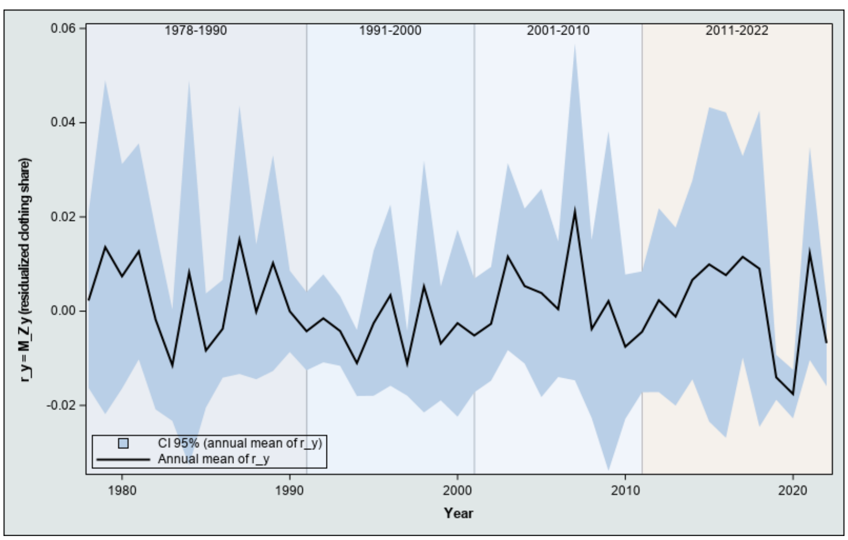

Figure 4: Annual Means of $M_Z\omega$ with Confidence Intervals and Decade Segments— Men

Note: See the notes in Figure 3.

Table 8: Results for budget share equations

	Women						Men					
Parameters	I	II	III	IV	V	VI	I	II	III	IV	V	VI
Parent variables												
$\alpha_{intercept}$	0.166***	0.166***	0.146***	0.227***	0.149***	0.180***	0.134***	0.135***	0.135***	0.133***	0.146***	0.139***
	(0.015)	(0.014)	(0.015)	(0.028)	(0.023)	(0.014)	(0.013)	(0.013)	(0.013)	(0.013)	(0.026)	(0.013)
$\alpha_{education}$	0.000	0.000	0.000	0.000	0.000	−0.001**	0.000	0.000	0.000	0.000	0.000	0.000
	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)
α_{age}	−0.004***	−0.003***	−0.003***	−0.004***	−0.003***	−0.003***	−0.004***	−0.004***	−0.004***	−0.004***	−0.004***	−0.004***
	(0.001)	(0.001)	(0.001)	(0.001)	(0.001)	(0.001)	(0.001)	(0.001)	(0.001)	(0.001)	(0.001)	(0.001)
α_{age^2}	0.004***	0.003***	0.003***	0.003***	0.003***	0.003***	0.004***	0.004***	0.004***	0.004***	0.004***	0.004***
	(0.001)	(0.001)	(0.001)	(0.001)	(0.001)	(0.001)	(0.001)	(0.001)	(0.001)	(0.001)	(0.001)	(0.001)
$\alpha_{labor\ participation}$	0.002	0.001	0.001	−0.003	0.002	0.000	0.001	0.001	0.002	0.001	0.000	0.000
	(0.003)	(0.003)	(0.003)	(0.004)	(0.003)	(0.002)	(0.002)	(0.002)	(0.002)	(0.002)	(0.002)	(0.002)
$\alpha_{house\ owner}$	0.016**	0.013**	0.014***	0.008	0.014***	−0.001	−0.002	−0.003	−0.001	−0.003	−0.003*	−0.003*
	(0.007)	(0.006)	(0.005)	(0.007)	(0.004)	(0.002)	(0.002)	(0.002)	(0.002)	(0.002)	(0.002)	(0.002)
Decades												
$\alpha_{decade\ dummy\ 1}$	0.006	0.006	0.021***	−0.056**	0.033***	0.006	0.002	0.002	0.001	0.003	−0.009	0.001
	(0.004)	(0.004)	(0.005)	(0.025)	(0.012)	(0.004)	(0.004)	(0.004)	(0.004)	(0.006)	(0.011)	(0.004)
$\alpha_{decade\ dummy\ 2}$	0.002	0.002	0.014***	−0.024*	0.019***	0.001	0.001	0.002	0.002	0.003	−0.002	0.001
	(0.003)	(0.003)	(0.003)	(0.013)	(0.007)	(0.003)	(0.003)	(0.003)	(0.003)	(0.003)	(0.006)	(0.003)
$\alpha_{decade\ dummy\ 3}$	0.008***	0.008***	0.009***	0.026***	−0.001	0.008***	0.002	0.002	0.002	0.001	0.004	0.002
	(0.002)	(0.002)	(0.003)	(0.005)	(0.007)	(0.002)	(0.002)	(0.002)	(0.002)	(0.002)	(0.006)	(0.002)
Region												
$\alpha_{Northern}$	−0.003	−0.002	−0.002	−0.004	−0.002	−0.001	−0.008*	−0.008*	−0.009*	−0.007	−0.007	−0.009*
	(0.006)	(0.006)	(0.005)	(0.006)	(0.005)	(0.004)	(0.005)	(0.005)	(0.005)	(0.005)	(0.005)	(0.005)
$\alpha_{York\ \&\ Humberside}$	−0.006	−0.005	−0.004	−0.007	−0.004	0.000	−0.016***	−0.016***	−0.016***	−0.015***	−0.014***	−0.016***
	(0.005)	(0.005)	(0.004)	(0.006)	(0.004)	(0.003)	(0.005)	(0.005)	(0.005)	(0.005)	(0.004)	(0.004)
$\alpha_{East\ Midlands}$	−0.006	−0.005	−0.004	−0.002	−0.005	0.003	−0.018***	−0.018***	−0.018***	−0.017***	−0.017***	−0.018***
	(0.005)	(0.005)	(0.005)	(0.006)	(0.005)	(0.003)	(0.005)	(0.005)	(0.005)	(0.005)	(0.004)	(0.004)
$\alpha_{East\ Anglia}$	−0.008	−0.007	−0.005	−0.008	−0.005	−0.001	−0.015***	−0.016***	−0.016***	−0.015***	−0.014***	−0.016***
	(0.006)	(0.005)	(0.005)	(0.006)	(0.005)	(0.003)	(0.005)	(0.005)	(0.005)	(0.005)	(0.005)	(0.005)
$\alpha_{Greater\ London}$	−0.006	−0.005	−0.002	−0.010*	−0.001	0.000	−0.015***	−0.015***	−0.014***	−0.014***	−0.014***	−0.015***
	(0.005)	(0.005)	(0.004)	(0.005)	(0.004)	(0.003)	(0.005)	(0.005)	(0.005)	(0.005)	(0.004)	(0.004)
$\alpha_{South-East}$	−0.006	−0.005	−0.005	−0.008	−0.005	−0.002	−0.019***	−0.019***	−0.019***	−0.018***	−0.018***	−0.019***
	(0.005)	(0.005)	(0.004)	(0.005)	(0.004)	(0.003)	(0.004)	(0.004)	(0.004)	(0.004)	(0.004)	(0.004)
$\alpha_{South-West}$	−0.011**	−0.011**	−0.010**	−0.011*	−0.009**	−0.003	−0.022***	−0.022***	−0.022***	−0.021***	−0.020***	−0.021***
	(0.005)	(0.005)	(0.004)	(0.006)	(0.004)	(0.003)	(0.005)	(0.005)	(0.005)	(0.005)	(0.004)	(0.004)
α_{Wales}	−0.010*	−0.009	−0.006	−0.007	−0.006	−0.004	−0.013**	−0.013**	−0.014***	−0.013**	−0.012**	−0.012**
	(0.006)	(0.006)	(0.005)	(0.006)	(0.005)	(0.004)	(0.005)	(0.005)	(0.005)	(0.005)	(0.005)	(0.005)
$\alpha_{West-Midlands}$	−0.008	−0.007	−0.005	−0.008	−0.005	0.000	−0.014***	−0.014***	−0.013***	−0.012***	−0.012***	−0.014***
	(0.005)	(0.005)	(0.005)	(0.006)	(0.004)	(0.003)	(0.005)	(0.005)	(0.005)	(0.005)	(0.004)	(0.004)
$\alpha_{North-West}$	−0.002	−0.001	0.001	−0.003	0.002	−0.001	−0.018***	−0.018***	−0.018***	−0.017***	−0.016***	−0.018***
	(0.005)	(0.005)	(0.004)	(0.006)	(0.004)	(0.003)	(0.004)	(0.004)	(0.004)	(0.004)	(0.004)	(0.004)
$\alpha_{Scotland}$	−0.003	−0.002	−0.001	−0.006	0.000	−0.003	−0.014***	−0.014***	−0.015***	−0.013***	−0.013***	−0.014***
	(0.005)	(0.005)	(0.004)	(0.005)	(0.004)	(0.003)	(0.005)	(0.005)	(0.005)	(0.005)	(0.004)	(0.004)
Financial variables												
β	0.015***	0.014***	0.002	0.082***	−0.028*	0.011***	0.005	0.004	0.005	0.001	0.019*	0.005
	(0.004)	(0.004)	(0.003)	(0.026)	(0.016)	(0.004)	(0.005)	(0.005)	(0.005)	(0.009)	(0.011)	(0.005)
γ	0.005	0.009	0.004	0.002	0.023	0.023***	0.016**	0.017***	0.011***	0.014*	0.015	0.015**
	(0.010)	(0.010)	(0.009)	(0.012)	(0.016)	(0.008)	(0.007)	(0.007)	(0.003)	(0.007)	(0.016)	(0.006)
η	−0.012*	−0.011*	−0.014**	−0.021***	−0.012	−0.010*	0.002	0.003	0.000	0.000	0.001	0.001
	(0.006)	(0.006)	(0.006)	(0.007)	(0.007)	(0.006)	(0.005)	(0.005)	(0.000)	(0.005)	(0.001)	(0.005)
Number of observations	24754	24754	24754	24754	24754	24754	15649	15649	15649	15649	15649	15649
(Number of free parameters, Instruments)	(47,59)	(46,58)	(48,58)	(49,58)	(50,58)	(35,58)	(47,59)	(46,58)	(48,58)	(49,58)	(50,58)	(35,58)

Notes: * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$. Standard errors are in parentheses. This table presents partial results on individual from the demand equations for women - columns (I-VI) - and men - columns (I-VI). See the notes in Table 3 in the main text.

C Budget Share Equations

Table 8 reports estimates of the clothing budget share equations for men and women. The demand estimates reveal that socio-demographic characteristics play a limited role, with age being the primary determinant of clothing budget allocation for both men and women. For both genders, age exhibits a nonlinear relationship with clothing expenditure. Specifically, the budget share decreases with age but increases at older ages, as indicated by the significantly negative coefficient on age and the positive coefficient on age squared.

D Additional Figures

Figure 5 plots the non-parametric distributions of the estimated fraction of each parent's resources devoted to children. The modal mass for mothers lies close to the lower part of the support (roughly 0.10–0.25), whereas the father distribution peaks at higher shares and appears to be more evenly distributed across a broader range. There is sizeable overlap between 0.15 and 0.60, but the father density allocates more probability in the 0.35–0.70 range, while the mother density concentrates below 0.35. This pattern is consistent with the regression evidence reported in the main text.

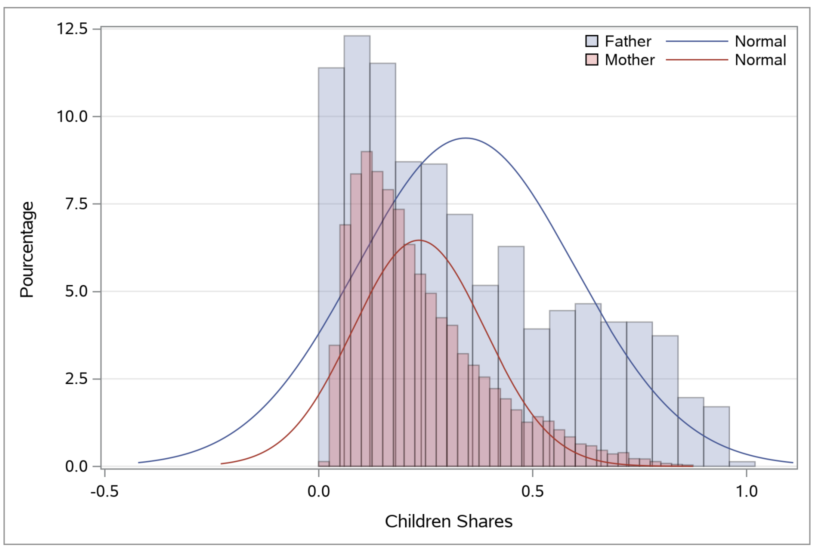

Figure 5: Non-parametric distribution of parents' resource shares allocated to children

Notes: This figure plots the density of the cost of children across parents. Based on the sharing rule estimates, the mean share of resources devoted to children is 0.28 and 0.40 respectively for mothers and fathers. The father distribution is shifted to the right and is more dispersed than the mother distribution; both depart from normality due to the bounded support $[0, 1]$.

Figure 6 plots the relative price trends for male and female clothing, with both indices set to 100 in the years 1978 and 2022, respectively. Both categories exhibit a marked decline between 1978 and around 2010, though the fall is substantially sharper for female clothing. By the mid-1980s, the female clothing index had dropped below 60, compared to above 70 for male clothing. The gap widened further through the 2000s, with female clothing prices reaching a trough of about 10 in

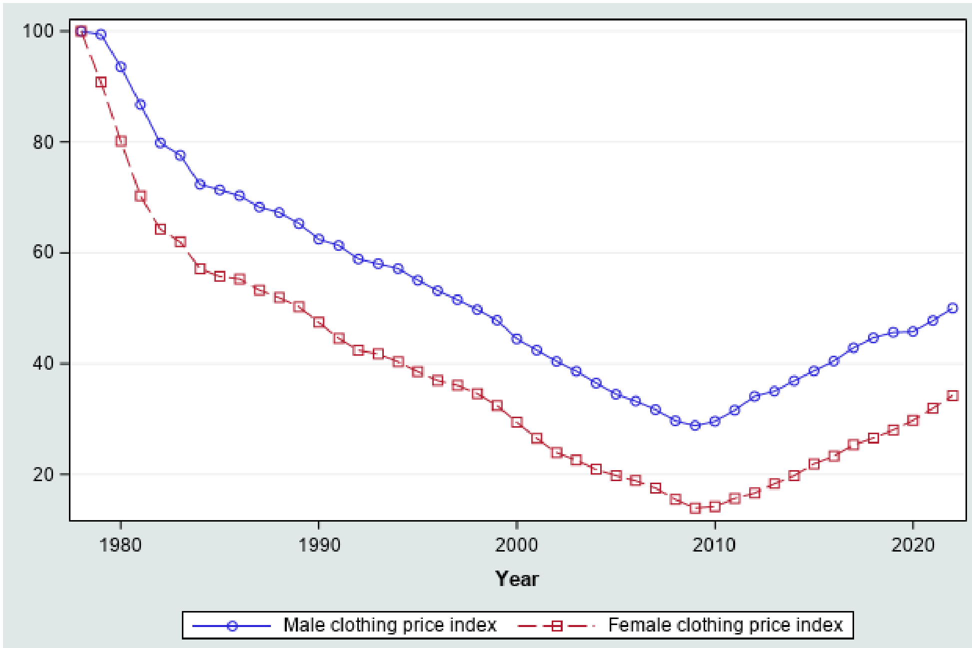

Figure 6: Relative price indexes for male and female clothing between 1978 and 2022

Notes: This figure illustrates the relative prices of male and female clothings from 1978 to 2022, represented by indices set to 100 in both 1978 and 2022, respectively.

2009, relative to roughly 30 for male clothing. Since 2010, both series have rebounded, with male and female clothing prices showing a gradual recovery.