Why Are Older Women Missing in India?
The Age Profile of Bargaining Power and Poverty

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Abstract

The ratio of women to men is particularly low in India relative to developed countries. It has recently been argued that close to half of these *missing women* are of post-reproductive ages (45 and above), but what drives this phenomenon remains unclear. I provide an explanation for this puzzle that is based on intra-household bargaining and resource allocation. I employ both reduced-form and structural modeling to establish the relevant connections between women's bargaining position within the household, their health, and their age. First, using amendments to the Indian inheritance law as a natural experiment, I demonstrate that improvements in women's bargaining position within the household lead to better health outcomes. Next, with a collective model of Indian households, I show that women's bargaining power and their ability to access household resources deteriorate at post-reproductive ages. Thus, at older ages poverty rates are significantly higher among women than men. I argue that gender inequality within the household and the consequent gender asymmetry in poverty account for a considerable fraction of missing women of post-reproductive ages and that policies promoting intra-household equality can have a large impact on female poverty and mortality.

Keywords: missing women, intra-household bargaining power, women's health, Hindu Succession Act, collective model, resource shares, poverty, elderly.

JEL codes: D1, K36, I12, I31, I32, J12, J14, J16.

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1 Introduction

There are far more men than women in India relative to developed countries. Following seminal work by Amartya Sen (1990; 1992), this fact has been dubbed the *missing women* phenomenon.[1] Sex-selective abortion and excess female mortality at early ages related to parental preferences for sons have been identified as important determinants of missing women and biased sex-ratios.[2] Recent work by Anderson and Ray, however, indicates that excess female mortality in India persists beyond childhood and that the majority of missing Indian women die in adulthood. While they do not dispute the presence of a severe gender bias at young ages and the role played by maternal mortality, Anderson and Ray (2010) demonstrate that close to half of missing women in India are of post-reproductive ages, i.e., 45 and above (see figure 1).[3] Unlike the missing girls phenomenon, excess female mortality at older ages in India has not received much attention and remains a puzzle.

I seek to explain this puzzle by examining the critical connections between women's age, intra-household bargaining, and health, while taking the link between health and mortality as given. Using reduced-form and structural methods, I identify one crucial mechanism – the decline in women's bargaining position during post-reproductive ages – that can account for up to 89 percent of the missing women in the 45-79 age group. The decrease in women's bargaining power is reflected in their diminished ability to access household resources. As a consequence, at older ages poverty rates are significantly higher among women than men. I call this fact *excess female poverty* and show that the age profile of excess female poverty matches the profile of excess female mortality in figure 1 nearly exactly at post-reproductive ages.

My analysis proceeds in two steps. First, by using amendments to the Indian inheritance law as a natural experiment, I analyze the relationship between women's bargaining power and their health. I focus on the Hindu Succession Act (HSA) amendments that equalized women's inheritance rights to men's in several Indian states between 1976 and 2005. Using data from the 2005-2006 National Family Health Survey (NFHS-3), I show that women's exposure to these reforms increases their body mass index and reduces the probability of them being anaemic or underweight by strengthening their bargaining power. Next, I examine whether older women are missing in India because their bargaining position weakens at post-reproductive ages. To test my hypothesis, I set out a household model with efficient bargaining to structurally estimate women's bargaining power and investigate its determinants. At post-reproductive ages, I find evidence of a substantial decline in women's bargaining position and in their ability to access household resources.

I model Indian households using the *collective* framework, where each family member has a sep-

[1] Bongaarts and Guilmoto (2015) estimates that over 126 million of Indian women who should be alive are missing. This number is expected to peak at 150 million in 2035.

[2] See e.g., Sen (1990) and DasGupta (2005).

[3] They estimate a total of 1.7 million excess female deaths in year 2000 alone (0.34 percent of the total female population), 45 percent of which occurred at the age of 45 and above. In contrast, they find that close to 44 percent of China's missing women are located "around birth". For each age category, Anderson and Ray (2010) compare the actual female death rate in India to a reference female death rate. The latter is one that would be obtained if the death rate of males in India were to be rescaled by the relative death rates for males and females (in the same category) in developed countries. They compute *missing women* as the product between the difference between the actual and reference death rates and the female population in each age group. Figure 1 displays estimates from Anderson and Ray (2010), table 3, p. 1275. These results are further investigated and confirmed in Anderson and Ray (2012). These findings are consistent with a striking non-monotonic pattern of sex-ratios over age in India. See section B.1 in the online Appendix for more details.

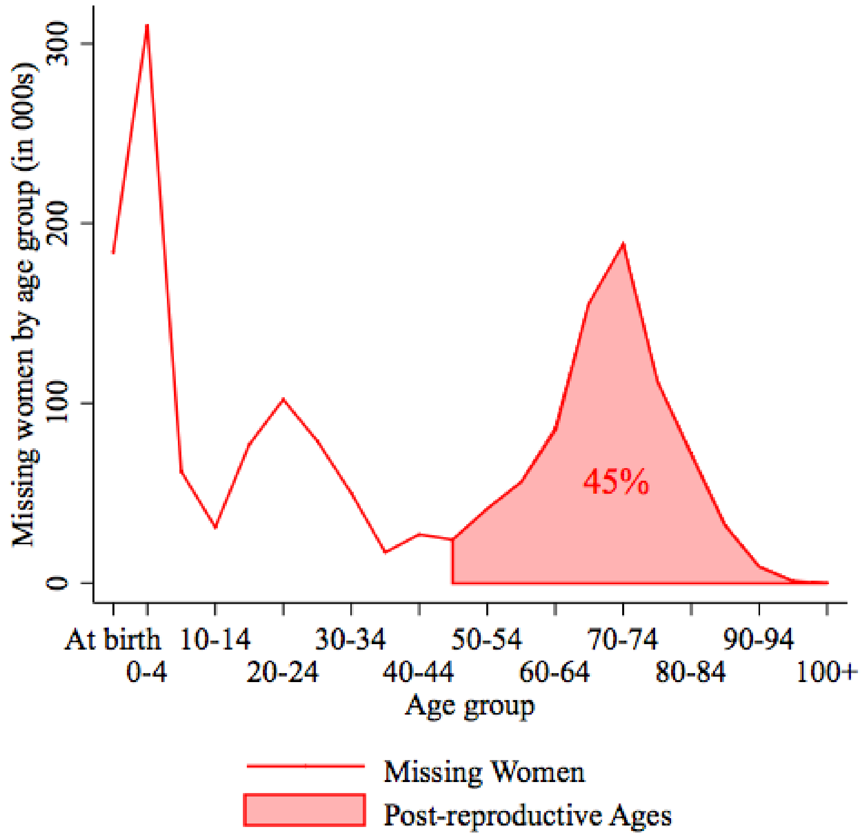


Figure 1: Missing Women by Age Group (Anderson-Ray, 2010)

arate utility function over goods and the intra-household allocation of goods is Pareto efficient. In line with the Indian family structure, I consider both nuclear and non-nuclear households. I measure women's bargaining power as their *resource share*, i.e., the fraction of the household's total expenditure consumed by women.[4] I identify household members' resource shares through Engel curves (demand equations holding prices constant) of clothing items that are consumed exclusively by women, men or children, using a methodology developed in Dunbar et al. (2013). I estimate the model with detailed expenditure data from the 2011-2012 National Sample Survey (NSS) of Consumer Expenditure and use these structural estimates to outline the profile of women's bargaining power over the life-cycle. During women's core reproductive ages the allocation of resources between women and men is symmetric. However, women's resource shares relative to men's decline steadily at post-reproductive ages, when women get as low as 60 percent of men's resources.

Due to the lack of NSS data on health outcomes, I cannot examine the relationship between resource allocation and mortality in the structural model. As morbidity and mortality rates are higher in poverty, however, I indirectly explore this link by studying how gender inequality within the household affects aggregate measures of welfare.[5] I use the model predictions to compute individual level expenditures that take into account unequal intra-household allocation. I compare these per-capita expenditures to poverty thresholds to calculate gender and gender-age specific poverty rates. By contrast, standard per capita poverty measures assume equal sharing and ignore intra-household inequality. My poverty estimates indicate that at all ages there are more women

[4]See e.g., Lewbel and Pendakur (2008), Browning et al. (2013), and Dunbar et al. (2013).

[5]The World Health Statistics report (2015) lists poverty as one of the major determinants of health. Moreover, according to WHO (1999) the likelihood of death before at adult ages is about 2.5 times higher among the poor than among the non-poor.

living in poverty than men, but the gap between female and male poverty rates widens dramatically at post-reproductive ages. For individuals aged 45 to 79, poverty rates are on average 80 percent higher among women than men. This result provides additional support for my hypothesis that intra-household inequality can explain excess female mortality. Using a simple model to relate my findings to the Anderson and Ray's estimates, I then demonstrate that a considerable proportion of missing women at older ages can be attributed to intra-household gender inequality.

My structural estimates are consistent with the existing reduced-form evidence of the importance of inheritance rights in shaping women's position within the household.[6] I find that exposure to the Hindu Succession Act amendments increases women's resource shares by 0.2 standard deviations. These reforms were enacted in different states at different times between 1976 and 2005 and only applied to Hindu, Buddhist, Sikh and Jain women who were not married at the time of implementation. A large fraction of Indian women, especially of older ages, is therefore excluded. I perform a counterfactual exercise and calculate women's resource shares in the hypothetical scenario of all women (of all religions and ages) benefiting from these reforms and demonstrate that granting universal equal inheritance rights can reduce the number of women living in extreme poverty by 9 percent and the number of excess female deaths at post-reproductive ages by up to 24 percent.

The contribution of this paper is fourfold. First, this study is the first to show that excess female mortality at older ages in India can be explained by asymmetries in intra-household bargaining and resource allocation. While Anderson and Ray (2010, 2012) raise awareness about this phenomenon, little work has been done to understand possible channels generating excess female mortality in adulthood in India, especially at older ages. Milazzo (2014) argues that excess mortality among women aged 30 to 49 could be partly explained by son preference, while the most recent paper by Anderson and Ray (2015) shows that excess female mortality between the ages of 20 and 65 is particularly severe among unmarried women and widows. My analysis departs from previous works in that it provides an original explanation for missing women at older ages (45 to 79) while employing both reduced-form analysis and structural modeling.[7] The structural model allows me to perform counterfactual experiments and to examine the roles of son preference and widowhood within a full model of household bargaining. Second, while the effect of changes in Indian inheritance laws on women's outcomes has been studied previously (e.g., Roy (2008, 2013), Deininger et al. (2013), Heath and Tan (2014)), no work focused explicitly on women's health and access to household resources. Third, this is the first attempt to formally outline the profile of women's intra-household bargaining power over the life-cycle in a developing country. Fourth, to the best of my knowledge, this paper is the first to provide measures of gender-age specific poverty that take into account unequal resource allocation within the household.

Important policy implications may be drawn from this analysis. As the population in India and in other developing countries ages, gender asymmetries among the elderly need to be further investi-

[6]See e.g., Roy (2008) and Heath and Tan (2014), who show that women's exposure to the HSA reforms improves self-reported measures of autonomy and negotiating power within their marital families.

[7]In this regard, this study relates to recent contributions in the econometrics literature highlighting the benefits of combining structural and the causal modeling approaches (e.g., Heckman (2008, 2010) and Lewbel (2016)).

gated and promptly addressed. Moreover, policies aimed at promoting equality within households, such as improving women's rights to inherit property, can have positive spillovers on female health, poverty and mortality. Finally, intra-household inequalities should be taken into account when measuring poverty and evaluating the effect of policies to alleviate it.

The rest of the paper is organized as follows. Section 2 provides an overview of the related literature and discusses further the contributions of this paper. Section 3 presents the reduced-form results and establishes a positive causal link between women's intra-household bargaining power and their health. Section 4 discusses the household model, the identification of resource shares and the structural estimation results. Section 5 outlines the age profiles of female bargaining power and poverty, and relates them to the phenomenon of excess female mortality at post-reproductive ages. Section 6 presents the counterfactual policy analysis. Section 7 concludes.

2 Related Literature

This paper relates to several strands of literature: the previous research on the missing women phenomenon, the existing studies on poverty among the elderly in South Asia, the work on inheritance rights and the Hindu Succession Act amendments in particular, and the literature on intra-household allocation and bargaining power.

Since it was first addressed by Amartya Sen in 1990, the phenomenon of missing women has been widely studied. It refers to the fact that in parts of the developing world, especially in India and China, the ratio of women to men is particularly skewed. Coale (1991) estimates a total of 60 million missing females in the world at the beginning of the nineties, with India accounting for more than one third of them. In 2010, 126 million women were missing from the global population, with China and India accounting for 85 percent of this bias in sex ratios (Bongaarts and Guilmoto (2015)). The literature has traditionally related this fact to son preference and several works have provided empirical evidence of sex-selective abortion, female infanticide and excess female mortality in childhood (see DasGupta (2005) for an overview of this literature). Jha et al. (2006), for example, find strong evidence of selective abortion of female fetuses in India. Moreover, the introduction of ultra-sound technologies at the end of the 1980s has been found to be associated with even more skewed sex-ratios and preferential prenatal treatment for boys (e.g., Bhalotra and Cochrane (2010), Bharadwaj and Lakdawala (2013)).[8] Finally, Oster (2009) and Jayachandran and Kuziemko (2011) show that gender differences in child mortality are associated with differential health investment between genders.[9]

[8]While access to ultra-sound has reduced gender gaps in post-neonatal child mortality, Anukriti et al. (2015) demonstrate that this decline is not large enough to compensate for the increase in the male-female sex ratio at birth due to sex-selective abortions.

[9]Discrimination against girls in India has also being investigated by directly looking at how household consumption patterns varies with the gender composition of children. Most of these works use the so called *Engel curve approach* - not to be confused with the structural approach of identification of resource shares based on Engel curves estimation that I implement in this paper -, which consists on regressing budget shares of a set of goods on log per-capita expenditure, log household size, the shares of various age-sex groups and other relevant household characteristics. Among others, Subramanian and Deaton (1990) find evidence of gender discrimination in rural Maharashtra for 10-14 year olds, Lancaster et al. (2008) find empirical evidence of gender bias in rural Bihar and Maharashtra for the 10-16 age group, while Zimmermann (2012) finds that gender discrimination in education expenditures between boys and girls increases with age.

A notable exception to this literature is the recent work by Anderson and Ray (2010; 2012; 2015), who indicate that close to half of missing women in India die at older ages. The plight of widows in the Indian subcontinent has been previously investigated by Jean Drèze and coauthors (e.g., Drèze et al. (1990), Chen and Drèze (1995), Drèze and Srinivasan (1997)). Moreover, the study on the conditions of the elderly in South Asia suggests that women's bargaining power and access to household resources may indeed be key to explain the phenomenon of missing women at post-reproductive ages. Kochar (1999), for example, finds that medical expenditure on the elderly in rural Pakistan is negatively affected by their declining economic contribution to the household, while Roy and Chaudhuri (2008) show that older Indian women report worse self-rated health status, higher prevalence of disabilities, and lower healthcare utilization than men. They show that health disadvantage and lower utilization among women cannot be explained by demographics, but that gender differentials disappear when controlling for economic independence.

Women's intra-household bargaining power and its changes are difficult to measure and often unobservable. Legal reforms aimed at improving women's property rights - inheritance rights, in particular - have been widely used in the literature to assess the relationship between bargaining power and women's outcomes. The present paper is the first to focus on adult women's health outcomes and on their access to household resources. Deininger et al. (2013), for example, find evidence of an increase of women's likelihood of inheriting land following the introduction of Hindu Succession Act (HSA) amendments that equalized women's inheritance rights to men's in several Indian states between 1976 and 2005. Moreover, Roy (2008) documents that women's exposure to the HSA reforms improves their bargaining power and autonomy within their marital families, while Roy (2013), Deininger et al. (2013), and Bose and Das (2015) indicate that it increases female education. Jain (2014) shows that HSA reforms mitigate son preference, and might be effective in reducing mortality differences between boys and girls in rural India. Finally, Heath and Tan (2014) argue that the HSA amendments increase women's labor supply, especially into high-paying jobs.[10]

A remarkably diverse literature has focused on intra-household resource allocation and bargaining power. On one hand, several papers have tested empirically whether households behave in accordance with the *unitary* model, which assumes that the household acts as a single decision unit maximizing a common utility function.[11] On the other hand, a number of papers have focused on developing techniques by which household level consumption data may be used to recover information about individual household members. Building on Chiappori (1988, 1992) and Apps and Rees

[10]While analyzing possible underlying mechanisms, they also find some preliminary evidence of health improvements following the implementation of HSA amendments. Recent works also document some unintended negative consequences of HSA reforms. Rosenblum (2015), for example, shows that the HSA amendments increase female child mortality, which is consistent with parents wanting to maximize their bequest per son, while Anderson and Genicot (2015) argue that HSA reforms are associated with a decrease in the difference between female and male suicide rates, but with an increase in both male and female suicides. Legal reforms in other countries have been studied as well. La Ferrara and Milazzo (2014), for example, exploit an amendment to Ghana's Intestate Succession Law and compare differential responses of matrilineal and patrilineal ethnic groups, finding that parents substitute land inheritance with children's education. Harari (2014) analyzes a law reform meant to equalize inheritance rights for Kenyan women and shows that women exposed to the reform are more educated, less likely to undergo genital mutilation, and have higher age at marriage and at first child.

[11]Most of this empirical literature focus on testing the *income pooling hypothesis*, i.e., that only household income matters for choice outcomes and not the source of the income. See e.g., Attanasio and Lechene (2002), who examine the effect of large cash transfers in rural Mexico (PROGRESA/*Oportunidades* conditional cash transfers), and Duflo (2003), who analyzes a reform in the South Africa social pension program for the elderly. They both find empirical evidence against the unitary model.

(1988), the vast majority of these studies concentrate on the estimation of *collective* household models, in which the household is characterized as a collection of individuals, each of whom has a well defined objective function and who interact to generate Pareto efficient allocations. In this framework, the exact intra-household bargaining protocol is left unspecified. Identification of individuals' *resource shares* (or *sharing rule*), defined as each member's share of total household consumption, is particularly appealing, as it provides a measure of individuals' intra-household bargaining power. Although a series of papers focus on the identification of *changes* in resource shares as functions of factors affecting bargaining power (e.g., Browning et al. (1994), Browning and Chiappori (1998), Vermeulen (2002)), a more recent strand of the literature deals with the identification of the *level* of resource shares, which is my main object of interest (e.g., Lewbel and Pendakur (2008), Browning et al. (2013), Dunbar et al. (2013)). Specifically, Dunbar et al. (2013) identify individuals' resource shares using Engel curves of assignable goods and imposing semiparametric restrictions on individual preferences. Using data from Malawi, they argue that children have higher rates of poverty than their parents, despite commanding a quite large share of household resources. With few exceptions, limited work has used this type of approach to investigate intra-household allocation at older ages.[12] To my knowledge, no previous work has investigated the age profile of women's intra-household bargaining power and its implications for poverty in a developing country.

3 A Reduced-Form Analysis of Bargaining Power and Health

While plausible, that an increase in the bargaining power of women inside the household positively affects their health outcomes is not an obvious fact. Women, for example, may divert resources to children when their position improves, so that no effect could be detected on their own health.

A woman's right to inherit land and other property is often claimed to be a significant determinant of women's economic security and position within the household (World Bank (2014)). I investigate the existence of a positive causal effect of intra-household bargaining power on women's health by exploiting legal reforms equalizing women's inheritance rights to men's. I compare health outcomes of women who were exposed to these reforms to those of women who were not. Research in the medical field indicates that low body mass index (hereinafter BMI), especially BMI below the underweight cutoff of 18.5, and anaemia are associated with an increased risk of mortality.[13] Evidence of improvements in these health outcomes following an increase in women's bargaining power would provide a first empirical validation to my hypothesis that intra-household resource allocation can explain excess female mortality.

Inheritance rights in India differ by religion and, for most of the population, are governed by the Hindu Succession Act (hereinafter HSA). The HSA was first introduced in 1956 and applied to all states other than Jammu and Kashmir and only to Hindus, Buddhists, Sikhs and Jains. It

[12] To analyze how intra-household allocation is affected by retirement and health status of elderly in the US, Bütikofer et al. (2010) estimate a collective model with data on married couples and widows/widowers between ages 50 and 80. Cherchye et al. (2012b) analyze consumption pattern of Dutch elderly households between 1978 and 2004 and find that traditional poverty rates seem to underestimate poverty among widows.

[13] See e.g., Visscher et al. (2000), Thorogood et al. (2003), and Zheng et al. (2011).

therefore did not apply to individuals of other religions, such as Muslims, Christians, Parsis, Jews, and other minority communities.[14] It aimed at unifying the traditional Mitakhshara and Dayabhaga systems, which were completely biased in favor of sons (Agarwal (1995)), and established a law of succession whereby sons and daughters would enjoy (almost) equal inheritance rights, as would brothers and sisters. Gender inequalities, however, persisted even after the introduction of the HSA. On one hand, in case of a Hindu male dying intestate, i.e., without leaving a will, all his *separate* or *self-acquired* property, devolved equally upon sons, daughters, widow, and mother. On the other hand, the deceased's daughters had no direct inheritance rights to *joint family property*, whereas sons were given direct right by birth to belong to the coparcenary.[15]

The Indian constitution states that both federal and state governments have legislative power over inheritance. In the decades following the introduction of the HSA, state governments enacted amendments (hereinafter HSAA) equalizing inheritance rights for daughters and sons. Kerala in 1976, Andhra Pradesh in 1986, Tamil Nadu in 1989, and Maharashtra and Karnataka in 1994 passed reforms making daughters coparceners. These reforms only applied to Hindu, Buddhist, Sikh or Jain women, who were not yet married at the time of the amendment. A national-level ratification of the amendments occurred in 2005.

I consider the following baseline specification:

$$y_{irsc} = \beta HSAA\ Exposed_{irsc} + X'_{irsc}\gamma + \alpha_r + \alpha_c + \alpha_s + \alpha_{rs} + \alpha_{rc} + \alpha_{sc} + \epsilon_{irsc} \tag{1}$$

where y_{irsc} is the outcome of interest for woman i, of religion r, living in state s and born in year c (i.e., BMI or an indicator variable for being underweight or severely, moderately or mildly anaemic). $HSAA\ Exposed_{irsc}$ is an indicator variable equal to one if woman i got married after the amendment in state s and is Hindu, Buddhist, Sikh or Jain. X_{irsc} is a vector of individual and household level covariates, including women's education, number of children in the household, a household wealth index, and indicator variables for having worked in the past year, for living in rural areas and for being part of disadvantaged social groups. The model includes cohort and state fixed effects, a religion dummy equal to 1 if a woman is Hindu, Buddhist, Sikh or Jain, and zero otherwise, and religion-cohort, cohort-state, and state-religion fixed effects.[16] β is the parameter of interest and represents the treatment effect of being exposed to HSA amendments, i.e., to an exogenous variation in women's intra-hous'ehold bargaining power. In the baseline specification, standard errors are clustered at the primary sampling unit (village) level. Results are robust to clustering the standard errors at the state, cohort-state, and at the cohort-state-religion level.

I estimate the model with OLS using a sample of married women of age 15 to 49 from the

[14] While most laws for Christians formally grant equal rights from 1986, gender equality is not the practice, as the Synod of Christian Churches has being arranging legal counsel to help draft wills to disinherit female heirs. The inheritance rights of Muslim women in India are governed by the Muslim Personal Law (Shariat) Application Act of 1937, under which daughters inherit only a portion of what the sons do (Agarwal (1995)).

[15] All persons who acquired interest in the joint family property by birth are said to belong to the *coparcenary*. The Hindu Women's Right to Property Act of 1937 enabled the widow to succeed along with the son and to take a share equal to that of the son. The widow was entitled only to a limited estate in the property of the deceased with a right to claim partition. A daughter, however, had virtually no inheritance rights.

[16] The fixed effects are based on 28 states, 35 cohorts and 2 religious categories.

Table 1: HSA Amendments and Women's Health

	Body Mass Index				Pr(Anaemia)		
	All Sample	$BMI \leq 23$	$BMI > 23$	Pr($BMI \leq 18.5$)	Severe	Moderate	Mild
	(1) OLS	(2) OLS	(3) OLS	(4) OLS	(5) OLS	(6) OLS	(7) OLS
HSAA Exposed	0.205***	0.264***	-0.0846	-0.0446***	-0.0123***	-0.0304***	-0.0316***
	(0.0776)	(0.0556)	(0.132)	(0.0102)	(0.00316)	(0.00897)	(0.0110)
N	81,534	57,607	23,927	81,534	77,777	77,777	77,777
Mean Dependent Variable	21.42	19.24	26.69	0.2648	0.0154	0.1559	0.5298

Note: $^{*}p < 0.10$, $^{**}p < 0.05$, $^{***}p < 0.01$. NFHS-3 data. Married women of age 15 to 49 included in the sample. All specifications include a religion dummy, equal to 1 if a woman is Hindu, Buddhist, Sikh or Jain, state and cohort fixed effects, and state-religion, state-cohort, and religion-cohort fixed effects. Individual controls include number of children under 5 in the household, a household wealth index, and indicator variables for having completed 12 years of school, having worked in the past 12 months, for living in rural areas and for being part of Scheduled Castes, Scheduled Tribes or Other Backward Classes. Mild anaemia includes moderate and severe anaemia, and moderate anaemia includes severe anaemia. Robust standard errors in parentheses. Standard errors clustered at the primary sampling unit (village) level (3,753). Sampling weights applied.

2005-2006 National Family Health Survey (NFHS-3). The average BMI in the estimation sample lies within the normal range (18.5 to 23). Nonetheless, 26 percent of the women in the sample is underweight. Moreover, anaemia appears to be an endemic problem, with 50 percent of women in the sample suffering from mild anaemia, 16 percent from moderate anaemia and 2 percent from severe anaemia (see table A1 in the Appendix).[17] Finally, one out of six women in the sample have been exposed to the HSA amendments.

Table 1 presents the estimation results. The first three columns focus on BMI outcomes, over the full sample (column 1), a sample restricted to women considered underweight or normal weight according to the WHO cutoffs (column 2), and a sample restricted to women overweight and obese (column 3). The sample breakdown aims at addressing potential concerns related to the non-monotonicity of the relationship between BMI and health: while increases in BMI correspond to better health for individuals below the overweight cutoff, the opposite holds true for individuals above it. Exposure to HSA amendments is associated with an increase in women's BMI by 0.21 when the entire sample is considered, and by 0.26 when only underweight and normal weight range individuals are included. As expected, no significant effect can be detected when only overweight and obese women are included in the analysis. Column 4 to 7 report the linear probability models estimation results. All specifications indicate that women exposed to HSA amendments have better health outcomes, as they are 4 percent less likely to be underweight, about 1 percent less likely to be severely anaemic, 3 percent less likely to be moderately anaemic, and 3 percent less likely to be mildly anaemic.[18]

Robustness Checks. I perform a series of robustness checks to test the sensitivity of the reduced-form results. Section B.2 in the online Appendix contains the complete set of results. First, as exposure to the HSA amendments is determined by each woman's year of marriage, I address concerns about the potential endogeneity of treatment by excluding from the analysis women who

[17]Section B.2 in the online Appendix contains more details on the data and on the use of BMI and anaemia as health measures.

[18]Results are robust to estimating maximum-likelihood probit models. Table A3 in the Appendix shows the marginal effects of being exposed to HSA amendments on binary health outcomes.

got married right around the reforms (the year of, before and after), by estimating an intent-to-treat effect using a measure of eligibility to HSA amendments that exploits variation in women's year of birth, religion and state, and by using an instrumental variable approach. I show that the potential endogeneity of the time of marriage is not the driving force behind my findings. Second, I assess how exposure to HSA amendments by the wife of the head of household affects expenditure patterns. I demonstrate that my results are not driven by changes in unearned income or wealth as the overall level of expenditure is unaffected. I find, however, that exposure to HSA amendments increases food budget shares, which represents an empirical rejection of the income pooling hypothesis and, in turn, of the unitary model (Attanasio and Lechene (2010)). Finally, I perform two falsification tests and confirm the validity of the identification strategy.

4 A Structural Analysis of Intra-household Bargaining

Is it possible to measure bargaining power? What are the determinants of intra-household resource allocation? How does women's position vary over the life-cycle? I provide answers to these questions in a structural framework. In the spirit of Dunbar et al. (2013), I model the households in the collective framework and include children in the analysis. To better capture the Indian family structure, I extend their model to include households with more than one adult man and one adult woman. In a slight abuse of terminology, I define *nuclear* households with only one male of age 15 and above and one female of age 15 and above, and *non-nuclear* those with more than one adult male or more than one adult female.[19]

4.1 A Collective Model of Indian Households

Let households consist of individuals of three different types t: adult males, m, adult females, f, and children c. F, M and C are the number of adult females, adult males and children, respectively. Households differ according to a set of observable attributes, such as composition, age of household members, location, and other socio-economic characteristics. Household characteristics may affect both preferences and bargaining power within the household. Any characteristics affecting bargaining power and how resources are allocated within the household, but neither preferences nor budget constraints, are defined as *distribution factors* (Browning et al. (2014)). For simplicity of notation, I omit household characteristics and distribution factors while discussing the model and identification.

Each household consumes K types of goods with prices $p = (p^1, ..., p^K)$. Household total expenditure y is set equal to household income. $h = (h^1, ..., h^K)$ is the vector of observed quantities of goods purchased by each household, while $x_t = (x_t^1 ..., x_t^K)$ is the vector of unobserved quantities of goods consumed by an individual of type $t = f, m, c$. I allow for economies of scale in consumption through a linear consumption technology, which converts purchased quantities by the household,

[19] In contrast, a nuclear family is usually defined a family group consisting of a pair of adults and their children, independent of the age of household members.

h, in *private good equivalents*, x. This specific technology assumes the existence of a $K \times K$ matrix A such that $h = A(Fx_f + Mx_m + Cx_c)$.[20]

Each member has a monotonically increasing, continuously twice differentiable and strictly quasi-concave utility function over a bundle of K goods. Let $U_t(x_t)$ be the sub-utility function of individual of type t over her consumption. I assume $U_t(x_t)$ to be the same for all household members of type t, i.e., common to all men, all women and all children, respectively. As further discussed in section 4.2, the choice of restricting the utility functions among individuals of the same type to be the same is data driven. For the same reason, I assume that within a household individuals of the same type are treated equally.[21] Each type t individual's total utility may depend on the utility of other household members, but I assume each type's utility function to be weakly separable over the sub-utility functions for goods, i.e. $\tilde{U}_t = \tilde{U}_t[U_t(x_t), U_{-t}(x_{-t})]$.

Each household maximizes the Bergson-Samuelson social welfare function, $\tilde{U}$, featuring the relative weights of the utility functions of its members:

$$\tilde{U}(U_f, U_m, U_c, p/y) = \sum_{t\in\{f,m,c\}} \mu_t(p/y)\tilde{U}_t \tag{2}$$

where $\mu_t(p/y)$ are the *Pareto weights*. The household program is as follows:

$$\max_{x_f, x_m, x_c, h} \tilde{U}(U_f, U_m, U_c, p/y) \quad \text{such that} \quad \begin{aligned} h &= A(Fx_f + Mx_m + Cx_c) \\ y &= h'p \end{aligned} \tag{3}$$

The solutions to this program provide the bundles of private good equivalents, x_t. Pricing those at the shadow prices $A'p$ gives the *resource share* $\lambda_t = \frac{\Lambda_t}{T}$, that is the fraction of household total resources that are devoted to each individual of type t.

Pareto weights are traditionally interpreted as measures of intra-household bargaining power: the larger is the value of μ_t, the greater is the weight that type t members' preferences receive in the household program. Browning et al. (2013) show that there exists a monotonic correspondence between Pareto weights and resource shares. Moreover, they argue that the latter is a more tractable measure of bargaining power, as it is invariant to unobservable cardinalizations of the utility functions.

Following the standard characterization of collective models, I assume the intra-household allocation to be Pareto efficient.[22] Thus, the household program can be decomposed in two steps:

[20] Suppose that the two members of a nuclear household with no children ride together a motorcycle and, therefore, share the consumption of gasoline, half of the time. Then the consumption of gasoline in private good equivalents is 1.5 times the purchased quantity of gasoline at the household level. Assuming the consumption of gasoline does not depend on consumption of other goods, the kth row of A would consist of 2/3 in the kth column and zero otherwise, such that $h^k = 2/3(x_f + x_m)$. 2/3 represents the level of *publicness* of good k within the household. If the two members ride the motorcycle together all the time, $A_k = 1/2$. For a private good, which is never jointly consumed, $A_k = 1$.

[21] This is an admittedly strong assumption. Distinguishing between individuals of the same type would be possible only if private assignable goods were observable for each individual within types. To the best of my knowledge, no such data is available for India. In estimation, however, all the preference parameters and the resource shares are allowed to vary with a set of household characteristics, including family composition. Thus, everything else equal, women's resource shares in a household where the wife of the head of household and a daughter in law coexist would be different from the resource shares in a household where a wife of the head of household and an unmarried daughter above 15 live together.

[22] See e.g., Chiappori (1988, 1992), and then Browning et al. (1994), Browning and Chiappori (1998), Vermeulen (2002), Lewbel and

the optimal allocation of resources across members and the individual maximization of their own utility function. Conditional on knowing λ_t, each household member chooses x_t as the bundle maximizing U_t subject to a Lindahl type shadow budget constraint $\sum_k A_k p^k x_t^k = \lambda_t y$.[23] By substituting the indirect utility functions $V_t(A'p, \lambda_t y)$ in equation (3), the household program simplifies to the choice of optimal resource shares subject to the constraint that total resources shares must sum to one.

I define a *private* good to be a good that does not have any economies of scale in consumption - e.g., food - and a *private assignable* good to be a private good consumed exclusively by household members of *known* type t - e.g., women, men or children clothing. The household demand functions for the private assignable goods, W_t, are given by:

$$W_t(y, p) = T\lambda_t w_t(A'p, \lambda_t y) = \Lambda_t w_t(A'p, \lambda_t y) \tag{4}$$

where $t = f, m, c$, $T = F, M, C$, and w_t is the demand function of each type t household member when facing her personal shadow budget constraint.

Women's total resource share, $\Lambda_f = F\lambda_f$, is my main object of interest, as it represents the share of total household expenditure consumed by women and provides a measure of the overall bargaining power of adult females.

4.2 Identification of Resource Shares

I identify type t individuals' resource shares using Engel curves of assignable clothing and the methodology developed in Dunbar et al. (2013). An Engel curve describes the relationship between the proportion of household expenditure spent on a good (budget share) and total expenditure, holding prices constant. Dunbar et al. (2013) demonstrate that resource shares are identified under observability of private assignable goods, semi-parametric restrictions imposing similarity of preferences over the private assignable goods, and the assumption that resource shares are independent of expenditure (at least at low levels of y).[24] As I observe type-specific assignable goods, I am only able to retrieve type-specific resource shares.

For simplicity, I assume that each household member has Muellbauer's Piglog preferences over assignable clothing at all levels of expenditure.[25] Under this assumption, the Engel curves for these goods are linear in the logarithm of household expenditure. In a slight abuse of notation, the demand functions for assignable clothing can be written in Engel curve form. In each household

Pendakur (2008), Browning et al. (2013), and Dunbar et al. (2013). While some papers provide evidence in favor of the collective model (e.g., Attanasio and Lechene (2014)), some others works have cast doubt on the assumption that households behave efficiently (e.g., Udry (1996), Angelucci and Garlick (2015) and Tommasi (2016)). In section B.4 of the online Appendix, I use auxiliary data on singles to show that the assumption of Pareto efficiency cannot be rejected in this context.

[23]This result follows directly from the second welfare theorem in an economy with public goods. See Browning et al. (2013) and Browning et al. (2014) for more details.

[24]Menon et al. (2012) show that for Italian households resource shares do not exhibit much dependence on household expenditure, therefore supporting identification of resource shares based on this particular assumption. Moreover, Cherchye et al. (2012a) use detailed data on Dutch households to show that revealed preferences bounds on women's resource shares are independent of total household expenditure. Finally, this restriction still permits resource shares to depend on other variables related to expenditure, such as measures of wealth (in this case land ownership and presence of a salary earner in the household).

[25]See Dunbar et al. (2013) for a discussion of identification of resource shares in a more general framework.

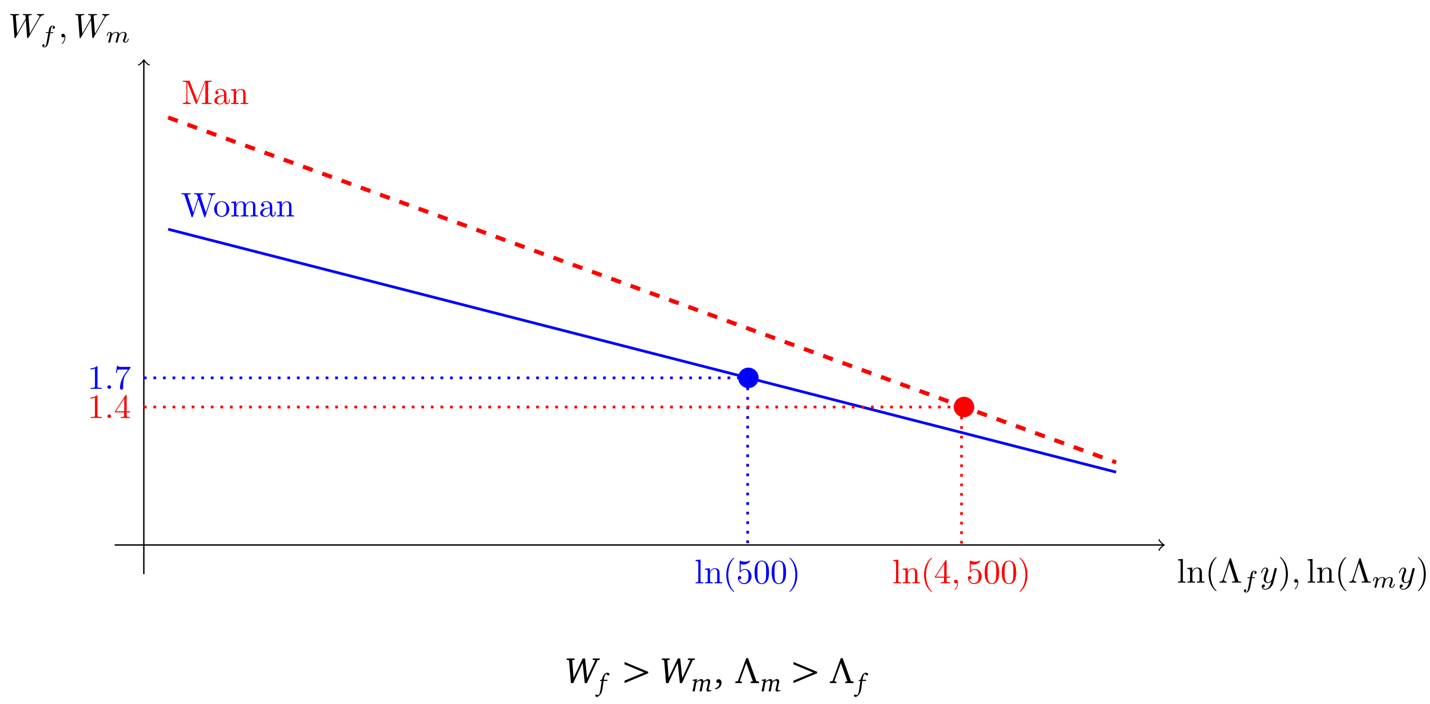


Figure 2: Assignable Clothing Engel Curves: An Illustrative Example

with children they are as follows:

$$\begin{cases} W_f(y) = \alpha_f \Lambda_f + \beta_f \Lambda_f \ln\left(\frac{\Lambda_f y}{F}\right) \\ W_m(y) = \alpha_m \Lambda_m + \beta_m \Lambda_m \ln\left(\frac{\Lambda_m y}{M}\right) \\ W_c(y) = \alpha_c \Lambda_c + \beta_c \Lambda_c \ln\left(\frac{\Lambda_c y}{C}\right) \end{cases} \tag{5}$$

where $W_t(y)$ is the budget share spent on type t's assignable clothing and y is the total household expenditure. α_t and β_t are combinations of underlying preference parameters, while Λ_t is the share of overall resources devoted to type t members ($t = f, m, c$). In the case of households without children, the system contains only two Engel curves, one for women's assignable clothing and one for men's assignable clothing.

Identification of resource shares is obtained by imposing similarities of preferences on private assignable goods across household members and across households. These restrictions allow to identify individual resource shares by comparing household demands for private assignable goods across people within households and across households. In particular, provided that $\beta_f = \beta_m = \beta_c = \beta$, the slopes of the Engel curves in equation (5) can be identified by a linear regression of the household budget shares W_t on a constant term and $\ln y$. $\beta\Lambda_t$ is the slope of type t's private assignable good Engel curve. The slopes are proportional to the unknown resource shares, with the factor of proportionality set by the constraint that the resource shares must sum to one ($\Lambda_f + \Lambda_m + \Lambda_c = 1$).

It is important to note that *budget shares* on assignable clothing and *resource shares* are different objects. Moreover, the relative magnitude of assignable clothing budget shares does not necessarily determine the relative magnitude of resource shares. In particular, that $W_f > W_m > W_c$ does not imply that $\Lambda_f > \Lambda_m > \Lambda_c$, and vice-versa. The following example may help clarifying this point.

An Illustrative Example. Consider the simple case of a nuclear household with no children ($F = M = 1$ and $C = 0$), with total household expenditure equal to 5,000 Rupees and observable budget shares for female and male clothing equal to 1.7 and 1.4, respectively. Let the Engel curves for assignable clothing be as in figure 2. The relationship between assignable clothing budget shares (W_f and W_m, on the vertical axis) and the logarithm of the total expenditure devoted to each type t household member ($\Lambda_t y$, on the horizontal axis) is linear under the functional form assumptions discussed above.[26] By inverting these Engel curves, I can identify two points on the horizontal axis, equal to $\ln(500)$ (≈ 6.21) and $\ln(4,500)$ (≈ 8.41). These, together with the constraint that the resource shares must sum to one, make it possible to compute individuals' resources shares at any level of y. At a total household expenditure of 5,000 Rupees, $\Lambda_f = 0.1$ and $\Lambda_m = 0.9$.

The graph depicts a situation where $W_m < W_f$ and $\Lambda_f < \Lambda_m$. In this specific numerical example, resources are split extremely unequally between the two household members, with the woman getting only 10 percent of the total household expenditure, whereas the budget share spent on female assignable clothing W_f is about 20 percent larger than the share spent on male clothing W_m.

4.3 Data

The 2011-2012 National Sample Survey (NSS) of Consumer Expenditure (68^{th} round) contains detailed data on expenditure of about 102,000 households, together with information on household characteristics, and demographic and other particulars of household members (about 465,000 individuals). Unfortunately, it does not include any health outcome, which prevents the direct investigation of the relationship between individuals' resource shares and their health.

Households are asked to report how much they spent on food, clothing, bedding, and footwear, during the month prior to the survey. The detailed breakdown of clothing expenditure allows me to identify the expenditure on clothing items that are assignable to specific types of household members, i.e., women, men and children. I define expenditure on women assignable clothing as the sum of expenditures on saree, shawls, chaddar, and kurta-pajamas suits for females. For men assignable clothing, I combine expenditure on dhoti, lungi, kurta-pajamas suits for males, pajamas, and salwar. For children, I use expenditure on school uniforms and infant clothing.

The survey also provides information about women's year of birth (but not year of marriage), state of residence, and religion. I construct a variable capturing women's *eligibility* to the Hindu Succession Act (HSA) amendments as the interaction between an indicator variable for being Hindu, Buddhist, Sikh or Jain, and an indicator variable equal to one if a woman was 14 or younger at the time of the amendment in her state and to zero if she was 23 or older.[27] For simplicity, I focus on the eligibility of the wife of the head of household and any effect should be interpreted as an intent-to-treat effect.

[26] The Engel curves displayed in figure 2 feature the average intercepts and slopes obtained by estimating the model using data on households without children under 15. In estimation, intercepts and slopes of the private assignable goods Engel curves are allowed to vary with several observable household characteristics (see section 4.4). Table A4 in the Appendix reports descriptive statistics of the predicted Engel curve slopes. While the estimated slopes are, on average, negative, the maximum estimated slopes are positive.

[27] I use 14 and 23 as they are the 10th and 90th percentiles of women's age at marriage in the NHFS-3 sample discussed in section 3. This variable is therefore fully determined by each woman's religion, year of birth and state. See Heath and Tan (2014).

Table 2: NSS Consumer Expenditure Survey - Descriptive Statistics

	Obs.	Mean	Median	St. Dev.
Expenditure (Rupees):				
Total Expenditure	87,373	8,108.98	6,775.00	5,042.64
Expenditure On Non-Durable Goods	87,373	7,694.28	6,538.33	4,579.95
Expenditure On Durable Goods	87,373	414.70	106.85	1,156.44
Budget Shares:				
Food	87,373	39.24	39.26	9.62
Female Assignable Clothing	87,373	1.37	1.17	1.16
Male Assignable Clothing	87,373	1.68	1.41	1.42
Children Assignable Clothing	87,373	0.51	0.00	0.76
Household Characteristics:				
No. Adult Females	87,373	1.68	1.00	0.85
No. Adult Males	87,373	1.76	1.00	0.92
Fraction of Female Children	57,158	0.45	0.50	0.39
Number of Children Under 5	87,373	1.32	1.00	1.26
$\mathbb{I}$(Daughter in Law)	87,373	0.20	0.00	0.40
$\mathbb{I}$(Unmarried Daughter Above 15)	87,373	0.23	0.00	0.42
$\mathbb{I}$(Widow)	87,373	0.15	0.00	0.35
Avg. Age Men 15 to 79	87,089	37.77	36.00	10.52
Avg. Age Women 15 to 79	87,263	36.96	35.00	10.15
Avg. Age Gap 15 to 79 (Men - Women)	87,005	0.88	3.00	11.15
Avg. Age Children 0 to 14	57,158	7.57	8.00	3.97
$\mathbb{I}$(HSAA Eligible)	74,127	0.12	0.00	0.33
$\mathbb{I}$(Hindu, Buddhist, Sikh, Jain)	87,373	0.79	1.00	0.41
$\mathbb{I}$(Sch. Caste, Sch. Tribe or Other Backward Classes)	87,373	0.69	1.00	0.46
$\mathbb{I}$(Salary Earner)	87,373	0.30	0.00	0.46
$\mathbb{I}$(Land Ownership)	87,373	0.89	1.00	0.31
$\mathbb{I}$(Female Higher Education)	87,373	0.12	0.00	0.32
$\mathbb{I}$(Male Higher Education)	87,373	0.19	0.00	0.39
$\mathbb{I}$(Rural)	87,373	0.61	1.00	0.49
$\mathbb{I}$(North)	87,373	0.31	0.00	0.46
$\mathbb{I}$(East)	87,373	0.20	0.00	0.40
$\mathbb{I}$(North-East)	87,373	0.14	0.00	0.35
$\mathbb{I}$(South)	87,373	0.22	0.00	0.41
$\mathbb{I}$(West)	87,373	0.12	0.00	0.33

Note: Budget shares are multiplied by 100. Budget share on food includes expenditures on cereals, cereals substitutes, pulses and products, milk and products, egg, fish and meat, vegetables, fruits, and processed food. Tea, coffee, mineral water, cold beverages, fruit juices and shake and other beverages, salt and sugar, edible oil and spices are not included. Women's assignable clothing includes expenditures on saree, shawls, chaddar, and kurta-pajamas suits for females; men's assignable clothing includes expenditures on dhoti, lungi, kurta-pajamas suits for males, pajamas, salwar, and cloth for coats, trousers, and suit and for shirt, pajama, kurta, and salwar; children's assignable clothing includes expenditures on expenditure on school uniforms and infant clothing. $\mathbb{I}$(Female Higher Education) and $\mathbb{I}$(Male Higher Education) are indicator variable for higher education (diploma or college) completed by at least one woman or man in the household. North India includes Jammu & Kashmir, Himachal Pradesh, Punjab, Chandigarh, Uttaranchal, Haryana, Delhi, Rajasthan, Uttar Pradesh, and Madhya Pradesh. East India includes West Bengal, Bihar, Jharkhand, Orissa, A & N Islands, and Chattisgarh. North-East India includes Sikkim, Arunachal Pradesh, Assam, Manipur, Meghalaya, Mizoram, Nagaland, and Tripura. South India includes Karnataka, Tamil Nadu, Andhra Pradesh, Kerala, Lakshadweep, and Pondicherry. West India includes Gujarat, Goa, Maharashtra, Daman & Diu, and D & N Haveli.

From these data, I select a sample of 87,373 households as follows. I exclude households with no women or no men above 15 years of age (10 percent of the full sample), households in the top 1 percent of expenditure to eliminate outliers, and households with head or head of household wife under 15 (0.3 percent). For simplicity, I also exclude households with more than 5 women, more than 5 men, or more than 5 children under 15 (1.4 percent) and polygamous households (0.3 percent). Finally, as unusual purchases of clothing items and non-standard expenditure patterns may occur for festivities and ceremonies, I exclude households reporting to have performed any ceremony during the past month (1.7 percent). For most of the items, e.g., food, the data refer to expenditure occurred during the month prior to the survey. For a few items, e.g., clothing, the survey contains information about expenditure occurred over the year prior to the survey. I convert annual into monthly figures for ease of comparison. Unless noted otherwise, budget shares are computed as percentage of total household expenditure, including durables.

Table 2 contains some descriptive statistics. On average, a household's total expenditure is equal to 8,109 Rupees (approximately 120US$). Food represents more than one third of the total expenditure, while assignable clothing budget shares are much smaller.[28] 12 percent of households are eligible to the HSA amendments (HSAA), according to the definition of eligibility discussed above. The average number of adult females and males is 1.68 and 1.76 respectively. Daughters in law are present in 1 out of 5 households; unmarried daughters above 15 and widows are present in 23 percent and 15 percent of households, respectively. Nuclear households represent only 35 percent of the sample; about 1 out of 3 households has no children under age 15. Table A2 in the Appendix presents descriptive statistics for the subsamples of households with and without children under 15 separately.

4.4 Estimation Strategy

I implement the model empirically by adding an error term to each equation in system (5) and by imposing similarity of preferences over private assignable goods, $\beta = \beta_f = \beta_m = \beta_c$. Although not required for the identification of the resource shares, I augment the system of Engel curves of private assignable goods with the inclusion of the household level Engel curve for food. The inclusion of this extra equation has a double motivation. On one hand, as the error terms are likely correlated between the equations, it may improve efficiency. On the other hand, it makes it possible to test whether the food Engel curve is downward sloping in this context.[29] Since the error terms may be correlated across equations, I estimate the system using non-linear Seemingly Unrelated Regression (SUR) method. Non-linear SUR is iterated until the estimated parameters and the covariance matrix settle. Iterated SUR is equivalent to maximum likelihood with multivariate normal errors.

[28]Assignable clothing budget shares for men and women are comparable in magnitude to those in Dunbar et al. (2013) based on expenditure data from the Malawi Integrated Household Survey (IHS2).

[29]This prediction is known as the Engel's law. Although the underlying preference parameters for food cannot be separately identified, both the intercept and the slope of the additional equation in the system can.

I take the following system of equations to the data:

$$\begin{cases} W_{food} = \tilde{\alpha}_{food} + \tilde{\beta}_{food} \ln y + \epsilon_{food} \\ W_f = \alpha_f \Lambda_f + \beta \Lambda_f \ln\left(\frac{\Lambda_f}{F}\right) + \beta \Lambda_f \ln y + \epsilon_f \\ W_m = \alpha_m \Lambda_m + \beta \Lambda_m \ln\left(\frac{\Lambda_m}{M}\right) + \beta \Lambda_m \ln y + \epsilon_m \\ W_c = \alpha_c \Lambda_c + \beta \Lambda_c \ln\left(\frac{\Lambda_c}{C}\right) + \beta \Lambda_c \ln y + \epsilon_c \end{cases} \tag{6}$$

where $\Lambda_c = 1 - \Lambda_f - \Lambda_m$. y is the total household expenditure (in Rupees) reported for the month prior to the survey, and W_t and W_{food} are the budget shares spent on assignable clothing and food, respectively. For households without children under 15, the system includes only the first three equations and $\Lambda_m = 1 - \Lambda_f$.

I account for observable heterogeneity across households by specifying α_t ($t = f, m, c$) and β as linear functions of observable household characteristics (*preference factors*, X). Moreover, Λ_t, $\tilde{\alpha}_{food}$ and $\tilde{\beta}_{food}$ depend linearly on X and one *distribution factor*, d.[30] This characterization renders the system of Engel curves non-linear. The vector $X = (X_1, ..., X_n)$ includes, among other variables, details about the composition of the household, socio-economic characteristics, such as demographic group, religion and land ownership, and polynomials in women's age and in the age gap between genders.[31] It also contains region fixed effects (South, East, West, North-East, North, with West being the excluded category) and a dummy variable for living in rural areas, which may capture unobserved geographical heterogeneity and area specific characteristics, such as price levels.[32] Although distribution factors are not required for identification, I include the eligibility of the wife of the head of the household to the HSA amendments as a factor affecting resource allocation but not preferences.[33]

I estimate models for households with and without children below 15, jointly and separately. Robust standard errors are clustered at the first sampling unit (2001 Census villages in rural areas and 2007-2012 Urban Frame Survey blocks in urban areas). Results are robust to clustering standard errors at the district level.

[30]Since resource shares cannot be disentangled from preference parameters in the food equation, intercept and slope are allowed to depend on d as well. For each type $t = f, m, c$ and $T = F, M, C$, total resource shares are specified as $\Lambda_t = l_{t,0} + l_{t,1}X_1 + ... + l_{t,n}X_n + \tilde{l}d$, where $n = 22$ for households without children, and $n = 25$ for households with children. The same holds true for $\tilde{\alpha}_{food}$ and $\tilde{\beta}_{food}$. α_t, $t = f, m, c$ and β are specified as linear functions of X where again $n = 22$ for households without children, and $n = 25$ for households with children.

[31]Women's age, children's age and age differences are divided by 100 for ease of computation.

[32]The choice of including the region instead of state fixed effects is due to computational tractability.

[33]Legal reforms have been used in the literature as distribution factors. Chiappori et al. (2002), for example, use US divorce laws as distribution factors to study intra-household bargaining and labor supply, while Voena (2015) examines how divorce laws affect couples' intertemporal choices in a dynamic model of household decision-making. Despite being permitted by the Indian legislation, there is a strong social stigma of divorce in India, which renders it an inadequate distribution factor. As HSA amendments reforms only applied to women who got married after the implementation, it is sensible to assume that they do not determine shifts in bargaining power over time and that their effects can be analyzed using a static framework.

4.5 Estimation Results

Table 3 reports the estimated coefficients of the covariates $(X_1, ...X_n, d)$.[34] I refer to these variables as the *possible* determinants of women's resource share, as they are related to bargaining power, but not necessarily in a causal sense. Column 1 reports the estimation results obtained when all households are considered in estimation. In columns 2 and 3, I present the results obtained by estimating separate models for households with and without children under 15.

As expected, household composition matters. Women's resource shares increase with the number of women in the household, and decrease with the number of men. Everything else equal, the presence of an additional woman increases women's resource shares by 4 percentage points in the overall sample, by 3.2 percentage points in households with children and by 5.5 percentage points in households without children. The number of children marginally increases resource shares and the fraction of female children is positively related to Λ_f: if all children are girls, women's resource shares are 1.1 percentage points larger. This result is in line with the findings in Dunbar et al. (2013) and can be attributed to the fact that adult women may be willing (or required) to forgo a higher fraction of household resources in presence of male children, due to son preference. Moreover, the presence of a widow in the household is associated with a smaller resource share for women, especially in households without children, which provides additional support to the previous works documenting the plight of widows in South Asia.[35] Finally, despite the coefficients are not statistically significant, the higher is women's age the lower is the fraction of household's total expenditure devoted to women. This finding is particularly relevant in households without children, suggesting that women's bargaining position inside the household may be tightly related to child rearing.

Household socio-economic characteristics play an important role, too. In particular, being part of Scheduled Caste, Scheduled Tribes, and other disadvantaged social classes is associated with higher women's bargaining power. The same holds true for residing in the North-East states, which is consistent with the presence of a number of matrilineal societies and cultures in these regions (Khasi and Garo societies, for example). In contrast, North Indian women seem to have a much lower bargaining power. Finally, households with more educated women and men devote a larger fraction of their resources to women, while the presence of a salary earner (male, in most cases) is associated with lower women's bargaining power.[36]

The estimated model confirms the importance of the Hindu Succession Act amendments (HSAA) in shaping women's bargaining position within the household. In households where the wife of the head of household is eligible to these reforms, women's resource shares are larger by 1.1 to 2.2 percentage points, depending on the model considered for estimation. These results align with the findings in Roy (2008) and Heath and Tan (2014) on the effects of HSA amendments on self-reported measures of women's autonomy and bargaining power. Specifically, Heath and Tan (2014)

[34]The estimated coefficients of the covariates for men's and children's resource shares and for the preference parameters $\tilde{\alpha}_{food}$, α_t, $t = f, m, c$, $\tilde{\beta}_{food}$, and β are available upon request.

[35]See e.g., Drèze et al. (1990), Chen and Drèze (1995), Drèze and Srinivasan (1997), and the recent work by Anderson and Ray (2015) on missing unmarried women.

[36]While male labor force participation is almost universal, only one woman out of three in India does any non-domestic work and an even smaller fraction is formally employed and work for salary (Fulford (2014) and Heath and Tan (2014)).

Table 3: Determinants of Women's Resource Shares

	Women's Resource Share		
	All Households	With Children < 15 Only	Without Children < 15 Only
	(1) NLSUR	(2) NLSUR	(3) NLSUR
No. Adult Women	0.0396***	0.0319***	0.0552***
	(0.00406)	(0.00473)	(0.00908)
No. Adult Men	-0.0283***	-0.0217***	-0.0267***
	(0.00315)	(0.00364)	(0.00660)
No. Children	0.00553**	0.00592**	-
	(0.00219)	(0.00246)	-
Fraction of Female Children	0.0205***	0.0108*	-
	(0.00563)	(0.00554)	-
$\mathbb{I}$(Daughter in Law)	0.0139**	0.00727	0.0126
	(0.00658)	(0.00714)	(0.0179)
$\mathbb{I}$(Unmarried Daughter above 15)	0.00403	0.00717	-0.00253
	(0.00715)	(0.00803)	(0.0169)
$\mathbb{I}$(Widow)	-0.0136*	-0.0316***	-0.0168
	(0.00814)	(0.00972)	(0.0174)
$\mathbb{I}$(HSAA Eligible)	0.0117***	0.0124**	0.0218**
	(0.00402)	(0.00507)	(0.00932)
$\mathbb{I}$(Hindu, Buddhist, Sikh, Jain)	-0.0362***	-0.00978	-0.0167
	(0.00960)	(0.00808)	(0.0150)
$\mathbb{I}$(SC, ST, Other Backward Caste)	0.0567***	0.0613***	0.0555***
	(0.00802)	(0.00873)	(0.0123)
$\mathbb{I}$(Salary Earner)	-0.0283***	-0.0225***	-0.0126
	(0.00479)	(0.00502)	(0.00995)
$\mathbb{I}$(Land Ownership)	0.00764	0.00432	-0.0155
	(0.00899)	(0.00912)	(0.0180)
$\mathbb{I}$(Female Higher Education)	0.0302***	0.0277***	0.0368**
	(0.00732)	(0.00867)	(0.0159)
$\mathbb{I}$(Male Higher Education)	0.0303***	0.0387***	0.0813***
	(0.00562)	(0.00673)	(0.0126)
$\mathbb{I}$(Rural)	-0.0353***	-0.0300***	-0.0402***
	(0.00667)	(0.00707)	(0.0116)
Avg. Age Diff. (Men 15 to 79 - Women 15 to 79)	0.00202	-0.115**	0.0514
	(0.0404)	(0.0485)	(0.0805)
Avg. Age Women 15 to 79	-0.572	-0.208	-1.632
	(0.597)	(0.801)	(1.144)
(Avg. Age Diff. (Men 15 to 79 - Women 15 to 79))^2	-0.199*	0.129	-0.504***
	(0.112)	(0.139)	(0.188)
(Avg. Age Women 15 to 79)^2	0.959	0.374	2.912
	(1.437)	(2.027)	(2.658)
(Avg. Age Diff. (Men 15 to 79 - Women 15 to 79))^3	0.0456	0.478	-0.705
	(0.514)	(0.741)	(0.762)
(Avg. Age Women 15 to 79)^3	-0.354	-0.262	-1.623
	(1.110)	(1.666)	(1.970)
Avg. Age Children 0 to 14	-0.0710	-0.0151	-
	(0.0488)	(0.0681)	-
$\mathbb{I}$(North)	-0.0785***	-0.0984***	-0.0652***
	(0.0150)	(0.0168)	(0.0232)
$\mathbb{I}$(East)	-0.0141	-0.0171	-0.0234
	(0.0164)	(0.0180)	(0.0254)
$\mathbb{I}$(North-East)	0.0512**	0.0374	0.168***
	(0.0229)	(0.0241)	(0.0284)
$\mathbb{I}$(South)	-0.00814	-0.0254	-0.0537**
	(0.0163)	(0.0181)	(0.0235)
Constant	0.438***	0.298***	0.715***
	(0.0835)	(0.105)	(0.161)
N	73,759	47,262	26,497
LL	-575,246.1	-381,414.4	-185,500.7
No. Parameters	318	318	188

Note: $^*p < 0.10$, $^{**}p < 0.05$, $^{***}p < 0.01$. NSS 68^{th} Round Consumer Expenditure data. Robust standard errors in parentheses. Standard errors clustered at the first sampling unit level. Women's age and age differences are divided by 100 to ease computation. West India is the excluded region.

find that exposure to HSA amendments is associated with a 6.6 percentage point decrease in the probability that a woman has no say in household decisions, a 8.2 percentage point increase in the probability that a woman can go to the market alone, a 6.9 percentage point increase in the probability that a woman can go to a health facility alone, and a 8.3 percentage point increase in the probability that a woman can travel outside the village alone.

I test the hypothesis of equality of coefficients between the models in columns 2 and 3 and find the likelihood ratio test statistics to be larger than the χ^2 critical value. As the null hypothesis is rejected, for the remainder of the paper I focus on the results obtained estimating the two models separately.

Robustness Checks. I perform a series of robustness checks to test the sensitivity of the structural estimates. All results are included in section A.1 of the Appendix. I estimate the system in (6) excluding the food Engel curve, which follows more closely the specification in Dunbar et al. (2013), and show that the results are unchanged (columns 1 to 2 of table A5). Moreover, while the findings discussed in this section are obtained using Engel curves in terms of total expenditure (durable and non-durable goods), I show that they are confirmed when estimating the system of Engel curves in terms of expenditure on non-durable goods only (columns 3 to 4 of table A5). In addition, I demonstrate that similar conclusions can be drawn when estimating the model using data for households with one individuals for each type ($t = f, m, c$) only (columns 1 to 2 of table A6), or when estimating a model for married couples with children (column 3 of table A6). In the latter case, I include dummies for the number of children as shifters of resource shares and preference parameters, which mirrors the specification in Dunbar et al. (2013).[37] Finally, I use auxiliary data on singles to empirically test the assumption of Pareto efficiency and check the validity of the theoretical framework (see section B.4 in the online Appendix).

5 Why Are Older Women Missing?

That health and mortality risk are related is indisputable. Section 3 demonstrates that changes in intra-household bargaining power affect women's health. In this section, I argue that older women are missing in India because their bargaining power diminishes at post-reproductive ages. First, I use the parameter estimates presented in the previous section to predict resource shares and to trace out the age profile of women's bargaining power and access to household resources. Next, using these predictions, I compute poverty rates that account for intra-household inequalities and outline the age distribution of female poverty. Finally, I relate my findings to the age distribution of missing women as estimated by Anderson and Ray (2010) and determine what proportion of their estimates is attributable to intra-household gender inequality and to the consequent gender asymmetry in poverty.

[37] They use data from the Malawi Integrated Household Survey (IHS2) and look at married couples with up to four children.

Table 4: Predicted Resource Shares: Descriptive Statistics

	Reference Households		All Households				
	Estimate (1)	Sd. Error (2)	Mean (3)	Sd. Dev. (4)	Median (5)	Min. (6)	Max. (7)
Panel A: Without Children < 15 Only							
Women's Resource Share $\hat{\Lambda}_f$	0.3710	0.0221	0.4593	0.1136	0.4388	0.1626	1.0000
Men's Resource Share $\hat{\Lambda}_m$	0.6290	0.0221	0.5407	0.1136	0.5612	0.0000	0.8374
Panel B: With Children < 15 Only							
Women's Resource Share $\hat{\Lambda}_f$	0.2275	0.0160	0.3015	0.0726	0.3057	0.0732	0.5873
Men's Resource Share $\hat{\Lambda}_m$	0.3795	0.0339	0.4784	0.1604	0.5147	0.0000	0.7548
Children's Resource Share $\hat{\Lambda}_c$	0.3834	0.0333	0.2200	0.1129	0.1793	0.0100	0.5489

Note: Reference households are nuclear households for which all other covariates are equal to their median values; see table A2.

5.1 Intra-household Allocation, Gender and Age

Using the estimates in table 3, and the analogous estimates for men and children, I predict women's, men's and children's resource shares for each household.[38] Panel A and B of table 4 contain descriptive statistics for the predicted resource shares obtained estimating the two models, without and with children under 15, respectively. In column 1 and 2, I report the prediction and the corresponding standard error for the reference household in each sample (nuclear households for which all other covariates are equal to their median values; see table A2). Column 3 to 5 show the mean, the standard deviation, and the median of the predicted values. These take into account the empirical distributions of the covariates $(X_1, ..., X_n, d)$. All predicted resource shares fall within the 0 to 1 interval.

In both specifications, the resource share for women is lower than that for men: $\hat{\Lambda}_f$ is slightly more than half of $\hat{\Lambda}_m$ in reference households without children, and slightly less than two thirds of $\hat{\Lambda}_m$ in reference households with children. Moreover, the distribution of households characteristics in the sample matters. While the mean predicted shares confirm the presence of gender inequality within households, some differences emerge. In households with children, women' resource shares are on average 64 percent of men's. However, asymmetries are on average less extreme in household without children under 15 with women's resource shares being on average 85 percent of men's. In figure A1 in the Appendix, households are sorted left to right by total expenditure and the estimated women's resource shares are plotted against y. Both in households with and without children, shares look uncorrelated to expenditure. This finding lends empirical support to the assumption that resource shares do not vary with the logarithm of total expenditure, which is required for identification.

I exploit the cross-sectional variation in women's age to investigate how female bargaining power varies with age.[39] For each $a = 15, ..., 79$, I compute $\hat{\Lambda}_f^a$ as the mean predicted women's resource

[38] For each type $t = f, m$ and $T = F, M$, total resource shares are computed as $\hat{\Lambda}_t = \hat{l}_{t,0} + \hat{l}_{t,1}X_1 + ... + \hat{l}_{t,n}X_n + \hat{\hat{l}}d$, where $n = 22$ for households without children, and $n = 25$ for households with children. $\hat{\Lambda}_m = 1 - \hat{\Lambda}_f$ in households without children, and $\hat{\Lambda}_c = 1 - \hat{\Lambda}_f - \hat{\Lambda}_m$ in households with children.

[39] Browning et al. (2013) show that there is a monotonic relationship between standard measures of bargaining power, i.e., the Pareto weights,

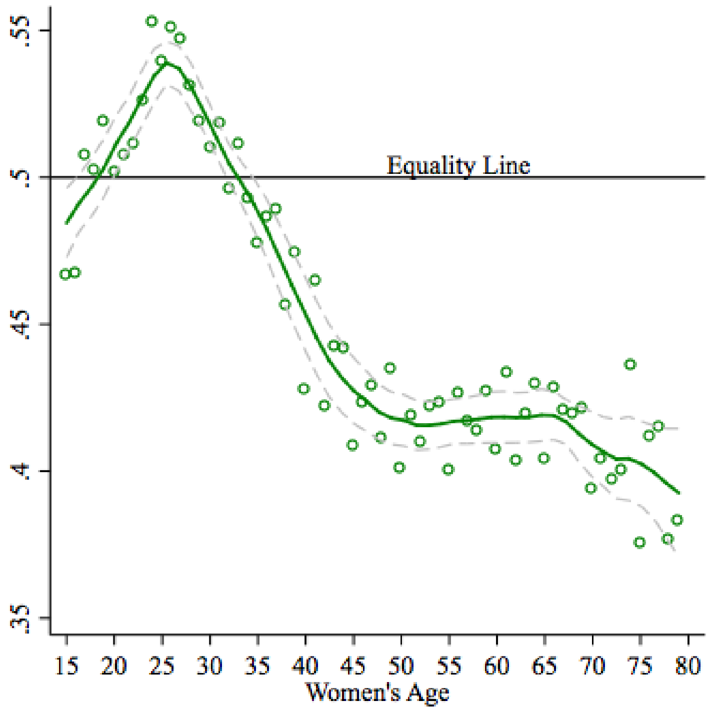
share among all households with women's average age equal to a. Figure 3 shows the average predicted women's resource share against women's age in households without children (panel A) and with children (panel B). The solid line is a running mean, while the dashed lines display the 95 percent confidence intervals for the smoothed values.

The reader should note the different vertical axis scales when comparing the two graphs: total resources are divided among three types of individuals in households with children, while they are shared among two types in households without children. In both cases, women experience a decay in their resource shares over the life-cycle, but the timing seems to differ between the two groups. The presence of children smoothes out the decline in women's bargaining power, which is consistent with the traditional view of women's main purpose of caregiving and child rearing. At post-reproductive ages, the model predictions indicate that women's resource shares are as low as 0.37 in households without children and 0.2 in households with children.

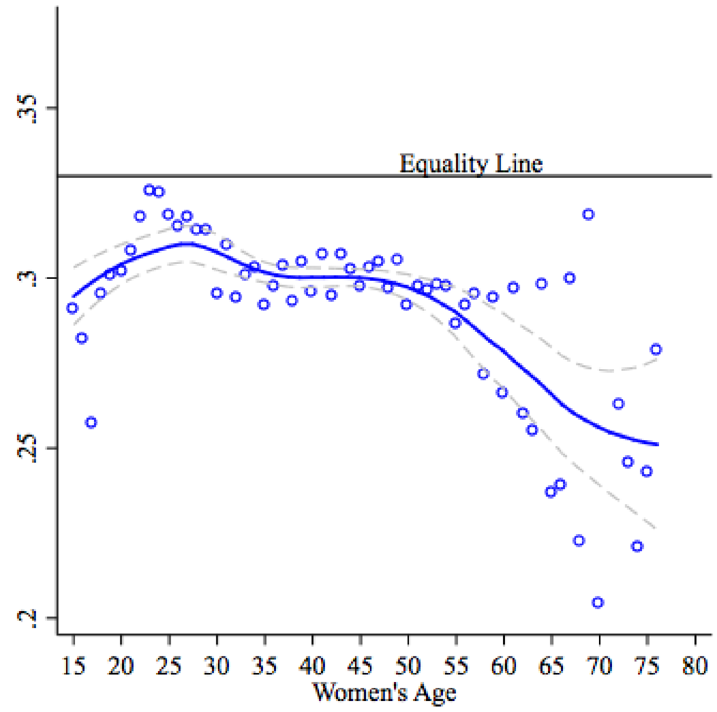
Women and men living in the same household, however, are often not of the same age. A perhaps more insightful exercise is the investigation of the relationship between female bargaining power *relative* to that of males over women's age. In panel C of figure 3, I combine the estimation results for the two models (with and without children) and plot the ratio between predicted resource shares for women and men, $\hat{\Lambda}_f/\hat{\Lambda}_m$, against women's age. The combination of the two sets of estimates accounts for the fact that the distribution of women's age is different in households with and without children. Moreover, as total resources are divided among three types of individuals in the former, while they are shared among two types in the latter, it simplifies the interpretation of the results. A *resource share ratio* equal to 1 indicates no gender asymmetry in intra-household resource allocation, independent of the presence of children: e.g., the ratio will take the same value in households with children where men and women receive both 30 percent of total resources, and in households without children where men and women receive both 50 percent. During women's core reproductive ages, allocation of resources between adult females and males is symmetric, even slightly biased towards women. However, women's resource shares relative to men's decline steadily at post-reproductive ages. In households with women's average age above 45, the resource shares ratio is particularly skewed, with women getting as low as 60 percent of men's resources.

Robustness Checks. I check the robustness of these patterns in several ways. First, I trace out the age profile of women's bargaining power in reference households, i.e., nuclear households with covariates set at their median values, and show that it leads to similar conclusions (see figure A2 in the Appendix). This result is reassuring, as it shows that my findings are not driven by any correlation between women's age and other household characteristics. Second, I repeat my calculations focusing on nuclear households only. For these households women's average age equals the age of the unique woman in the family (see figure A3). Since my results are confirmed, I argue that the aggregation of adult females within a single category is not the main driver of my findings. Finally, to confirm that I am indeed capturing changes in bargaining power across ages, I exploit

and resource shares. In order to interpret changes in resource shares across ages as changes in bargaining power, I assume this relationship to be age invariant.



(A) Households Without Children, $\hat{\Lambda}_f$



(B) Households With Children, $\hat{\Lambda}_f$

(C) Resource Share Ratio ($\hat{\Lambda}_f/\hat{\Lambda}_m$)

Note: Mean predicted women's resource share among households with women's average age equal to $a = 15, ..., 79$ in panels A and B. Mean predicted resource share ratio among households with women's average age equal to $a = 15, ..., 79$ in panel C. The solid line is a running mean. The dashed lines are the 95 percent pointwise confidence interval for the smoothed values.

Figure 3: Average Predicted Women's Resource Shares and Age

the variation in women's eligibility to the Hindu Succession Act reforms. For ease of interpretation, I focus on nuclear households only. I compare households where the woman is eligible to the HSA amendments with those where the woman is not. As expected, women's resource shares are higher for women exposed to the reforms and this difference is particularly significant at younger ages, where a larger fraction of women in the sample are eligible to the amendments (see figure A4).[40]

I also investigate differences between genders by comparing the age profile of women's and men's resource shares. Figure A5 in the Appendix displays the mean predicted women's resource share among all households with women's average age equal to $a = 15, ..., 79$, together with the mean predicted men's resource share among households with men's average age equal to $a = 15, ..., 79$. The comparison between the profile of women's and men's resource shares over the life-cycle indicates that intra-household allocation is biased towards men at all ages. Moreover, this asymmetry becomes more prominent at post-reproductive ages. These findings hold when focusing on nuclear households only (see figure A6).

5.2 Poverty, Gender and Age

Understanding how intra-household gender inequality affects aggregate measures of well-being is of primary policy interest. Moreover, as morbidity and mortality rates are higher in poverty, evidence of a higher poverty incidence among women than men at older ages would provide additional support for my hypothesis that intra-household inequality is responsible for excess female mortality. I use the model estimates to construct poverty rates that take into account *unequal* resource allocation within the household. These are different from standard poverty measures in that they do not assume that each household member gets an *equal* share of household resources. In other words, I allow for the presence, within each household, of individuals who live in poverty and of individuals who do not.

I compute person level expenditures as the product between total household expenditure and the individual resource shares predicted by the model: $\hat{\lambda}_t y = \frac{\hat{\Lambda}_t y}{T}$ $(t = f, m, c,\ T = F, M, C)$. I then construct poverty head count ratios by comparing person level expenditures to poverty lines. I consider the thresholds set by the World Bank for *extreme* poverty (1.90 US$/day) and *average* poverty (3.10 US$/day).[41] As in Dunbar et al. (2013), I set the poverty lines for children to 60 percent of adults' to take into account the fact that children may have lower needs, while I use the same poverty lines for men and women. Section B.5 of the online Appendix investigates deviations from this last assumption.

By definition, the ratio of female to male poverty rates provides a measure of female poverty relative to that of males. I call this measure the *poverty sex ratio*. A ratio equal to 1 indicates no gender asymmetry in poverty, while a ratio larger than 1 indicates that female poverty is higher

[40] As the amendments were enacted in different states at different times and applied only to women who got married after their implementation, no woman above 50 is eligible in both samples (nuclear households with and without children).

[41] Since October 2015, the World Bank uses updated international poverty lines of US$1.90/day and US$3.10/day, which incorporate new information on differences in the cost of living across countries (the PPP exchange rates). The new lines preserve the real purchasing power of the previous lines of 1.25US$/day in 2005 prices and US$2/day in 2005 prices in the world's poorest countries. The poverty rate at US$1.25 (US$2) is the proportion of the sample population living on less than US$1.25 (US$2) per day, adjusted for purchasing power parity (PPP).

Table 5: Poverty Head Count Ratios

	Model Predictions (Unequal Sharing)				Equal Sharing
	Women	Men	Children	Poverty Sex Ratio	All
	(1)	(2)	(3)	(4)	(5)
Panel A: Household Level Poverty Rates					
1.90 US$/day	0.1932	0.1306	0.3094	1.4793	0.1486
3.10 US$/day	0.4767	0.3049	0.5381	1.5635	0.4694
Panel B: Individual Level Poverty Rates					
1.90 US$/day	0.2442	0.1621	0.4222	1.5065	0.1778
3.10 US$/day	0.5444	0.3764	0.6518	1.4463	0.5239

than that of males, and can therefore be interpreted as *excess female poverty*.

Table 5 reports the poverty estimates. Panel A shows the fraction of households in the sample living in poverty. Column 1 to 3 report the proportions of households with women, men, or children living below poverty line when the model predictions are used to compute person level expenditures. Column 4 shows the implied poverty sex ratio as defined above. Column 5 reports the poverty rates obtained under the assumption that each household member gets an equal share of household resources. In panel B, I present the poverty rates estimated at the individual level, i.e., the fraction of all individuals, or women, men and children, separately, who live in poverty. This distinction is quite crucial, as poor and non poor households may have systematically different size and composition.[42]

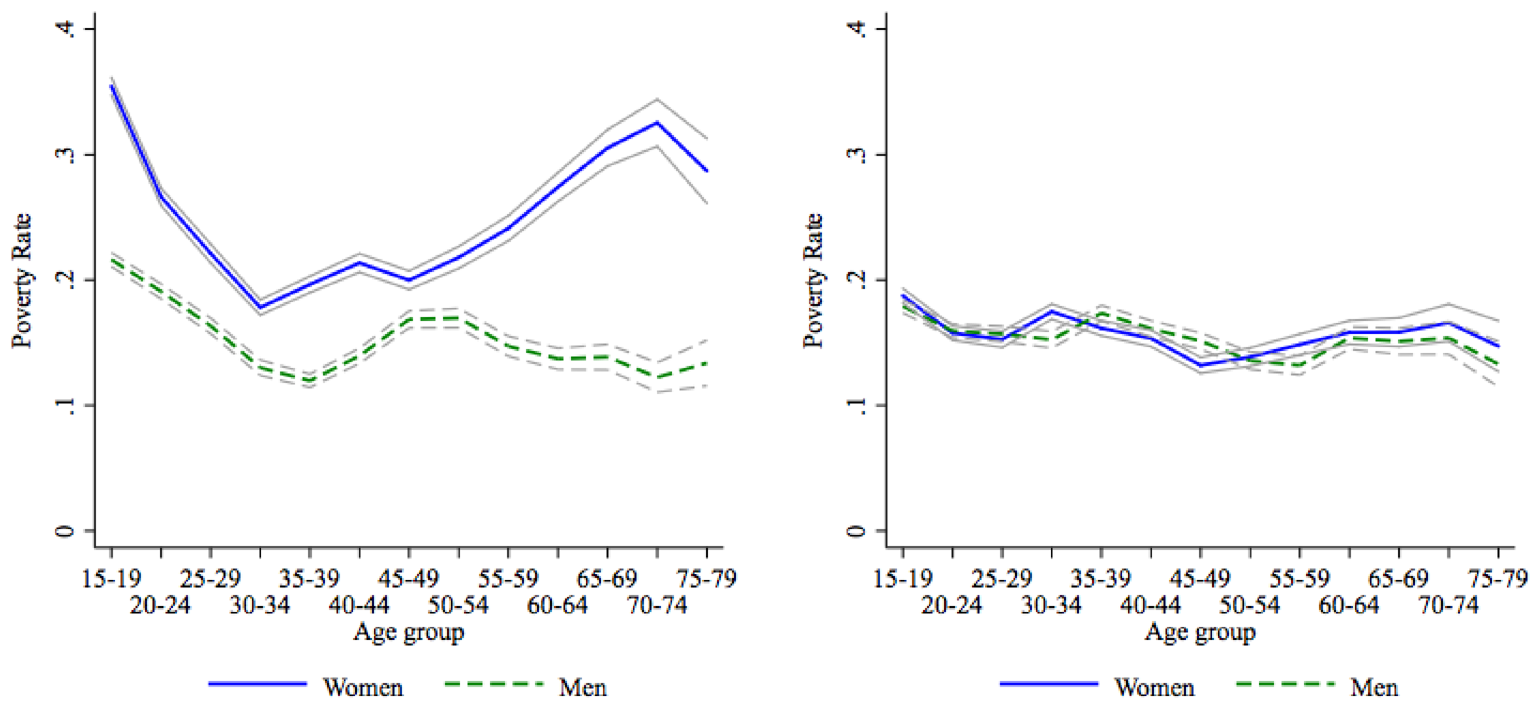
The poverty estimates indicate that there are more households with women living below poverty line than with poor men and that there are more women living below poverty line than men. In particular, 19 percent of households in my sample have women living in extreme poverty, while 13 percent of households have men living on less than US$1.90/day. In terms of individual head count ratios, 24 percent of women live below the extreme poverty cutoff, while 16 percent of men do. Similar gender patterns hold when the alternative poverty line is considered: about 54 percent of women live with less than US$3.10/day, while about one man out of three lives below this threshold. Children poverty rates are the highest, with 42 percent of children living in extreme poverty.[43] Finally, my poverty estimates provide evidence of excess female poverty, as the poverty sex ratios are largely above 1.

While I do not formally present robustness analysis along this dimension, these patterns are robust to the use of national poverty lines.[44] I do not wish to stress the absolute levels of poverty too much as they depend on the relative needs of each household member and on the modeling assumptions discussed in section 4. However, my findings suggest that intra-household inequalities should be taken into account when measuring poverty and evaluating the effect of policies to

[42] As a comparison, the latest World Bank estimates (2011) poverty head count ratios are 21.3 and 58 percent of individuals live below the US$1.90/day and US$3.10/day poverty lines at 2005 international prices, respectively. Any difference is not surprising since a selected sample is used in estimation. See section 4.3 for more details.

[43] This is in line with the findings by Dunbar et al. (2013). Understanding the mechanisms driving this phenomenon, however, goes beyond the scope of this paper and is an open area of research.

[44] See Planning Commission (2014).



(A) Model Predictions (Unequal Sharing)

(B) Equal Sharing

Note: The graphs show the fraction of females or males in each age group living below poverty line. Individuals from all households in the sample are used for calculations. Per capita expenditures are computed using the model predictions in panel A. In panel B, I assume that household expenditure is split equally among household members. Per capita expenditures are compared to the US$1.90/day poverty line.

Figure 4: Poverty Rates By Gender and Age (US$1.90/day)

alleviate it.

To investigate the distribution of male and female poverty across ages, I compute gender-specific head count ratios within thirteen 5-year age groups, from 15-19 to 75-79. As above, I use the model estimates to calculate per capita expenditures. I then compute gender-age specific poverty rates as the fraction of females or males in each age group living below poverty line. Figure 4 shows gender specific poverty rates across age groups, together with the corresponding 95 percent confidence intervals. Poverty rates are based on the World Bank US$1.90/day poverty threshold.

There are at least three features to note. First, the gender-age specific poverty estimates confirm that poverty calculations are drastically affected by the inclusion of intra-household gender asymmetries: female poverty rates are higher at all ages when unequal distribution is accounted for (panel A), whereas almost no difference can be detected when equal distribution of resources is assumed (panel B). In this case, any difference between female and male poverty rates is due to different household and age group compositions. Second, standard poverty estimates may suggest that male and female poverty rates are relatively stable across ages. In contrast, my estimates unveil an interesting pattern: while male poverty is roughly constant over age, the relationship between female poverty and age is U-shaped, with peaks in the 15-19 and 70-74 age groups.[45] Finally, the gap between female and male poverty rates widens dramatically at ages 45-49 and above, indi-

[45]The empirical evidence in the existing literature on the relationship between poverty and age is mixed. Some studies suggest the existence of a U-shaped relationship of age and poverty, with elderly population facing a higher incidence of poverty compared to other groups (Barrientos et al. (2003)). Other studies document that poverty among elderly households is lower than that of non-elderly households, mainly due to survival bias (Deaton and Paxson (1995)). Finally, Gasparini et al. (2007) show that in countries with weak social security systems, there is no significant difference between old age poverty and the overall poverty rates, while in countries with a well developed pension system, poverty rates are lower for the elderly than for other groups.

cating that female poverty relative to males' is particularly high at post-reproductive ages. These patterns are confirmed when the US$3.10/day poverty threshold is considered (see figure A7 in the Appendix).

5.3 Excess Female Poverty and Mortality

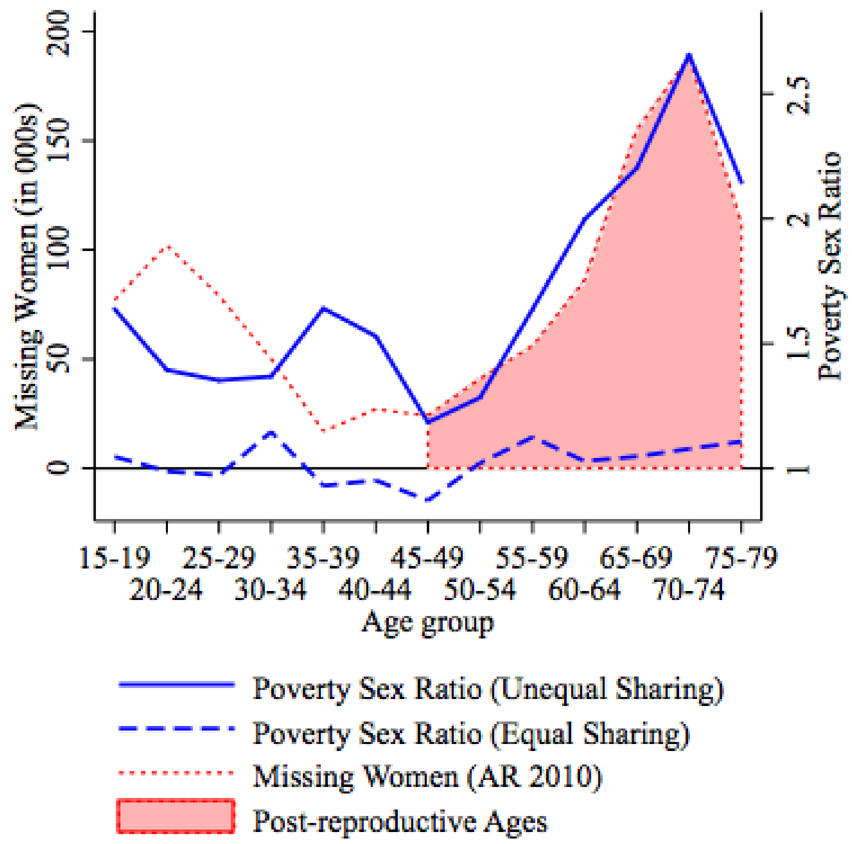
The poverty analysis discussed above indicates that gender inequality in resource allocation inside the household implies gender asymmetries in poverty. The fraction of women in the sample living in extreme poverty (under US$1.90/day) is 0.24 when the model predictions are used and unequal distribution of resources within the household is taken into account. Under the assumption of equal intra-household sharing, 16 percent of women live below the extreme poverty threshold, indicating that granting equal access to resources across genders could reduce the number of women living below US$1.90/day by one third. When focusing on post-reproductive ages, equal sharing of household resources between genders is associated with a reduction in the number of extremely poor women by more than 40 percent.

As poverty and mortality are tightly linked, these findings corroborate my hypothesis that excess female mortality at older ages in India can be explained by the decrease in women's bargaining power and access to resources over the life-cylce. Figure 4 provides a graphical illustration of this claim. The solid line displays the poverty sex ratio, that is the ratio between female and male poverty rates in each age group. As above, a ratio larger than 1 indicates that female poverty is higher than that of males and can be interpreted as excess female poverty. The dotted line plots the age distribution of missing women estimated by Anderson and Ray (2010). As in figure 1, the shaded area represents missing women at post-reproductive ages. For simplicity and consistency with the model estimates, I here focus on age groups 15-19 to 75-79.

When unequal allocation of resources within the households is taken into account and the model predictions are used for calculation (solid line), the age profile of excess female poverty matches the age distribution of missing women remarkably well. This match is nearly exact at post-reproductive ages, when the correlation coefficient between excess female mortality and excess female poverty is equal to 0.96. Not surprisingly, there is almost no evidence of excess female poverty when equal sharing of household resources is assumed (dashed line). The area between the solid and the dashed lines represents the reduction in female poverty relative to males' that is achievable by granting equal allocation of household resources between genders. At post-reproductive ages, an up to 94 percent reduction in excess female poverty in the sample can be obtained by removing intra-household asymmetries.[46]

One should be cautious when trying to translate these results in terms of excess female mortality. While several alternative approaches could be used, I here proceed in the possibly simplest way. I take the correlation shown in figure 4 seriously and assume the relationship between excess female mortality and excess female poverty to be linear and independent of age (see figure A8 in the

[46]The difference between the areas below the solid and the dashed line at post-reproductive ages is equal to 5.2216, which is about 94 percent of the total area between 1 and the solid line (5.5265).



Note: The graph shows the fraction of females poverty rate to male poverty rate in each age group. Individuals from all households in the sample are used for calculations. The underlying gender-age specific poverty rates are displayed in figure 4 and are calculated using the US$1.90/day poverty line.

Figure 5: Poverty Sex Ratio and Missing Women by Age (US$1.90/day)

Appendix).[47] For each age category, I predict how many missing women would be there in absence of excess female poverty as the intercept of a regression line of excess female deaths on excess female poverty. A one unit reduction in excess female poverty is associated with a decrease in the number of missing women by about 97,465. The R-squared of the simple linear regression model is equal to 0.68 and the estimated intercept is 10,237. The predicted number of excess female deaths in the absence of excess female poverty is therefore about 71,659, while the total number of excess female deaths at ages 45 to 79 estimated by Anderson and Ray (2010) is 662,000.[48] Thus, this back-of-the-envelope calculation suggests that up to 89 percent of missing women at post-reproductive ages can be attributed to the decrease in women's bargaining power over the life-cycle, and to the consequent increase in female poverty at older ages relative to that of males. Alternatively, it is possible to calculate the number of excess female deaths in a situation with no gender inequalities in resource allocation in all households (where the poverty sex ratio is equal to the dashed line in figure 5). In this alternative scenario, the predicted number of missing women at post-reproductive ages is about 99,186, that is 85 percent lower than the estimated excess female deaths.

As other forces may obviously be playing a role and the link between poverty and mortality is likely to vary with age, these magnitudes should be interpreted with caution. Nonetheless, my analysis indicates that a potentially sizable reduction in excess female mortality at older ages can be achieved by alleviating the problem of gender asymmetries within the household and the conse-

[47]Further work however should investigate in more details the link between relative poverty and mortality.
[48]10,237 × 7, where 7 is the number of post-reproductive age groups.

quent asymmetries in poverty.

6 Counterfactual Policy Experiment

As they reinforce individuals' bargaining power and improve access to household resources, inheritance rights may affect female poverty and, in turn, female mortality. The Hindu Succession Act amendments that equalized inheritance rights between genders were enacted in different states at different times and applied only to women who got married after their implementation. A large fraction of women in the sample is therefore excluded. Especially after the nation-wide amendment of 2005, understanding the role of these reforms in shaping female poverty and risk of mortality is of primary policy interest.

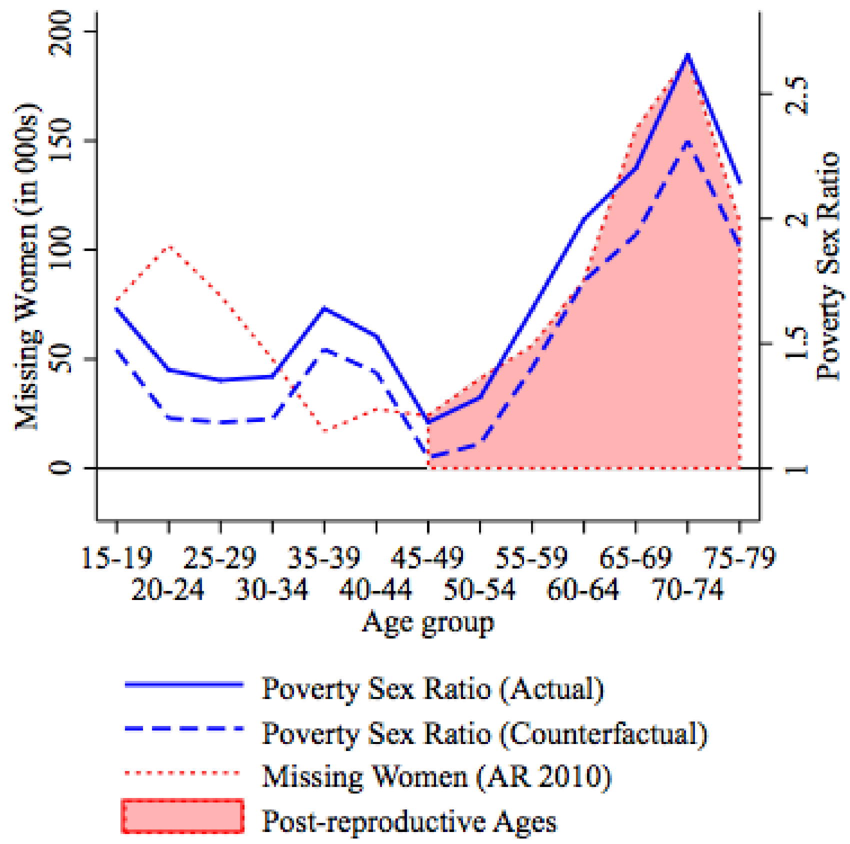
I consider the counterfactual scenario of all women (of all religious affiliations and ages) being exposed to the reforms. More precisely, due to the data limitations discussed above, this hypothetical scenario refers to a situation in which the wife of the head of household is *eligible* to the HSA amendments in all households. The model estimates indicate that eligibility of the wife of the head of household to the HSA amendments increases women's overall resource shares by 1.2 percentage points in households with children and 2.2 percentage points in households without children.[49] I compute counterfactual female and male resource shares and use these predictions to calculate counterfactual poverty rates for each age group 15-19 to 75-79. Figure 6 shows the actual (solid line) and counterfactual (dashed line) poverty sex ratios across age groups and the age distribution of missing women estimated by Anderson and Ray (2010). The reader should note that only one resource share is retrievable for each household member type. Moreover, the distribution factor is not interacted with any other covariate, for simplicity. Thus, HSAA eligibility affects individuals of the same type by the same amount and the difference between the actual and the counterfactual poverty sex ratios is therefore fairly constant at all age groups.

I find that granting access to equal inheritance rights for all women in the sample significantly reduces female poverty. The fraction of women in poverty is 9 percent lower in the counterfactual scenario. Moreover, analogously to figure 5, the area between the solid and the dashed lines can be interpreted as the reduction in female poverty relative to males' that can be achieved by granting equal inheritance rights to all women in the sample. At post-reproductive ages, a 27 percent reduction in excess female poverty can be obtained by equalizing inheritance rights across genders.[50]

When taking all the caveats discussed in section 5.3 into account, it is possible to predict the number of missing women in this counterfactual scenario (when the poverty sex ratio is equal to the dashed line in figure 6). The predicted number of excess female deaths under equal inheritance rights for all women is about 504,000, suggesting that up to a 24 percent reduction in the number

[49] I calculate counterfactual individual resource shares for women equal to $(\hat{\Lambda}_f + 0.0218)/F$ in households without children and to $(\hat{\Lambda}_f + 0.0124)/F$ in household with children. Counterfactual individual resources shares for men are equal to $(1 - \hat{\Lambda}_f - 0.0218)/M$ in households without children and to $(\hat{\Lambda}_m + 0.0031)/M$ in household with children, where 0.0031 is the estimated effect of HSAA eligibility on Λ_m. Figure A9 in the Appendix shows the age profiles of the actual and counterfactual women's resource shares in households with and without children.

[50] The difference between the areas below the sold and the dashed line at post-reproductive ages is equal to 1.4983, which is about 27 percent of the total area between 1 and the solid line (5.5265).



Note: The graph shows the fraction of females poverty rate to male poverty rate in each age group. Individuals from all households in the sample are used for calculations. The underlying actual and counterfactual poverty rates are calculated using the US$1.90/day poverty line.

Figure 6: Counterfactual Experiment: HSA Amendment and Poverty Sex Ratio

of excess female deaths at post-reproductive ages could be obtained by granting equal inheritance rights to all women, of all ages and religions.

7 Conclusion

> *"At older ages, excess female deaths may stem from unequal treatment, but the notion needs to be amplified."* (Anderson and Ray (2012), p. 94)

The present paper focuses on gender asymmetries in intra-household bargaining power and access to household resources as one form of unequal treatment. I show that a large portion of the missing women at post-reproductive ages estimated by Anderson and Ray (2010) can be explained by inequalities in intra-household resource allocation and by the consequent gender asymmetries in poverty. The emphasis on intra-household allocation is motivated by the existence of a positive causal link between women's bargaining power and their health. I demonstrate this fact by analyzing the effect of amendments in the Indian inheritance legislation on a set of women's health outcomes. These reforms equalized inheritance rights between genders and therefore represent a source of exogenous variation in women's position inside the household.

I provide a structural model for estimating women's bargaining power, defined as the fraction of total household expenditure that is consumed by women, and for analyzing its determinants. The model predictions indicate that the allocation of resources between women and men is symmetric during women's core reproductive ages, while the share of household resources devoted to women declines significantly at post-reproductive ages. One consequence of this fact is that at older ages poverty rates are significantly higher among women than among men. Standard per capita poverty measures, which by construction ignore intra-household inequality, are unable to unveil this pattern.

As documented by the reduced-form analysis and the counterfactual exercises, policies aimed at promoting equality within households, such as improving inheritance rights for women, can have a large impact on female health, poverty and, in turn, mortality. Future research should focus on identifying alternative mechanisms generating excess female mortality at post-reproductive ages and on evaluating policies to successfully tackle the problem of excess female poverty, especially among the elderly. Moreover, subsequent work should investigate the effects of the HSA amendments on marital sorting and partner matching. As standard practice in the collective literature, I take the match as given, but interesting insight might arise from relaxing this assumption. Finally, in the spirit of Mazzocco (2007), Mazzocco et al. (2014), and Voena (2015), a promising avenue of research is the investigation of the age profile of women's bargaining power using an inter-temporal model in which household members cannot commit to the future allocation of resources.

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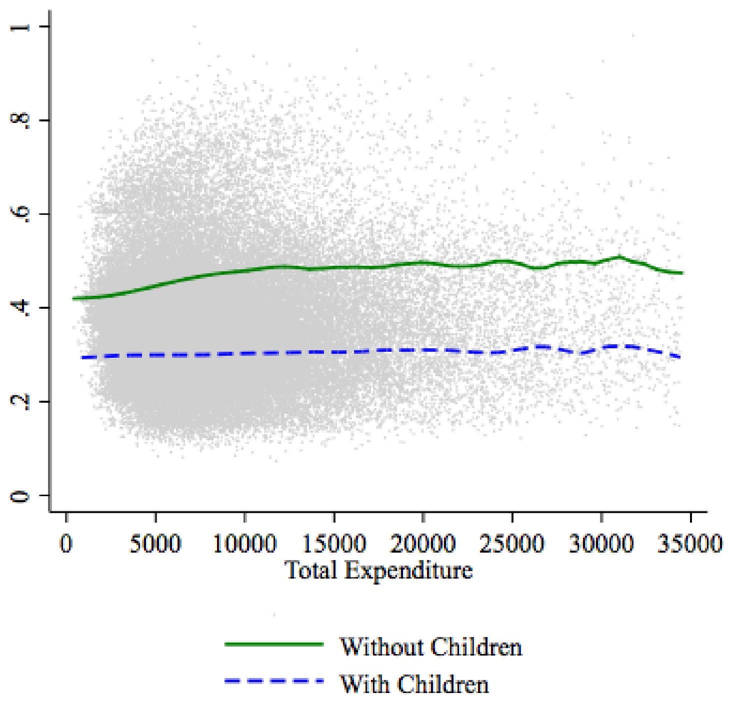
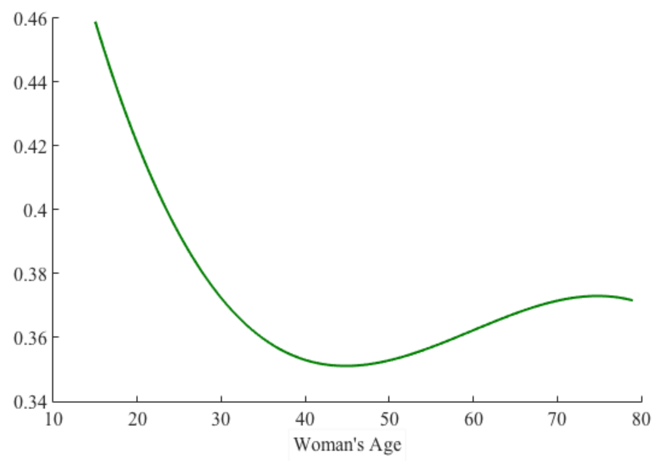
A Appendix

A.1 Additional Figures and Tables

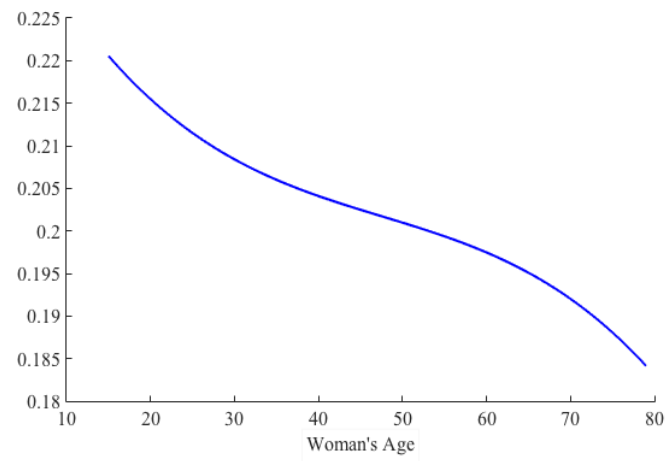


Note: Local mean smoothing (kernel regression).

Figure A1: Women Resource Shares and Total Expenditure



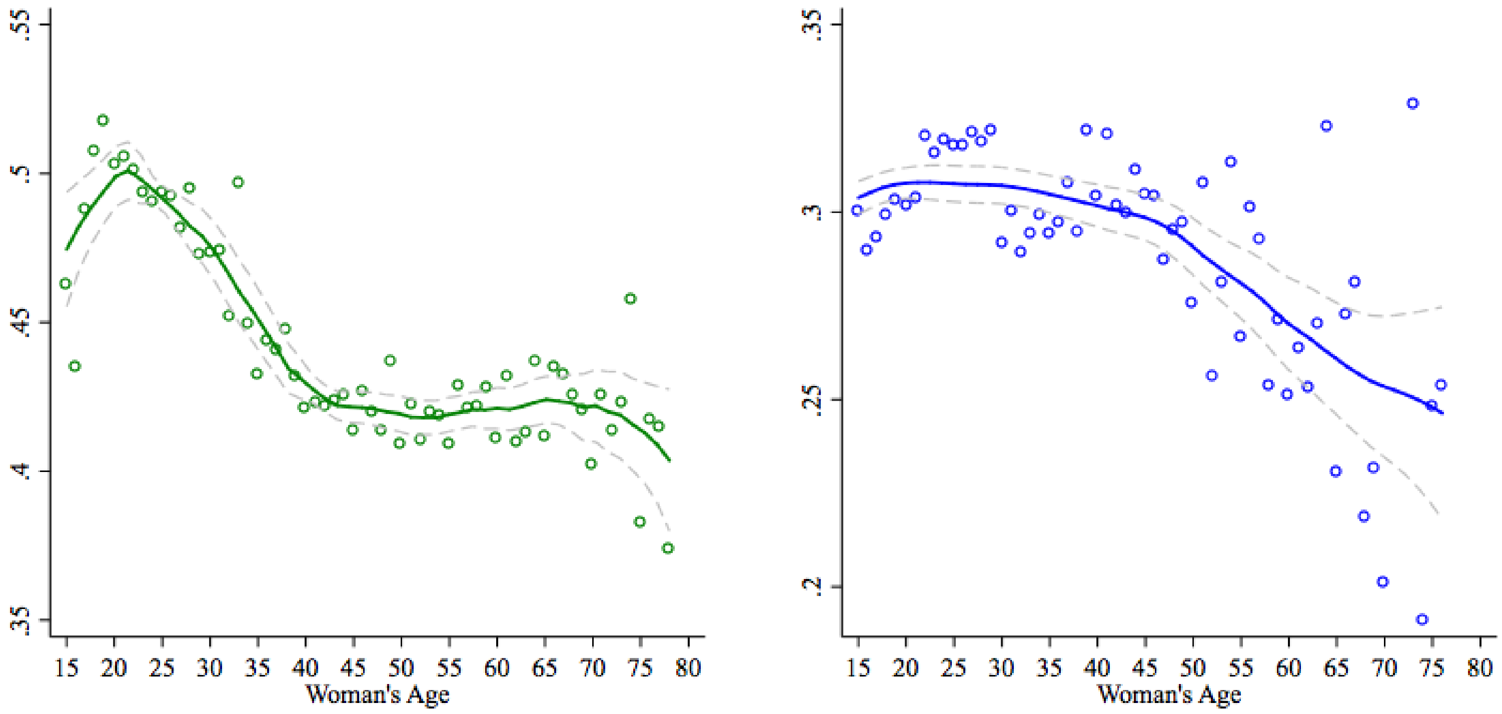
(A) Households Without Children, $\hat{\Lambda}_f$



(B) Households With Children, $\hat{\Lambda}_f$

Note: Third order polynomials in the woman's age. All covariates set to their median values. Nuclear households.

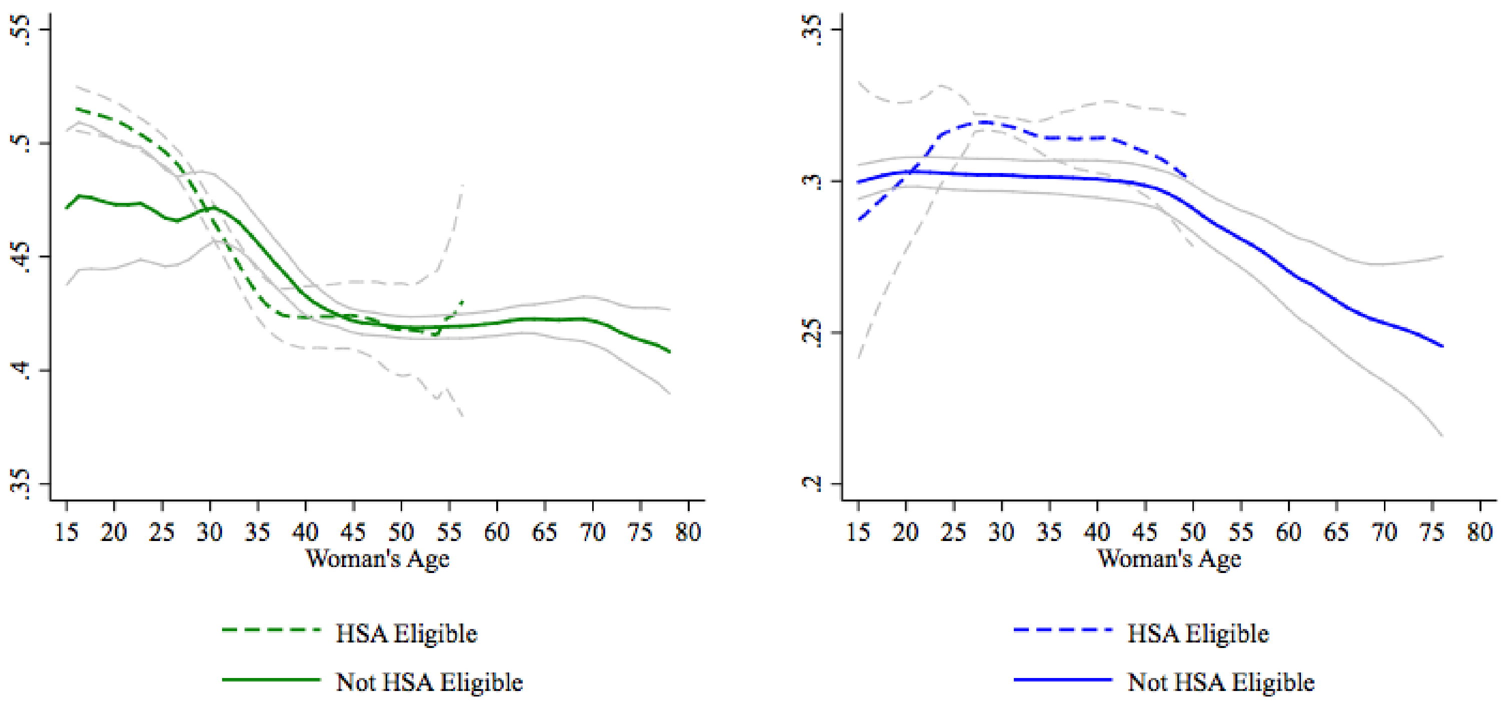
Figure A2: Predicted Women's Resource Shares and Age (Reference Households)



(A) Without Children, $\hat{\Lambda}_f$ (B) With Children, $\hat{\Lambda}_f$

Note: Nuclear households only. Mean predicted women's resource share among households with woman's age equal to $a = 15, \ldots, 79$. The solid line is a running mean. The dashed lines are the 95 percent pointwise confidence interval for the smoothed values.

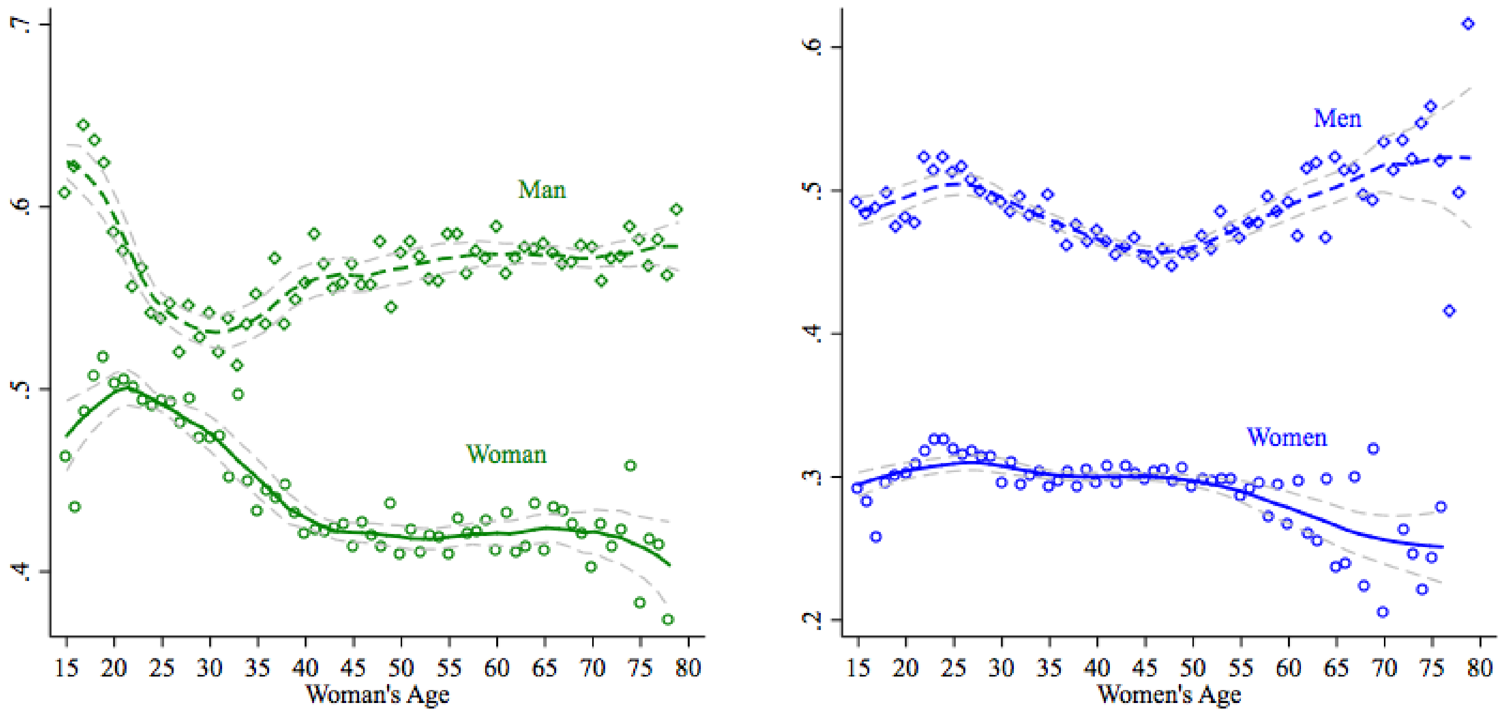
Figure A3: Average Predicted Women's Resource Shares and Age (Nuclear Households Only)



(A) Without Children, $\hat{\Lambda}_f$ (B) With Children, $\hat{\Lambda}_f$

Note: Nuclear households only. Mean predicted women's (men's) resource share among households with woman's age equal to $a = 15, \ldots, 79$. The solid line is a running mean. The dashed lines are the 95 percent pointwise confidence interval for the smoothed values.

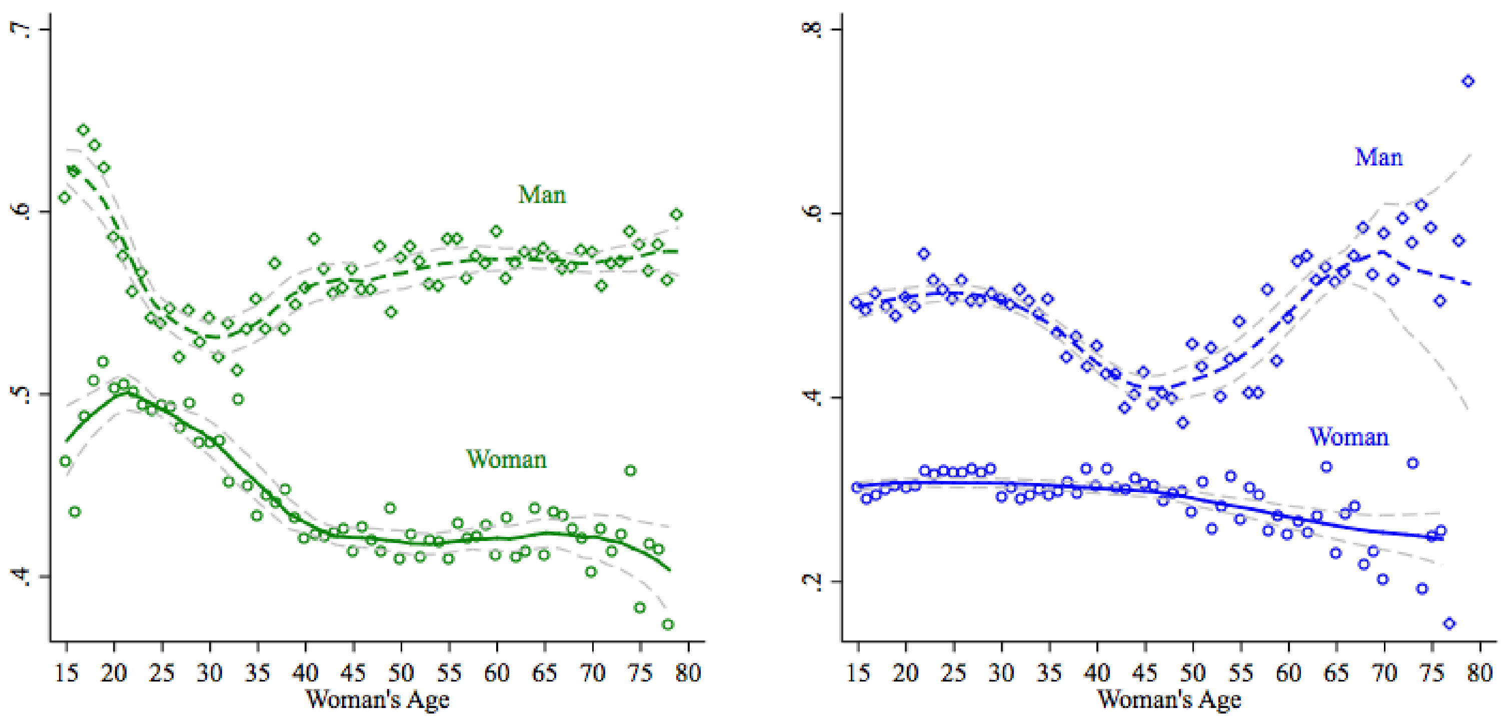
Figure A4: Predicted Resource Shares, Age and HSA Amendments (Nuclear Households Only)



(A) Without Children, $\hat{\Lambda}_f$ and $\hat{\Lambda}_m$ (B) With Children, $\hat{\Lambda}_f$ and $\hat{\Lambda}_m$

Note: Mean predicted women's (men's) resource share among households with women's (men's) average age equal to $a = 15, ..., 79$. The solid line is a running mean.

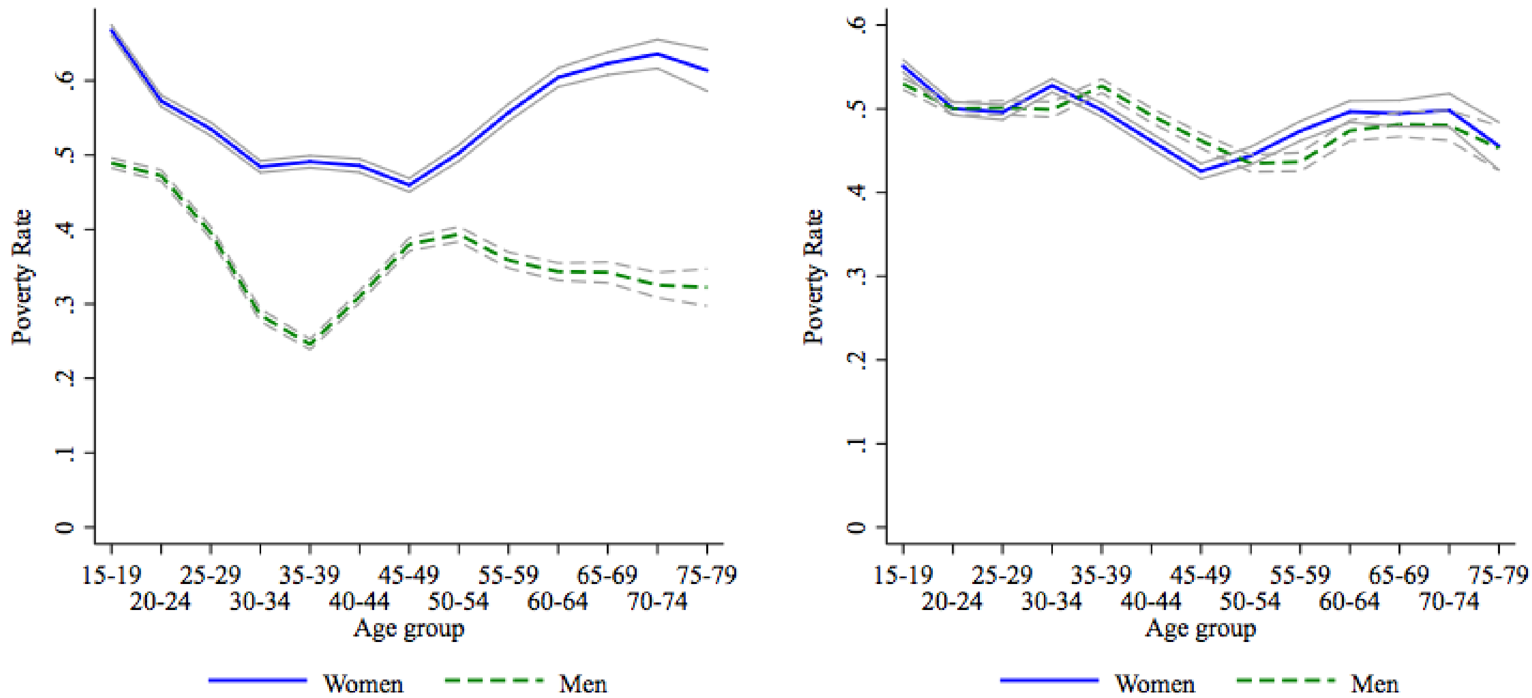
Figure A5: Average Predicted Resource Shares and Age



(A) Without Children, $\hat{\Lambda}_f$ and $\hat{\Lambda}_m$ (B) With Children, $\hat{\Lambda}_f$ and $\hat{\Lambda}_m$

Note: Nuclear households only. Mean predicted women's (men's) resource share among households with woman's (man's) age equal to $a = 15, ..., 79$. The solid line is a running mean.

Figure A6: Average Predicted Resource Shares and Age (Nuclear Households Only)

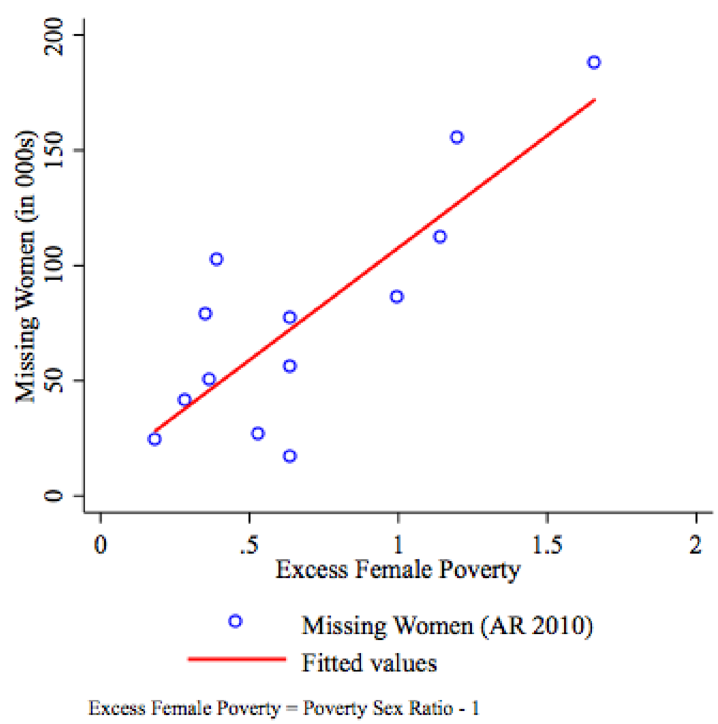


(A) Model Predictions (Unequal Sharing)

(B) Equal Sharing

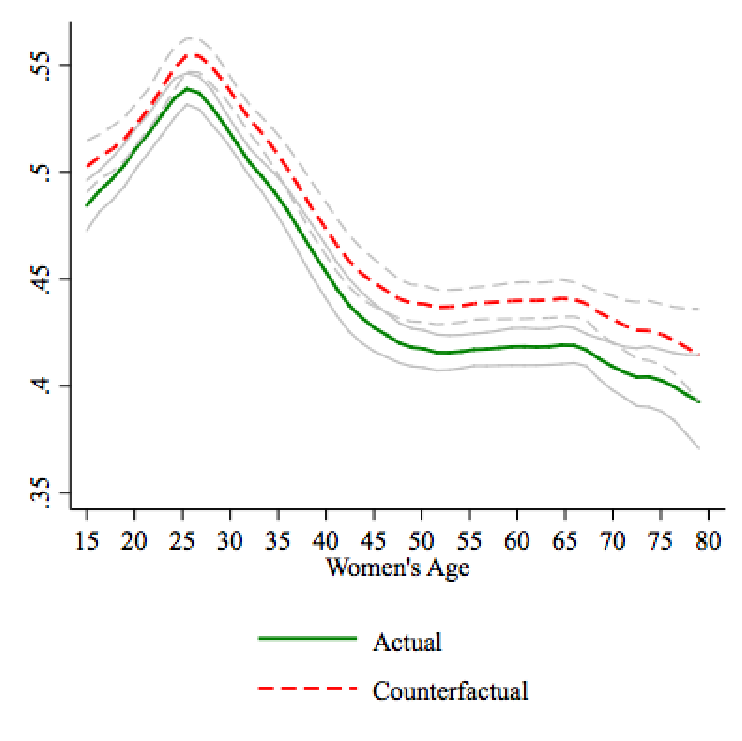
<u>Note:</u> The graphs show the fraction of females or males in each age group living below poverty line. Individuals from all households in the sample are used for calculations. Per capita expenditures are computed using the model predictions in panel A. In panel B, I assume that household expenditure is split equally among household members. Per capita expenditures are compared to the US$3.10/day poverty line.

Figure A7: Poverty Rates By Gender and Age (US$3.10/day)

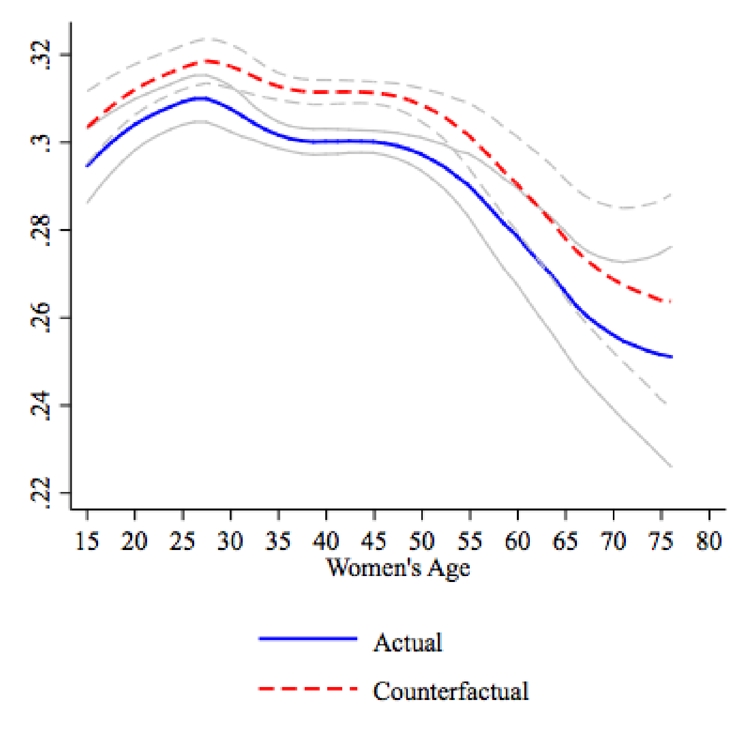


Age Groups 15-19 to 75-79

Figure A8: Excess Female Mortality and Excess Female Poverty



(A) Without Children



(B) With Children

Note: Running means of the actual and counterfactual average women's resource shares among households with women's average age equal to $a = 15, \ldots, 79$.

Figure A9: Counterfactual Experiment: HSA Amendment and Women's Resource Shares

Table A1: NFHS-3 Descriptive Statistics

	Obs.	Mean	Median	St. Dev.
Body Mass Index (BMI)	82,303	21.42	20.61	4.22
$\mathbb{I}(BMI \leq 18.5)$	82,303	0.26	0.00	0.44
$\mathbb{I}$(Severe Anaemia)	78,521	0.02	0.00	0.12
$\mathbb{I}$(Moderate Anaemia)	78,521	0.16	0.00	0.36
$\mathbb{I}$(Mild Anaemia)	78,521	0.53	1.00	0.50
$\mathbb{I}$(HSAA Exposed)	85,881	0.15	0.00	0.36
$\mathbb{I}$(HSAA Eligible)	69,561	0.11	0.00	0.31
$\mathbb{I}$(Hindu, Buddhist, Sikh, Jain)	85,881	0.79	1.00	0.41
$\mathbb{I}$(Sch. Caste, Sch. Tribe or Other Backward Classes)	85,222	0.63	1.00	0.48
$\mathbb{I}$(Rural)	85,881	0.56	1.00	0.50
$\mathbb{I}$(Worked in Past 12 Months)	85,869	0.40	0.00	0.49
Wealth Index	85,742	0.27	0.20	0.24
Age	85,881	31.95	31.00	8.42
No. Children Under 5	85,881	0.81	0.00	1.02
Years of Schooling	85,876	5.33	5.00	5.18
$\mathbb{I}$(Higer Education)	85,876	0.09	0.00	0.28

Note: Only married women of age 15 to 49 included. Three levels of severity of anaemia are distinguished in the 2005 National Family Health Survey: mild anaemia (10.0-10.9 grams/deciliter for pregnant women, 10.0-11.9 g/dl for non-pregnant women, and 12.0-12.9 g/dl for men), moderate anaemia (7.0-9.9 g/dl for women and 9.0-11.9 g/dl for men), and severe anaemia (less than 7.0 g/dl for women and less than 9.0 g/dl for men). The wealth index is constructed by combining information on a set of household assets (radio, refrigerator, television, bicycle, motorcycle, car, and land) using principal component analysis. Higher education equals 1 if the number of years of schooling is greater than 12.

Table A2: NSS - CES Descriptive Statistics

	Households With Children < 15				Households Without Children < 15			
	Obs.	Mean	Median	St. Dev.	Obs.	Mean	Median	St. Dev.
Expenditure (Rupees):								
Total Expenditure	57,158	8,226.58	6,908.00	4,911.55	30,215	7,886.53	6,481.00	5,274.64
Expenditure On Non-Durable Goods	57,158	7,849.90	6,695.14	4,492.67	30,215	7,399.88	6,206.62	4,726.76
Expenditure On Durable Goods	57,158	376.67	106.85	1,022.50	30,215	486.65	107.26	1,371.71
Budget Shares:								
Food	57,158	40.46	40.41	9.42	30,215	36.95	37.06	9.58
Female Assignable Clothing	57,158	1.31	1.13	1.09	30,215	1.49	1.25	1.29
Male Assignable Clothing	57,158	1.62	1.36	1.38	30,215	1.78	1.51	1.48
Children Assignable Clothing	57,158	0.69	0.51	0.81	-	-	-	-
Household Characteristics:								
No. Adult Females	57,158	1.69	1.00	0.86	30,215	1.67	1.00	0.83
No. Adult Males	57,158	1.67	1.00	0.90	30,215	1.91	2.00	0.93
Fraction of Female Children	57,158	0.45	0.50	0.39	-	-	-	-
Number of Children Under 5	57,158	2.01	2.00	1.01	-	-	-	-
$\mathbb{I}$(Daughter in Law)	57,158	0.24	0.00	0.43	30,215	0.11	0.00	0.32
$\mathbb{I}$(Unmarried Daughter Above 15)	57,158	0.17	0.00	0.38	30,215	0.33	0.00	0.47
$\mathbb{I}$(Widow)	57,158	0.14	0.00	0.35	30,215	0.16	0.00	0.37
Avg. Age Men 15 to 79	57,109	36.94	36.00	8.76	29,980	39.37	36.00	13.10
Avg. Age Women 15 to 79	57,137	34.84	34.00	8.20	30,126	40.98	40.00	12.09
Avg. Age Gap 15 to 79 (Men - Women)	57,090	2.10	3.00	9.93	29,915	-1.44	1.50	12.86
Avg. Age Children 0 to 14	57,158	7.57	8.00	3.97	-	-	-	-
$\mathbb{I}$(HSAA Eligible)	47,330	0.15	0.00	0.35	26,797	0.08	0.00	0.28
$\mathbb{I}$(Hindu, Buddhist, Sikh, Jain)	57,158	0.77	1.00	0.42	30,215	0.83	1.00	0.38
$\mathbb{I}$(Sch. Caste, Sch. Tribe or Other Backward Classes)	57,158	0.71	1.00	0.45	30,215	0.65	1.00	0.48
$\mathbb{I}$(Salary Earner)	57,158	0.29	0.00	0.46	30,215	0.32	0.00	0.47
$\mathbb{I}$(Land Ownership)	57,158	0.89	1.00	0.31	30,215	0.90	1.00	0.30
$\mathbb{I}$(Female Higher Education)	57,158	0.10	0.00	0.30	30,215	0.14	0.00	0.35
$\mathbb{I}$(Male Higher Education)	57,158	0.17	0.00	0.37	30,215	0.24	0.00	0.43
$\mathbb{I}$(Rural)	57,158	0.63	1.00	0.48	30,215	0.57	1.00	0.50
$\mathbb{I}$(North)	57,158	0.33	0.00	0.47	30,215	0.28	0.00	0.45
$\mathbb{I}$(East)	57,158	0.21	0.00	0.41	30,215	0.19	0.00	0.39
$\mathbb{I}$(North-East)	57,158	0.16	0.00	0.36	30,215	0.12	0.00	0.33
$\mathbb{I}$(South)	57,158	0.19	0.00	0.39	30,215	0.27	0.00	0.45
$\mathbb{I}$(West)	57,158	0.12	0.00	0.32	30,215	0.13	0.00	0.34

Note: See table 2 for details.

Table A3: HSA Amendments and Women's Health: Probit Model

	Body Mass Index	Pr(Anaemia)		
	Pr($BMI \leq 18.5$)	Severe	Moderate	Mild
	(1) Probit	(2) Probit	(3) Probit	(4) Probit
HSAA Exposed	-0.0476***	-0.00902***	-0.0269***	-0.0328***
	(0.0106)	(0.00189)	(0.00812)	(0.0113)
N	81,115	56,742	77,529	77,755
Mean Dependent Variable	0.2648	0.0154	0.1559	0.5298

Note: *$p < 0.10$, **$p < 0.05$, ***$p < 0.01$. NFHS-3 data. Married women of age 15 to 49 included in the sample. Individual and household controls defined as in table 1. Robust standard errors in parentheses. Standard errors clustered at the primary sampling unit (village) level (3,753). Sampling weights applied.

Table A4: Predicted Engel Curve Slopes: Descriptive Statistics

	Mean (1)	Sd. Dev. (2)	Median (3)	Min. (4)	Max. (5)
Panel A: Without Children < 15 Only					
Women Assignable Clothing	-0.2547	0.0809	-0.2488	-0.5505	0.0175
Men Assignable Clothing	-0.3270	0.1448	-0.3294	-1.2825	0.0069
Food	-6.4665	1.5405	-6.4864	-13.1397	-2.4183
Panel B: With Children < 15 Only					
Women Assignable Clothing	-0.1169	0.0821	-0.0952	-0.4726	0.0559
Men Assignable Clothing	-0.1905	0.1240	-0.1802	-0.8466	0.0550
Children Assignable Clothing	-0.0711	0.0405	-0.0663	-0.2707	0.0678
Food	-6.8114	1.5491	-6.7840	-13.2749	-1.4069

Table A5: Determinants of Women's Resource Shares: Robustness Checks

	Women's Resource Share			
	Food Engel Curve Excl.		Exp. on Non-Durables	
	With Children < 15 Only	Withou Children < 15 Only	With Children < 15 Only	Withou Children < 15 Only
	(1) NLSUR	(2) NLSUR	(3) NLSUR	(4) NLSUR
No. Adult Women	0.0320***	0.0552***	0.0292***	0.0594***
	(0.00440)	(0.00904)	(0.00400)	(0.00883)
No. Adult Men	-0.0215***	-0.0268***	-0.0191***	-0.0271***
	(0.00347)	(0.00659)	(0.00313)	(0.00618)
No. Children	0.00406*	-	0.00279	-
	(0.00229)	-	(0.00203)	-
Fraction of Female Children	0.0117**	-	0.0114**	-
	(0.00526)	-	(0.00473)	-
$\mathbb{I}$(Daughter in Law)	0.00677	0.0125	0.00710	0.0177
	(0.00678)	(0.0179)	(0.00614)	(0.0182)
$\mathbb{I}$(Unmarried Daughter above 15)	0.00939	-0.00256	0.00855	-0.0116
	(0.00756)	(0.0170)	(0.00678)	(0.0161)
$\mathbb{I}$(Widow)	-0.0354***	-0.0166	-0.0351***	-0.0150
	(0.00928)	(0.0174)	(0.00839)	(0.0165)
$\mathbb{I}$(HSAA Eligible)	0.0108**	0.0217**	0.00784*	0.0216**
	(0.00505)	(0.00931)	(0.00460)	(0.00925)
$\mathbb{I}$(Hindu, Buddhist, Sikh, Jain)	-0.0168**	-0.0168	-0.0130*	-0.0164
	(0.00773)	(0.0150)	(0.00680)	(0.0143)
$\mathbb{I}$(SC, ST, Other Backward Caste)	0.0699***	0.0556***	0.0664***	0.0564***
	(0.00857)	(0.0123)	(0.00786)	(0.0119)
$\mathbb{I}$(Salary Earner)	-0.0218***	-0.0128	-0.0216***	-0.0169*
	(0.00474)	(0.00994)	(0.00424)	(0.00936)
$\mathbb{I}$(Land Ownership)	0.00724	-0.0154	0.00854	-0.00459
	(0.00880)	(0.0179)	(0.00770)	(0.0166)
$\mathbb{I}$(Female Higher Education)	0.0276***	0.0366**	0.0266***	0.0404***
	(0.00816)	(0.0159)	(0.00747)	(0.0156)
$\mathbb{I}$(Male Higher Education)	0.0413***	0.0814***	0.0407***	0.0803***
	(0.00625)	(0.0126)	(0.00581)	(0.0125)
$\mathbb{I}$(Rural)	-0.0294***	-0.0402***	-0.0289***	-0.0388***
	(0.00648)	(0.0116)	(0.00584)	(0.0108)
Avg. Age Diff. (Men 15 to 79 - Women 15 to 79)	-0.115**	0.0521	-0.113***	0.0725
	(0.0464)	(0.0804)	(0.0409)	(0.0764)
Avg. Age Women 15 to 79	-0.624	-1.584	-0.812	-1.652
	(0.760)	(1.146)	(0.658)	(1.091)
(Avg. Age Diff. (Men 15 to 79 - Women 15 to 79))^2	0.115	-0.503***	0.121	-0.477***
	(0.131)	(0.187)	(0.113)	(0.180)
(Avg. Age Women 15 to 79)^2	1.661	2.802	2.155	2.755
	(1.914)	(2.664)	(1.640)	(2.527)
(Avg. Age Diff. (Men 15 to 79 - Women 15 to 79))^3	0.460	-0.706	0.513	-0.685
	(0.704)	(0.760)	(0.597)	(0.740)
(Avg. Age Women 15 to 79)^3	-1.469	-1.544	-1.891	-1.403
	(1.563)	(1.976)	(1.325)	(1.871)
Avg. Age Children 0 to 14	-0.0506	-	-0.0423	-
	(0.0644)	-	(0.0579)	-
$\mathbb{I}$(North)	-0.102***	-0.0651***	-0.102***	-0.0722***
	(0.0162)	(0.0232)	(0.0149)	(0.0221)
$\mathbb{I}$(East)	-0.0210	-0.0232	-0.0241	-0.0145
	(0.0177)	(0.0253)	(0.0156)	(0.0241)
$\mathbb{I}$(North-East)	0.0518**	0.169***	0.0501**	0.181***
	(0.0240)	(0.0283)	(0.0216)	(0.0280)
$\mathbb{I}$(South)	-0.0274	-0.0538**	-0.0218	-0.0540**
	(0.0177)	(0.0234)	(0.0161)	(0.0226)
Constant	-0.0274	-0.0538**	-0.0218	-0.0540**
	(0.0177)	(0.0234)	(0.0161)	(0.0226)
N	47,262	26,497	47,262	26,497
LL	-194,649.7	-84,369.0	-376,811.8	-187,727.3
No. Parameters	264	140	318	188

Note: $*p < 0.10$, $**p < 0.05$, $***p < 0.01$. NSS 68th Round Consumer Expenditure data. Robust standard errors in parentheses. Standard errors clustered at the first sampling unit level. Women's age and age differences are divided by 100 to ease computation. West India is the excluded region.

Table A6: Determinants of Women's Resource Shares: Restricted Samples

	Women's Resource Share		
	$F = M = 1$, $C = \{0, 1\}$		Married Couples
	With Children < 15 Only	Without Children < 15 Only	With Children < 15 Only
	(1) NLSUR	(2) NLSUR	(3) NLSUR
𝕀(Child is Female)	0.0466**	-	-
	(0.0218)	-	-
𝕀(Daughter in Law)	-0.0482	0.0797	-
	(0.169)	(0.131)	-
𝕀(Unmarried Daughter above 15)	-0.000739	-0.103	-
	(0.204)	(0.0674)	-
𝕀(Widow)	-0.0179	-0.0316	-
	(0.106)	(0.0300)	-
𝕀(Two Children)	-	-	-0.0125**
	-	-	(0.00603)
𝕀(Three Children)	-	-	0.00142
	-	-	(0.00784)
𝕀(Four Children)	-	-	0.00450
	-	-	(0.0109)
𝕀(Five Children)	-	-	0.00475
	-	-	(0.0169)
Fraction of Female Children	-	-	0.0160**
	-	-	(0.00655)
𝕀(HSAA Eligible)	0.0383**	0.00166	0.0169**
	(0.0187)	(0.0151)	(0.00688)
𝕀(Hindu, Buddhist, Sikh, Jain)	-0.0554	0.0340*	-0.00120
	(0.0363)	(0.0176)	(0.00854)
𝕀(SC, ST or Other Backward Caste)	0.0727**	-0.0309**	0.105***
	(0.0343)	(0.0140)	(0.0126)
𝕀(Salary Earner)	-0.0223	-0.00403	-0.0250***
	(0.0231)	(0.0144)	(0.00591)
𝕀(Land Ownership)	0.00933	0.0359**	-0.00172
	(0.0278)	(0.0169)	(0.00942)
𝕀(Female Higher Education)	-0.0197	-0.0306	0.0213**
	(0.0398)	(0.0226)	(0.0101)
𝕀(Male Higher Education)	0.0571*	0.0488***	0.0488***
	(0.0309)	(0.0175)	(0.00795)
𝕀(Rural)	-0.0634**	-0.00315	-0.0285***
	(0.0295)	(0.0130)	(0.00729)
Age Diff. (Man 15 to 79 - Woman 15 to 79)	0.0852	-0.0872	0.0611
	(0.230)	(0.103)	(0.0539)
Age Woman 15 to 79	-0.925	-2.300**	-1.865*
	(2.176)	(1.137)	(1.028)
(Age Diff. (Man 15 to 79 - Woman 15 to 79))^2	0.277	0.0835	-0.273
	(0.651)	(0.211)	(0.180)
(Age Woman 15 to 79)^2	2.679	4.194	4.537*
	(5.448)	(2.623)	(2.697)
(Age Diff. (Man 15 to 79 - Woman 15 to 79))^3	0.377	-0.243	0.781
	(1.801)	(0.620)	(1.064)
(Age Woman 15 to 79)^3	-2.537	-2.761	-3.349
	(4.360)	(1.912)	(2.280)
Age Child 0 to 14	-0.0606	-	-0.177**
	(0.322)	-	(0.0784)
𝕀(North)	-0.119**	0.134***	-0.117***
	(0.0569)	(0.0475)	(0.0193)
𝕀(East)	0.00936	0.0530	-0.0108
	(0.0607)	(0.0470)	(0.0198)
𝕀(North-East)	0.0422	0.0854	0.0562**
	(0.0653)	(0.0641)	(0.0260)
𝕀(South)	-0.00339	-0.167***	0.0403*
	(0.0500)	(0.0428)	(0.0216)
Constant	0.529*	0.911***	0.557***
	(0.282)	(0.160)	(0.129)
N	4,172	6,967	32,622
ll	-33,507.4	-47,276.4	-262,003.1
No. Parameters	282	172	294

Note: $^*p < 0.10$, $^{**}p < 0.05$, $^{***}p < 0.01$. NSS 68^{th} Round Consumer Expenditure data. Robust standard errors in parentheses. Standard errors clustered at the first sampling unit level. Woman's age and age differences are divided by 100 to ease computation. West India is the excluded region.