Essays on Health and Family Economics in India

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Boston College Electronic Thesis or Dissertation, 2016

Boston College

Morrissey College of Arts and Sciences

Graduate School

Department of Economics

ESSAYS ON HEALTH AND FAMILY ECONOMICS IN INDIA

a dissertation

by

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submitted in partial fulfillment of the requirements

for the degree of

Doctor of Philosophy

April 2016

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by

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Abstract

A person's health not only influences her chance of surviving to adulthood and her life expectancy, but also her economic decisions, her productivity, and her well-being. Since a healthy population is a major factor in economic development, it is important to understand the determinants of individuals' health-related decisions and outcomes. The three essays that comprise this dissertation make advancements in this direction and focus on the Indian subcontinent. The first and second essays analyze how intra-household decision making affects individuals' health outcomes and welfare, with a special attention towards within family gender inequality. The third essay studies how exposure to historical medical facilities affects individual health outcomes across generations.

From a methodological point of view, this dissertation highlights the advantages of combining economic models with data from a wide range of sources, theory with empirics. I employ both quasi-experimental and structural estimation methods, using the former to uncover relevant causal links and policy levers, and the latter to estimate deep parameters, overcome data limitations and perform counterfactual policy analysis. More broadly, with this work I stress the importance of research in development economics being open to a variety of methodologies and empirical approaches.

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. In the first essay, titled **Why Are Older Women Missing in India? The Age Profile of Bargaining Power and Poverty**, I provide an explanation for this puzzle that is based on intra-household bargaining and resource allocation. I use both reduced-form and structural modeling to establish the critical 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 structural 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. The analysis indicates that gender inequality within the household and the consequent gender asymmetry in poverty can account for a substantial fraction of missing women of post-reproductive ages. Finally, I demonstrate that policies aimed at promoting intra-household equality, such as improving women's rights to inherit property, can have a large impact on female poverty and mortality.

The first essay contributes to a wide literature showing that a relevant determinant of the household decisions and outcomes is the relative bargaining position of the decision makers. Although this link is well-accepted in this literature, intra-household bargaining power is *de facto* an unobserved variable. In the second essay, joint with Arthur Lewbel and Denni Tommasi and titled **Women's Empowerment and Family Health: A Two-Step Approach**, we propose a novel two-step approach to overcome this data limitation and to directly assess the causal link between women's empowerment and family health. In the first stage, we structurally recover a dollar-based measure of women's intra-household empowerment, with a clear interpretation provided by economic theory; in the second stage, we identify the causal effect of women's decision power relative to men's on household members' health. We demonstrate that women's bargaining power improves their own health outcomes, while not affecting their spouses'. When we turn to children, we find that improvements in women's position within the family does not affect their weight or height, but it increases their likelihood to receive vaccinations.

The determinants of individuals' health, however, go beyond the family, and trace back to historical developments. In the third essay, joint with Federico Mantovanelli and titled **Long-Term Effects of Access to Health Care: Medical Missions in Colonial India**, we examine the long-term effect of access to historical health facilities on current individual health outcomes. To this aim, we construct a novel and fully geocoded dataset that combines contemporary individual-level data with historical information on Protestant medical missions. We exploit variation in the activities of missionary societies and use an instrumental variable approach to show that proximity to a Protestant medical mission has a causal effect on individuals' health status. The investigation of potential transmission channels indicates that the long-run effect of access to health care is not driven by persistence of infrastructure, but by changes in individual habits regarding hygiene, preventive care and health awareness, which have been bequeathed over time.

Important policy implications can be drawn. First, policies aimed at promoting gender equality within families, such as improving women's property and inheritance rights, can have positive spillovers on women's health, poverty and mortality, and can boost health investments in children. Second, as the population in India and in other developing countries ages, gender asymmetries among the elderly need to be further investigated and promptly addressed by the development practitioners. Third, intra-household inequalities, between genders and across ages, should be taken into account when measuring poverty and evaluating the effect of policies to alleviate it. Finally, in light of the existence of long-run effects, the expansion of health care access in India should become an even more prominent goal for policy makers, as it can beneficially affect both current and future generations.

To Tina and Anita

Acknowledgments

This work would not be possible without the invaluable help and continued support of many others. I am forever grateful to Arthur Lewbel, Scott Fulford and Andrew Beauchamp, for their constant guidance and precious advice, and for teaching me an incredible amount about economics and being an academic. I thank the Department of Economics, the Graduate School of the Morrissey College of Arts and Sciences and the Institute On Aging at Boston College for the exceptional support and the many opportunities afforded me throughout my doctoral studies. I appreciate greatly the helpful advice by S Anukriti and Claudia Olivetti, and the many insightful discussions with my colleagues and friends in the department. I thank them all.

I would also like to thank all the people who have been part of this incredible journey and with whom I shared laughter and tears, inspiring discussions, delicious dinners and bottles of wine, especially Laura Bonacorsi, Giacomo Candian, Rossella Greco, Kate Griwert, Sylvia Hristakeva and Denni Tommasi. I thank in particular Federico Mantovanelli, a great economist and a true friend. I thank all the Bg Dancers in Italy, especially Lara Breviario, Valentina Colombi and Michela Parisi, for their timeless friendship and their support from afar. I also owe an incredible debt of gratitude to Samson Alva, for teaching me the true meaning of scholarship and a lot of what I know now, and for guiding me the many times I got lost during the writing of this dissertation. I could not have asked for a better partner in this journey.

Finally, to my parents, Emanuela and Enzo, goes my eternal gratitude. I thank them for their unique support, encouragement and sacrifices made throughout my life, and for believing in me, always and unconditionally. None of this could have happened without them. Grazie di cuore.

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Chapter 1

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

1.1 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

[1] Coale (1991) estimates that over 23 million of Indian women who should be alive are missing.
[2] See e.g., Sen (1990) and DasGupta (2005).

above (see figure 1.1).[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 82 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 1.1 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-

[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 1.1 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. See section A.1 in the Appendix for more details.

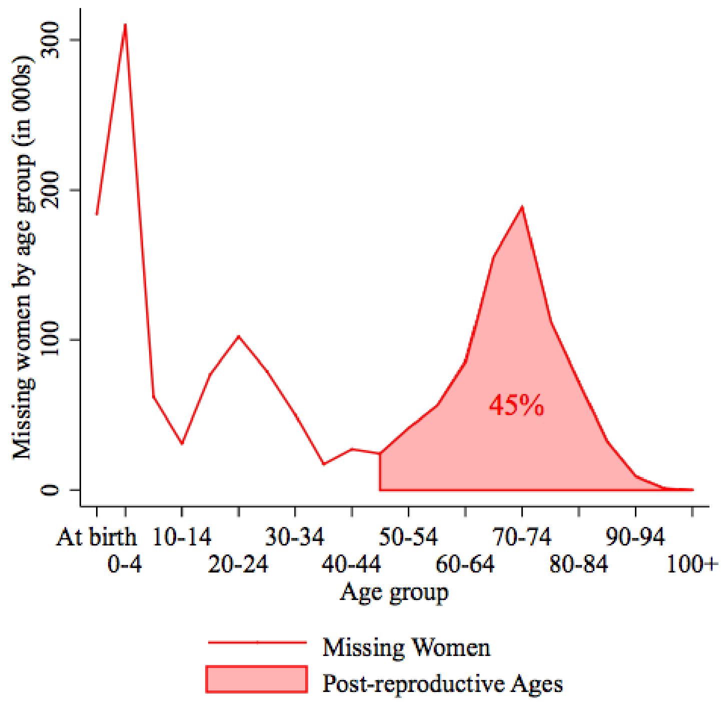


Figure 1.1: Missing Women by Age Group (Anderson-Ray, 2010)

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 separate 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 struc-

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

tural 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.[5] As morbidity and mortality rates are higher in poverty, however, I indirectly explore the link between intra-household allocation, health, and mortality by studying how gender inequality within the household affects aggregate measures of welfare.[6] 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 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.[7] I find that exposure to the Hindu Succession Act amendments increases women's resource shares by 0.2 standard deviations. These

[5] In Chapter 2, we address the NSS data limitations on health outcomes by performing an out-of-sample prediction on the NFHS data to directly investigate how resource shares affect health production.
[6] 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.
[7] 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.

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. I demonstrate that granting universal equal inheritance rights can reduce the number of women living in extreme poverty by 10 percent and the number of excess female deaths at post-reproductive ages by up to 26 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 a recent working 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. 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), and 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 investigated 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 1.2 provides an overview of the related literature and discusses further the contributions of this paper. Section 1.3 presents the reduced-form results and establishes a positive causal link between women's intra-household bargaining power and their health. Section 1.4 discusses the household model, the identification of resource shares and the structural estimation results. Section 1.5 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 1.6 presents the counterfactual policy analysis. Section 1.7 concludes.

1.2 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% 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) 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 (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]

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 in a series of papers (Drèze et al. (1990); Chen and Drèze (1995); Drèze and Srinivasan (1997)). Moreover, previous work 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. Moreover, 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. While the health disadvantage and

[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.

lower utilization among women cannot be explained by demographics, they find that gender differentials disappear when controlling for economic independence.

Women's 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. Mine is the first paper that focuses on health outcomes, measures of undernourishment and anaemia in particular, and on access to household resources. Deininger et al. (2013) 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. Moreover, Roy (2008) shows that women's exposure to the HSA reforms improves their bargaining power and autonomy within their marital families and Roy (2013), Deininger et al. (2013), and Bose and Das (2015) indicate that it increases female education.[10] Heath and Tan (2014) show that the HSA amendments increase women's labor supply, especially into high-paying jobs.[11] Finally, Jain (2014) shows that HSA reforms mitigate son preference, and might be effective in reducing mortality differences between boys and girls in rural India.[12]

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.[13] On the other hand, a number of papers have focused on developing tech-

[10] Unintended negative consequences of these reforms have been studied as well. 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. Anderson and Genicot (2015) show 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.

[11] While analyzing possible underlying mechanisms, they also find some preliminary evidence of health improvements following the implementation of HSA amendments.

[12] 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.

[13] 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

niques by which household level consumption data may be used to recover information about individual household members. Building on Chiappori (1988a, 1992) and Apps and Rees (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, while 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 (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 (Lewbel and Pendakur (2008), Browning et al. (2013), Dunbar et al. (2013)). Dunbar et al. (2013), for example, identify individuals' resource shares using Engel curves of assignable clothing and find that Malawian 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.[14] 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.

1.3 Bargaining Power and Health: A Reduced-Form Analysis

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

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.

[14] 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.

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 causal positive 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 (BMI), especially BMI below the underweight cutoff of 18.5, and anaemia are associated with an increased risk of mortality.[15] Evidence of improvements in these health outcomes following an increase in women's bargaining power would provide 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 (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 therefore did not apply to individuals of other religions, such as Muslims, Christians, Parsis, Jews, and other minority communities.[16] 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.[17]

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

[16] 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)).

[17] 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 Indian constitution states that both federal and state governments have legislative power over inheritance. In the decades following the introduction of Hindu Succession Act, state governments enacted amendments 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.

To study the link between women's bargaining position and health, I consider the following baseline specification:

$$y_{irsc} = \beta HSAA\ Exposed_{irsc} + X'_{irsc}\gamma + \alpha_r + \alpha_c + \alpha_s + \alpha_{rs} + \alpha_{rc} + \alpha_{sc} + \epsilon_{irsc} \tag{1.1}$$

where y_{irsc} is the outcome of interest for woman i, of religion r, living in state s and born in year c (BMI or an indicator variable for being underweight or severely, moderately or mildly anemic). $HSAA\ Exposed_{irsc}$ 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.[18] 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-household bargaining power. In the baseline specification, standard errors are clustered at the primary sampling unit (village) level. Results

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.

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

are robust to clustering the standard errors at the state, cohort-state, and at the cohort-state-religion level.

I estimate the model via OLS using a sample of married women of age 15 to 49 from the 2005-2006 National Family Health Survey (NFHS-3). The average body mass index 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.[19] Finally, one out of six women in the sample have been exposed to the HSA amendments.

Table 1.1 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 0.15 when the entire sample is considered, and by 0.25 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.[20]

[19] Section A.3 contains more details on the data and on the use of BMI and anaemia as health measures. Table A.4 presents some descriptive statistics.

[20] To address the well known limitations of linear probability models, I estimate a maximum-likelihood probit model. Table A.2 in the Appendix shows the marginal effects of being exposed to HSA amendments on binary health outcomes. The marginal effects are computed at the average value of the independent variables. As the number of observation is much larger than the number of fixed effects the incidental parameter problem associated with estimating a probit model with fixed effects may be limited (Greene et al. (2002), Fernández-Val (2009)). Results are unchanged.

Table 1.1: 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	0.149*	0.250***	-0.161	-0.0395***	-0.0125***	-0.0306***	-0.0309***
	(0.0774)	(0.0556)	(0.132)	(0.0101)	(0.00315)	(0.00897)	(0.0109)
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

*$p < 0.10$, **$p < 0.05$, ***$p < 0.01$. 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 women's years of schooling, number of children under 5 in the household, a household wealth index, and indicator variables for 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.

I perform a series of robustness checks to test the sensitivity of the reduced-form 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 got married right around the reforms, 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 show 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 my identification strategy. Section A.3 in the Appendix contains the full set of results, together with additional details about the empirical strategies.

1.4 A Structural Analysis of Bargaining Power

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.[21]

1.4.1 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, so there is no borrowing or lending. $h = (h^1, ..., h^K)$ 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

[21] 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.

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, h, in *private good equivalents*, x. This specific technology assumes the existence of a $K \times K$ matrix A such that $h = A(Fx_f + Mx_m + Cx_c)$.[22]

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 adult men, all adult women and all children, respectively. As further discussed in section 1.4.2, 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.[23] 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:

$$\tilde{U}(U_f, U_m, U_c, p/y) = \sum_{t \in \{f,m,c\}} \mu_t(p/y)\tilde{U}_t \tag{1.2}$$

[22] 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, $A_k = 1$.

[23] 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 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 with a wife of the head of household and an unmarried daughter above 15.

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

$$\max_{x_f, x_m, x_c, h} \tilde{U}(U_f, U_m, U_c, p/y) \quad \text{such that} \quad h = A(Fx_f + Mx_m + Cx_c)$$
$$y = h'p \tag{1.3}$$

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.[24] Thus, the household program can be decomposed in two steps: 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 $\sum_k A_k p^k x_t^k = \lambda_t y$.[25] By substituting the indirect utility functions $V_t(A'p, \lambda_t y)$ in equation (1.3), 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.,

[24] See e.g., Chiappori (1988a, 1992), Browning et al. (1994), Browning and Chiappori (1998), Vermeulen (2002), Lewbel and 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)). In Appendix A.6, I use auxiliary data on singles to show that the assumption of Pareto efficiency cannot be rejected in this context.

[25] 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.

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:

$$W_t(y, p) = T\lambda_t w_t(A'p, \lambda_t y) = \Lambda_t w_t(A'p, \lambda_t y) \tag{1.4}$$

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.

1.4.2 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 assumptions that resource shares are independent of expenditure (at least at low levels of y).[26] 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

[26]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).

clothing at all levels of expenditure.[27] 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 with children they are as follows:

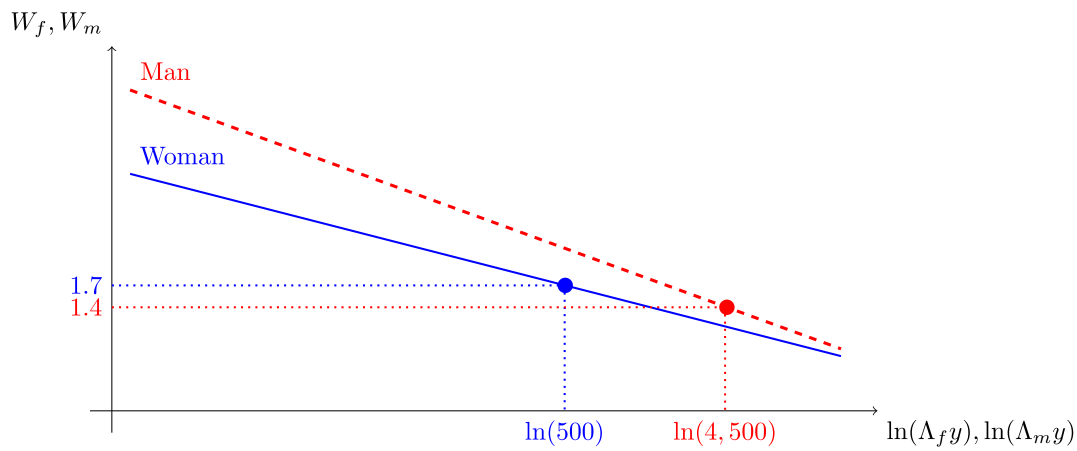
$$\begin{cases} W_f(y) = \alpha_f \Lambda_f + \beta_f \Lambda_f \ln\left(\frac{\Lambda_f y}{F}\right) \\ W_m(y) = \alpha_m \Lambda_m + \beta_m \Lambda_m \ln\left(\frac{\Lambda_m y}{M}\right) \\ W_c(y) = \alpha_c \Lambda_c + \beta_c \Lambda_c \ln\left(\frac{\Lambda_c y}{C}\right) \end{cases} \tag{1.5}$$

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 adult women's assignable clothing and one for adult 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 (1.5) 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

[27] Dunbar et al. (2013) discuss identification of type t individuals' resource shares using Engel curves in a more general framework.



$W_f > W_m$, $\Lambda_m > \Lambda_f$

Figure 1.2: Assignable Clothing Engel Curves: An Illustrative Example

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 1.2. 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.[28] 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 $\Lambda_m = 0.9$.

[28] The Engel curves displayed in figure 1.2 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 1.4.4). Table A.3 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.

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.

1.4.3 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 about household characteristics and demographic and other particulars of household members (about 463,000 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 expenditure on clothing items that are assignable to specific types of household members, i.e., adult women, adult 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.[29] For simplicity, I focus on the eligibility of the wife of the

[29] 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 1.3. This variable is therefore fully determined by each woman's religion, year of birth and state. See Heath and Tan (2014).

Table 1.2: NSS Consumer Expenditure Survey - Descriptive Statistics

	Obs.	Mean	Median	St. Dev.
Expenditure (Rupees):				
Total Expenditure	87,450	8,105	6,775	5,034
Expenditure On Non-Durable Goods	87,450	7,692	6,539	4,575
Expenditure On Durable Goods	87,450	413	107	1,148
Budget Shares:				
Food	87,450	37.83	37.89	9.84
Female Assignable Clothing	87,376	1.31	1.12	1.11
Male Assignable Clothing	87,376	1.61	1.34	1.37
Children Assignable Clothing	87,376	0.50	0	0.76
Household Characteristics:				
$\mathbb{I}$(HSAA Eligible)	74,228	0.12	0	0.33
No. Adult Females	87,450	1.68	1	0.85
No. Adult Males	87,450	1.76	1	0.92
$\mathbb{I}$(Children Under 15)	87,450	0.65	1	0.48
No. Children Under 15	87,450	1.32	1	1.26
Fraction Of Female Children	57,202	0.45	0.5	0.39
$\mathbb{I}$(Daughter in Law)	87,450	0.20	0	0.40
$\mathbb{I}$(Unmarried Daughter)	87,450	0.23	0	0.42
$\mathbb{I}$(Widow)	87,450	0.15	0	0.35
Avg. Age Men 15 to 79	87,166	37.78	36	10.52
Avg. Age Women 15 to 79	87,340	36.96	35	10.15
Avg. Age Gap 15 to 79 (Men - Women)	87,082	0.89	3	11.15
Avg. Age Children 0 to 14	57,202	7.57	8	3.97
$\mathbb{I}$(Hindu, Buddhist, Sikh, Jain)	87,450	0.79	1	0.41
$\mathbb{I}$(Sch. Caste, Sch. Tribe or Other Backward Classes)	87,450	0.69	1	0.46
$\mathbb{I}$(Salary Earner)	87,450	0.30	0	0.46
$\mathbb{I}$(Land Ownership)	87,450	0.89	1	0.31
$\mathbb{I}$(Female Higher Education)	87,450	0.12	0	0.32
$\mathbb{I}$(Male Higher Education)	87,450	0.19	0	0.39
$\mathbb{I}$(Rural)	87,450	0.61	1	0.49
$\mathbb{I}$(North India)	87,450	0.31	0	0.46
$\mathbb{I}$(East)	87,450	0.20	0	0.40
$\mathbb{I}$(North-East)	87,450	0.14	0	0.35
$\mathbb{I}$(South)	87,450	0.22	0	0.41
$\mathbb{I}$(West)	87,450	0.12	0	0.33

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. $\mathbb{I}$(Higher Education Women) and $\mathbb{I}$(Higher Education Men) 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.

head of household and any effect should be interpreted as an intent-to-treat effect.

From these data, I select a sample of 87,450 households as follows. I exclude households with no women or no men above 15 years of age (1.4 percent of the full sample), households in the top 1 percent of expenditure to eliminate outliers, households with a female head (3 percent) and households with head or head of household wife under 15 (0.2 percent). Moreover, for simplicity I exclude households with more than 5 women, more than 5 men, or more than 5 children under 15 (10 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 otherwise specified, budget shares are computed as percentage of total household expenditure, including durables.

Table 1.2 contains some descriptive statistics. On average, household total expenditure is equal to 8,105 Rupees (approximately 125US$). Food represents more than one third of the total expenditure, while assignable clothing budget shares are much smaller. 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 A.1 in the Appendix presents descriptive statistics for the subsamples of households with and without children under 15.

1.4.4 Estimation Strategy

I implement the model empirically by adding an error term to each equation in system (1.5) 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.[30] 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.

I take the following system of equations to the data:

$$\begin{cases} W_{food} = \tilde{\alpha}_{food} + \tilde{\beta}_{food} \ln y + \epsilon_{food} \\ W_f = \alpha_f \Lambda_f + \beta \Lambda_f \ln\left(\frac{\Lambda_f}{F}\right) + \beta \Lambda_f \ln y + \epsilon_f \\ W_m = \alpha_m \Lambda_m + \beta \Lambda_m \ln\left(\frac{\Lambda_m}{M}\right) + \beta \Lambda_m \ln y + \epsilon_m \\ W_c = \alpha_c \Lambda_c + \beta \Lambda_c \ln\left(\frac{\Lambda_c}{C}\right) + \beta \Lambda_c \ln y + \epsilon_c \end{cases} \tag{1.6}$$

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

[30] 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.

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.[31] 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. 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 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.

1.4.5 Estimation Results

Table 1.3 reports the estimated coefficients of the covariates $(X_1, ...X_n, d)$. I refer to these variables as the potential determinants of women's resource share, as they are related to bargaining power, but not

[31] 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.

[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.

necessarily in a causal sense.[34] 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, households 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 3.9 percentage points in the overall sample, by 3.3 percentage points in households with children and by 6.2 percentage points in households without children. While the number of children per se does not seem to play a significant role, the fraction of female children is positively related to Λ_f: if all children are girls, women's resource shares are 1.3 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 confirms the previous work on the plight of widows in South Asia.[35] Finally, the higher is women's age the lower is the fraction of household's total expenditure devoted to women, everything else equal.[36] This finding is particularly relevant in households without children, which suggest that women's bargaining position inside the household may be tightly related to child rearing.

Household socio-economic characteristics also play an important role. 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 (Khasi and Garo societies, for example). In contrast, North Indian women seem to have a much lower bargaining power. Households

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

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.[37]

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 higher by 1.2 to 2.3 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 women's autonomy and bargaining power. Heath and Tan (2014) find that exposure to HSA amendments is associated with a decrease of 6.6 percentage points in the probability that a woman has no say in household decisions, an increase of 8.2 percentage points 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. As their results rely on self-reported measures of independence and decision-making power within the household, a direct comparison of the magnitudes is not straightforward.

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, in the rest of the paper I focus on the results obtained estimating the two models separately.

I perform a series of robustness checks to test the sensitivity of the structural estimates. All results are included in section A.5 of the Appendix. While the findings discussed in this section are obtained using total expenditure (on 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 (table A.11). Moreover, I estimate the system in (1.6) without the Engel curve for food, which follows more closely the

[37] 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 1.3: Determinants of Women's Resource Shares

	Adult Women's Resource Share		
	All Households Sample	With Children < 15 Only	Without Children < 15 Only
	(1) NLSUR	(2) NLSUR	(3) NLSUR
No. Adult Women	0.0386***	0.0334***	0.0617***
	(0.00579)	(0.00469)	(0.00956)
No. Adult Men	-0.0248***	-0.0229***	-0.0331***
	(0.00414)	(0.00359)	(0.00652)
No. Children	-0.00209	0.00368	
	(0.00314)	(0.00240)	
Fraction of Female Children	0.0134*	0.0131**	
	(0.00787)	(0.00540)	
$\mathbb{I}$(Daughter in Law)	0.0125	0.00976	0.0218
	(0.00886)	(0.00704)	(0.0175)
$\mathbb{I}$(Unmarried Daughter above 15)	-0.0000940	0.00971	-0.00647
	(0.00911)	(0.00794)	(0.0164)
$\mathbb{I}$(Widow)	-0.00339	-0.0363***	-0.0235
	(0.0110)	(0.00966)	(0.0178)
$\mathbb{I}$(HSAA Eligible)	0.0170***	0.0120**	0.0225**
	(0.00437)	(0.00507)	(0.00978)
$\mathbb{I}$(Hindu, Buddhist, Sikh, Jain)	-0.0136	-0.0184**	-0.0113
	(0.0116)	(0.00812)	(0.0151)
$\mathbb{I}$(SC, ST, Other Backward Caste)	0.0335***	0.0681***	0.0590***
	(0.00885)	(0.00844)	(0.0125)
$\mathbb{I}$(Salary Earner)	-0.0195***	-0.0192***	-0.0170*
	(0.00664)	(0.00492)	(0.00972)
$\mathbb{I}$(Land Ownership)	-0.00719	0.00953	-0.0191
	(0.0127)	(0.00918)	(0.0182)
$\mathbb{I}$(Female Higher Education)	0.0368***	0.0317***	0.0381**
	(0.0123)	(0.00884)	(0.0155)
$\mathbb{I}$(Male Higher Education)	0.0153**	0.0415***	0.0800***
	(0.00774)	(0.00656)	(0.0121)
$\mathbb{I}$(Rural)	-0.0389***	-0.0314***	-0.0398***
	(0.0107)	(0.00685)	(0.0115)
Avg. Age Diff. (Men 15 to 79 - Women 15 to 79)	0.0140	-0.127***	0.00839
	(0.0456)	(0.0475)	(0.0807)
Avg. Age Women 15 to 79	-0.376	-0.882	-2.064*
	(0.760)	(0.758)	(1.151)
(Avg. Age Diff. (Men 15 to 79 - Women 15 to 79))^2	-0.165	0.110	-0.453**
	(0.145)	(0.134)	(0.193)
(Avg. Age Women 15 to 79)^2	0.455	2.409	4.037
	(1.778)	(1.894)	(2.722)
(Avg. Age Diff. (Men 15 to 79 - Women 15 to 79))^3	0.00222	0.550	-0.772
	(0.423)	(0.689)	(0.789)
(Avg. Age Women 15 to 79)^3	0.106	-2.138	-2.501
	(1.338)	(1.534)	(2.069)
Avg. Age Children 0 to 14	0.0331	-0.0133	
	(0.0698)	(0.0678)	
$\mathbb{I}$(North)	-0.0691***	-0.0977***	-0.0778***
	(0.0169)	(0.0168)	(0.0232)
$\mathbb{I}$(East)	-0.0116	-0.0241	-0.0168
	(0.0181)	(0.0184)	(0.0252)
$\mathbb{I}$(North-East)	-0.00613	0.0581**	0.177***
	(0.0220)	(0.0248)	(0.0297)
$\mathbb{I}$(South)	-0.0237	-0.0304*	-0.0520**
	(0.0171)	(0.0181)	(0.0244)
Constant	0.368***	0.392***	0.748***
	(0.111)	(0.101)	(0.161)
N	73,642	47,159	26,546
LL	-617,866.9	-376,832.2	-183,904.7
No. Parameters	318	318	188

*$p < 0.10$, **$p < 0.05$, ***$p < 0.01$. 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.

analysis in Dunbar et al. (2013), and show that my results hold (table A.12). In addition, I demonstrate that similar conclusions are obtained when estimating the model using data for households with one individuals for each type ($t = f, m, c$) only (columns 1 to 3 of table A.13), and when estimating a model for married couples with children, where dummies for the number of children are included as shifters of resource shares and preference parameters (column 4 of table A.13).[38] Finally, to check the validity of the theoretical framework, I use auxiliary data on singles to empirically test the assumption of Pareto efficiency (see section A.6 in the Appendix).

1.5 Why Are Older Women Missing?

That health and mortality risk are related is indisputable. Section 1.3 demonstrates that changes in intra-household bargaining power affect women's health, therefore validating my hypothesis that a weak bargaining position can explain excess female mortality. 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 discussed 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 estimated by Anderson and Ray (2010) and determine what proportion of these missing women is attributable to intra-household gender inequality and to the consequent gender asymmetry in poverty.

[38] This specification mirrors that estimated in Dunbar et al. (2013) using data from the Malawi Integrated Household Survey (IHS2). They look at married couples with up to four children.

1.5.1 Intra-household Allocation, Gender and Age

Using the estimates in table 1.3, and the analogous estimates for men and children, I predict women's, men's and children's resource shares for each household.[39] Panel A and B of table 1.4 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 independent variables are equal to their median values).[40] 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 more 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 67 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 80 percent of men's.

In figure A.2 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.

[39] 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. In the same vein, the structural estimates allow me to predict $\hat{\beta}$, $\hat{\alpha}_{food}$, $\hat{\beta}_{food}$, $\hat{\alpha}_t$, $t = f, m, c$ and therefore to investigate the shape of the Engel curves. Table A.3 in the Appendix contains some descriptive statistics of the predicted Engel curve slopes. The average assignable clothing Engel curve slope is negative for all member types - women, men, and children, while the maximum prediction is positive. For food, the model predicts downward sloping Engel curves with no exceptions.

[40] See table A.1.

Table 1.4: Predicted Resource Shares: Descriptive Statistics

	Reference Household		All households		
	Estimate (1)	Sd. Error (2)	Mean (3)	Sd. Dev. (4)	Median (5)
Panel A: Without Children < 15 Only					
Women's Resource Share $\hat{\Lambda}_f$	0.3479	0.0218	0.4449	0.1167	0.4250
Men's Resource Share $\hat{\Lambda}_m$	0.6521	0.0218	0.5551	0.1167	0.5750
Panel B: With Children < 15 Only					
Women's Resource Share $\hat{\Lambda}_f$	0.2365	0.0164	0.3113	0.0727	0.3113
Men's Resource Share $\hat{\Lambda}_m$	0.3795	0.0339	0.4698	0.1545	0.5082
Children's Resource Share $\hat{\Lambda}_c$	0.3839	0.0347	0.2189	0.1061	0.1818

I exploit the cross-sectional variation in women's age to investigate how female bargaining power varies with age.[41] For each $a = 15, ..., 79$, I compute $\hat{\Lambda}_f^a$ as the mean predicted women's resource share among all households with women's average age equal to a. Figure 1.3 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 decrease 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 0.36 in households without

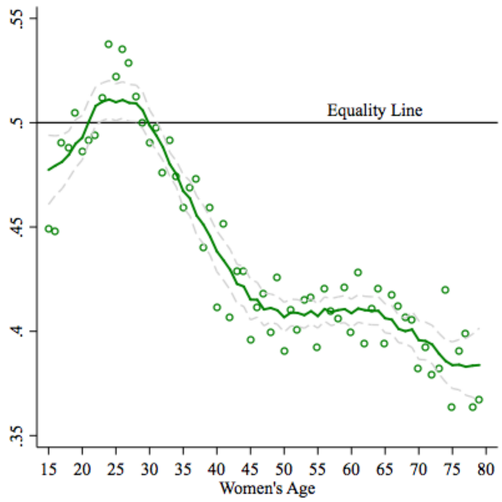
[41] Browning et al. (2013) show that there is a monotonic relationship between standard measures of bargaining power, i.e., the Pareto weights, and resource shares. In order to interpret changes in resource shares across ages as changes in bargaining power I assume this relationship to be age invariant. The interpretation in terms of intra-household resource allocation, however, does not require any additional assumption.

children and 0.2 in households with children.

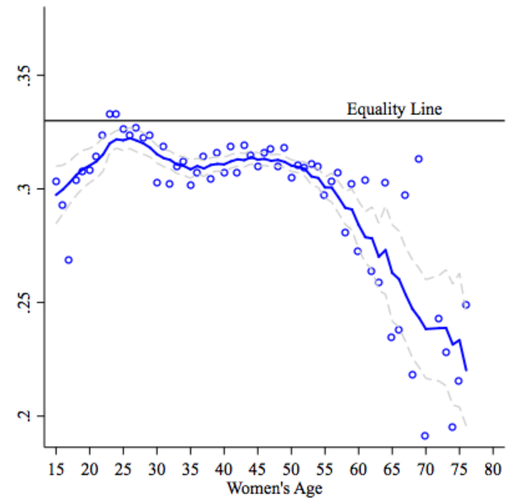
I check the robustness of my findings in several ways. First, I trace out the age profile of women's bargaining power in reference households, i.e., nuclear households with covariates set at their median values, and show that it leads to similar conclusions (see figure A.3 in the Appendix). This result is reassuring, as it shows that my findings are not driven by the correlation between women's age and other household characteristics. Second, I repeat my calculations focusing only on nuclear households. For these households the women's average age equals the age of the unique woman in the family (see figure A.4). As 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 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 on average 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 A.5).[42]

I investigate differences between genders by comparing the age profile of women's and men's resource shares. Figure A.6 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 all 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 A.7).

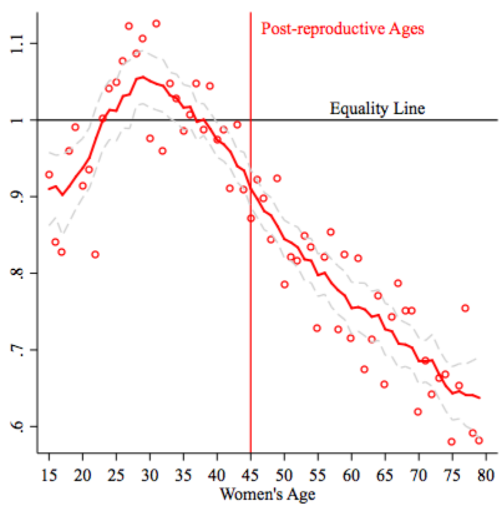
[42] As the amendments were enacted in different states at different times and applied only to women who got married after their implementation, no woman above 50 is eligible in both samples (nuclear households with and without children).



(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 1.3: Average Predicted Women's Resource Shares and Age

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 1.3, 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 35 percent of total resources, and in households without children where men and women receive both 50 percent. During the core reproductive ages of women, 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.

1.5.2 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 can explain 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. Thus, 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, as 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 (1.25 US$/day) and *average* poverty (2 US$/day).[43] As in Dunbar et al. (2013), I set the poverty lines for children to 60 percent of adults', in order to take into account the fact that children may have lower needs, while I use the same poverty lines for men and women. In section A.7 of the Appendix I investigated 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*. While a ratio equal to 1 indicates no gender asymmetry in poverty, a ratio larger than 1 indicates that female poverty is higher than that of males, and can therefore be interpreted as *excess female poverty*.

Table 1.5 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.[44]

The poverty estimates indicate that there are more households with women living below poverty line

[43] 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).

[44] The poverty head counts in the estimation sample are lower than the latest World Bank estimates (2011), which indicate that 24 and 59 percent of individuals live below the 1.25 US$/day and 2 US$/day poverty lines at 2005 international prices, respectively. This is not surprising since a selected sample is used in estimation.

Table 1.5: 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					
1.25 US$/day	0.1657	0.1177	0.2822	1.4078	0.1280
2 US$/day	0.4313	0.2759	0.5058	1.5632	0.4250
Panel B: Individual Level Poverty Rates					
1.25 US$/day	0.2098	0.1463	0.3925	1.4340	0.1540
2 US$/day	0.4965	0.3422	0.6231	1.4509	0.4784

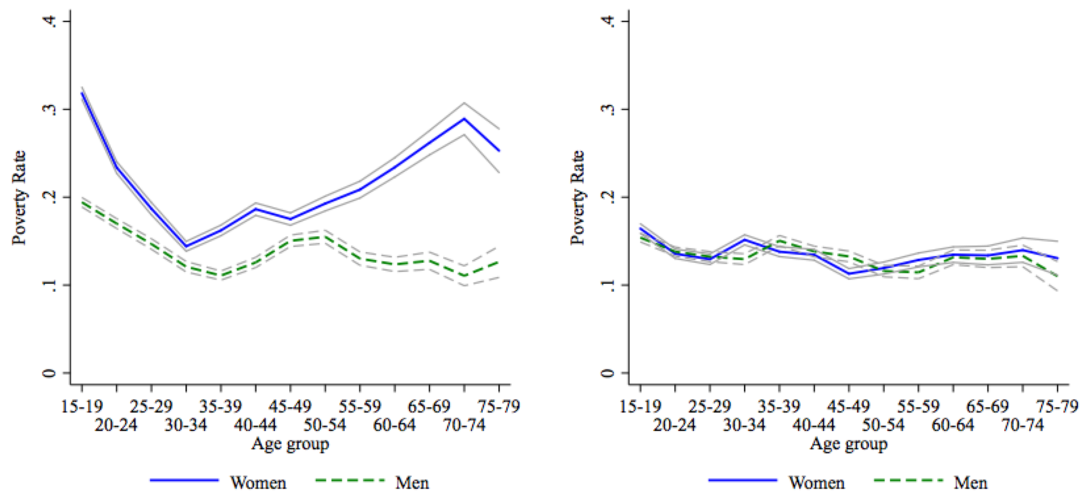
than with poor men and that there are more women living below poverty line than men. In particular, 17 percent of households in my sample have women living in extreme poverty, while less than 12 percent of households have men living on less than US$ 1.25/day. In terms of individual head count ratios, 21 percent of women live below the extreme poverty cutoff, while 15 percent of men do. Similar gender patterns hold when the alternative poverty line is considered: about 50 percent of women live with less than US$ 2/day, while one man out of three lives below this threshold. Children poverty rates are the highest, with 40 percent of children living in extreme poverty.[45] Finally, my poverty estimates provide evidence of excess female poverty, as the poverty sex ratios are above 1.

While I do not formally present robustness analysis along this dimension, these patterns are robust to the use of national poverty lines.[46] 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 1.4. However, my findings suggest that intra-household inequalities should be taken into account when measuring poverty and evaluating the effect of policies to alleviate it.

To investigate the distribution of male and female poverty across ages, I compute gender-specific head

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

[46] 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 1.25 US$/day poverty line.

Figure 1.4: Poverty Rates By Gender and Age (1.25 US$/day)

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 1.4 shows gender specific poverty rates across age groups, together with the corresponding 95 percent confidence intervals. Poverty rates are based on the World Bank 1.25 US$/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.[47] Finally, the gap between female and male poverty rates widens dramatically at ages 45-49 and above, indicating that female poverty relative to males' is particularly high at post-reproductive ages. These patterns are confirmed when the 2 US$/day poverty threshold is considered (see figure A.8 in the Appendix).

1.5.3 Explaining Excess Female Mortality

The poverty analysis discussed above indicates that gender asymmetries in resource allocation inside the household imply gender asymmetries in poverty. The fraction of women in the sample living in extreme poverty (under 1.25 US$/day) is 0.21 when the model predictions are used and unequal distribution of resources within the household is taken into account. Under the assumption of equal intra-household sharing, 14 percent of women live below the extreme poverty threshold, indicating that granting equal access to resources across genders could reduce the number of women living below 1.25 US$/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 1.4 provides a graphical illustration of this claim. The solid line represents the poverty sex ratio, obtained as the ratio between female and male poverty rates in each age group. As above, a ratio

[47] The empirical evidence in the existing literature on the relationship between poverty and age is mixed. Some empirical 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)). Third, 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 age groups.

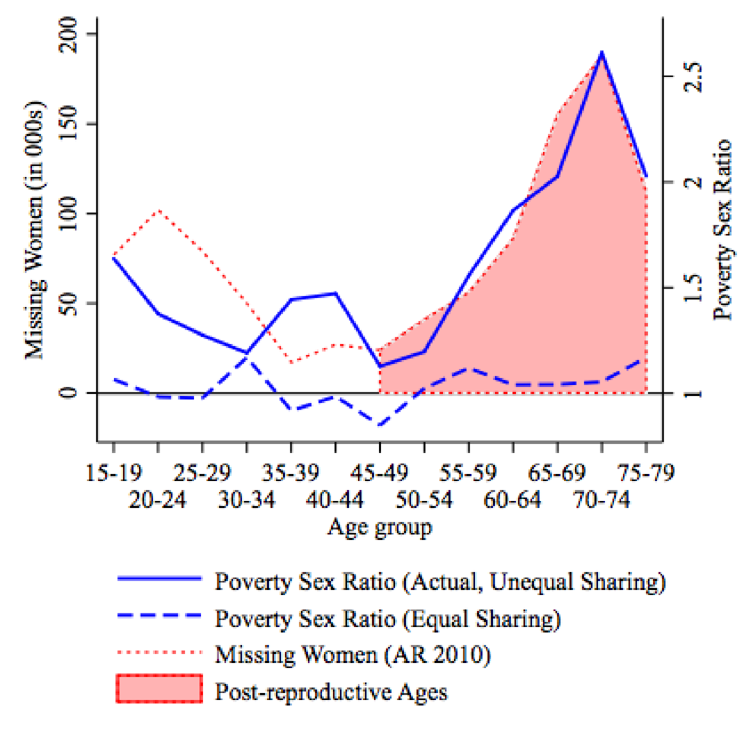
larger than 1 indicates that female poverty is higher than that of males and can be interpreted as excess female poverty. The dotted line plots the age distribution of missing women estimated by Anderson and Ray (2010). As in figure 1.1, the shaded area represents missing women at post-reproductive ages. 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 close to perfect 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, a 94 percent reduction in excess female poverty in the sample can be obtained by removing intra-household asymmetries.[48]

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 take the correlation shown in figure 1.4 seriously and assume the relationship between excess female mortality and excess female poverty to be linear and independent of age (see figure A.9 in the Appendix).[49] 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 the excess female poverty is associated with a decrease in the number of missing women by about 101,287. The R-squared of the simple linear regression model is equal to 0.69 and the estimated intercept is 17,118, while the total number of excess female deaths at ages 45 to 79 estimated by Anderson and Ray (2010) is 662,000. The

[48]The difference between the areas below the sold and the dashed line at post-reproductive ages is equal to 4.6317, which is about 94 percent of the total area between 1 and the solid line (4.9283).

[49]Whereas this is beyond the scope of this paper, further work should investigate in more details the link between relative poverty and mortality.



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 1.4 and are calculated using the 1.25 US$/day poverty line.

Figure 1.5: Poverty Sex Ratio and Missing Women by Age (1.25 US$/day)

predicted number of excess female deaths in the absence of excess female poverty is about 120,000.[50] Thus, up to 82 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.[51]

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

[50] 17,118 × 7, where 7 is the number of post-reproductive age groups.

[51] 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 1.5). In this alternative scenario, the predicted number of missing women at post-reproductive ages is about 150,000, which is 77 percent lower than the Anderson and Ray (2010) estimates.

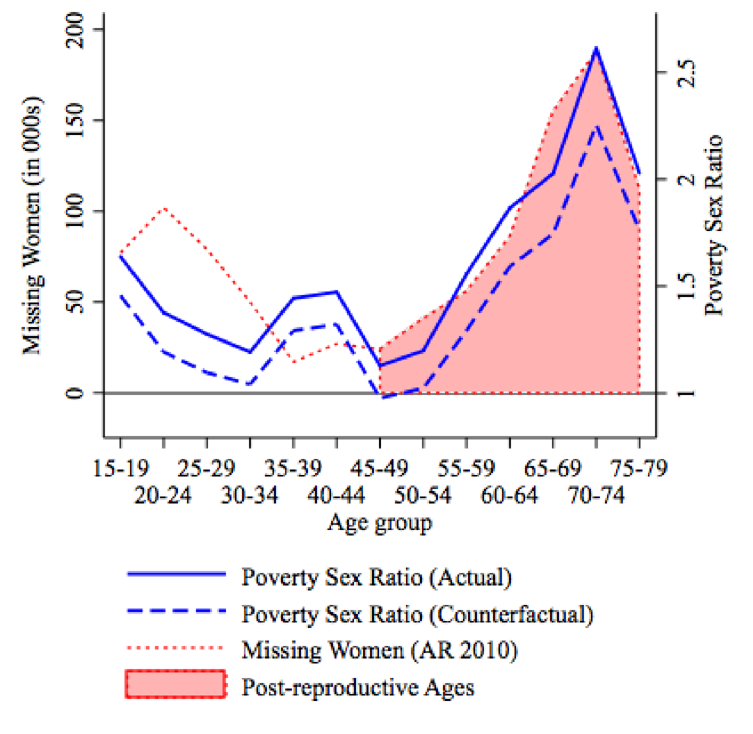
that a potentially sizable reduction in excess female mortality at older ages can be achieved by alleviating the problem of gender asymmetries within the household and the consequent asymmetries in poverty.

1.6 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 religions and ages) being exposed to the reforms. More precisely, 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 2.3 percentage points in households without children.[52] 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 1.6 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). As only one resource share is retrievable for each household member type, HSAA eligibility affects individuals of the same type by the same amount. The difference between the actual and the counterfactual poverty sex ratios

[52] I calculate counterfactual individual resource shares for women equal to $(\hat{\Lambda}_f + 0.0225)/F$ in households without children and to $(\hat{\Lambda}_f + 0.0120)/F$ in household with children. Counterfactual individual resources shares for men are equal to $(1 - \hat{\Lambda}_f - 0.0225)/M$ in households without children and to $(\hat{\Lambda}_m + 0.0016)/M$ in household with children, where 0.0016 is the estimated effect of HSAA eligibility on Λ_m. Figure A.10 in the Appendix shows the age profiles of the actual and counterfactual women's resource shares in households with and without children.



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 1.25 US$/day poverty line.

Figure 1.6: Counterfactual Experiment: HSA Amendment and Poverty Sex Ratio

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 10 percent lower in the counterfactual scenario. Moreover, analogous to figure 1.5, the area between the solid and the dashed lines represents 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 32 percent reduction in excess female poverty can be obtained by equalizing inheritance rights across genders.[53]

While taking the caveats discussed in section 1.5.3 into account, it is possible to predict the number of

[53]The difference between the areas below the sold and the dashed line at post-reproductive ages is equal to 1.5791, which is about 32 percent of the total area between 1 and the solid line (4.9283).

missing women in this counterfactual scenario (when the poverty sex ratio is equal to the dashed line in figure 1.6). The predicted number of excess female deaths under equal inheritance rights for all women is about 489,000, suggesting that up to a 26 percent reduction in the number of excess female deaths at post-reproductive ages could be obtained by granting equal inheritance rights to all women, of all ages and religions.

1.7 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)

In this paper, I focus 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 h'ealth outcomes. These reforms equalized inheritance rights between genders and therefore represent a source of exogenous variation in women's position inside the household.

I then 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 I demonstrate with my reduced-form analysis and counterfactual experiments, 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, risk of 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.

Chapter 2

Women's Empowerment and Family Health: A Two-Step Approach

2.1 Introduction

What is the effect of women's empowerment on their health outcomes? What about their children's and their spouses'? Several papers reject empirically the hypothesis that the household acts as a single decision unit maximizing a common utility function, showing that a relevant determinant of the household decisions and outcomes is the relative position of the decision makers.[1] In particular, there is consensus that the bargaining power of the decision makers leads to allocations that are closer to the preference structure of the "stronger" individual. Although this link is by now well-accepted, bargaining power is *de facto* an unobserved variable.[2] The present paper shows that the causal effect of this *unobserved*

[1]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. empirical evidence against the unitary model can be found e.g. in Lundberg et al. (1997), Attanasio and Lechene (2002) and Duflo (2003).

[2]For instance, in the case of the well cited Mexican PROGRESA program, one of the main objectives of the government was to provide a large amount of non-labor income in the hands of the mother to improve her bargaining position within the household. Although there is a large literature showing that indeed PROGRESA shifted the balance of power, virtually all the results rely on indirect evidence of power shift.

treatment on observable outcomes can be identified by combining features of *structural* and *causal* (or *reduced-form*) analyses in a *two-step* approach. We use the former to recover a measure of women's bargaining power (*step 1*), and the latter to establish the causal relationship of interest (*step 2*). We apply this approach to study the effect of intra-household empowerment on the health status of family members in India.

Within the camp of the reduced-form approach, two main strategies have been used to account for the bargaining power of household members and to study its effects on household's outcomes: either to separately estimate the parameters of an ad-hoc power function and then to add such estimates into a regression equation, or to rely on exogenous policies or programs targeting women. Quisumbing and Maluccio (2003), for instance, use datasets from Bangladesh, Indonesia, Ethiopia, and South Africa to construct a proxy of bargaining power based on the assets controlled by women evaluated at the time of marriage, and study its impact on expenditure allocations towards children. Reggio (2011), instead, uses self-reported questions about who in the household makes decisions about different assets to construct an ad-hoc measure of women's empowerment to investigate how the mother's bargaining power affects child labor in Mexico. Furthermore, legal reforms aimed at improving women's property rights, inheritance rights in particular, have been used in the literature to assess the relationship between bargaining power and women's or children's outcomes. In the Indian context, Roy (2008) shows that women's exposure to legal reforms that equalized inheritance laws between genders improves self-reported measures of bargaining power and autonomy. Roy (2013) and Deininger et al. (2013) indicate that these reforms increase female education, while Calvi (2016) finds that they improve women's health, measures of undernourishment and anaemia in particular, by increasing their bargaining position and their access to household resources.[3]

[3]Legal reforms in other countries have been studied as well. La Ferrara and Milazzo (2014) exploit an amendment to Ghana's Intestate Succession Law, finding that parents substitute land inheritance with children's education, while 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.

In contrast to this line of work, we rely on the *collective households* framework to structurally recover a non-ad-hoc measure of bargaining (or decision) power. In this framework, which was pioneered by Chiappori (1988b, 1992) and Apps and Rees (1988), and subsequently elaborated by Browning et al. (1994), Browning and Chiappori (1998), Blundell et al. (2005) and Chiappori and Ekeland (2006), a 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.[4] Recent advances in the collective households literature permit the recovery of *resource shares* (or a *sharing rule*), defined as each member's share of total household consumption. These are particularly appealing, as they provide measures of individuals' intra-household bargaining power. There are two mainstream approaches to recover the level of resource shares. On one hand, Browning et al. (2013) (BCL) provide a model that non-parametrically identifies the levels of resource shares of both adult household members based on the observation of the demand functions of single men and single women living alone.[5] An obvious limit of this identification strategy is the impossibility to identify resource shares of children, since they cannot be observed as singles. On the other hand, an alternative and powerful methodology has been proposed by Dunbar et al. (2013) (DLP). They extend the BCL model to include children and obtain identification of resource shares using Engel curves (demand equations holding prices constant) of goods that are consumed exclusively by women, men or children, e.g., clothing.

In general, however, researchers remain agnostic about how intra-household bargaining power affects the health of family members. In our empirical application, we assume that the parents are the decision makers and that they obtain utility from investing in someone's health, albeit not necessarily to the same degree. As a consequence, we allow for disagreement over this decision. We then proceed in two

[4]The efficiency assumption is quite convenient since the choice of a particular bargaining protocol to model household behavior has a strong limitation: if its empirical implications are rejected, then it is not possible to determine whether the particular choice itself is rejected or the bargaining setting in general (Vermeulen, 2002).

[5]The same strategy is followed in the applications by Lise and Seitz (2011) and Cherchye et al. (2012b). A similar approach is represented by Lewbel and Pendakur (2008). They propose some restrictions on BCL that permit the identification of the levels of adult's resource shares in a model based on comparing the Engel curves, which does not require price variation, of single men and women and men and women in childless married couples. Bargain and Donni (2012) is the first contribution that, by extending BCL, is able to identify the resource shares also of children.

steps. First, using the DLP approach, we estimate the household demand model using data with detailed expenditure data from the 2011-2012 National Sample Survey (NSS) of Consumer Expenditure. Due to the lack of NSS data on health outcomes, we use the structural estimates to perform an out-of-sample prediction on the 2005-2006 National Family Health Survey (NFHS). The latter includes the same socio-economic characteristics of individuals and households as the NSS dataset and detailed information about women's (aged 15 to 49), men's (aged 15 to 54) and children's (aged 0 to 5) health indicators. This first step allows us to recover the unobserved treatment of interest, i.e., the fraction of parents' resources that is allocated to the mother. This is a non-ad-hoc measure of women's intra-household empowerment, with a clear interpretation that is provided by the collective household model. Second, we make use a Two Stages Least Square (2SLS) approach to study the causal effect of this structurally-recovered measure of women's decision power on household members' health status.[6]

Our results are threefold. First, when we focus on women's health outcomes, we find positive and significant effects of women's decision power relative to their male counterparts. An increase in our structurally-recovered measure of bargaining power improves women's weight and their body mass index, and decreases their probability of being underweight or anemic. Moreover, our estimates indicate that more empowered mothers face higher birth intervals (or birth spacing), which may improve maternal health and reduce the risk of anemia. Second, when we focus on men's health outcomes, we do not find any significant effect of women's decision power. Finally, when we look at children, our estimates indicate that women's bargaining position does not significantly affect their anthropometric indicators such as weight and height, while it affects the likelihood of children being vaccinated. We show that this result is mainly driven by improvements in the vaccinations against polio.

Important implications can be drawn from this analysis. From a policy perspective, policies aimed at empowering women within households, such as improving women's rights to inherit property, can

[6] We here impose some structure in the second step as well and consider a linear regression model. In ongoing work, we relax these restrictions and focus on identifying the local average treatment effect.

have positive spillovers on women's health and, to a certain extent, can boost investments in children's health. From a methodological perspective, our approach shows the advantages of combining *structural* and *causal* features in conducting empirical research. By using a household model to recover a structural measure of bargaining power, we are able to answer a wider range of important research and policy questions and to provide a more accurate assessment of the effect of women's empowerment on family member's outcomes.

Proponents of causal modeling are often concerned about the complexity of the structural modeling and question the validity of its underlying assumptions. Advocates of the structural approach, instead, stress the much richer insights that can be obtained when one allows economic theory to guide the empirical work (Wolpin, 2013). Recent contributions in the econometrics literature have started to formally unify the two camps in order to overcome these divisions.[7] Heckman (2010), for example, proposes to combine the best features of both the structural and the causal modeling approaches in what he calls a "third way" of policy analysis. Along the same line, Lewbel (2016) argues that "the conflict between causal and structural modeling is foolish, even harmful", and that in some instances the best strategy for identification is to combine the strengths of both approaches. With this work, we make a further step in this direction and show that a combination of structural and causal analyses can significantly expand the set of questions that researchers can answer.

The remainder of this paper is organized as follows. Section 2.2 sets out the structural model and discusses how intra-household bargaining power may affect health decisions. Section 2.3 presents the data, the estimation strategy, and our instrumental variable, while section 2.4 discusses the empirical results. Section 2.5 concludes.

[7] See e.g., Vytlacil (2002), Heckman et al. (2006), Heckman and Vytlacil (2007), and Pearl (2009).

2.2 Intra-Household Bargaining Power and Health

Our goal is to study the effect of women's empowerment on the health status of family members. To motivate our causal analysis, we set out a theoretical framework, which augments a standard collective model of the household to include the decision of investment on the health of each family member and allows us to study such decision jointly with the intra-household resource allocation process. This same framework allows for the estimation of our unobserved treatment of interest.

2.2.1 The Model

We consider the static framework developed by DLP and extend it to include health investment decisions. We use subscripts to index goods and superscripts to index individuals. We consider three types of individuals: $t \in \{m, f, c\}$ indicating father, mother, and children, and focus only on households composed by one mother, one father, and one to four children aged 0 to 14. The two parents are the relevant *decision makers* in the household, whereas children receive a share of the resources but do not participate in the decision making process. Households differ in many other attributes such as age of members, residency, and several socioeconomic characteristics that may affect preferences. We suppress these arguments in the theoretical section but we will take this heterogeneity into account when we estimate the model empirically.

Each household consumes K types of goods with prices $p = (p_1, ..., p_K)$. y is the household income. $z = (z_1, ..., z_K)$ is the vector of observed quantities of goods purchased by each household, while $q^t = (q_1^t ..., q_K^t)$ is the vector of unobserved quantities of goods consumed by an individual t. We allow for economies of scale in consumption through a linear consumption technology, which converts purchased quantities by the household, h, in *private good equivalents*, x.[8] Let $u^t(q^t)$ denote the utility that an individual of type t would attain if she consumed the bundle of goods q^t. We refer to the latter as

[8]This specific technology assumes the existence of a $K \times K$ matrix A such that $h = A(q^f + q^m + sq^c)$.

material welfare. Each decision maker has caring preferences toward her children and her spouse, that is, her individual total utility depends, in a weakly separable manner, on the material welfare of the other household members, i.e. the utility they derive from consumption ($u^{-t}(q^{-t})$). For children, $u^c(q^c)$ can be interpreted as the child's actual utility function over the bundle of goods q^c that the child consumes, or the utility function that parents believe the child has. For simplicity, it is assumed that each child of the same family is assigned the same utility function.[9]

In addition, we assume that each parent attaches some value to the health status of household member's $j \in \{m, f, c^1, \ldots, c^s\}$, where j now identifies either herself, her spouse, or the children, and therefore to the decision of investing in j's health. These decisions are summarized by the J-dimensional vector H ($J = s + 2$), whose elements indicate whether the parents decide to invest or not to invest in j's health. We model the utility parents get from each member's health as an expenditure in a public good. Hence H is interpreted as the outcome of a joint decision of expenditure on a good which has public good features. For simplicity, parents are assumed to make single period (myopic) decisions about to what extent to invest in j's health. This is a restrictive assumption in general, but it is quite sensible in our framework as we focus on poor or constrained households. Moreover, we assume that each health investment decision is independent of another. In other words, given any two family members, the decision of investing in one's health affects the decision of investing in the other one only through the budget constraint.

In line with the collective modeling of household behavior, we assume that household decisions result from a cooperative bargaining process among the decision makers, which leads to a Pareto efficient outcome. The collective framework implies that households will choose consumption levels and health investments in order to maximize a weighted sum of parents' utilities:

$$U^H = RU^m(u^m, u^f, u^c, H) + (1-R)U^f(u^f, u^m, u^c, H) \tag{2.1}$$

[9] It is straightforward to extend the model to allow each child to have a different utility function. What we need is information on a separate private assignable good for each child to achieve identification of each child's resource share.

where $R = R(p, y, D)$ is the *Pareto weight*, which may depend on prices p, total expenditures y, and on a vector of distribution factors D. The latter are defined as variables with no direct impact on preferences or budget constraint, but that may influence the decision process. From a bargaining point of view, R can be seen as a measure of the mother's influence in the decision process: the larger R, the greater is the weight that m's preferences receive in the resulting household's allocation of resources.

Each parent's preferences in (2.1) are assumed to be separable in both the material well-being of the other family members and the utility derived from H. For simplicity, we parametrize the individual parents' utilities as follows:

$$U^m = u^m(q^m) + \delta^m_f u^f(q^f) + \delta^m_c u^c(q^c) + H'\alpha^m$$

$$U^f = u^f(q^f) + \delta^f_m u^m(q^m) + \delta^f_c u^c(q^c) + H'\alpha^f$$

where α^m and α^f are vector of parameters, whose elements, α^m_j and α^f_j, describe how much additional utility each parent m or f attaches to the health investment on j. We can rearrange equation (2.1) as follows:

$$U^H = U^{Cons} + RH'\alpha^m + (1-R)H'\alpha^f \tag{2.2}$$

where U^{Cons} is the total utility obtained from consumption by the household. Choosing H consists of deciding whether to invest or not in each family member's health. The choice must be optimal in the sense that it maximizes (2.2) subject to the following budget constraint:

$$y \leq z'p + K_H \tag{2.3}$$

where K_H is the total cost of health investment. Under the assumption that each decision maker's power $R(p, y, D)$ is fixed, we can reinterpret the household program as that of a unique entity that maximizes a common welfare index. Alternatively, we can assume that the decision power is in the hands of one individual (i.e., $R(p, y, D) = 0$) and that this benevolent dictator makes optimal decisions for all members of the household. These are the common restrictions implied by the *unitary* model of the household.[10] Our *collective* model framework is more general and allows us to analyze the role played by the decision makers' bargaining power in the health investment decisions, especially when there is disagreement. The following example may help clarifying this point.

2.2.2 Bargaining Power and Health Decisions: An Illustrative Example

H is a $J-$dimensional vector, whose the elements are $h_j \in \{0, 1\}$, and indicate whether there is an investment in j's health. Given that the choice of H is optimal, it is in particular preferred to $\tilde{H}$, which differs from H only in the health investment decision on member j. Without loss of generality, assume that parent's decide to invest in j's health, so that $h_j = 1$ and $\tilde{h}_j = 0$. Then it must be that:

$$\begin{aligned} U_0^H &< U_1^H \\ U_0^{Cons} + \tilde{H}' \left(R\alpha^m + (1-R)\alpha^f \right) &< U_1^{Cons} + H' \left(R\alpha^m + (1-R)\alpha^f \right) \\ U_0^{Cons} + \tilde{h}_j \left(R\alpha_j^m + (1-R)\alpha_j^f \right) &< U_1^{Cons} + h_j \left(R\alpha_j^m + (1-R)\alpha_j^f \right) \\ U_0^{Cons} &< U_1^{Cons} + R\alpha_j^m + (1-R)\alpha_j^f \end{aligned} \tag{2.4}$$

This allows us to set up a latent function that determines the decision of investing in j's health, conditional on the other health investment decisions. The parents will decide to invest in j's health if the latent variable h_j^* is greater than the threshold zero and not to invest otherwise:

$$h_j^* = U_1^{Cons} - U_0^{Cons} + R\alpha_j^m + (1-R)\alpha_j^f \tag{2.5}$$

[10]Note that both restrictions generate the "income-pooling hypothesis" according to which, any distributional variable should have no effect on household decisions.

Since the bargaining weights do not depend on the health investment decisions, the sharing rule for the household is, in some sense, fixed with respect to the decision of H. Therefore, each household member can compute her own overall utility in both status. The following table summarizes the possible scenarios that our model allows for:

h_j	m	f
0	U^m	U^f
1	$\tilde{U}^m$	$\tilde{U}^f$

Here U^j are the parents' utilities in case there is no health investment on individual j, whereas $\tilde{U}^j$ refers to the utilities when there is a positive health investment decision. Three situations may emerge:

$$\begin{aligned} \tilde{U}^m > U^m \text{ and } \tilde{U}^f > U^f \quad &\rightarrow \quad \text{Agreement, Investment in } j \\ \tilde{U}^m < U^m \text{ and } \tilde{U}^f < U^f \quad &\rightarrow \quad \text{Agreement, No Investment in } j \\ \tilde{U}^m < U^m \text{ and } \tilde{U}^f > U^f \text{ or } \tilde{U}^m > U^m \text{ and } \tilde{U}^f < U^f \quad &\rightarrow \quad \text{Disagreement} \end{aligned}$$

In the first two cases, there is agreement among the household decision makers concerning the investment in j's health. In particular, in the first instance both parents gain by investing in j's health, whereas in the second situation they both lose. In these instances, the decision of the household is trivial: either they invest in the first case, or do not invest in the second. In the third case, however, there is a disagreement with respect to the health investment decision. In this instance, the relative power of the parents in the household decision process may play a crucial role.

An important remark is in order. In our setup, we assume that the Pareto weights of the parents are independent of the health investment decisions. This allows us to obtain a simple expression for the impact of decision power on the health status choice made by the household. However, this also restricts our ability to fully exploit the mechanisms of compensation and transfers within the collective household literature.[11] While this is a quite interesting generalization of our approach, we leave the relaxation of this assumption to future work.

2.2.3 Recovering the Unobserved Treatment

We rely on Dunbar et al. (2013) (DLP) and Browning et al. (2013) (BCL) to recover a measure of mother's bargaining power relative to the father, i.e., the *unobserved treatment*. With the former, we obtain resource shares for fathers, mothers, and children. With the latter, we show that, under some functional form assumption, the decision power weight R can be measured as the fraction of parents' resources that are devoted to the mother. The total household utility from consumption can be rewritten as follows:

$$\begin{aligned} U^{Cons} &= R\left(u^m(q^m)+\delta^m_f u^f(q^f)+\delta^m_c u^c(q^c)\right)+(1-R)\left(u^f(q^f)+\delta^f_m u^m(q^m)+\delta^f_c u^c(q^c)\right) \\ &= R\tilde{u}^m+(1-R)\tilde{u}^f+(R\delta^m_c+(1-R)\delta^f_c)u^c \end{aligned} \tag{2.6}$$

Equation (2.6) shows that, conditional on the health investment decisions, the household problem can equivalently be represented as a maximization of a weighted sum of the three utility functions from consumption. Moreover, given the separability assumptions made in section 2.2.1, conditional on the

[11] In particular, it has been shown by Chiappori (1992) that the assumption of Pareto optimal household decisions allows for a representation in terms of autonomous (individual) decisions, where first household income is divided among the household members and then private consumption decisions are taken, given the respective shares obtained from the first stage. This representation has been referred to as the sharing rule representation and it is by now a classic result in the literature on collective household models. Given the heterogeneity of tastes between parents on health status, it seems therefore very sensible to allow for compensatory transfers at the level of the household, through the sharing rule. This kind of reasoning would require us, however, to drop the assumption that Pareto weights are independent of the health status decision.

health investment decision, the household program is as follows:

$$
\begin{aligned}
&\max_{q_f, q_m, q_c, z} \mu^f u^f + \mu^m u^m + (1 - \mu^f - \mu^m) u^c \\
&\quad \text{such that} \\
&\quad z = A(q^f + q^m + sq^c) \\
&\quad y = z' p + K_H
\end{aligned}
\tag{2.7}
$$

The solutions to this program provide the bundles of private good equivalents, q^t. Pricing those at the shadow prices $A'p$ gives the *resource shares*, η^m, η^f and $\eta^c = 1 - \eta^f - \eta^m$, that are the fractions of household total resources that are devoted to individuals of type t. BCL show that there exists a monotonic correspondence between the weights μ^t 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.

The household program (2.7) can be decomposed in two steps: 1) the optimal allocation of resources across members and 2) the individual maximization of their own utility function. This results follows directly from the second welfare theorem, as any Pareto efficient allocation can be implemented as an equilibrium of this economy, possibly after lump sum transfers between members. Conditional on knowing η^t, each household member chooses q^t as the bundle maximizing u^t subject to the Lindahl type shadow budget constraint $\sum_k A_k p^k q_t^k = \eta^t y$. Substituting the indirect utility functions $V^t(A'p, \eta^t y)$ in the program above simplifies the optimization problem to the choice of optimal resource shares subject to the constraint that total resources shares must sum to one. The household program can be therefore written as:

$$\max_{\eta^f,\eta^m,\eta^c} \mu^f V^f + \mu^m V^m + (1-\mu^f-\mu^m)V^c$$
$$\text{such that} \tag{2.8}$$
$$\eta^f + \eta^m + s\eta^c = 1$$

An exact relationship between the shares η and the weights μ can be obtained from the three-member analogue of Proposition 2 in BCL, which states that:

$$\kappa\mu^m = \frac{\dfrac{\partial V^c(A'p,\eta^c y)}{\partial \eta^c}}{\dfrac{\partial V^m(A'p,\eta^m y)}{\partial \eta^m}} \text{ and } \kappa\mu^f = \frac{\dfrac{\partial V^c(A'p,\eta^c y)}{\partial \eta^c}}{\dfrac{\partial V^f(A'p,\eta^f y)}{\partial \eta^f}} \tag{2.9}$$

where V^t is the indirect utility function of household member t.[12] Assume each household member t has a subutility given by Muellbauer's Price Independent Generalized Logarithmic (Piglog) model. In this framework, the indirect utility functions can be written as follows:

$$V^t(A'p,\eta^t y) = \psi^t\left[\ln\left(\ln\left(\frac{\eta^t y}{G^t(A'p)}\right)\right) + F^t(A'p)\right] \tag{2.10}$$

where $A'p$ and $\eta^t y$ are the relevant (Lindahl) prices and income for the household member's own optimization program, respectively. ψ^t is differentiable and strictly monotonically increasing and in this case does not depend on prices for the private assignable goods of other household members. Under a convenient choice of ψ^t, we can obtain the following relation:

$$R = \frac{\mu^m}{\mu_m + \mu^f} = \frac{\eta^m}{\eta^m + \eta^f} \tag{2.11}$$

which allows us to use the fraction of parents' resources devoted to the mother as an index of mother's

[12]Notice that, using BCL notation, $\kappa = (\lambda/y)\sum_{i=1}^n x_i \sum_{j=1}^n p_j \partial F^j(x)/\partial x_i$, where λ is the Lagrange multiplier coming from Program (2.8).

decision power of in the household.[13]

2.2.4 Health Production

A fully specified structural model would require a theory of health production within the household, stemming from the underlying health investment decisions. Our approach, instead, does not require this further step, and allows us to focus directly on the *causal* relationship of interest in a reduced-form setting, while remaining agnostic on the exact health production mechanism. In other words, we use economic theory to recover our unobserved treatment, but we avoid the complexity of structural modeling and estimation when investigating the relationship of interest, taking full advantage of clear and effective tools of causal inference. For simplicity, in order to study the effect of R on Y, we now use a linear regression model. In ongoing work, we relax this assumption.

2.3 Estimation

We use a two-step estimation procedure to assess the role woman' bargaining power plays in affecting the health status of household members. First, we structurally estimate a measure of intra-household bargaining power. Second, we use this measure to assess the causal effect of women's empowerment on family health in a reduced-form setting.

[13] $\psi(.)$ is set to exp(.), which simplifies the expression for $\kappa\mu^m$ described above:

$$\frac{\partial V^t}{\partial \eta^t} = \exp F^t \frac{\partial}{\partial \eta^t}\left[\ln\left(\frac{\eta^t y}{G^t}\right)\right] = \frac{\exp F^t}{\eta^t} \tag{2.12}$$

This results in the following expressions for the parents' decision power weights:

$$\kappa\mu^t = \frac{\eta^t \exp F^c}{\eta^c \exp F^t} \tag{2.13}$$

where $F^t(A'p) = A'p^t\phi(A'p)$ with $\phi(A'p)$ independent of household member t. This form is implied by the identifying assumption of similar preferences across people (SAP). See Dunbar et al. (2013). Thus, an assumption of proportionality between prices of the exclusive goods for the different members t is needed to simplify further: $\kappa\mu^t = \frac{\eta^t}{\eta^c}\frac{p^c}{p^t}$ and $\kappa'\mu^t = \frac{\eta^t}{\eta^c}$. This assumption may not be as strong as it looks since any price variation in the exclusive goods used in this application, clothing and shoes, is very likely to stem solely from differences in remoteness of the different locations. Such variation should affect the prices for all clothes in very similar measure.

2.3.1 Data

We implement our empirical analysis by means of two Indian datasets: the 2011-2012 National Sample Survey (NSS) of Consumer Expenditure (68^{th} round) and the 2005-2006 National Family Health Survey (NFHS-3).

NSS data. The NSS Consumer Expenditure Survey contains detailed data on household expenditure and details about household socio-economic characteristics, and other particulars of household members. We select a sample of households with one woman and one man above age 15, and up to 4 children under 15. Moreover, we exclude households with no women or no men above 15 years of age, households in the top 1 percent of expenditure, households with a female head and households with head or head of household wife under 15. Finally, we exclude households reporting to have performed any ceremony during the month prior to the survey, as unusual purchases of clothing items and non-standard expenditure patterns may occur for festivities and ceremonies.

Among other items, households are asked to report how much they spent on clothing, bedding, and footwear. Given the detailed breakdown of clothing expenditure, it is possible to identify the expenditure on items of clothing that can be assigned to women, men, and children. Observability of *private assignable* goods (goods that are consumed exclusively by one known member or by household members of the same type) is necessary for the identification and the estimation of our measure of parent's relative bargaining power. We follow Calvi (2016) and 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, we combine expenditure on dhoti, lungi, kurta-pajamas suits for males, pajamas, and salwar. For children, we use expenditure on school uniforms and infant clothing.[14] Unfortunately, the NSS Sur-

[14] Notice that Tommasi and Wolf (2016) have recently shown that when the data exhibit relatively flat Engel curves in the consumption of the private assignable goods, the DLP model is weakly identified and induces high variability and an implausible pattern in least squares estimates of resource shares. However, the advantage of following Calvi (2016) is that households in her dataset have a large variation in the consumption of private assignable goods, which facilitates identification. Hence the problem of weak identifiability in this context is very mild and negligible.

vey does not include any information on individuals' health status. Table 2.1 contains some descriptive statistics. For clothing items, the survey entry refers to the expenditure occurred in the past 365 days. For simplicity, we convert annual into monthly figures. Budget shares are computed as percentage of total household expenditure, including durables.

NFHS-3 data. The 2005-2005 National Family Health Survey includes detailed information about women's (aged 15 to 49) and men's (aged 15 to 54) health indicators, such as anthropometrics and measures of anaemia, together with socio-economic characteristics of individuals and households that render the out-of-sample prediction of resource shares feasible. For children born in the 5 years prior to the date of interview, we observe anthropometric indicators and vaccination records. To investigate the causal effect of women's intra-household empowerment on the health outcomes of family members, we consider the NHFS women, men and children datasets separately. Table 2.2 presents some descriptive statistics.

For women, men and children, we observe anthropometric indicators, such as body mass index (BMI), weight, height, and levels of anaemia.[15] BMI is defined as weight in kilograms divided by height in meters squared. A cut-off point of 18.5 is used to define thinness or acute undernutrition, and a BMI of 23 or above indicates overweight or obesity for Asian Indians (Shiwaku et al., 2004). Anaemia, instead, is a condition in which the number of red blood cells or their oxygen-carrying capacity is insufficient. Although its primary cause is iron deficiency, it coexists frequently with a number of other causes, such as malaria, parasitic infection, and nutritional deficiencies; 90 percent of anaemia sufferers live in developing countries. Prevalence of anaemia in South Asian countries is among the highest in the world.[16] For women, we also compute their average birth spacing, i.e., the average number of months between births. The adverse health consequences of a short spacing between births are well known: as

[15] See Schultz (2009) for more details.
[16] See Kaur (2014).

Table 2.1: NSS Consumer Expenditure Data and NFHS Data

	NSS Sample				NFHS Sample			
	Obs.	Mean	Median	St. Dev.	Obs.	Mean	Median	St. Dev.
Total Expenditure (Rupees)	15,173	7,333	6,002	5,928				
Woman's Assignable Clothing Budget Share	15,164	1.22	1.06	0.99				
Man's Assignable Clothing Budget Share	15,164	1.42	1.18	1.26				
Children's Assignable Clothing Budget Share	15,164	0.75	0.59	0.83				
$\mathbb{I}$(1 child)	15,173	0.27	0.00	0.44	23,699	0.26	0.00	0.44
$\mathbb{I}$(2 children)	15,173	0.44	0.00	0.50	23,699	0.39	0.00	0.49
$\mathbb{I}$(3 children)	15,173	0.21	0.00	0.41	23,699	0.24	0.00	0.42
$\mathbb{I}$(4 children)	15,173	0.08	0.00	0.27	23,699	0.11	0.00	0.31
Fraction of Female Children	15,173	0.44	0.50	0.37	23,699	0.47	0.50	0.37
$\mathbb{I}$(Hindu, Buddhist, Jain, Sikh)	15,173	0.79	1.00	0.41	23,695	0.77	1.00	0.42
Woman's Age	15,173	32.80	32.00	6.54	23,697	30.22	30.00	7.32
Gender Age Gap	15,173	4.51	5.00	5.93	23,662	5.80	5.00	5.01
Children's Avg. Age	15,173	7.70	8.00	3.46	23,699	6.46	6.33	3.58
Wealth Index	15,173	0.38	0.37	0.17	23,699	0.39	0.36	0.23
$\mathbb{I}$(Woman's Higher Education)	15,173	0.19	0.00	0.39	23,699	0.09	0.00	0.28
$\mathbb{I}$(Man's Higher Education)	15,173	0.28	0.00	0.45	23,699	0.13	0.00	0.34
$\mathbb{I}$(Rural)	15,173	0.59	1.00	0.49	23,699	0.54	1.00	0.50
$\mathbb{I}$(North India)	15,173	0.27	0.00	0.44	23,699	0.30	0.00	0.46
$\mathbb{I}$(East India)	15,173	0.18	0.00	0.39	23,699	0.16	0.00	0.37
$\mathbb{I}$(North-East India)	15,173	0.16	0.00	0.37	23,699	0.20	0.00	0.40
$\mathbb{I}$(South India)	15,173	0.28	0.00	0.45	23,699	0.21	0.00	0.41
$\mathbb{I}$(West India)	15,173	0.11	0.00	0.31	23,699	0.13	0.00	0.34
$\mathbb{I}$(HSA Eligible)	15,173	0.29	0.00	0.46	21,110	0.06	0.00	0.24

Budget shares are multiplied by 100. Woman's assignable clothing includes expenditures on saree, shawls, chaddar, and kurta-pajamas suits for females; man'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. The household wealth index is obtained using principle component analysis. $\mathbb{I}$(Higher Education Women) and $\mathbb{I}$(Higher Education Men) 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.

the mother is not allowed sufficient time to restore the adequate supply of nutrients, short spacing is found to affect maternal health and increase the risk of anemia (Conde-Agudelo and Belizán (2000) and Milazzo (2014)). Finally, for children we also observe whether a child received a BCG vaccine (against tuberculosis), one to three DPT vaccines (against diphtheria, pertussis, and tetanus), and one to four polio vaccines (at birth and one to three after). For simplicity, we consider whether or not a child has ever received a BCG, DPT or polio vaccine, or any vaccine to prevent diseases.

Table 2.2: Causal Inference, NFHS Data

	Women (age 15 to 49)			Men (age 15 to 54)			Children (age 0 to 5)		
	Obs.	Mean	St. Dev.	Obs.	Mean	St. Dev.	Obs.	Mean	St. Dev.
Weight (kg)	16,608	48.24	10.08	8,587	56.86	10.82	11,063	10.67	3.29
BMI (kg/m^2)	16,593	20.89	3.94	8,584	21.14	3.50	11,037	15.13	5.07
$\mathbb{I}$(BMI $\leq$18.5)	16,593	0.30	0.46	8,584	0.24	0.43			
$\mathbb{I}$(Anemic)	15,614	0.16	0.37	7,989	0.16	0.36			
Avg. Birth Gap (months)	14,082	35.12	17.93						
Height (cm)							11,037	83.95	13.93
$\mathbb{I}$(BCG)							12,022	0.78	0.42
$\mathbb{I}$(DPT)							11,959	0.75	0.43
$\mathbb{I}$(Polio)							12,063	0.91	0.29
$\mathbb{I}$(Any Vaccination)							12,078	0.92	0.28
Woman's Decision Making Index	17,227	28.53	34.99						
Age	17,241	29.85	6.01	9,067	35.43	6.73	39,134	7.43	4.92
$\mathbb{I}$(Worked in Past 12 Months)	17,239	0.42	0.49	9,066	0.44	0.50	43,410	0.45	0.50
Education (years)	17,240	5.05	5.12	9,067	5.21	5.13	43,415	4.24	4.85
$\mathbb{I}$(Female)							43,416	0.48	0.50
$\mathbb{I}$(Birth Ord. = 1)							43,416	0.39	0.49
$\mathbb{I}$(Birth Ord. = 2)							43,416	0.33	0.47
$\mathbb{I}$(Birth Ord. = 3)							43,416	0.19	0.39
$\mathbb{I}$(Birth Ord. = 4)							43,416	0.09	0.29
$\mathbb{I}$(Rural)	17,241	0.55	0.50	9,067	0.53	0.50	43,416	0.59	0.49
$\mathbb{I}$(Sch. Caste, Sch. Tribe, OBC)	17,241	0.68	0.47	9,067	0.71	0.45	43,416	0.71	0.45
Wealth Index	17,241	0.38	0.23	9,067	0.38	0.22	43,416	0.35	0.22
Number of children (under 15)	17,241	2.26	0.96	9,067	2.27	0.95	43,416	2.57	0.95
$\mathbb{I}$(Hindu, Buddhist, Jain, Sikh)	17,241	0.80	0.40	9,067	0.78	0.42	43,416	0.78	0.41
$\mathbb{I}$(BPL card)	17,140	0.19	0.39	9,016	0.19	0.39	43,173	0.21	0.41
Fraction of Female Children	17,241	0.47	0.36	9,067	0.47	0.36	43,416	0.47	0.33
$\mathbb{I}$(HSA Eligible)	14,433	0.09	0.29	7,307	0.15	0.36	37,457	0.07	0.26

Age is defined in years for men and women, and in months for children. For children, $\mathbb{I}$(Worked in Past 12 Months) and education refer to the mother.

2.3.2 Indian Inheritance Rights Reforms

A woman's right to inherit land and other property is often claimed to play a significant role in determining women's position within the household (World Bank, 2014). Inheritance rights in India differ by religion and, for most of the population, are governed by the Hindu Succession Act (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 therefore did not apply to individuals of other religions, such as Muslims, Christians, Parsis, Jews, and other minority communities.[17] 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. Gender inequalities, however, remained 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.[18] In the decades following the introduction of Hindu Succession Act, state governments passed amendments that equalized 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 amendments 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.

Both the NSS and the NFHS survey contain information about women's year of birth, state of resi-

[17] 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).

[18] 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.

dence, and religion. We 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.[19] Our choice to focus on women's eligibility (rather than their actual exposure) to the inheritance rights amendments has a double motivation. First, while the NFHS data include information about each woman's year at marriage and is therefore suitable for the identification of women who are Hindu, Buddhist, Sikh or Jain women and who were not yet married at the time of the amendment, the NSS does not. Second, as exposure to the HSA amendments is determined by each woman's year of marriage, we focus on an intent-to-treat effect and address the concerns about a potential endogeneity of treatment. As shown in Table 2.1, in both samples about 80 percent of the sample is Hindu, Buddhist, Jain, or Sikh. However, the percentage of households where the woman is eligible to the Hindu Succession Act amendments is much higher in the NSS. This is mainly due to the timing of the surveys, 2011-2012 for the NSS and 2005-2006 for the NHFS, as the former includes a larger number of women who were unmarried at the time of the national amendment in 2005.

2.3.3 Empirical Strategy

Step 1: Structural Estimation. To recover the unobserved treatment R, we estimate the resource shares η^t for all household members. To this aim, we set up an Engel curve system as in DLP. This system is based on the Piglog preferences, whose indirect utility function is defined in Equation (2.10). Under these functional form assumption, we obtain a set of three budget share equations for assignable

[19]We use 14 and 23 as they are the 10th and 90th percentiles of women's age at marriage in the NHFS-3 sample. This variable is therefore fully determined by each woman's religion, year of birth and state.

clothing that are linear in the total household expenditure:

$$\begin{cases} W^f = \eta^f[\delta^f + \beta^f(lny + ln\eta^f)] \\ W^m = \eta^m[\delta^m + \beta^m(lny + ln\eta^m)] \\ W^c = s\eta^c[\delta^c + \beta^c(lny + ln\eta^c - lns)] \end{cases} \tag{2.14}$$

where $s\eta^c = 1 - \eta^f - \eta^m$. y is the total household expenditure (in Rupees) reported for the month prior to the survey, and w^t are the budget shares spent on assignable clothing and food, respectively.

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 (2.14) can be identified by a linear regression of the household budget shares W_t on a constant term and $\ln y$. $\beta\eta^t$ is the slope of type t's assignable clothing 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.

We account for observable heterogeneity across households by specifying δ^t $(t = f, m, c)$ and β as linear functions of observable household characteristics (Z). Moreover, η^t depend linearly on X and one *distribution factor*, D, i.e.,

$$\eta^t = \theta_0^t + \theta_1^t Z_1 + ... + \theta_n^t Z_n + \theta_D^t D \tag{2.15}$$

This characterization renders the system of Engel curves non-linear. The vector $Z = (Z_1, ..., Z_n)$ includes, among other variables, details about the composition of the household, socio-economic characteristics, such as demographic group, religion, measures of parental education, and a household wealth index, and the mother's, the father's, and the average age of children. It also contains region fixed effects

(South, East, West, North-East, North, with West being the excluded category) and dummy variables identifying the number of children under 15, for living in rural areas, and for belonging to a disadvantaged demographic group (Scheduled Caste, Scheduled Tribe or other backward classes). Although distribution factors are not required for identification of the resource shares, it is going to be crucial for the identification of a causal effect of interest in the reduced form model. Thus, we include the mother's eligibility to the HSA amendments as a factor affecting resource allocation but not preferences.

We implement the model empirically by adding an error term to each equation in system (2.14) and by imposing similarity of preferences over private assignable goods. Since the error terms may be correlated across equations, we 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. We estimate the model using the NSS sample discussed in section 2.3.1. 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).

Once we obtain estimates $\widehat{\theta}^t = (\widehat{\theta}_0^t, \widehat{\theta}_1^t, ..., \widehat{\theta}_n^t, \widehat{\theta}_D^t)$ for the underlying θ, we predict the resource shares parameters and we construct our measure of the mother's decision power in the household, i.e., the fraction of parents' resources that are devoted to the mother, as:

$$\widehat{R}(D, Z, \widehat{\theta}) = \frac{\widehat{\eta}^m(D, Z, \widehat{\theta}^m)}{\widehat{\eta}^m(D, Z, \widehat{\theta}^m) + \widehat{\eta}^f(D, Z, \widehat{\theta}^f)} \tag{2.16}$$

Since we observe $Z = (Z_1, ..., Z_n)$ and D in the NSS and in the NFHS-3 sample, we can compute $\widehat{R}(D, Z, \widehat{\theta})$ in both samples. As the NSS sample does not include health information about family members, we investigate the effect of mothers' decision power on family member's health outcomes using the NFHS-3 data only.

Step 2: Causal Inference. We study the relationship between R on H by imposing more structure and regressing Y on $\widehat{R}$ and X. Our preferred specification is as follows:

$$Y_{ircs} = \lambda \ln \widehat{R}_{ircs} + X'_{ircs}\gamma + \alpha_r + \alpha_c + \alpha_s + \alpha_{rs} + \alpha_{rc} + \delta_s c + \epsilon_{ircs} \tag{2.17}$$

where Y is a measure of the health status of individual j, either a health outcome or a binary decision of health investment, and X is a vector of covariates that include variables that characterize the individual (e.g., age and gender), the household (e.g., wealth, demographics, number of children), and the environment of the household (e.g., fixed effects and time trends).[20]

To allow a causal interpretation to our estimate, we instrument $\ln \widehat{R}$ with women's *eligibility* to the Hindu Succession Act (HSA) amendments. The first stage of our Two Stages Least Squares (2SLS) approach is defined as:

$$\ln \widehat{R}_{ircs} = \tilde{\lambda} D_{ircs} + X'_{ircs}\gamma + \alpha_r + \alpha_c + \alpha_s + \alpha_{rs} + \alpha_{rc} + \delta_s c + \epsilon_{ircs} \tag{2.18}$$

where D is defined 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.[21] The second stage consists of estimating equation (2.17), where now the regressor of interest $\ln \widehat{R}$ is replaced by the fitted values from the first-stage regression. λ is the parameter of interest and measures the average effect of a one percent increase in the unobserved treatment $\widehat{R}$, for those households in which the relative bargaining position of the decision makers is influenced by women's exposure to the HSA amendments.

Identification of a causal effect comes from the assumption that D is uncorrelated with the error term

[20] We consider the logarithm of the $\widehat{R}$, instead of the variable in levels, as we believe that percentage changes matter more than a unit change, i.e., we believe that an increase in $\widehat{R}$ from 0.20 to 0.30 does not have the same effect of an increase from 0.60 to 0.70.

[21] See section 2.3.1 for details.

in our outcome equation (2.17), while adequately correlated with $\ln\widehat{R}$. The exclusion restriction needed for identification is that the mother being eligible to the Hindu Succession Act Amendments has an effect on the health outcomes of interest only through the increase in mother's decision power. To strengthen the argument that this is the case, we include state, cohort, religion, religion-cohort, and state-religion fixed effects, and state specific linear trends in all specifications.

2.4 Results

In this section we present the results of both the structural (step 1) and causal (step 2) estimates. Unless otherwise indicated, in all 2SLS regressions we present standard errors calculated from a bootstrap with 2,000 replications. Each replicate draws a sample of primary sampling units (village) to allow for intra-cluster correlation in the standard errors.[22] We estimate the causal model in the women, men, and children separately, which allows us to investigate the impact of mother's decision power on her, her spouse's, and her children's health outcomes.

2.4.1 Step 1: Structural Estimation

Table 2.3 in the Appendix reports the estimated marginal effects $\widehat{\theta}$ for the covariates $(D, Z_1, ..., Z_n)$. Households composition and socio-economic characteristics play an important role. The mother's share of resource increases the lower is the number of children, while there does not seem to be any significant correlation with the share of resources devoted to the father. In line with Dunbar et al. (2013) and Calvi (2016), the fraction of female children is positively related to η_m: if all children are girls, the mother's resource share is 3.6 percentage points larger. This result 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

[22]Note that if in step 1 the unobserved treatment was specified as linear in the set of covariates Z, and Z coincided with X, our IV strategy in step 2 would simplify to a least square regression of Y on $\widehat{R}$ and X, provided that standard errors are appropriately adjusted.

children, due to son preference.

The estimates also indicate that the wealthier the household the higher is the fraction of household's total expenditure devoted to the father, everything else equal. In addition, being part of Scheduled Caste, Scheduled Tribes, and other disadvantaged social classes is associated with higher bargaining power of the mother relative to the father, while the opposite holds true for living in rural areas. The fraction of resources devoted to the mother are higher in the North-East states and lower for the father, which is consistent with the presence of a number of matrilineal societies and cultures (Khasi and Garo societies, for example). In contrast, North Indian women seem to have a much lower bargaining power relative to their male counterparts. Finally, a higher likelihood to be exposed to the Hindu Succession Act amendments - women's eligibility to the reforms - is associated with a stronger women's bargaining position within the household. These results confirm existing reduced-form findings in on the effects of HSA amendments on self-reported measures of women's autonomy and bargaining power (e.g., Roy (2008) and Heath and Tan (2014)).

Using the estimates $\widehat{\theta}$ in Table 2.3, we predict the share of resources devoted to the mother, the father and the children, as in equation (2.15). We obtain two sets of predictions: one *in-sample*, using the NSS data that we use for the estimation of the Engel curve system, and one *out-of-sample*, using the NFHS-3 data, which contains all the relevant covariates $(D, Z_1, ..., Z_n)$, together with individual's health outcomes that we will use in step 2. Table 2.4 contains descriptive statistics for the predicted resource shares obtained in the two samples and for the implied unobserved treatment, $\widehat{R}$. These summary statistics take into account the empirical distributions of the covariates $(D, Z_1, ..., Z_n)$, as they average over all the values of demographic factors observed in the population. That the minima and maxima of estimated resource shares do not fall outside the zero to one range for any person in any household in the two sample is reassuring. Moreover, the standard deviations of resource shares are larger for men than for women, suggesting that the covariates induce more variation for fathers than for mothers. In

Table 2.3: Estimated DLP parameters, $\widehat{\theta}^m$ and $\widehat{\theta}^f$

	Mother's Resource Share (η_m)	Father's Resource Share (η_f)
	(1)	(2)
$\mathbb{I}$(1 child)	0.0396*	0.0232
	(0.0239)	(0.0364)
$\mathbb{I}$(2 children)	0.0124	0.00833
	(0.0207)	(0.0335)
$\mathbb{I}$(3 children)	0.0183	0.00534
	(0.0211)	(0.0343)
Fraction of Female Children	0.0358**	-0.0164
	(0.0166)	(0.0195)
$\mathbb{I}$(HSA Eligible)	0.0295***	-0.0156
	(0.0102)	(0.0125)
$\mathbb{I}$(Hindu, Buddhist, Jain, Sikh)	-0.0194	0.0469**
	(0.0170)	(0.0216)
$\mathbb{I}$(Sch. Caste, Sch. Tribe, OBC)	0.0545***	-0.0516***
	(0.0181)	(0.0200)
Wealth Index	-0.0539	0.155***
	(0.0435)	(0.0591)
$\mathbb{I}$(Woman's Higher Education)	0.0216	-0.00904
	(0.0172)	(0.0227)
$\mathbb{I}$(Man's Higher Education)	0.0163	0.00611
	(0.0141)	(0.0209)
$\mathbb{I}$(Rural)	-0.0386**	-0.0133
	(0.0154)	(0.0213)
Gender Age Gap	0.0312	-0.0687
	(0.134)	(0.204)
Woman's Age	-0.420	-0.140
	(1.874)	(2.156)
Gender Age Gap^2	-0.263	-0.138
	(0.405)	(0.580)
Woman's Age^2	0.909	-0.593
	(4.626)	(5.256)
Gender Age Gap^3	0.891	-0.226
	(1.459)	(2.180)
Woman's Age^3	-0.501	0.953
	(3.673)	(4.085)
Children's Avg. Age	-0.269	0.423*
	(0.202)	(0.257)
$\mathbb{I}$(North)	-0.0913***	0.0638*
	(0.0323)	(0.0365)
$\mathbb{I}$(East)	0.0119	-0.0194
	(0.0374)	(0.0403)
$\mathbb{I}$(North-East)	0.0562	-0.208***
	(0.0454)	(0.0436)
$\mathbb{I}$(South)	0.0119	-0.0444
	(0.0330)	(0.0354)
Constant	0.352	0.476*
	(0.250)	(0.288)
Observations	15,164	15,164

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

Table 2.4: Estimated Resource Shares and Woman's Power ($\widehat{R}$)

	Mean	St. Dev.	Min.	Max.
Panel A: NSS Sample (Obs:15,164)				
Woman's Resource Share ($\widehat{\eta}_m$)	0.2993	0.0743	0.0671	0.5368
Man's Resource Share ($\widehat{\eta}_f$)	0.4484	0.1055	0.1122	0.7461
Children's Resource Share ($\widehat{\eta}_c$)	0.2522	0.0762	0.0000	0.5105
Each Child's Resource Share ($\widehat{\eta}_c/s$)	0.1375	0.0641	0.0000	0.4484
Woman's Power ($\widehat{R}$)	0.4052	0.1136	0.930	0.8271
Panel B: NFHS Sample (Obs: 20,131)				
Woman's Resource Share ($\widehat{\eta}_m$)	0.2860	0.0755	0.0829	0.4950
Man's Resource Share ($\widehat{\eta}_f$)	0.4531	0.1263	0.1060	0.7313
Children's Resource Share ($\widehat{\eta}_c$)	0.2609	0.0847	0.0568	0.5131
Each Child's Resource Share ($\widehat{\eta}_c/s$)	0.1372	0.0712	0.0374	0.4684
Woman's Power ($\widehat{R}$)	0.3960	0.1301	0.1102	0.8005

Summary statistics of the estimated resource shares for both samples used.

both samples, the resource share for mothers is lower than that for fathers: on average, mothers receive about 30 percent of household resources, while close to half of the family expenditure is allocated to the fathers. On average only one fourth of resources is allocated to children, and the per child amount of resources declines as the number of children increases.[23] The mean estimated unobserved treatment is equal to about 40 percent in both prediction samples, which indicates that on average the mother gets 40 percent of the total resources allocated to parents (the decision makers).

How does our structurally-motivated measure of bargaining power compare with other more standard proxies? As discussed in details in section 2.2.1, the larger is $\widehat{R}$, the greater is the weight of the mother's preferences in the household program, i.e., the higher is her say in the household decision process. We confirm this interpretation by analyzing how $\widehat{R}$ relates to self-reported measures of women's participation in household's decisions. From the NFHS data, we assess a woman's decision-making ability using her answers to the question: "Who makes decisions about [X] in your household?". Using principal

[23]These numbers are in line with the results obtained in Dunbar et al. (2013) using data from the Malawi Integrated Household Survey (IHS2). They look at married couples with up to four children.

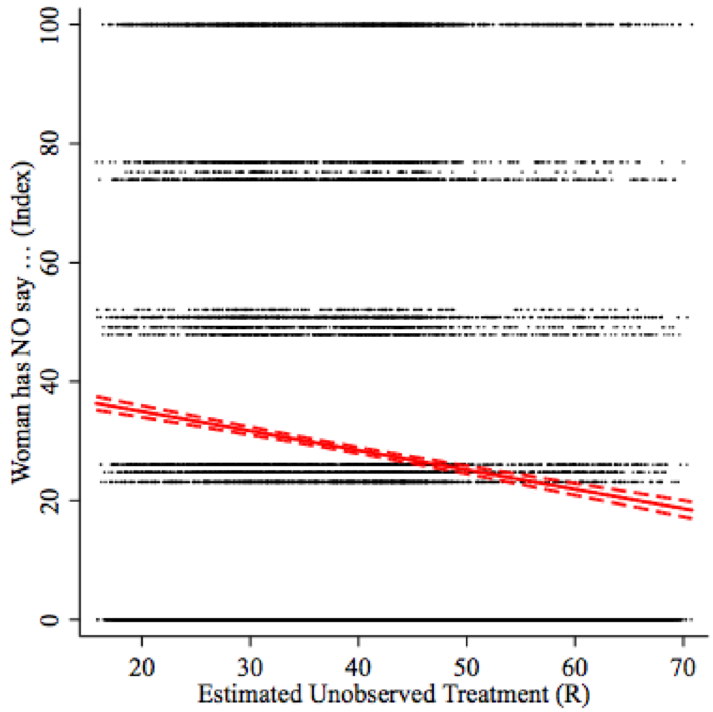


Figure 2.1: $\widehat{R}$ and "NO say" Index

component analysis, we create an index combining answers related to the decisions about the woman's own health care, large household purchases, visits to family or relatives, and purchases for daily needs. Figure 2.1 displays the clear negative relationship between our "NO say" Index - rescaled to range from 0 to 100 - and the unobserved treatment, corroborating that the larger is $\widehat{R}$, the higher is her say in the household decision process. These findings are confirmed in Table B.1 in the Appendix, which shows the correlation coefficients between the respondent's answers, "NO say" Index and $\widehat{R}$.

2.4.2 Step 2: Causal Inference

We adopt an instrumental variable approach to ensure the causal interpretation of the effect of the unobserved treatment $\widehat{R}$ on the health outcomes of household members. In this section, we report the

Table 2.5: First Stage, OLS Estimates

	Woman's Power ($\ln\widehat{R}$)					
	Women		Men		Children	
	(1)	(2)	(3)	(4)	(5)	(6)
$\mathbb{I}$(HSA Eligible)	0.167^{***}	0.0211^{***}	0.212^{***}	0.155^{***}	0.174^{***}	0.0187^{***}
	(0.00702)	(0.00646)	(0.00860)	(0.00886)	(0.00728)	(0.00662)
Observations	14,340	14,340	7,258	7,258	35,487	35,487
State Fixed Effects	No	Yes	No	Yes	No	Yes
Cohort Fixed Effects	No	Yes	No	Yes	No	Yes
State-Religion Fixed Effects	No	Yes	No	Yes	No	Yes
Cohort-Religion Fixed Effects	No	Yes	No	Yes	No	Yes
State Linear Trends	No	Yes	No	Yes	No	Yes

$^*p < 0.10$, $^{**}p < 0.05$, $^{***}p < 0.01$. Estimates are obtained using the NFHS-3 data. All specifications include the number of children below 15, the fraction of female children, a household wealth index, indicator variables for owning a BPL card, for living in rural areas, and for being part of Scheduled Castes, Scheduled Tribes or Other Backward Classes. Specifications in Columns 2-4-6 include religion, state, cohort (woman), state-religion, and cohort-religion fixed effects, and state time trends. Specifications in columns 1-2 and 5-6 also include an indicator variable equal to 1 if the woman worked in the past 12 months and the woman's years of education; in columns 3-4, they include an indicator variable equal to 1 if the man worked in the past 12 months and the man's years of education. Bootstrap standard errors in parentheses (2,500 repetitions). Standard errors clustered at the primary sampling unit (village) level.

results obtained with a 2SLS approach, while the least squares estimates are in tables in the Appendix.

Table 2.5 contains the first stage estimation results for women, men, and children models separately. Household level and individual level covariates are included in columns (2)-(4)-(6). Clearly, the sample sizes and the sets of covariates are slightly different in the three models: all specifications include the number of children below 15, the fraction of female children, a household wealth index, indicator variables for owning a BPL card, for living in rural areas, and for being part of Scheduled Castes, Scheduled Tribes or Other Backward Classes; columns (1)-(2) and (5)-(6) also include an indicator variable equal to 1 if the woman worked in the past 12 months and the woman's years of education, while columns (3)-(4) include an indicator variable equal to 1 if the man worked in the past 12 months and the man's years of education. Moreover, the specifications in columns (5)-(6) also include the child's age, sex, and a dummy variable for his/her birth order (up to four). To ensure the exclusion restriction to hold, specifications in columns (3)-(5)-(7) include also state, cohort (mother's), state-religion, cohort-religion, and state-specific linear trends. Even conditioning on these several sources of unobserved heterogeneity, mothers' eligibility to the Hindu Succession Act amendments is positively and significantly correlated with their bargaining position relative to the fathers, providing some evidence of the fact that weak instrumentation is not an issue in our setting.[24] In all samples, the unobserved treatment $\widehat{R}$ is significantly higher for women who are exposed to these amendments.

The 2SLS results indicate that the relative decision power of parents matters, however not for everyone. When we focus on women's health outcomes (Table 2.6), we find positive and significant effects of women's decision power relative to their male counterparts. A one percent increase in $\widehat{R}$ determines significant increases in women's weight and in their body mass index (BMI) - by about 1.2 kilograms and 0.5 units respectively - and a decrease in their probability of being anemic by 0.04. Moreover, our estimates indicate that more empowered mothers face higher birth intervals (or birth spacing). The adverse health consequences of a short spacing between births are well documented: as the mother is not allowed sufficient time to restore the adequate supply of nutrients, short spacing is found to affect maternal health and mortality both directly and indirectly through increased risk of anemia (e.g., Conde-Agudelo and Belizán (2000) and Milazzo (2014)). In contrast, we do not find any significant effect on men's health outcomes (Table 2.7), suggesting intra-household health investment is not necessarily a zero-sum game and, thus, that women's empowerment inside the family does not result in a decline in men's outcomes. Moreover, that women's bargaining position has a much more important effect on their health relative to their spouses might indicate that the household decision makers disagree more strongly (and more often) on investment on women's rather than men's health.

Table 2.8 contains the 2SLS estimation results for the impact of mothers' decision power on children's health. Our estimates indicate that women's bargaining position does not significantly affect children's anthropometric indicators, such as weight and height. When we consider the impact on mother's decision power relative to the father's on the likelihood of children being vaccinated, however, the conclusion

[24] This is also confirmed by the Wald tests for joint significance of the instruments, both included and excluded.

Table 2.6: Women's Health, 2SLS Estimates

	Weight (kg)	Body Mass Index	Pr(BMI≤18.5)	Pr(Anemic)	Avg. Birth Space (months)
	(1)	(2)	(3)	(4)	(5)
Woman's Power ($\ln \widehat{R}$)	121.1*	51.87*	-3.744	-4.874**	321.5**
	(68.48)	(27.03)	(2.613)	(2.453)	(151.6)
Observations	13,838	13,824	13,824	12,952	12,028
Mean Dependent Variable	48.24	20.89	0.30	0.16	35.11

*$p < 0.10$, **$p < 0.05$, ***$p < 0.01$. Estimates are obtained using the NFHS-3 data. All specifications include state, cohort (woman), state-religion, and cohort-religion fixed effects, and state time trends. The sample includes married women of age 15 to 49 included in nuclear households with up to 4 children. All specifications include a religion dummy, equal to 1 if a woman is Hindu, Buddhist, Sikh or Jain, number of children below 15, the fraction of female children, woman's years of education, a household wealth index, indicator variables for owning a BPL card, for living in rural areas, and for being part of Scheduled Castes, Scheduled Tribes or Other Backward Classes, for having worked in the past 12 months. Anemia includes mild and moderate anemia. Bootstrap standard errors in parentheses (2,500 repetitions). Standard errors clustered at the primary sampling unit (village) level.

Table 2.7: Men's Health, 2SLS Estimates

	Weight (kg)	Body Mass Index	Pr(BMI≤18.5)	Pr(Anemic)
	(1)	(2)	(3)	(4)
Woman's Power ($\ln \widehat{R}$)	70.78	33.55	-0.436	-1.268
	(112.9)	(37.00)	(4.188)	(3.412)
Observations	6,886	6,884	6,884	6,230
Mean Dependent Variable	56.85	21.13	0.24	0.10

*$p < 0.10$, **$p < 0.05$, ***$p < 0.01$. Estimates are obtained using the NFHS-3 data. All specifications include state, cohort (woman), state-religion, and cohort-religion fixed effects, and state time trends. The sample includes married men of age 15 to 54 included in nuclear households with up to 4 children. All specifications include a religion dummy, equal to 1 if a woman is Hindu, Buddhist, Sikh or Jain, number of children below 15, the fraction of female children, man's age and years of education, a household wealth index, indicator variables for owning a BPL card, for living in rural areas, and for being part of Scheduled Castes, Scheduled Tribes or Other Backward Classes, for having worked in the past 12 months. Anemia includes mild and moderate anemia. Bootstrap standard errors in parentheses (2,500 repetitions). Standard errors clustered at the primary sampling unit (village) level.

Table 2.8: Children's Health, 2SLS Estimates

	Weight (kg)	Height (cm)	Pr(Any Vaccine)	Pr(BCG)	Pr(DPT)	Pr(Polio)
	(1)	(2)	(3)	(4)	(5)	(6)
Woman's Power ($\ln \widehat{R}$)	-14.35	-109.2	5.428*	6.032	7.671	6.551**
	(40.01)	(149.6)	(3.290)	(4.124)	(5.273)	(2.876)
Observations	8,761	8,739	9,549	9,511	9,459	9,538
Mean Dependent Variable	10.67	83.94	0.91	0.78	0.75	0.91

*$p < 0.10$, **$p < 0.05$, ***$p < 0.01$. Estimates are obtained using the NFHS-3 data. All specifications include state, cohort (woman), state-religion, and cohort-religion fixed effects, and state time trends. The sample includes children 0 to 5 in nuclear households with up to 4 children. All specifications include a religion dummy, equal to 1 if a woman is Hindu, Buddhist, Sikh or Jain, number of children below 15, the fraction of female children, mother's years of education, a household wealth index, indicator variables for owning a BPL card, for living in rural areas, and for being part of Scheduled Castes, Scheduled Tribes or Other Backward Classes, for having a mother who worked in the past 12 months. Bootstrap standard errors in parentheses (2,000 repetitions). Standard errors clustered at the primary sampling unit (village) level.

is different. We find that a one percent increase in the unobserved treatment $\widehat{R}$ leads to a 0.05 increase in the probability of children being vaccinated. This result seem to be driven by the likelihood of children being vaccinated against polio. Due to the larger standard errors, we do not find a significant effect on the probability of having received the BCG (against tuberculosis) or the DPT (against diphtheria, pertussis, and tetanus) vaccines.

2.5 Conclusion

We apply a novel two-step approach to study the effect of intra-household women's empowerment on the health status of family members in India. In a first step, we rely on the structural model to recover a measure of the bargaining power of each member, which is unobservable in practice. In the second step, we use this structurally-motivated proxy of power and study its causal effect on household member's health status in a reduced-form framework.

To motivate our empirical analysis, we augment the standard Dunbar et al. (2013) (DLP) collective model to include health investment decisions on each household's member. We show that in this framework the relative bargaining position of the decision makers matters, especially in situations of disagreement. Results of the empirical application indicate that the relative decision power of parents matters, however not for everyone. We find that the effect of women's empowerment on their own health outcomes is positive and significant. An increase in their bargaining power (relative to their spouse) improves both their weight and body mass index, decreases their probability of being anemic, and increases birth spacing. When we turn to estimate the effect of women's relative decision power on the health outcomes of their spouse and children, we find, respectively, no significant effect and no significant effect on anthropometric indicators such as weight and height of the children, but a significant increase in the likelihood of children being vaccinated.

Aside from the quantitative findings of the application, our analysis highlights the advantages of combining *structural* and *causal* features in conducting empirical analysis. By using a household model to recover a structural measure of bargaining power, we are able to answer a wider range of important research and policy questions and to provide a more accurate assessment of the effect of women's empowerment on family members' outcomes.

Chapter 3

Long-Term Effects of Access to Health Care: Medical Missions in Colonial India

3.1 Introduction

According to the most recent World Health Organization (WHO) estimates, India is ranked 150^{th} in the world in terms of life expectancy and one out of three Indian adults is underweight. These country-level figures hide a considerable regional variation in health conditions: while South Indian states are doing relatively well in terms of health performance, the same cannot be said for some of the most populous states, such as Uttar Pradesh and Bihar, whose health indicators are worse than those of many Sub-Saharan African countries.[1] Moreover, recent evidence suggests that differences in health performance persist even at the sub-state level and across areas within the same district.[2]

This paper studies one of the possible sources of this geographical heterogeneity in health conditions by examining the long-term effect of access to historical health facilities on current individual health

[1] See Pal and Ghosh (2007).
[2] See Srinivisan (2010). India comprises 28 states and 7 union territories. These are divided into 593 districts (Census 2001).

outcomes. We exploit information about the diffusion, location and activities of Protestant medical missionaries in colonial India.[3] In the late nineteenth century, missionaries became aware of the role that medicine could have in facilitating evangelization among non-Christian populations and the number of doctors and medical personnel sent to India by Protestant societies increased dramatically.[4]

We investigate whether the proximity to a Protestant medical mission in colonial India predicts individual health outcomes today. To this aim, we construct a novel and unique dataset that combines contemporary individual-level data with historical information on Protestant medical missions.[5] Consistent with the previous literature, we measure health status with individuals' anthropometric indicators.[6] We use geocoding tools to calculate the distance between the location of individuals today and that of Protestant health facilities in nineteenth century India. We find that the proximity to a Protestant medical mission is positively and significantly associated with current individuals' body mass index (BMI) and other health indicators.

To give a causal interpretation to our estimates, we account for two types of selection bias. First, historical and geographic characteristics may have driven the missionaries' location decisions. Second, Protestant missions may have endogenously selected themselves into building a health facility and into providing health care. We show that these potential biases are not the the driving force behind our results. To correct for possible systematic differences between regions with and without Protestant missions, we restrict our sample of analysis to individuals living within a certain distance from a Protestant mission. By limiting the analysis to more concentrated areas we minimize the risk of our findings being driven by unobserved geographical differences. In addition, we control for the minimum distance from

[3] Our decision of focusing on Protestant medical missions as opposed to Catholic missions is mainly data driven. To the best of our knowledge, there is no document that systematically records medical activities of Catholic missions. Moreover, India was the preferred destination of the Protestant missionary venture due to its British element. Protestant were especially committed to operate among non-elites and socially disadvantaged groups, while Catholic missionaries, at least until the Second Vatican Council in 1965, strategically decided to focus their efforts on the elite groups of society (Mathew (1999)).

[4] See Mathew (1999) and Fitzgerald (2001).

[5] Throughout the paper we use the terms medical missions, medical facilities, medical settlements, health facilities, and hospitals, interchangeably.

[6] See Schultz (2002, 2003).

generic Protestant and Catholic missions. These variables should capture factors that also determined the location decisions of Protestant medical missionaries. If our results were purely driven by unobserved heterogeneity across locations, we would find a significant correlation between these variables and individuals' current health outcomes. We address the second type of selection bias by developing an instrumental variable approach in the spirit of Cagé and Rueda (2014). We exploit the fact that Protestant missionary societies differ in their inclination to undertake medical activities. We use these differences to compute each society's share of medical missions in all regions of the world *outside* of India and then use those shares to construct our instrument. We show that our results are robust to the instrumental variable approach. We find that halving the distance from the location of a Protestant medical facility increases current individuals' BMI by 0.4.

We analyze three potential transmission channels: persistence of infrastructure, improvements in health potential (health status of previous generations and nutrition), and improvements in health habits (hygiene and health awareness). First, consistent with the previous literature, we show that proximity to current health facilities does not play a role.[7] Second, we find that the long-term effect of historical medical missions is partially, but not entirely, driven by improvements in health potential. Finally, we provide evidence of the fact that proximity to a health care facility in the past may have affected individual habits regarding hygiene, preventive care and health awareness, which have been bequeathed over time and have influenced health outcomes of later generations.

This paper represents the first attempt to investigate how the exposure to historical medical missions may have determined advancements in current individual health outcomes. With this contribution, we add to the growing empirical literature that looks at historical institutions as important determinants of current outcomes and investigates path persistence in developing countries.[8] More specifically, we contribute to previous literature that recognizes the importance of the activities of religious organizations as

[7] See e.g. Chaudhury and Hammer (2004) and Banerjee et al. (2004).

[8] See e.g. Acemoglu et al. (2001), Glaeser and Shleifer (2002), Nunn (2009), and Alesina et al. (2013).

possible fundamental sources of economic and social outcomes.[9] Previous works have identified Protestantism as an important determinant of advancements in human capital by focusing on its contribution to literacy and education.[10]

Recent works study how religious institutions and affiliation may help explaining differences in health outcomes. Almond and Mazumder (2011) analyze births to Arab parents in Michigan and find that prenatal exposure to Ramadan results in lower birth weight. While Bhalotra et al. (2010) show that Muslim children in India have a significantly higher probability of survival in infancy than Hindu children, Brainerd and Menon (2013) demonstrate that this advantage does not persist beyond infancy. Finally, Menon and McQueeney (2015) find that Christian infants in India have higher health outcomes compared to infants of other religious identities. We differ from this strand of literature in that we focus on the long-term effect of religious institutions.

Important policy implications may be drawn from our analysis. On one hand, understanding historical determinants of regional variation in health conditions may provide useful guidance in addressing issues related to health inequality in modern India. On the other hand, in light of the existence of long-run positive spillovers, the expansion of healthcare access becomes an even more critical priority, as it provides benefits for both current and future generations. Health can influence not only individuals' early survival and life expectancy, but also cognitive performance, schooling decisions, and productivity as adults. Understanding accumulation of health capital in developing countries means understanding one of the foremost determinants of economic development.[11]

The rest of the paper is organized as follows. Section 3.2 provides a brief historical description of the Protestant medical enterprise in colonial India. Section 3.3 describes our novel dataset, and discusses the use of anthropometric indicators as a measure of health. Section 3.4 presents our empirical strategy.

[9] See e.g. Barro and McCleary (2003, 2005).

[10] See e.g. Weber (1930), Becker and Woessmann (2009), Gallego and Woodberry (2009), Bai and Kung (2014), Cagé and Rueda (2014), and Mantovanelli (2014).

[11] See Schultz (2009).

Section 3.5 discusses the estimation results, while section 3.6 checks the robustness of our findings. Section 3.7 investigates potential transmission channels. Finally, section 3.8 concludes.

3.2 Medical Missions in Colonial India

The period between 1800-1914 represents an era of great missionary expansion led by the Protestants of Britain and the United States. The intensity of the missionary evangelization was such that that period is known as the *great century* of missions. India was the preferred destination of the Protestant missionary venture due to its British element. In the mid-nineteenth century more than a quarter of all the Protestant missionaries were stationed in India. On the eve of the first World War 5,200 Protestant missionaries (of whom 2,500 British and 1,800 American) were posted in the Indian sub-continent. At that time, only China rivaled India in terms of Protestant presence.[12]

At first, missionary societies showed very little interest in medicine. Protestant organizations did not consider the establishment of hospitals and the provision of medical care as crucial for the scope of their missions. The missionary movement was convinced that the biblical command of evangelizing every nation was limited only to preaching and teaching the Gospel.[13]

Evangelizing strategies started to change from the 1860s onwards in correspondence with the growing awareness of the deficiencies of the orthodox missionary methods. Anxiety over the lack of tangible results provided a powerful trigger for a change in official mission policies and practices. The belief was that the missionary sphere could rightfully include a range of activities beyond those of preaching and teaching:

> *"I feel convinced that medical missions are amongst the means best calculated* [...] *to give a fresh impetus to the cause of the Gospel, to help lift the chariot out of the rut in which sometimes it*

[12]For more details refer to Richter (1908).

[13]The constitution of the Baptist Missionary Society, for example, stated that the propagation of Christianity consisted of "the preaching of the Gospel, the translation of the Holy Scriptures and the establishment of Schools."

> *seems for a moment set fast and its progress retarded."*
>
> (Rev. Thomas V. French, Bishop of Lahore, quoted by Bishop McDougal in a speech at the Oxford Missionary Conference of 1877)

Medical missions were finally recognized as valuable auxiliaries to the work of propagating Christianity,

> *"The work of the doctor is to open the door that the evangelist may enter in."*
>
> (Medical Missionary Journal [1874], 9, 58-59)

In contrast to official colonial medicine, Protestant mission medicine sought to place itself within non-European social and institutional milieus and reached out to embrace all classes of the *native* society, regardless of caste or gender.[14]

In light of the recognition of the importance of medicine as an effective weapon in the evangelization process, the last decades of the nineteenth century witnessed an unprecedented growth of the medical mission enterprise. The number of doctors and medical agents sent to India by the Protestant societies rose from 28 in 1870 to 335 in 1912. This was accompanied by a rapid increase in the number of mission hospitals and dispensaries. By 1912, approximately three million patients were treated in Protestant medical facilities annually. The quality of medical missionary medicine increased as well in the closing years of the nineteenth century. At the Bombay Missionary Conference of 1892-93 a resolution was passed stating that medical missionaries in India should invariably possess a medical degree or diploma sufficient to qualify as a licensed medical practitioner in the West. Fitzgerald (2001) indicates that, despite initial fears, it was surgery that determined acceptance of Western medicine among the Indian people and their greater willingness to use the resources of medical missions. Missionary reports document that over time patients became increasingly willing to enter a mission hospital for periods of in-patient care and were generally applying for medical assistance at an earlier stage in their illness.[15] In addition to foreign staff, medical missions trained and employed Indians working as assistants (around

[14] See Hardiman (2008).
[15] See Fitzgerald (2001).

700 in 1912), creating another vital contribution to the local communities.

3.3 Data

3.3.1 Data Sources

One of the main contributions of this paper is the creation of a unique and fully geocoded dataset that combines historical, geographic, and contemporary information.

Contemporary Data. Current individual data are from the *2003 India World Health Survey* (WHS). The World Health Surveys (WHS) were launched by the World Health Organization to monitor and evaluate critical health outcomes and health systems through cross-country comparable household surveys. The WHS were implemented between 2002 and 2004 in 70 countries selected to represent all regions of the world. The 2003 India WHS covers individuals and households residing in seven states: Assam, Karnataka, Madhya Pradesh, Maharashtra, Rajasthan, Uttar Pradesh and West Bengal. While the survey is centered around health-specific issues, it also provides demographic information of the respondents and data about household composition and educational level. Crucially for the purpose of our research, the survey discloses the geographic coordinates of where the respondents live and some anthropometric indicators.[16] Figure 3.1 shows the locations of survey respondents.

To investigate possible transmission channels of the long-term effect of access to Protestant medical missions, we use data from the Village Directory of the 2001 Census of India and the 2007-2008 District Level Household & Facility Survey (DLHS-3). We obtain information about the location of current health facilities from the former, and data about health and preventive care practices from the latter. The 2001 Village Directory contains detailed information about population and infrastructure/amenities -

[16] The decision of not using the Demographic Health Survey, known as National Family Health Survey for India, is data driven, as geographic coordinates are not available for this country.

including hospitals and several types of health facilities - of all villages and towns in India. The DLHS-3, one of the largest demographic and health surveys ever implemented in India, allows us to assess the utilization of services provided by health care facilities and people's health practices.[17] The survey does not provide any village level identifier, but it contains information about the village population as per Census 2001. We follow Banerjee and Sachdeva (2014) and use the village population to match the DLHS-3 data set with the Census 2001 data. The matching provides information on census village codes and villages names for all of the DLHS-3 sample.[18] We successfully match and geocode 21,151 villages.[19] To the best of our knowledge, this paper is the first one exploiting DLHS-3 information to construct a fully geocoded version of the survey.[20]

Historical Data. Information about the location of Protestant missions in colonial India comes from the *Statistical Atlas of Christian Missions*, published in 1910 in correspondence with the World Missionary Conference held in Edinburgh (Scotland). We geocode the maps of India contained in the Atlas to identify the exact location of Protestant missions as of 1908. Figure C.1 in the Appendix shows an example of these maps. To identify missionary medical infrastructures we rely on the *Centennial Survey of Foreign Missions* (Dennis (1902)) which provides a worldwide list of all the Protestant missions with medical facilities and a record of their basic characteristics. For example, we know the designation of the facility (hospital or dispensary), the foundation date, the name of the supporting Protestant society, and the number of patients and surgical cases. A reproduction of one page of the survey is provided in Figure C.2 in the Appendix. We digitize and geocode this information for the Indian subcontinent. We use the same data source to identify Protestant missions with educational institutions. Our final sample

[17]It consists of one village level, one household level, and two individual level questionnaires (ever married women of age 15-49 and unmarried women of age 15-24). While it covers all India and the sample size is much larger than in the WHS, the DLHS-3 does not contain anthropometric indicators.

[18]The matching is done generating village identifiers using state, district, and village population. We drop duplicates from the census data and all DLHS villages that have the same population at the tehsil/taluk (sub-district administrative units) level.

[19]The geocoding is based on Google coordinates for a string containing the names of *state*, *district*, and *village*.

[20]Banerjee and Sachdeva (2014) match the DLHS-3 data set with the Census 2001 data to analyze the impact of a nationwide road construction program on the usage, provision, and awareness of preventive health care. Their dataset does not include geographic coordinates. We are grateful to Ashish Sachdeva for suggesting this approach.

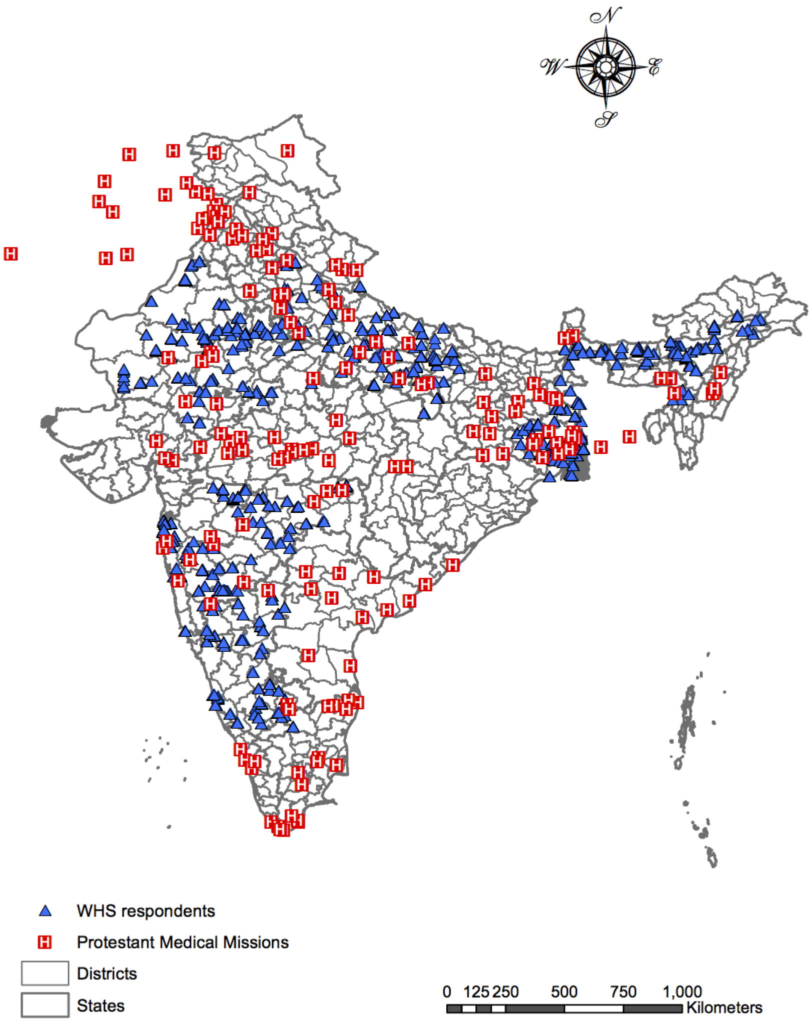


Figure 3.1: Protestant Medical Missions and WHS Respondents

includes 1,069 Protestant missions. Of these, 183 are equipped with either a hospital or dispensary. Figure 3.1 shows the locations of the Protestant medical missions.

Beginning in the middle of the nineteenth century, the Government of India started a great transportation infrastructure project aimed at developing a railroad network in the Indian sub-continent.

Railroads likely played a significant role in determining the missionaries' location decisions. Areas with a well-developed railways network were more likely to be targeted by missionaries as they were more easily accessible. To take this into account we digitize and geocode historical maps from the *Constable's Hand Atlas of India*. Figure C.3 in the Appendix reports a section of one of those maps showing the status of the Indian railways network as of 1891.

In our empirical analysis we also include information about the location of historical Catholic missions. We rely on maps contained in the *Atlas Hierarchicus*, an official document of the Vatican published in 1913 reporting the worldwide geographic distribution of Catholic missions as of 1911. Figure C.4 in the Appendix shows one of these maps.

Finally, we obtain additional historical information from the History Database of the Global Environment - HYDE 3.1.[21] This database provides rasters of gridded total (estimated) historical population (inhabitants/gridcell) and gridded cropland (km^2/gridcell) from 10,000 B.C. to 2005 A.D. at a 5 $\times$ 5 minutes resolution (about 85 km^2 around the equator). We extract the 1900 estimates for an area within a 5-kilometer radius from each WHS location.

Geographical Data. We control for geographic characteristics at the WHS location level. We use the *Global Gridded Population Database* from CIESIN to compute the population in year 2000 in an area within a 5-kilometer radius from each WHS location.[22] We also calculate the average elevation within the 5-kilometer buffer using the *SRTM 90m Digital Elevation Data* compiled by the CGIAR Consortium for Spatial Information.[23] Both population and elevation data are available at a 30 $\times$ 30 seconds resolution (about 0.8 km^2 around the equator). To account for the importance of access to water sources, we also include a variable measuring the number of water streams within the 5-kilometer buffer around the WHS locations. Data are from the *Digital Chart of the World*.[24]

[21] See Klein Goldewijk et al. (2010, 2011).
[22] http://www.ciesin.org/.
[23] http://srtm.csi.cgiar.org/.
[24] http://worldmap.harvard.edu/data/geonode:Digital_Chart_of_the_World.

3.3.2 Measuring Health

Measurement of health has recently evolved to rely on anthropometric indicators of physical development. Previous literature has used height, weight-for-height or body mass index to measure individual health.[25] In this paper we use body mass index (BMI) as our main measure of current individual health status.[26] We also consider height and a self-reported description of current health status as alternative measures of health.

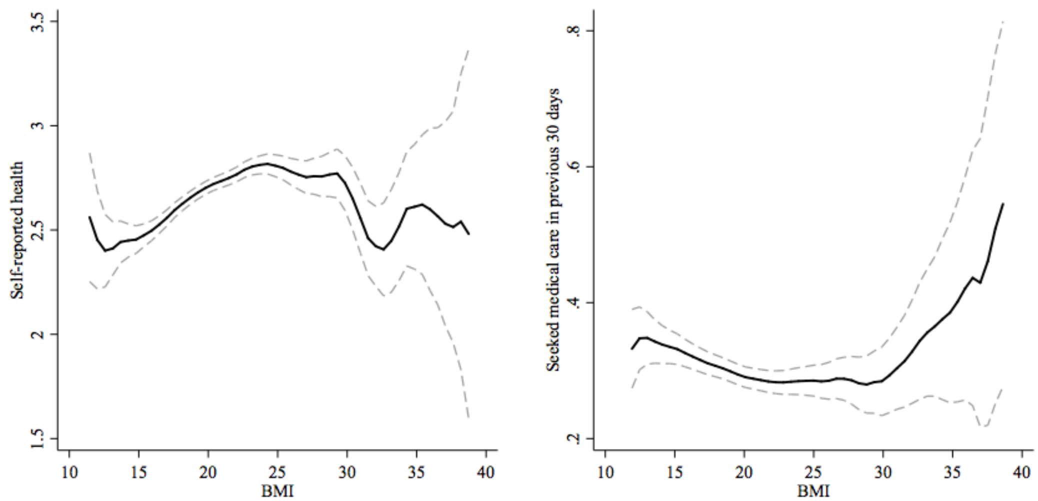
That the relationship between health status and BMI is nonlinear is a well known fact, both at the individual and at the aggregate level. Wailer (1984) clearly illustrates this non-linearity by showing a U-shaped relationship between relative risk of mortality and BMI using Norwegian data from the 1970s. Especially in low income countries, however, an increase in the caloric intake shifts the lower tail of the BMI distribution to the right. Fogel (1994, 2004) interprets this shift as an accumulation of the population's health human capital, which tends to be associated with both declines in mortality and increases in labor productivity.

Figure 3.2 shows the results of a non-parametric estimation of the relationship between health status and BMI. Panel A suggests the presence of a non-monotonic concave relationship between respondents' self-reported health status and BMI, while panel B indicates a convex relationship between the probability of seeking health care in the previous month and BMI.

In our main empirical analysis we focus on individuals aged 20 to 60 with BMI between 15, the cutoff for starvation, and 30, the obesity threshold. We therefore concentrate only on the monotonic part of

[25] See e.g. Wailer (1984), Fogel (1994), Schultz (2002), Schultz (2003), Weil (2005), Steckel (2008), Deaton (2008), and Schultz (2009).

[26] The BMI is defined as $\frac{weight(kg)}{height^2(m)} = \frac{weight(lb)}{height^2(in)} \times 703$. The WHO provides an international classification of adult underweight, overweight and obesity according to BMI and fixes 18.5 as the cut-off value between underweight and normal-range, 25 as the cut-off value between normal-range and overweight and 30 as the cut-off value between overweight and obese. A WHO expert consultation in 2004 (Shiwaku et al. (2004)) addresses the recent debate about interpretation of recommended body-mass index (BMI) cut-off points for determining overweight and obesity in Asian populations, and develops population-specific cut-off points for BMI. The current BMI cut-off values for Asian Indians are 18.5, 23, 25 and 30 for the thresholds between underweight and normal-range, between normal-range and overweight, between overweight and pre-obese and between pre-obese and obese, respectively.



(A) Self-reported health status (B) Health care needed in the past month

Figure 3.2: Non-parametric regression of health status on BMI

the BMI-health relationship. We exclude children, teenagers and elderly in order to avoid changes in individuals' BMI due to biological growth and aging. Figure 3.3 shows the distribution of BMI in the main estimation sample.[27] Consistent with the previous literature, BMI is log normally distributed.[28]

3.3.3 Descriptive Statistics

Table 3.1 contains some descriptive statistics for the variables used in our empirical analysis. Columns 1 to 3 present figures for the entire population in the survey, while columns 4 to 6 show figures for the baseline estimation sample (individuals of age 20 to 60, with BMI between 15 and 30). As there are no significant differences between the survey sample and our baseline estimation sample, we discuss here some descriptive statistics based on the latter.

The average BMI is 20.42, which falls in the normal range according to the WHO cutoffs, and the average height is 159 cm. Sixty-one percent of individuals rate their own health *good* or *very good*,

[27]Bandwidth: 0.75; Epanechnikov kernel function.
[28]See Burmaster and Crouch (1997) and Hjelmborg et al. (2008).

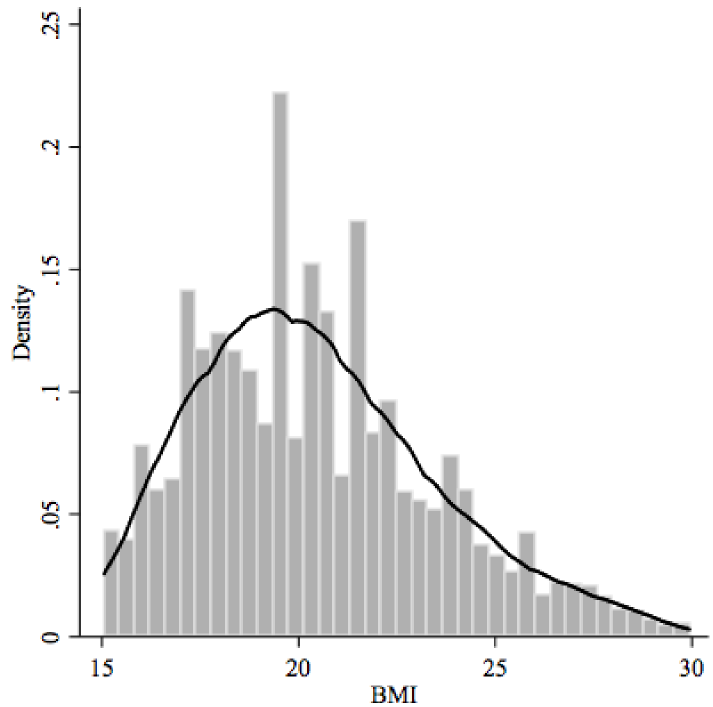


Figure 3.3: Distribution of BMI in the estimation sample

while only 13 percent rate their health *bad* or *very bad*.[29] About half of the sample consists of females and individuals who completed primary school; the average age is 37 years and households have on average six members. The wealth quartiles are based on a wealth index computed as the first principal component that combines information about a set of household assets.[30]

The mean distance from a Protestant medical mission is 81 kilometers, and the mean distance from a generic Protestant mission - with or without a medical facility - is equal to 37 kilometers. On average, respondents live 65 kilometers away from a Catholic mission, while the distance from the location of Protestant missions with an educational institution is larger. Data on population in year 2000 are available at at the 30" × 30" grid cell. For year 1900, they are available at the 5' × 5' resolution, which

[29]Self-reported health status is an ordinal variable (from 0, "very bad", to 4, "very good") determined by the response to the survey question "How would you rate your health today?".

[30]Number of rooms in dwelling, refrigerator, scooter, radio, and bike.

explains why the average population in 1900 is larger than that in year 2000.

The location of a Protestant mission is likely not random. A number of geographic and historical factors may have driven the settlement decisions of missionaries in colonial India, such as access to clean water, population density, altitude, accessibility by railways, land crop suitability, and the proximity to a Catholic mission. Previous literature shows that Protestant missionaries choose to locate in geographically favorable and more accessible regions.[31]

In Table 3.2, we investigate how geographic and historical controls vary as we restrict the sample to areas closer to a Protestant mission or to a Protestant medical mission. Columns 1-2 (3-4) (5-6) report means and standard deviations of historical and geographic controls when the sample is restricted to observations below the 25^{th} (50^{th}) (75^{th}) percentile of minimum distance from a mission (Panel A) or from a medical mission (Panel B).[32] Historical population in 5-kilometer radius increases as we restrict the sample to nearby areas, suggesting that missionaries settled in more densely populated areas. Moreover, areas close to missionary settlements have better access to colonial railways and water sources. The distance from a Catholic mission increases as farther away areas are included in the sample. On one hand, this suggests that the factors driving location decisions of Protestant missionaries are similar to those driving location decisions of Catholic missionaries. On the other hand, it indicates that the presence of a Catholic mission did not act as a deterrent to the settlement of Protestant missionaries.

[31] See e.g. Johnson (1967), Nunn (2010), and Cagé and Rueda (2014).

[32] The mission distance percentiles are 14 kilometers (25^{th}), 29 kilometers (50^{th}) and 49 kilometers (75^{th}); the medical mission distance percentiles are 28 kilometers (25^{th}), 75 kilometers (50^{th}) and 121 kilometers (75^{th}).

Table 3.1: Descriptive Statistics

	All sample			Estimation Sample		
	Obs. (1)	Mean (2)	St. Dev. (3)	Obs. (4)	Mean (5)	St. Dev. (6)
Body Mass Index (BMI)	9,305	20.59	6.16	7046	20.42	3.02
Height (cm)	9,305	158.73	11.35	7046	159.08	10.10
$\mathbb{1}$(Self-reported health = Very bad)	9,566	0.02	0.14	7046	0.01	0.12
$\mathbb{1}$(Self-reported health = Bad)	9,566	0.13	0.34	7046	0.12	0.32
$\mathbb{1}$(Self-reported health = Moderate)	9,566	0.26	0.44	7046	0.26	0.44
$\mathbb{1}$(Self-reported health = Good)	9,566	0.36	0.48	7046	0.38	0.49
$\mathbb{1}$(Self-reported health = Very good)	9,566	0.22	0.41	7046	0.23	0.42
$\mathbb{1}$(Female)	9,627	0.51	0.50	7046	0.52	0.50
$\mathbb{1}$(Married)	9,624	0.77	0.42	7046	0.82	0.38
$\mathbb{1}$(Major ethnic group)	9,625	0.81	0.39	7046	0.81	0.39
$\mathbb{1}$(Primary school completed)	9,627	0.52	0.50	7046	0.53	0.50
$\mathbb{1}$(Wealth 1^{st} quartile)	9,469	0.29	0.45	7046	0.29	0.45
$\mathbb{1}$(Wealth 2^{nd} quartile)	9,469	0.22	0.41	7046	0.22	0.41
$\mathbb{1}$(Wealth 3^{rd} quartile)	9,469	0.26	0.44	7046	0.25	0.44
$\mathbb{1}$(Wealth 4^{th} quartile)	9,469	0.24	0.43	7046	0.24	0.42
Age	9,626	38.85	15.22	7046	36.68	11.19
Household size	9,619	5.99	2.42	7046	5.99	2.36
Population in 2000 (000s)	9,181	5815.12	12851.37	7046	5745.19	12657.62
Altitude (meters)	9,181	264.72	250.02	7046	262.04	247.97
No. rivers/water sources	9,181	1.16	1.57	7046	1.17	1.59
No. colonial railways	9,181	0.19	0.43	7046	0.18	0.43
Population in 1900 (000s)	9203	33523.47	83538.05	7046	33288.43	82041.03
Cropland in 1900 (km^2)	9,203	38.48	33.15	7046	38.14	33.30
Latitude	9181	22.85	4.81	7,046	22.99	4.77
Longitude	9181	80.01	6.43	7,046	80.05	6.48
Distance from Protestant medical mission (km)	9,181	81.37	59.58	7046	81.49	59.61
Distance from Protestant mission (km)	9,203	36.51	34.06	7046	36.75	34.12
Distance from Catholic mission (km)	9,203	64.13	52.86	7046	64.65	53.18
Distance from Protestant mission with university (km)	9,203	225.09	177.98	7046	227.08	180.96
Distance from Protestant mission with boarding school (km)	9,203	113.65	103.04	7046	114.98	104.98
Distance from Protestant mission with high school (km)	9,203	156.36	173.67	7046	159.48	176.93
Distance from allopathic hospital from Census 2001 (km)	9,203	20.64	25.42	7046	20.55	25.12

Only individuals of age 20 to 60 and BMI 15 to 30 are included in the estimation sample. The wealth quartiles are based on a wealth index computed as the first principal component that combines information about a set of household assets: number of rooms in dwelling, and asset ownership, such as refrigerator, scooter, radio, and bike. Population, altitude, presence of rivers/water sources, access to colonial railways and cropland data are extracted in 5 km-radius buffers around the WHS locations. Population in year 2000 and elevation data are available at the 30" $\times$ 30" grid cell; original population in year 1900 and cropland data are available at the 5' $\times$ 5' resolution. Latitude and longitude are in decimal degrees.

Table 3.2: Geographical and Historical Controls

	Sample Restricted (Below Percentile of Distance)					
	25%		50%		75%	
	Mean (1)	St. Dev (2)	Mean (3)	St. Dev (4)	Mean (5)	St. Dev (6)
Panel A: Distance from Protestant mission						
Population in 2000 (000s)	12.22	18.99	9.91	17.21	7.37	14.50
Altitude (meters)	269.46	317.21	252.50	275.35	258.54	266.86
No. rivers/water sources within 5 km	1.31	1.81	1.31	1.73	1.21	1.61
Latitude	21.10	4.73	21.94	4.91	22.29	4.91
Longitude	80.17	6.31	80.41	6.31	80.35	6.21
Population in 1900 (000s)	77.07	133.91	55.13	110.97	41.75	94.58
Cropland in 1900 (km^2)	39.84	34.23	41.20	33.22	41.73	33.50
No. colonial railways within 5 km	0.38	0.60	0.29	0.52	0.22	0.46
Distance from Catholic mission (km)	37.43	39.36	44.29	43.42	50.36	42.25
Distance from Protestant mission with university (km)	179.74	150.30	190.08	160.31	194.29	163.11
Distance from Protestant mission with boarding school (km)	79.73	83.96	90.57	88.82	94.29	89.08
Distance from Protestant mission with high school (km)	9203	156.36	173.67	7046	159.48	176.93
Panel B: Distance from Protestant medical mission						
Population in 2000 (000s)	16.83	21.40	10.31	16.98	7.42	14.54
Altitude (meters)	199.48	281.18	219.16	259.60	252.04	257.92
No. rivers/water sources within 5 km	1.39	1.88	1.35	1.70	1.24	1.66
Latitude	22.44	4.15	22.83	4.25	23.04	4.47
Longitude	79.99	6.07	80.39	5.87	80.24	6.13
Population in 1900 (000s)	98.60	145.72	56.70	112.08	41.49	94.81
Cropland in 1900 (km^2)	39.72	35.21	45.33	33.69	41.77	33.63
No. colonial railways within 5 km	0.45	0.62	0.27	0.52	0.22	0.47
Distance from Catholic mission (km)	30.30	30.70	45.14	37.37	55.09	40.34
Distance from Protestant mission with university (km)	110.55	117.00	139.15	121.84	187.69	151.23
Distance from Protestant mission with boarding school (km)	44.28	41.66	58.94	48.17	83.78	69.46
Distance from Protestant mission with high school (km)	62.44	78.74	87.60	89.20	123.43	136.51

Only individuals of age 20 to 60 and BMI 15 to 30 are included in the estimation sample. The mission distance percentiles (based on all sample statistics) are 14 kilometers (25^{th}), 29 kilometers (50^{th}) and 49 kilometers (75^{th}); the medical mission distance percentiles (based on all sample statistics) are 28 kilometers (25^{th}), 75 kilometers (50^{th}) and 121 kilometers (75^{th}). Population, altitude, number of colonial railways and cropland data are extracted in 5 km buffers around the WHS locations. Original population in year 2000 and elevation data are available at the 30" × 30" grid cell; original population in year 1900 and cropland data are available at the 5' × 5' resolution. Latitude and longitude are in decimal degrees.

3.4 Empirical Strategy

Our baseline specification is as follows:

$$BMI_{ivd} = \alpha_d + \beta Hospital\ distance_{ivd} + X'_{ivd}\gamma_1 + W^{g\prime}{}_{vd}\gamma_2 + W^{h\prime}{}_{vd}\gamma_3 + \epsilon_{ivd} \tag{3.1}$$

where BMI_{ivd} is the body mass index of individual i, living in village v in district d.[33] $Hospital\ distance_{ivd}$ is the logarithm of the minimum aerial distance between the Protestant medical mission and survey respondent i.[34] $\beta/100$ measures the change in BMI following a 1 percent change in the distance from a location of a Protestant medical mission.

We control for a large set of covariates influencing both current individuals' BMI and the location of a Protestant hospital at the end of the nineteenth century. X_{ivd} contains individual and household level characteristics such as gender, marital status, level of education, ethnic group, age, household size, and a measure of wealth. The vector W^g_{vd} includes village specific geographic controls, such as population density (2000 population living within a 5-kilometer radius), number of rivers or lakes and average altitude in a 5-kilometer radius, latitude and longitude. W^h_{vd} contains additional historical characteristics, such as estimated population and area covered by cropland in 1900, and access to colonial railways in 1892. We include district level fixed effects, α_d, to account for unobserved heterogeneity across districts.

In order to identify the causal effect of the proximity to a historical Protestant medical facility on current individuals' health outcomes, we address two types of selection bias. First, as discussed in section 3.3.3, historical and geographic characteristics may have driven the missionaries' decision of settling down in specific locations. Moreover, Protestant missions may have endogenously selected themselves into building a health facility and into providing health care.

[33] The survey provides anthropometric information for one single individual per household, so the individual and household dimensions coincide in our dataset.

[34] We consider the logarithm of the distance, instead of the variable in levels, as we believe that percentage changes in distance from a Protestant medical mission matters more than level changes, i.e. we believe that an increase in distance from 10 km to 50 km does not have the same effect of an increase in distance from 110 km to 150 km.

We deal with the first potential selection issue in two ways. First, in order to correct for possible systematic differences between regions with and without Protestant missions, we restrict our sample of analysis to individuals living close to a Protestant mission. By limiting the analysis to more concentrated areas we attempt to minimize the risk of our findings being driven by (within district) unobserved differences. Second, we expand W_{vd}^{h} to include the minimum distance from a generic Protestant and Catholic missions. These variables, while potentially endogenous, should capture factors that also determine the location decisions of Protestant medical missionaries. If our results were purely driven by unobserved heterogeneity across locations, we would find a significant correlation between these variables and individuals' current health outcomes. Moreover, this approach allows us to check that our results are indeed driven by proximity to a Protestant medical mission and not by proximity to a generic mission.

We address the second potential source of selection bias by using an instrumental variable approach. The fact that Protestant missionaries built medical missions in historically and geographically favorable areas may bias the OLS estimates of β upwards. In contrast, potential downward bias of OLS estimates may arise due to measurement error, and to the fact that medical missions were built in areas where medical care was more needed.

Several associations of missionaries and evangelists started to emerge in the late eighteenth century with the purpose of "*administering funds,* [...] *sending out missionaries, initiating and conducting missionary operation, funding churches and institutions, and otherwise fulfilling the varied aims of mission effort*".[35] To construct our instrument we follow Cagé and Rueda (2014) and exploit the heterogeneity in the missionary societies' propensity to undertake medical activities. For each of the 183 Protestant medical missions in our dataset, we observe the supporting missionary society. For each missionary

[35] Dennis (1902).

society, we compute its share of medical missions *outside* India as:

$$\textit{Society hospital share} = \frac{\textit{No. medical missions outside India}}{\textit{No. missions outside India}} \tag{3.2}$$

By construction, all medical missions supported by the same missionary society are assigned the same share.[36] *Society hospital share* can be interpreted as a measure of the propensity of the supporting missionary society to be involved in medical activities, and, therefore, as a measure of the likelihood to have a health facility in their missions. We then construct our instrument as the sum of these shares for all medical missions, $m(k)$, located $k \leq K$ kilometers away from individual i living in village v:[37]

$$\textit{Hospital share}^{K}_{ivd} = \sum_{k \leq K} \textit{Society hospital share}^{m(k)}_{ivd} \tag{3.3}$$

Figure C.5 in the Appendix provides a graphical illustration of how we construct the instrument.[38] The first stage of our IV approach is defined as:

$$\textit{Hospital distance}_{ivd} = \alpha_d + \lambda \textit{Hospital share}^{K}_{ivd} + X'_{ivd}\gamma_1 + W^{g\prime}_{vd}\gamma_2 + W^{h\prime}_{vd}\gamma_3 + \epsilon_{ivd} \tag{3.4}$$

where $\textit{Hospital distance}_{ivd}$ is the logarithm of the minimum distance between the Protestant medical mission and the location of survey respondent i and $\textit{Hospital share}^{K}_{ivd}$ is our instrument as defined in

[36] Table C.1 in the Appendix contains information about the missionary societies we include in our main analysis. The table includes only societies affiliated with medical missions historically located less than 50 kilometers away from a WHS respondent. Data are from the *Centennial Survey of Foreign Missions*.

[37] Our findings do not change if we construct our instrumental variable as the sum of shares, defined as in equation (3.3), for all missions, both medical and non-medical. See section 3.6 for more details.

[38] The larger is the radius K, the weakly larger is the value of the instrument, by construction. We consider the sum of *Society hospital share* instead of the average, as we want to capture the fact that being close to missions affiliated with health oriented societies indeed is associated with a lower distance from a Protestant medical mission. Consider two individuals A and B, both located less than K kilometers away from a mission affiliated with the American Baptist Missionary Union, for which about 23 percent of missions outside of India are medical. If A and B only had one mission located less than K kilometers away, $\textit{Hospital share}^{K}_{ivd}$ would equal 23 for both of them. Now suppose individual A, instead, was located less than K kilometers away from two missions, one supported by the American Baptist Missionary Union and the other by the Wesleyan Missionary Society, for which less than 3 percent of missions outside of India are medical. Considering the average of *Society hospital share* would assign a value of 12.2 to individual A, and therefore make him *less likely* than individual B to live closer to a medical mission. This would be incorrect.

(3.3).

Identification comes from the assumption that *Hospital share*$^K_{ivd}$ is uncorrelated with the error term in our outcome equation (3.1), while adequately correlated with *Hospital distance*$_{ivd}$. The larger *Society hospital share*, the more likely it is for a mission associated with this society to be equipped with a hospital. The larger *Hospital share*$^K_{ivd}$, the more likely it is for an individual to be closer to a medical mission. The exclusion restriction requires *Hospital share*$^K_{ivd}$ to affect current individuals health outcomes only through the proximity to a Protestant medical mission. More specifically, *Society hospital share* must be uncorrelated with any other long-term determinants of individuals' health. In section 3.6, we address concerns about potential violations of this assumption.

3.5 Results

3.5.1 Baseline Model

Table 3.3 presents the baseline OLS estimates. Column 1 shows that BMI and the distance from a Protestant medical mission are significantly and negatively related. The estimated coefficient is equal to -0.271, indicating that a 1 percent increase in the distance is associated with a decrease in individual's BMI by 0.00271. In other words, doubling the distance is associated with a decrease in BMI by 0.19.[39]

Estimates in columns 2 to 5 indicate that the negative correlation between proximity to a medical mission and health outcome today is robust to the inclusion of district fixed effects, individual characteristics and geographic and historical village level controls. While village level controls do not seem to play a role in determining individuals' health status - with altitude and access to colonial railways being the only exceptions - some of the individual level variables are significantly associated with BMI. As expected, better educated individuals have higher levels of BMI, and the same holds for wealthier

[39]We compute the expected change in y associated with a p% increase in x as $\ln\left(\frac{100+p}{100}\right) \times \hat{\beta}$.

Table 3.3: OLS Estimates

	BMI				
	OLS (1)	OLS (2)	OLS (3)	OLS (4)	OLS (5)
Distance from Protestant medical mission (log)	-0.271***	-0.370***	-0.247***	-0.246***	-0.246***
	(0.0527)	(0.0975)	(0.0912)	(0.0921)	(0.0923)
𝟙(Female)			-0.103	-0.108	-0.108
			(0.0779)	(0.0779)	(0.0843)
𝟙(Married)			0.104	0.109	0.109
			(0.104)	(0.103)	(0.0953)
𝟙(Primary school completed)			0.287***	0.281***	0.281***
			(0.0890)	(0.0892)	(0.0873)
𝟙(Major ethnic group)			-0.135	-0.109	-0.109
			(0.105)	(0.105)	(0.105)
Age			0.156***	0.156***	0.156***
			(0.0221)	(0.0222)	(0.0225)
Age^2			-0.00148***	-0.00147***	-0.00147***
			(0.000280)	(0.000281)	(0.000284)
Household size			-0.149**	-0.151**	-0.151**
			(0.0697)	(0.0697)	(0.0650)
Household size^2			0.00784*	0.00807*	0.00807*
			(0.00448)	(0.00448)	(0.00420)
𝟙(Wealth 2^{nd} quartile)			0.287***	0.285***	0.285***
			(0.0971)	(0.0970)	(0.0967)
𝟙(Wealth 3^{rd} quartile)			0.536***	0.538***	0.538***
			(0.0965)	(0.0960)	(0.0941)
𝟙(Wealth 4^{th} quartile)			1.076***	1.083***	1.083***
			(0.112)	(0.113)	(0.107)
Population in 2000 (000s)				0.00725	0.00725
				(0.00549)	(0.00667)
Altitude (meters)				0.000853*	0.000853*
				(0.000458)	(0.000460)
No. rivers				-0.0261	-0.0261
				(0.0368)	(0.0368)
Latitude				0.209	0.209
				(0.219)	(0.219)
Longitude				-0.246	-0.246*
				(0.153)	(0.138)
Population in 1900 (000s)				0.00000992	0.00000992
				(0.000897)	(0.00147)
Cropland in 1900 (km^2)				0.00126	0.00126
				(0.00181)	(0.00191)
No. colonial railways				-0.307*	-0.307*
				(0.177)	(0.164)
N	7,046	7,046	7,046	7,046	7,046
Mean Dependent Variable	20.42	20.42	20.42	20.42	20.42
District FE	No	Yes	Yes	Yes	Yes
Individual controls	No	No	Yes	Yes	Yes
Geographical controls	No	No	No	Yes	Yes
Historical controls	No	No	No	Yes	Yes
Spatial standard errors	No	No	No	No	Yes

Only individuals of age 20 to 60 and BMI 15 to 30 are included in the estimation sample. Robust standard errors in parentheses. * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$. Standard errors clustered at the village (ward) level in columns 1 to 4. Standard errors corrected for spatial correlation in column 5.

individuals. Moreover, BMI increases with age, but at a decreasing rate. The magnitude and significance of the coefficient of interest are almost unchanged across all specifications. Once district fixed effects, individual, and village level controls are included, the OLS estimates indicate that a 50 percent reduction in the distance from a Protestant medical mission is associated with an increase in BMI by 0.17.[40] In columns 1 to 4, robust standard errors are clustered at the primary sampling unit level, which corresponds to villages in rural areas and to wards in urban areas. Column 5 reports standard errors corrected for spatial correlation (Conley (2008)).[41]

3.5.2 Alternative Measures of Health and Proximity

In this paper, BMI and the logarithm of the minimum distance are our preferred variables of interest. Before proceeding further with our empirical analysis, we demonstrate that our findings are robust to alternative measures of health and proximity.

Results are confirmed when using height or a self-reported description of current health status to measure individual health.[42] The higher is the distance from a Protestant medical mission, the lower is the individuals' height. In particular, doubling the distance is associated with a decrease in height by about 0.75 centimeters. In addition, we find that an increase in distance from a historical medical facility is associated with an increase in the probability that an individual describes her health as "moderate" or worse and with a decrease in the probability that an individual describes her health as "good" or "very good". Table 3.4 presents the estimated coefficients. Figure C.6 in the Appendix displays the ordered probit marginal effects.

When considering the logarithm of the distance, instead of the variable in levels, we assume that

[40] See footnote 39.

[41] Findings are robust to a change of the specification to a log-log model, where the coefficient can be interpreted as the the percentage change in BMI following a 1 percent change in the distance from a location of a Protestant medical mission. Results for this alternate specification are included in table C.2 in the Appendix.

[42] Height has been widely used in the literature and includes the long-run effect of fetal and childhood nutritional limitations and disease environment (Fogel (1994), Schultz (2002), Schultz (2003), Weil (2005), and Steckel (2008)).

Table 3.4: Alternative Measures of Health

	Height (cm)	Self-reported Health Status	
	OLS	OLS	Ordered Probit
	(1)	(2)	(3)
Distance from Protestant medical mission (log)	-0.754**	-0.0627**	-0.129*
	(0.339)	(0.0318)	(0.0678)
N	7,046	6,995	6,995
Mean Dependent Variable	159.08	2.69	2.69
District FE	Yes	Yes	Yes
Individual controls	Yes	Yes	Yes
Geographical controls	Yes	Yes	Yes
Historical controls	Yes	Yes	Yes

Only individuals of age 20 to 60 and BMI 15 to 30 are included in the estimation sample. Robust standard errors in parentheses. * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$. Standard errors clustered at the village (ward) level. Column 3 contains the estimated coefficient of the latent model. The estimated coefficient in column 4 suggests that the latent variable increases with the proximity to a Protestant medical mission.

percentage changes in distance from a Protestant medical mission matter more than level changes. We test against this assumption by estimating a linear model. We do not find evidence of a linear relationship between the minimum distance from a Protestant medical mission and current individuals' BMI when the entire sample is considered. However, the existence of a linear relationship cannot be rejected if only relatively small distances are considered.[43]

Our findings are confirmed when we consider non-parametric measures of proximity, such as the presence of a Protestant medical mission nearby. On average, individuals living less than 5 or 25 kilometers away from the location of a historical medical mission have higher BMI than individuals who do not (by about 0.78 and 0.48, respectively). We find that the presence of a medical mission within 50 kilometer plays a role only when interacted with the logarithm of the minimum distance. For individuals living less than 50 kilometers away, doubling the distance would decrease their BMI by 0.25. For those living more than 50 kilometers away, doubling the distance would decrease their BMI by 0.19. Table 3.5 presents the estimation results.[44]

[43]We restrict the sample to individuals living less than 50 kilometers from a historical medical facility. However, the existence of a linear relationship cannot be rejected if individuals living up to 68 kilometers away are considered.

[44]To take into account the fact that the minimum distance may not capture actual proximity due to geographic and historical

Table 3.5: Alternative Measures of Proximity

	BMI					
	Baseline sample	Distance ≤ 50 km	Baseline Sample			
	OLS (1)	OLS (2)	OLS (3)	OLS (4)	OLS (5)	OLS (6)
Distance from Protestant medical mission (km)	-0.00291	-0.0433***				
	(0.00208)	(0.00859)				
𝟙(Medical mission within 5 km)			0.775***			
			(0.224)			
𝟙(Medical mission within 25 km)				0.479**		
				(0.217)		
𝟙(Medical mission within 50 km)					0.0578	
					(0.161)	
Distance from Protestant medical mission (log)						-0.276***
						(0.0912)
𝟙(Medical mission within 50 km) × Distance from Protestant medical mission (log)						-0.0797*
						(0.0466)
N	7,046	2,666	7,046	7,046	7,046	7,046
Mean Dependent Variable	20.42	20.42	20.42	20.42	20.42	20.42
District FE	Yes	Yes	Yes	Yes	Yes	Yes
Individual controls	Yes	Yes	Yes	Yes	Yes	Yes
Geographical controls	Yes	Yes	Yes	Yes	Yes	Yes
Historical controls	Yes	Yes	Yes	Yes	Yes	Yes

Only individuals of age 20 to 60 and BMI 15 to 30 are included in the estimation sample. Robust standard errors in parentheses. * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$. Standard errors clustered at the village (ward) level.

3.5.3 Geographical and Historical Selection

As discussed in section 3.3.3, the non-random location of missionary stations may introduce bias in the OLS estimates presented in table 3.3.[45]

In columns 1 to 3 of table 3.6, we restrict the estimation sample to areas within 49 kilometers (75th percentile), 29 kilometers (50th percentile), and 14 kilometers (25th percentile) from a Protestant mis-

characteristics, we interact it with the presence of colonial railways, access to rivers and altitude, alternatively. The interaction term is not significant in any of the alternative specifications. Results are available upon request.

[45]Endogenous migration decisions of individuals may also introduce a bias in baseline estimates. While we do not observe the migration history of survey respondents and their ancestors, we can control for the mass migration occurred after the partition of India of 1947, and the secession of Bangladesh from Pakistan in 1971. Our findings are confirmed when we estimate our specification (3.1) in a subsample that excludes those Indian states that were more affected by these episodes of mass migration (Assam, West Bengal, and Rajasthan). Results are available upon request.

sion. Such restrictions represent an attempt to correct for possible differences between regions with and without Protestant missions. Our coefficient of interest, β remains statistically significant as the sample size decreases, suggesting that our results are not driven by endogenous missionary location. The estimated coefficients on the logarithm of the distance from a Protestant medical mission are -0.331, -0.272, and -0.403, respectively. For individuals living within the 25^{th} percentile of the minimum distance from a Protestant mission, a 50 percent increase in the distance decreases BMI by 0.16. If only individuals living within the median distance are included, doubling the distance from a medical mission decreases the BMI by about 0.19, while if only the upper quartile is excluded from the sample, we find that a 100 percent increase in the distance decreases BMI by 0.23.

We then augment our set of controls to include measures of proximity to a Protestant mission - with or without a medical facility - and a Catholic mission. This alternative specification allows us to address two issues. First, it provides an additional way of testing whether unobserved heterogeneity across locations is the main driver of our findings. Common factors may have determined the location decisions of Protestant and Catholic missionaries, independent of the missionary activities. If these factors were driving our results, we would find a significant correlation between individual health today and these additional variables. Second, it allows us to assess whether our previous results are driven by vicinity to a medical mission or by proximity to other missionary activities. Column 5 to 6 show this is not the case. Our results indicate that that proximity to a generic Protestant or Catholic mission and current individual BMI are not correlated. In contrast, our coefficient of interest β remains negative and statistically significant at the 5 percent level.

3.5.4 Instrumental Variable Estimation

We adopt an instrumental variable approach to deal with measurement error and the possible endogenous selection of Protestant missionaries into building a health facility. Table 3.7 shows the estimation

Table 3.6: Geographical and Historical Selection

	BMI				
	Distance ≤ percentile			Baseline	
	75%	50%	25%	Sample	
	OLS (1)	OLS (2)	OLS (3)	OLS (4)	OLS (5)
Distance from Protestant medical mission (log)	-0.331*** (0.120)	-0.272* (0.141)	-0.403*** (0.128)	-0.246** (0.102)	-0.216** (0.107)
Distance from Protestant mission (log)				-0.000107 (0.103)	0.0347 (0.111)
Distance from Catholic mission (log)					-0.131 (0.117)
N	5,230	3,476	1,825	7,046	7,046
Mean Dependent Variable	20.54	20.70	20.91	20.42	20.42
District FE	Yes	Yes	Yes	Yes	Yes
Individual controls	Yes	Yes	Yes	Yes	Yes
Geographical controls	Yes	Yes	Yes	Yes	Yes
Historical controls	Yes	Yes	Yes	Yes	Yes

Only individuals of age 20 to 60 and BMI 15 to 30 are included in the estimation sample. Robust standard errors in parentheses. * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$. Standard errors clustered at the village (ward) level. Mission distance percentiles based on all sample statistics. 14 kilometers: 25^{th}; 29 kilometers: 50^{th}; 49 kilometers: 75^{th}.

results.[46] Columns 1 to 3 report the estimated coefficients for the first stage regression, without and with individual, geographic, and historical controls. The distance from a Protestant medical mission is negatively correlated with the instrument *Hospital share*$^{50}_{ivd}$. In all specifications, the coefficient of interest in equation (3.4), λ, is significant at the 1 percent level. Columns 4 to 6 show the 2SLS estimation results.The first stage F statistics is largely above 10, suggesting that *Hospital share*$^{50}_{ivd}$ is a non-weak instrument for our endogenous regressor *Hospital distance*$_{ivd}$.[47]

The relationship between distance from a medical mission and individuals' BMI presented in section 3.5.1 is robust to the instrumental variable approach: proximity to a Protestant medical missions increases BMI. Even though the level of significance of β decreases to 5 percent in column 5, the magnitude of the estimated coefficient more than doubles if compared to the OLS estimates in table 3.3.[48]

[46] We build our instrument using a 50 kilometer radius. In section 3.6 we show that our results are robust to changes in the radius.
[47] We report the Kleibergen-Paap F statistics, which is robust to non-iid errors.
[48] This significant increase in the magnitude of the effect is likely due to measurement error in our measure of distance and

Table 3.7: IV Estimates

	Distance from Protestant medical mission (log)			BMI		
	OLS (1)	OLS (2)	OLS (3)	2SLS (4)	2SLS (5)	2SLS (6)
Hospital share (50km radius)	-1.138***	-1.012***	-0.774***			
	(0.211)	(0.238)	(0.215)			
Distance from Protestant medical mission (log)				-0.769***	-0.622**	-0.734*
				(0.249)	(0.300)	(0.401)
Kleibergen-Paap rk Wald F statistic	-	-	-	29.02	18.03	13.01
N	7,046	7,046	7,046	7,046	7,046	7,046
Mean dependent variable	20.42	20.42	20.42	20.42	20.42	20.42
District FE	Yes	Yes	Yes	Yes	Yes	Yes
Individual controls	No	Yes	Yes	No	Yes	Yes
Geographical controls	No	Yes	Yes	No	Yes	Yes
Historical controls	No	Yes	Yes	No	Yes	Yes
Distance from Protestant mission (log)	No	No	Yes	No	No	Yes
Distance from Catholic mission (log)	No	No	Yes	No	No	Yes

Only individuals of age 20 to 60 and BMI 15 to 30 are included in the estimation sample. Robust standard errors in parentheses. * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$. Standard errors clustered at the village (ward) level.

When individual and village level controls are included, a 50 percent reduction in the distance from a Protestant medical facility increases current BMI by 0.43.

3.6 Robustness Checks and Additional Results

We perform a series of robustness checks to test the sensitivity of our findings. In this section, all specifications are estimated with OLS, and, unless otherwise noted, contain the minimum distance from a Protestant mission and from a Catholic mission as additional controls.

the consequent attenuation bias of the OLS estimates. The geocoding of Protestant medical missions is based on Google coordinates for a string containing the names of *state*, *district*, and *village/city*, which may suffer from measurement issues. While we acknowledge that the increase in the magnitude of the effect may be driven by violations of the exclusion restriction, in section 3.6 we show that this is unlikely.

3.6.1 Exclusion Restriction

The exclusion restriction for the validity of our IV approach requires $\textit{Hospital share}^K_{ivd}$ to affect current individuals' health outcomes only through the proximity to a Protestant medical mission. This assumption may be problematic as societies with and without medical missionaries might differ along dimensions other than their likelihood to have a health facility. These differences may affect development and contemporary health outcomes independent of the health channel, which would represent a violation of the exclusion restriction.

We address this concern in three ways. First, we show that, once controlling for the number of missionary stations, societies involved and not involved in medical activities are not different from each other. Second, we follow Lewbel (2012) to supplement the model with external instruments and confirm the validity of our instrument by performing a test of overidentifying restrictions. Finally, we relax the exclusion restriction as in Conley et al. (2012) and show that the direct effect of the instrument on individual BMI would have to be quite large to render our IV results insignificant.

Societies With and Without Medical Missionaries. To test whether missionary societies involved in medical activities are systematically different from those that are not, we exploit information about 295 missionary society active worldwide in 1902.[49] Table C.5 in the Appendix compares income, diffusion, and investments of missionary societies outside of India depending on whether they had medical missionaries or physicians. On average, societies involved in medical activities were founded earlier and had a larger number of stations worldwide. However, when looking at activities, size and income of each missionary station affiliated with medical and non-medical societies, no significant difference can be detected.

[49] Data are from the *Centennial Survey of Foreign Missions*. For simplicity, we only focus on “societies directly engaged in conducting foreign missions" (*class 1 societies* according to Dennis (1902)).

Testing Overidentifying Restrictions. If the number of instruments is strictly larger than the number of endogenous regressors, the validity of instruments can be checked by performing a test of overidentifying restrictions. While this is not a test for exogenenity of the instruments, but rather that the additional restrictions imposed by having additional instruments are valid, a rejection of the null hypothesis would cast doubt on the validity of the instruments. In order to achieve overidentification, we follow Lewbel (2012) and supplement the model with heteroskedasticity based external instruments.[50] We estimate the model in equation (3.1) by 2SLS using *Hospital share*$^{50}_{ivd}$ and the generated variables as instruments for *Hospital distance*$_{ivd}$.[51] Table C.6 reports estimates obtained using only generated instruments and using both generated and excluded instruments, together with the Hansen J statistic, its degrees of freedom, and p-value. In both cases, the null hypothesis cannot be rejected.

Plausibly Exogenous Instrument. Following Conley et al. (2012), we relax the exclusion restriction and assume that the *potential* direct effect of the instrument on individual BMI is uniformly distributed in an interval $[0, \delta]$, with $\delta > 0$.[52] By varying δ, we identify the threshold at which the second-stage coefficient on the distance from a medical missions becomes insignificant at the 10% level. Figure C.7 shows how large the exclusion restriction violation would need to be in order to invalidate the reduced form results: as long the direct effect of *Hospital share*$^{50}_{ivd}$ on BMI is smaller than 0.24, our second stage is still significant at the 10% level.[53] To gauge magnitudes, we compare this to the overall reduced-form effect of our instrument on individual BMI, which is 0.63.[54] Therefore, to render our IV results insignificant, the direct effect of *Hospital share*$^{50}_{ivd}$ on BMI should be about as large as 40 percent of the overall effect.

[50]The Pagan-Hall test strongly rejects the null hypothesis of homoskedastic disturbances ($\chi^2(170) = 279.602$), so that Lewbel's method can be applied. We implement it in Stata using the `ivreg2h` command (Baum and Schaffer (2012)).

[51]For simplicity, we center all variables to remove district fixed effects. This allows us to generate 19 additional instruments.

[52]Standard IV estimation requires that $\gamma = 0$ in the following equation $y = X\beta + Z\gamma + \epsilon$. Conley et al. (2012) relax the restriction that $\gamma = 0$, and replace it with the assumption that gamma is close to, but not necessarily equal to, zero.

[53]Figure C.7 is produced using the `plausexog` code by Clarke (2014) and the *local to zero* approach (Conley et al. (2012)).

[54]This is obtained by estimating the following model by OLS: $BMI_{ivd} = \alpha_d + \beta Hospital\ share^{50}_{ivd} + X'_{ivd}\gamma_1 + W^{g\prime}_{vd}\gamma_2 + W^{h\prime}_{vd}\gamma_3 + \epsilon_{ivd}$.

3.6.2 Estimation sample

In section 3.3.2, we discuss the use of BMI as a measure of health status and our decision to restrict the estimation sample to individuals of age 20 to 60 and with BMI between 15 and 30. Given the non-monotonic relationship between individuals' BMI and health status, we expect our results to weaken when we include individuals in the far right tail of the BMI distribution (i.e., those who would fall into the obese category, as defined by the WHO). Moreover, as teenagers and elderly may experience changes in BMI due to biological growth and aging that are not directly related to their health status, we expect our results to be affected by the inclusion of individuals 15-20 and 60-65 in the sample. Our results are robust, but weaker, if we extend the estimation sample to individuals of age 15 to 65 and with BMI 18.5 to 30, 15 to 35, and 15 to 45.[55] The sample size decreases drastically once we include only people with BMI above the pre-obesity threshold. As hypothesized, however, the relationship between proximity to a Protestant medical mission and current individuals' BMI disappears. Table C.3 in the Appendix shows the results for these different estimations samples.

3.6.3 Instrument

We test whether our findings are robust to the way we construct the instrumental variable. First, we find our results from the IV estimation to be robust to variations in the radius used to construct the instrument. Second, we show that our findings do not change if we construct our instrumental variable as the sum of societies' medical shares, defined as in equation (3.3), for all missions, both medical and non-medical. To this end, we use data from the *World Atlas Of Christian Missions (1911)* which includes information about medical and non-medical missions active in 1911 and their affiliated missionary societies.[56] Table C.4 in the Appendix presents the estimation results.

[55] See footnote 26 for detailed description of the WHO cutoffs.

[56] We classify a mission to be a medical mission if the presence of a doctor is recorded. Non-medical missions were often affiliated with societies with relatively low *Society hospital share*. The correlation coefficient between the two versions of the instrument is 0.89 and statistically significant at the 1 percent level.

3.6.4 Proximity to Protestant Schools

Parallel to and in most cases preceding the development of medical activities, Protestant missionaries invested in several educational, cultural and philanthropic activities. Previous literature stressed the role of Protestantism in increasing human capital by focusing on its contribution to the advancement of literacy and education.[57]

To assess whether the long-term effect is transmitted via an educational, rather than medical, channel, we control for the minimum distance to different types of educational institutions. Figure C.8 in the Appendix shows the spatial distribution of Protestant boarding schools, high schools and universities operating in colonial India in 1902. While there is some overlapping between the locations of medical facilities and schools, it is far from perfect. We exploit this variation to disentangle the effect of proximity to educational from that related to proximity to medical facilities. To this end, we augment our empirical specification with the logarithm of the minimum distance between individual i and a Protestant educational institution. As shown in table C.7, we still find a negative and statistically significant relationship between hospital distance and current BMI. Proximity to Protestant educational institutions does not seem to play a role, both when excluding (column 4) and including (column 5) the distance from a Protestant medical mission.

3.6.5 Medical Mission Activities

We investigate whether the effect of proximity to a Protestant medical mission varies with the size of the medical facility and with the type of medical procedures performed.

First, we interact the logarithm of the minimum distance with an indicator variable equal to one if the closest medical mission had a total number of patients, treatments and surgeries above the median.[58]

[57] See e.g. Cagé and Rueda (2014) and Mantovanelli (2014).

[58] We collect additional information about activities and characteristics of Protestant medical missions from the *Centennial Survey of Foreign Mission*. In particular, for each medical mission we retrieve data about the total number of patients, treatments and surgeries in 1902 and we use their sum as a proxy for the size of the medical mission. When more than one hospital

Second, we test the hypothesis that it was surgery that determined the acceptance of Western medicine among the Indian people by interacting the distance with an indicator variable equal to one if surgeries were performed at the closest medical mission.[59]

Table C.8 shows the OLS estimates for these alternative specifications. The estimates indicate that there is no independent effect on individuals' health coming from the closest medical mission being sizable. Proximity to a Protestant medical mission, however, has a larger impact on individuals who live close to a mission with an above median number of treatments and patients. We show that individuals for which surgeries were performed at the closest medical mission have, on average, a higher BMI. However, when an interaction term is included, the independent effect of proximity to surgery disappears. The effect of proximity to a medical mission is larger when surgery was available. These differences are however not statistically significant.

3.7 Long-Term Transmission Channels

In this section, we look at the mechanisms underlying the long-term effect of Protestant medical missions on individual health. We identify three potential channels: infrastructure, health potential (health status of previous generations, income and nutrition), and health culture (hygiene and health awareness).[60]

3.7.1 Persistence of Infrastructure

In the previous sections, we omit current access to health care from our baseline specification to avoid identification issues. We include it now as we want to investigate the independent effects of long-term and short-term access to health care. We measure short-term access to health care today with the

or dispensary are present in the same village (city), the total number of patients, treatment and surgeries is computed at the village (city) level.

[59] Fitzgerald (2001) suggests it was surgery that determined acceptance of Western medicine among the Indian people and their greater willingness to use the resources of medical missions.

[60] In section C.2 in the Appendix, we illustrate these transmission channels in the context of a health production function.

minimum distance (logarithm) from a village with at least one hospital.[61] We first assess how proximity to a Protestant medical mission and access to current health facilities correlate. We then estimate the following model:

$$\begin{aligned} BMI_{ivd} = \alpha_d + \beta_1 Hospital\ distance^{LR}{}_{ivd} + \beta_2 Hospital\ distance^{SR}{}_{ivd} \\ + X'{}_{ivd}\gamma_1 + W^{g\prime}{}_{vd}\gamma_2 + W^{h\prime}{}_{vd}\gamma_3 + \epsilon_{ivd} \end{aligned} \tag{3.5}$$

Our coefficients of interest are β_1 and β_2. If the only relevant channel was the persistence of health infrastructure over time, we would expect the coefficient on $Hospital\ distance^{LR}{}_{ivd}$ to become smaller in magnitude and, potentially, insignificant once including $Hospital\ distance^{SR}{}_{ivd}$ in the model.

Table 3.8 presents the estimation results. The long-run effect of access to health care facilities on individuals' health is not driven by persistence of infrastructure. Our measure of short-term access to medical facilities is significantly and positively correlated with the distance from a Protestant medical mission, but it does not seem to independently affect individuals' health.[62] Even when controlling for current access to health care, the distance from a Protestant medical facility remains significantly and negatively associated with BMI. The magnitude of the coefficient on $Hospital\ distance^{LR}{}_{ivd}$ is practically unchanged compared to the results in table 3.6.

3.7.2 Health Potential and Nutrition

The cross-sectional nature of the WHS does not allow us to match survey respondents with health outcomes of previous cohorts. Height as an adult, however, includes the long-run effect of fetal and child-

[61] Using information from the Village Directory of the 2001 Census of India, we successfully geocoded 4,617 villages in Assam, Karnataka, Madhya Pradesh, Maharashtra, Rajasthan, Uttar Pradesh and West Bengal with at least one allopathic hospital and computed the aerial minimum distance between WHS respondents and the allopathic hospitals. For the sake of simplicity, we exclude Ayurvedic, Unani and homeopathic hospitals, dispensaries, maternity and child welfare centers, maternity homes, primary health centers, family welfare centers, T. B. clinics, nursing homes, private and subsidized medical practitioners and community health workers.

[62] This findings are in line with Banerjee et al. (2004) and Das and Hammer (2007), which show that health care provision in India is highly dysfunctional, due for example to the high levels of absenteeism, and not necessarily impacting individuals' health outcomes in a positive way.

Table 3.8: Transmission Channels: Persistence of Infrastructures

	Distance from all. hospital (log)	BMI	
	OLS (1)	OLS (2)	OLS (3)
Distance from Protestant medical mission (log)	0.244***		-0.210*
	(0.0518)		(0.108)
Distance from allopathic hospital (log)		-0.0767	-0.0532
		(0.0701)	(0.0703)
N	7,046	7,046	7,046
Mean Dependent Variable	2.54	20.42	20.42
District FE	No	Yes	Yes
Individual controls	No	Yes	Yes
Geographical controls	No	Yes	Yes
Historical controls	No	Yes	Yes
Distance from Protestant mission (log)	No	No	Yes
Distance from Catholic mission (log)	No	No	Yes

Only individuals of age 20 to 60 and BMI 15 to 30 are included in the estimation sample. Robust standard errors in parentheses. * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$. Standard errors clustered at the village (ward) level.

hood nutrition and exposure to diseases, together with the genetic height potential. Via the intergenerational transmission of health, access to health care and the health status of individuals in one generation could affect the height of individuals in the next generation. The results presented in table 3.4 indicate that proximity to a Protestant medical mission positively affects individual height, providing preliminary evidence of the fact that the health potential may indeed be an important transmission channel.

We now test whether these results are driven by access to better nutrition. To this end, we augment our baseline specification to include measures of quality and quantity of available food (expenditure on food and number of servings of fruit and vegetables consumed per day).[63] We also control for total household expenditure, as a proxy for income. Table 3.9 presents the estimation results. As expected, we find a positive and significant relationship between household expenditure on food and individual health. The same holds for the total household expenditure. We still find that the proximity to a Protestant medical

[63]Expenditure on food includes the value of any food produced and consumed by the household, and excludes alcohol, tobacco and restaurant meals. Moreover, respondents are asked how many servings of fruit and vegetables they consume per day. Unfortunately, no information about other aliments is included in the survey.

Table 3.9: Transmission Channels: Health Potential

	Height (cm)		BMI	
	OLS (1)	OLS (2)	OLS (3)	OLS (4)
Distance from Protestant medical mission (log)	-0.653**	-0.628*	-0.224**	-0.215*
	(0.330)	(0.328)	(0.113)	(0.113)
Fruit & vegetables (servings per day)	-0.127	-0.135	0.0121	0.00937
	(0.127)	(0.121)	(0.0431)	(0.0421)
Household expenditure on food (log)	0.833***	0.524*	0.349***	0.240**
	(0.263)	(0.299)	(0.0820)	(0.0990)
Fruit & vegetables (servings per day)	0.0138	0.0147	-0.00183	-0.00150
× Household expenditure on food (log)	(0.0172)	(0.0165)	(0.00590)	(0.00577)
Household total expenditure (log)		0.508*		0.180**
		(0.267)		(0.0875)
N	5,729	5,729	5,729	5,729
Mean Dependent Variable	158.83	158.83	20.47	20.47
District FE	Yes	Yes	Yes	Yes
Individual controls	Yes	Yes	Yes	Yes
Geographical controls	Yes	Yes	Yes	Yes
Historical controls	Yes	Yes	Yes	Yes
Distance from Protestant mission (log)	Yes	Yes	Yes	Yes
Distance from Catholic mission (log)	Yes	Yes	Yes	Yes

Only individuals of age 20 to 60 and BMI 15 to 30 are included in the estimation sample. Robust standard errors in parentheses. * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$. Standard errors clustered at the village (ward) level.

mission has an independent effect on individual health outcomes. Moreover, the magnitude of the estimated coefficients is not affected, suggesting that the long-term effect of access to health care on current individual height and BMI cannot be simply explained by better nutrition and higher income.

3.7.3 Health Culture and Practices

The quality of medical missionary medicine increased in the final years of the nineteenth century, especially after the Bombay Missionary Conference of 1892-93. At that time, European ideas of modern medicine spread worldwide through the activities of medical missionaries.[64] We use data from the

[64]See Fitzgerald (2001). New scientific discoveries and theories set the foundations of contemporary biomedicine and gained gradual acceptance in Europe and the United States. The pivotal role of the *germ theory of disease* in modern medicine is often described as “the most important single concept for the history of medicine" (Worboys (2000)).

DLHS-3 to investigate whether proximity to a health care facility in the past may have affected individual habits regarding hygiene, self and preventive care and health awareness.[65] The empirical specification is as follows:

$$Health\ culture_{ivd} = \alpha_d + \beta Hospital\ distance_{vd} + X'_{ivd}\gamma_1 + W^{g\prime}_{vd}\gamma_2 + W^{h\prime}_{vd}\gamma_3 + \epsilon_{ivd} \tag{3.6}$$

where $Hospital\ distance_{vd}$ is now the minimum distance between a DLHS-3 village v and a Protestant medical mission (logarithm). We focus on outcomes related to hygiene and health practices, and perinatal maternal and child care. In particular, $Health\ culture_{ivd}$ is, alternatively, an indicator variable equal to one if the respondent treats water to make it safer to drink, if she practices open defecation, or if she turns to traditional healers when she gets sick. To measure disease awareness, we combine the respondent's answers to seven questions regarding knowledge about the danger signs of pneumonia and create an index using a principal component analysis. Finally, to investigate practices related to maternal and child health, we use indicator variables for receiving or purchasing iron folic acid tablets or bottles during the last pregnancy, for receiving a full antenatal check-up (ANC) during last pregnancy, for having a safe delivery (i.e., assisted by skilled personnel, and for taking the newborn for a check-up after birth).

Table 3.10 presents the estimation results.[66] Our measures of hygiene, preventive care and health knowledge are significantly associated with the distance from a Protestant medical mission. In particular, living far from a village historically equipped with a Protestant medical mission decreases the probability of treating water and it increases the probability of practicing open defecation. Interestingly, we do not find any effect on the use of traditional medicine; the fraction of people reporting to turn to traditional healers and home treatments when sick is however quite small. Moreover, distance from a Protestant

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[65] See section 3.3 for more details. Table C.9 in the Appendix presents the descriptive statistics of the variables included in the analysis.

[66] In columns 1 to 3 we use information from the Household survey, while in columns 4 to 8 we consider responses from the Ever Married Women questionnaire. Reported are the marginal effects at the average value of the independent variables. As the number of observations is much larger than the number of fixed effects, the incidental parameter problem associated with estimating a probit model with fixed effects is not a concern (Greene et al. (2002), Fernández-Val (2009)).

medical mission is associated with a lower knowledge of the early signs of pneumonia.

The closer to a Protestant medical mission, the better the practices related to maternal and child health. We find that proximity to a historical medical facility increases women's probability of taking iron supplements while pregnant, of receiving a full antenatal check-up, and of having a safe delivery. Finally, it increases children's probability of receiving a medical check-up within 24 hours since birth.[67]

3.8 Conclusion

In this paper, we investigate the long-term effect of access to health care on current individual health outcomes by focusing on the medical missionary enterprise in India during the second half of the nineteenth century. We combine historical information about the location and activities of Protestant medical missions in colonial India with contemporary individual level data. We show that proximity to the location of a Protestant medical mission positively affects current individuals' body mass index (BMI) and other individual health indicators, such as height and self-reported health status. We verify that our findings are not driven by other non-medical missionary activities, and show that the vicinity to Protestant missions without health facilities does not affect current health outcomes. The investigation of potential transmission channels shows that persistence of infrastructure and current access to health care do not play a role, while improvements in health potential and changes in health culture and preventive care are the main drivers of the long-term effect. In particular, we find that proximity to a historical medical facility improves health habits related to hygiene, health awareness, and perinatal maternal and child health.

This paper represents a first attempt to analyze the persistence of health in developing countries and to investigate how exposure to Protestant medical missions may have determined advancements in

[67] All results are robust to estimating linear probability models and to including the presence of a health facility today in DLHS-3 village v as additional geographic control. Results are available upon request.

current individual health. In light of the existence of long-run effects, the expansion of healthcare access becomes an even more critical priority, as it beneficially affects both current and future generations. In a related project we investigate whether such patterns exist in the African continent. Further work should focus on explicitly connecting the long-term effect of access to health care, current health status, productivity, and economic outcomes.

Table 3.10: Transmission Channels: Health Culture and Practices

	Hygiene & Health Practices				Maternal & Child Health			
	Pr(Treat water)	Pr(Open defecate)	Pr(Traditional medicine)	Knowledge pneumonia	Pr(Iron suppl. while pregnant)	Pr(Full ANC)	Pr(Safe delivery)	Pr(Child checked within 24hrs)
	Probit	Probit	Probit	Tobit	Probit	Probit	Probit	Probit
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
Distance from Protestant medical mission (log)	-0.0362***	0.0264*	-0.0321	-0.00992***	-0.0170*	-0.0311*	-0.0296*	-0.0376**
	(0.0134)	(0.0157)	(0.0333)	(0.00383)	(0.00987)	(0.0171)	(0.0174)	(0.0186)
N	273,109	275,386	219,033	198,225	198,168	74,017	72,728	71,864
Mean Dependent Variable	0.30	0.66	0.04	0.22	0.81	0.13	0.42	0.39
District FE	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Individual controls	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Geographical controls	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Historical controls	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Distance from Protestant mission (log)	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Distance from Catholic mission (log)	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes

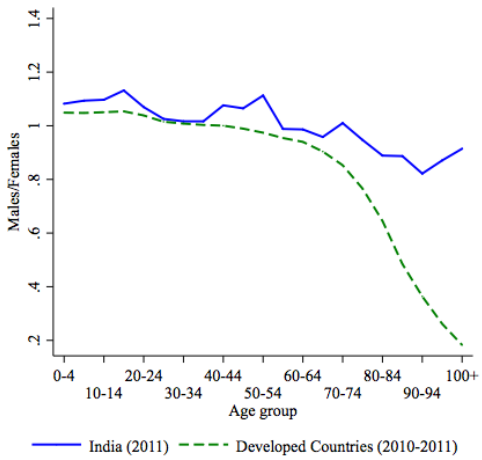
DHLS-3 data. Columns 1 to 3 uses information from the Household survey. Columns 4 to 8 uses responses from the Ever Married Women questionnaire. Respondents are ever married women 15-49. Robust standard errors in parentheses. * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$. Standard errors clustered at the village level. Marginal effects in columns 1 to 3 and 5 to 8.

Appendix A

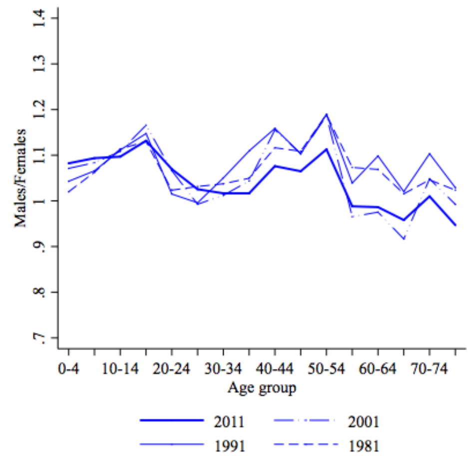
Why Are Older Women Missing in India? The Age Profile of Bargaining Power and Poverty

This Appendix contains seven main sections. Appendix A.1 shows that the recent findings about missing Indian women at older ages are consistent with the distribution of sex-ratios across ages. Appendix A.2 contains additional tables and figures. Appendix A.3 presents the full empirical analysis of the effect of the Hindu Succession Act amendments on women's health and details on the dataset used, the empirical strategy and the robustness checks. Appendix A.4 discusses a fully specified collective model of Indian households. While invoking additional functional forms that are not required for the empirical estimation of the resource shares, the fully specified model delivers explicit formulas for household demands and resource shares that satisfy the identification requirements. Appendix A.5 investigates the sensitivity of the model estimates. Appendix A.6 uses additional data on singles to empirically test the Pareto efficiency assumption in this framework. Appendix A.7 relaxes the assumption of equality of poverty thresholds between genders and simulates gender and gender-age poverty rates in this alternative scenario.

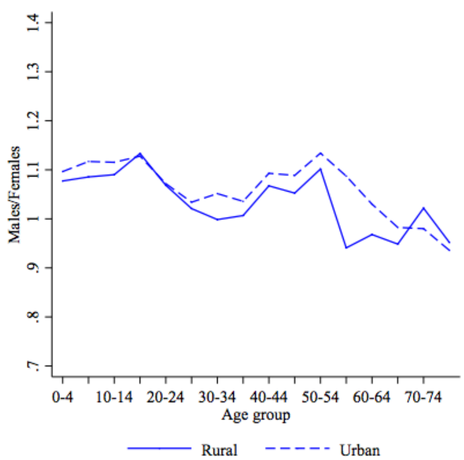
A.1 Sex Ratios and Age



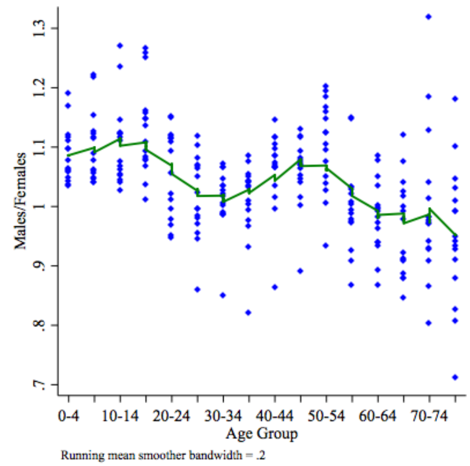
(A) Sex Ratio By Age Group



(B) Cohort Comparison



(C) Urban vs. Rural

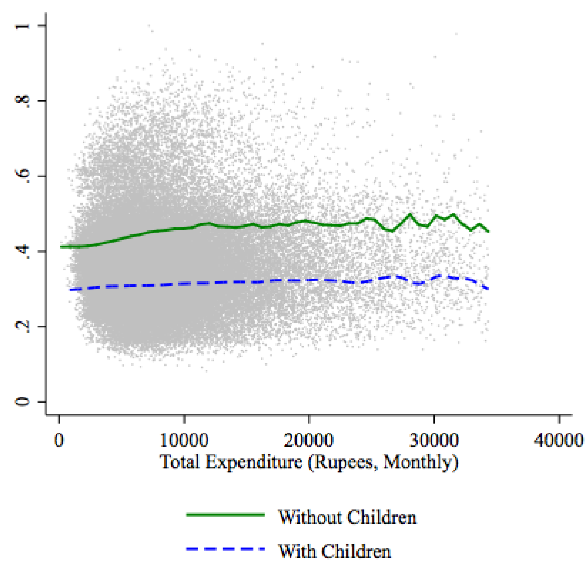


(D) Distribution Across States

Note: One must be cautious when comparing aggregate sex-ratios across countries due to different fertility and death rates, and different age distributions and disease composition (Anderson and Ray (2010)). However, it is interesting to note that the recent findings about missing Indian women at older ages are consistent with a striking non-monotonic pattern of the distribution of sex-ratios across ages: compared to developed countries, figures are particularly skewed at post-reproductive ages. Data are from the United Nations Statistics Division and Census of India. Panel A plots the sex-ratios for India and for selected developed countries across age groups. Developed countries include US, Canada, Germany, Italy, Japan, Spain, and Portugal. This pattern is robust to changes in the set of developed countries considered. For India the ratio does not decrease substantially with age, with the gap between India and developed countries actually increasing at post-reproductive ages. The comparison between data from different census years (1981, 1991, 2001 and 2011) indicates that this pattern is not driven by a specific cohort (panel B). Moreover, the non-monotonic pattern is common in rural and urban areas (panel C). Finally, with some notably exceptions such as Kerala, this pattern is confirmed by state level data (panel D).

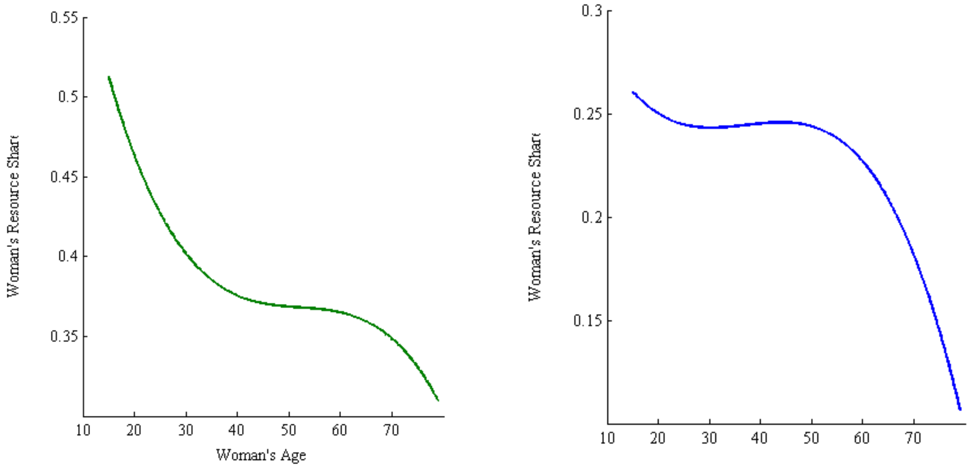
Figure A.1: The Age Distribution of Sex Ratios

A.2 Additional Figures and Tables



Note: Local mean smoothing obtained using Epanechnikov kernel regression.

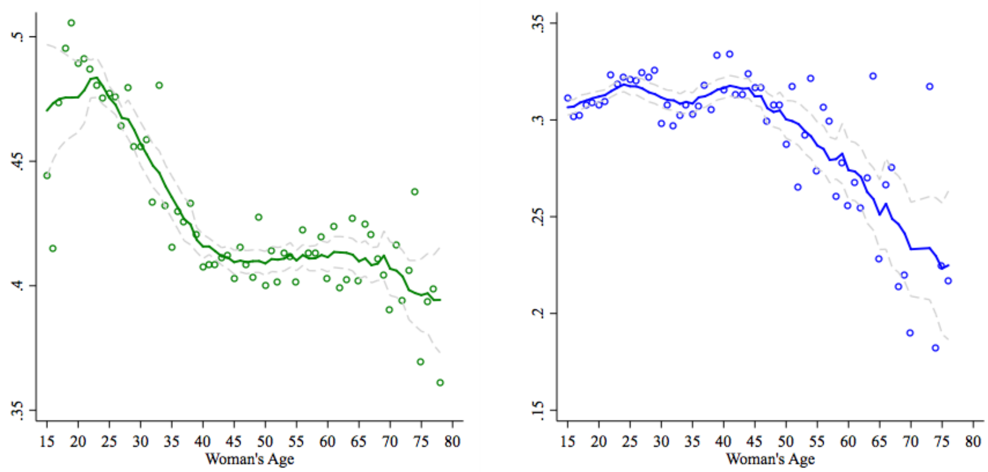
Figure A.2: Women Resource Shares and Total Expenditure



(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.

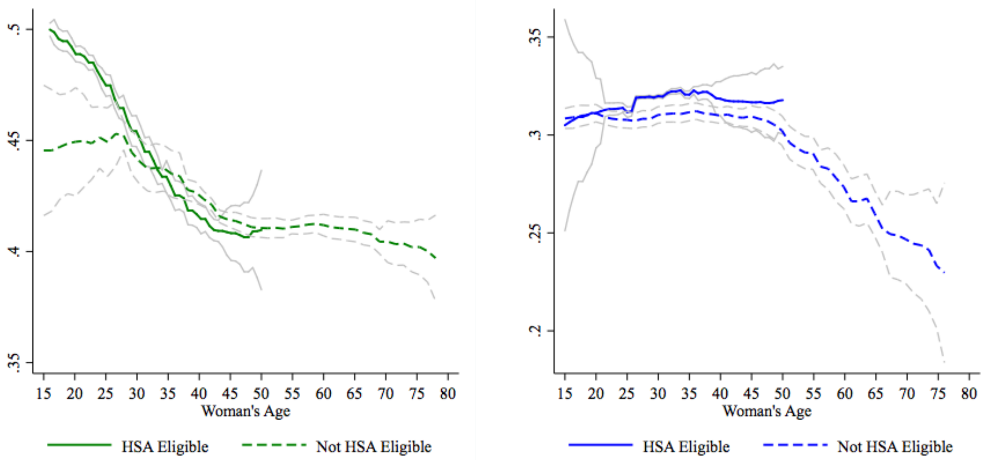
Figure A.3: Predicted Women's Resource Shares and Age (Reference Households)



(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.

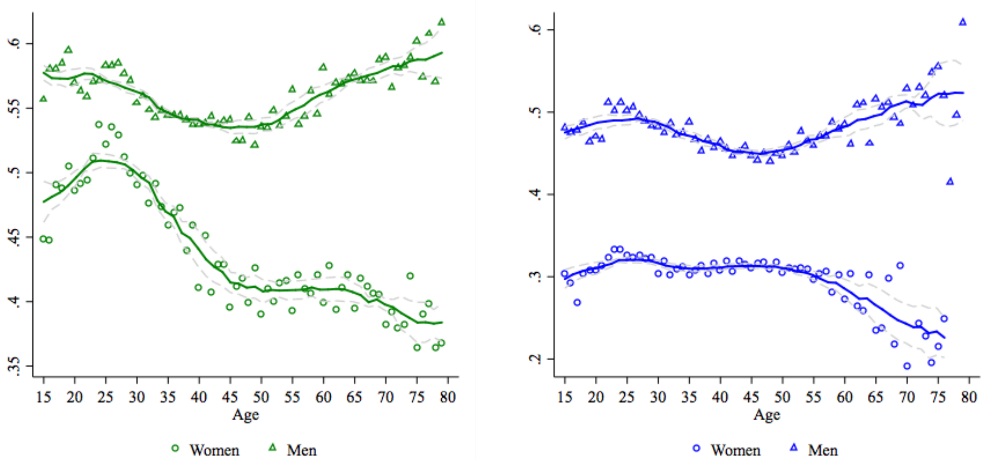
Figure A.4: Average Predicted Women's Resource Shares and Age (Nuclear Households Only)



(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 $a = 15, ..., 79$. The solid line is a running mean. The dashed lines are the 95 percent pointwise confidence interval for the smoothed values.

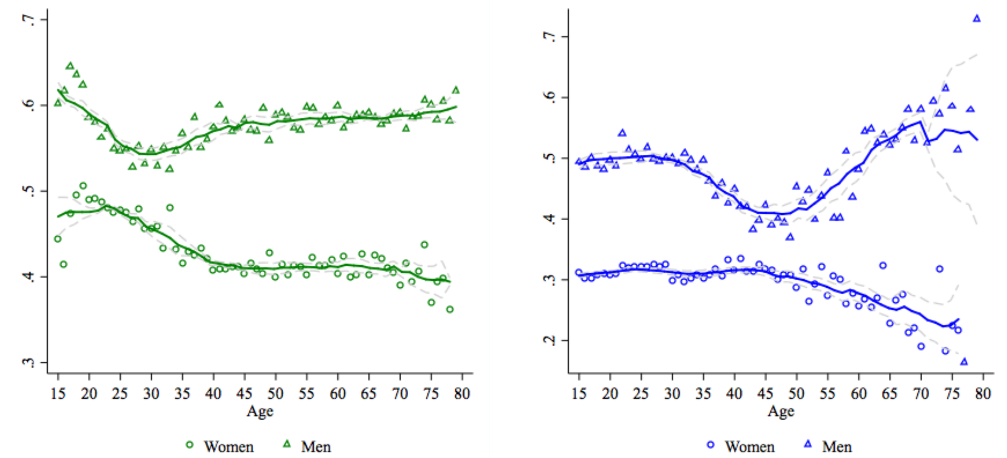
Figure A.5: Predicted Resource Shares, Age and HSA Amendments (Nuclear Households Only)



(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. The dashed lines are the 95 percent pointwise confidence interval for the smoothed values.

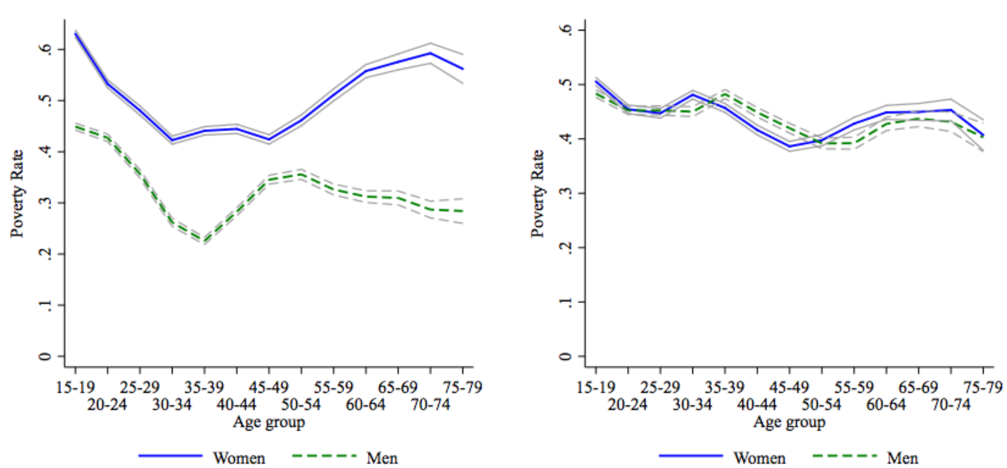
Figure A.6: Average Predicted Resource Shares and Age



(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. The dashed lines are the 95 percent pointwise confidence interval for the smoothed values.

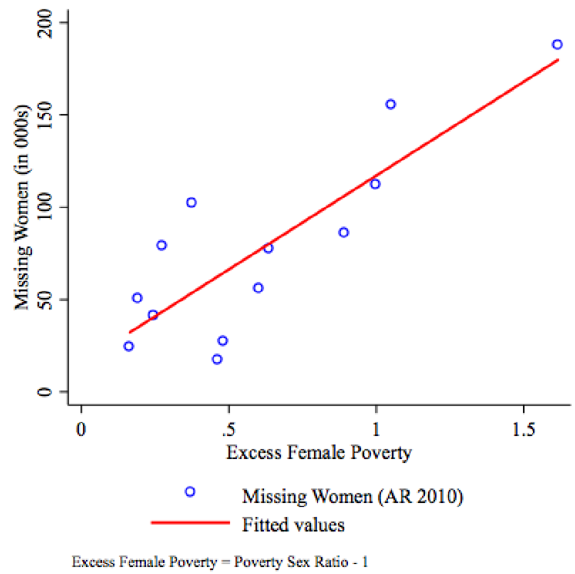
Figure A.7: Average Predicted Resource Shares and Age (Nuclear Households Only)



(A) Model Predictions (Unequal Sharing) (B) Equal Sharing

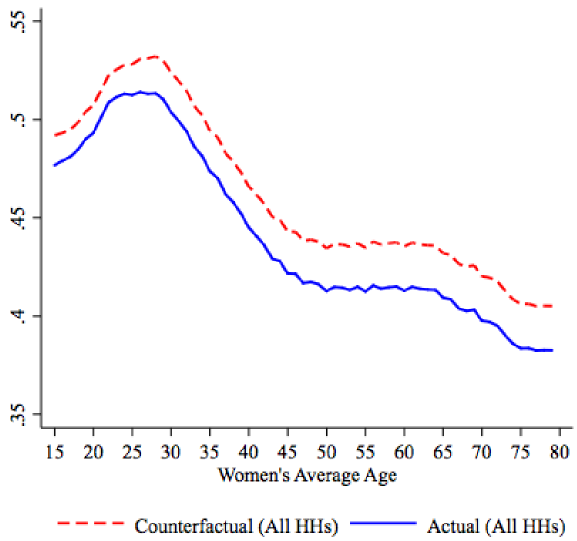
Note: The graphs show the fraction of females or males in each age group living below poverty line. Individuals from all households in the sample are used for calculations. Per capita expenditures are computed using the model predictions in panel A. In panel B, I assume that household expenditure is split equally among household members. Per capita expenditures are compared to the 2 US$/day poverty line.

Figure A.8: Poverty Rates By Gender and Age (2 US$/day)

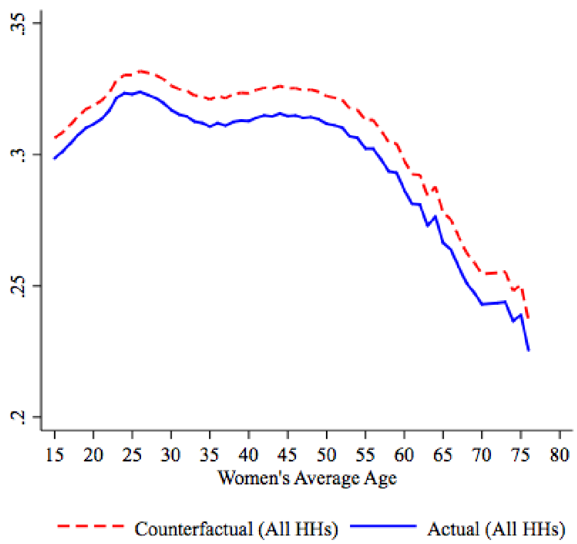


Age Groups 15-19 to 75-79

Figure A.9: Excess Female Deaths 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 A.10: Counterfactual Experiment: HSA Amendment and Women's Resource Shares

Table A.1: NSS - CES Descriptive Statistics

	Households Without Children < 15				Households With Children < 15			
	Obs.	Mean	Median	St. Dev.	Obs.	Mean	Median	St. Dev.
Expenditure (Rupees):								
Total Expenditure	30,248	7881	6479	5265	57,202	8224	6909	4903
Expenditure On Non-Durable Goods	30,248	7396	6206	4720	57,202	7848	6696	4489
Expenditure On Durable Goods	30,248	485	107	1367	57,202	375	107	1012
Budget Shares:								
Food	30,248	35.39	35.50	9.82	57,202	39.12	39.16	9.61
Female Assignable Clothing	30,219	1.41	1.19	1.22	57,157	1.26	1.09	1.04
Male Assignable Clothing	30,219	1.70	1.43	1.44	57,157	1.56	1.30	1.34
Children Assignable Clothing	-	-	-	-	57,157	0.67	0.49	0.80
Household Characteristics:								
$\mathbb{I}$(HSAA Eligible)	26,846	0.08	0	0.28	47,382	0.14	0	0.35
No. Adult Females	30,248	1.67	1	0.83	57,202	1.69	1	0.86
No. Adult Males	30,248	1.91	2	0.93	57,202	1.67	1	0.90
No. Children Under 15	-	-	-	-	57,202	2.02	2	1.01
Fraction Of Female Children	-	-	-	-	57,202	0.45	0.5	0.39
$\mathbb{I}$(Daughter in Law)	30,248	0.11	0	0.32	57,202	0.24	0	0.43
$\mathbb{I}$(Unmarried Daughter)	30,248	0.33	0	0.47	57,202	0.17	0	0.38
$\mathbb{I}$(Widow)	30,248	0.16	0	0.37	57,202	0.14	0	0.35
Avg. Age Men 15 to 79	30,013	39.37	36	13.11	57,153	36.93	36	8.75
Avg. Age Women 15 to 79	30,159	40.98	40	12.09	57,181	34.84	34	8.20
Avg. Age Gap 15 to 79 (Men - Women)	29,948	-1.43	-1.5	12.86	57,134	2.10	3	9.93
Avg. Age Children 0 to 14	-	-	-	-	57,202	7.57	8	3.97
$\mathbb{I}$(Hindu, Buddhist, Sikh, Jain)	30,248	0.83	1	0.38	57,202	0.77	1	0.42
$\mathbb{I}$(Sch. Caste, Sch. Tribe or Other Backward Classes)	30,248	0.65	1	0.48	57,202	0.71	1	0.45
$\mathbb{I}$(Salary Earner)	30,248	0.32	0	0.47	57,202	0.29	0	0.46
$\mathbb{I}$(Land Ownership)	30,248	0.90	1	0.30	57,202	0.89	1	0.31
$\mathbb{I}$(Female Higher Education)	30,248	0.14	0	0.35	57,202	0.10	0	0.30
$\mathbb{I}$(Male Higher Education)	30,248	0.24	0	0.43	57,202	0.17	0	0.37
$\mathbb{I}$(Rural)	30,248	0.57	1	0.50	57,202	0.63	1	0.48
$\mathbb{I}$(North India)	30,248	0.28	0	0.45	57,202	0.33	0	0.47
$\mathbb{I}$(East)	30,248	0.19	0	0.39	57,202	0.21	0	0.41
$\mathbb{I}$(North-East)	30,248	0.12	0	0.33	57,202	0.16	0	0.36
$\mathbb{I}$(South)	30,248	0.27	0	0.45	57,202	0.19	0	0.39
$\mathbb{I}$(West)	30,248	0.13	0	0.34	57,202	0.12	0	0.32

Table A.2: 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	-0.0448***	-0.00940***	-0.0278***	-0.0320***
	(0.0106)	(0.00188)	(0.00810)	(0.0112)
N	81,115	56,746	77,530	77,755
Mean Dependent Variable	0.2648	0.0154	0.1559	0.5298

*$p < 0.10$, **$p < 0.05$, ***$p < 0.01$. 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 women's years of schooling, number of children under 5 in the household, a household wealth index, and indicator variables for 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.

Table A.3: 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	-0.2623	0.0917	-0.2565	-0.6080	0.0571
Men Assignable Clothing	-0.3545	0.1589	-0.3546	-1.3401	0.0159
Food	-6.8129	1.6372	-6.8380	-13.9877	-2.4195
Panel B: With Children < 15 Only					
Women Assignable Clothing	-0.1163	0.0840	-0.0967	-0.5119	0.0922
Men Assignable Clothing	-0.1843	0.1213	-0.1780	-0.8767	0.1505
Children Assignable Clothing	-0.0678	0.0412	-0.0630	-0.2936	0.1062
Food	-6.6591	1.5287	-6.6465	-13.0420	-1.3406

A.3 Bargaining Power and Health: Details and Robustness Checks

A.3.1 Measuring Women's Health

Measurement of health has recently evolved to rely on anthropometric and physiologic indicators of physical development.[1] In section 1.3, I use body mass index (BMI) and levels of anaemia as measures of women's health. BMI is defined as weight in kilograms divided by height in meters squared. A cut-off point of 18.5 is used to define thinness or acute undernutrition, and a BMI of 23 or above indicates overweight or obesity for Asian Indians.[2]

That the relationship between health status and BMI is non-linear is a well known fact, both at the individual and at the aggregate level. Wailer (1984) clearly illustrates this non-linearity by showing a U-shaped relationship between relative risk of mortality and BMI using Norwegian data from the 1970s. In low income countries, however, increases in the caloric intake shifts the lower tail of the BMI distribution to the right. Fogel (1994, 2004) interprets this shift as an accumulation of the population's health human capital stock, which tends to be associated with a decline in mortality. Research work in the medical literature indicates that individuals with very low BMI (below 18.5) experience substantially increased mortality risk compared with those with a normal-range BMI (18.5 to 23), especially in relation to cardiovascular and respiratory diseases.[3]

Anaemia is a condition in which the number of red blood cells or their oxygen-carrying capacity is insufficient. Although its primary cause is iron deficiency, it is seldom present in isolation and more frequently it coexists with a number of other causes, such as malaria, parasitic infection, and nutritional deficiencies; 90 percent of anaemia sufferers live in developing countries, about 2 billion people in total. Prevalence of anaemia in South Asian countries is among the highest in the world.[4] While girls and

[1] See Schultz (2009) for more details.
[2] See Shiwaku et al. (2004).
[3] See e.g Visscher et al. (2000), Thorogood et al. (2003), and Zheng et al. (2011).
[4] See Kaur (2014).

Table A.4: NFHS-3 Descriptive Statistics

	Obs.	Mean	Median	St. Dev.
Body mass index	82,303	21.4232	20.61	4.2226
$\mathbb{I}(BMI \leq 18.5)$	82,303	0.2648	0	0.4412
$\mathbb{I}$(Severe Anaemia)	78,521	0.0154	0	0.1230
$\mathbb{I}$(Moderate Anaemia)	78,521	0.1559	0	0.3628
$\mathbb{I}$(Mild Anaemia)	78,521	0.5298	1	0.4991
$\mathbb{I}$(HSAA Exposed)	85,881	0.1528	0	0.3598
$\mathbb{I}$(HSAA Eligible)	69,558	0.1095	0	0.3123
$\mathbb{I}$(Hindu, Buddhist, Sikh, Jain)	85,881	0.7920	1	0.4059
$\mathbb{I}$(Sch. Caste, Sch. Tribe or Other Backward Classes)	85,222	0.6345	1	0.4816
$\mathbb{I}$(Rural)	85,881	0.5609	1	0.4963
$\mathbb{I}$(Worked in past 12 months)	85,869	0.4023	0	0.4904
Wealth index	85,742	0.2707	0.2015	0.2425
Age	85,881	31.9504	31	8.4235
No. Children Under 5	85,881	0.8075	0	1.0179
Years of Schooling	85,876	5.3261	5	5.1842

Only married women of age 15 to 49 included. 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. 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). Appropriate adjustments in these cutoff points were made for respondents living at altitudes above 1,000 meters and respondents who smoke, since both of these groups require more haemoglobin in their blood.

women at child-bearing ages are the most affected, there are longer term effects stemming from anaemic conditions, such as cardiovascular deficiencies, that may affect women's health at post-reproductive and older ages.[5] Medical studies have demonstrated that anaemia is associated with increased risk of death among older adults.[6]

The table includes summary statistics for all the variables included in the empirical analysis.

A.3.2 Robustness Checks

Endogeneity of Treatment. Exposure to the HSA amendments is determined by each woman's year of marriage, which may be an endogenous choice. On one hand, if more progressive families waited

[5] The World Health Organization estimated that about half of the global maternal deaths due to anaemia occur in South Asian countries, with India alone accounting for half of global maternal deaths (De Benoist et al. (2008)). Maternal short stature and iron deficiency anaemia increase the risk of death of the mother at delivery, accounting for at least 20 percent of maternal mortality (Black et al. (2008), Kalaivani et al. (2009)).

[6] See e.g., Culleton et al. (2006) and Patel (2008).

until the reform to marry their daughters, the treatment effect estimated by OLS could potentially be capturing unobserved heterogeneity in early childhood health investment for girls and therefore lead to upward biased estimates. On the other hand, if women's access to inheritance rights was seen as a compensation for a daughter's weak constitution or proneness to sickness, OLS estimates of β could underestimate the treatment effect of HSAA.

I address endogeneity concerns in three different ways. First, I exclude from the analysis women who got married right around the reforms: the year of, before and after. By excluding women whose marriage could be more easily scheduled before or after the amendment according to their - or their family's - preferences, I minimize the risk of results being driven by unobserved heterogeneity across households. Second, I focus on 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. I estimate the intent-to-treat effect by comparing girls who were 14 or younger at the time of the HSA amendment in their state to girls who were 23 or older:

$$y_{irsc} = \tilde{\beta} \mathit{HSAA\ Eligible}_{rsc} + X'_{irsc}\tilde{\gamma} + \tilde{\alpha}_r + \tilde{\alpha}_c + \tilde{\alpha}_s + \tilde{\alpha}_{rs} + \tilde{\alpha}_{rc} + \tilde{\alpha}_{sc} + \tilde{\epsilon}_{irsc} \tag{A.1}$$

where $\mathit{HSAA\ Eligible}_{rsc}$ is fully determined by each woman's religion, year of birth and state, as it is the interaction between an indicator variable equal to one if a woman was 14 or younger at the time of the reform in her state and to zero if she was 23 or older and an indicator variable for being Hindu, Buddhist, Sikh or Jain.[7] Finally, I estimate the effect of being exposed to HSA amendments on women's health outcome by instrumenting $\mathit{HSAA\ Exposed}_{irsc}$ with $\mathit{HSAA\ Eligible}_{rsc}$.[8] The exclusion restriction

[7] As in the cohort comparison analysis provided in Heath and Tan (2014), I use 14 and 23 as they are the 10th and 90th percentiles of women's age at marriage in the survey. Results are robust to alternative choice of thresholds. Duflo (2001) uses a similar strategy to estimate returns to education in Indonesia and compares the educational attainment and the wages of individuals who had little or no exposure to the INPRES program (12 to 17 at time of implementation) to those of individuals who were exposed all the time they were in primary school (2 to 6 in 1974).

[8] Heath and Tan (2014) also use an instrumental variable approach to address the endogeneity problem. In the first stage, however, they instrument woman's HSAA treatment status by 2,088 fixed effects for each religion-year of birth-state cell. Moreover, they do not include control variables in their analysis.

needed for identification is that, once controlling for religion-cohort, cohort-state, and state-religion fixed effects, being Hindu and younger than 14 at the time of implementation has an effect on women's health outcomes only through their actual exposure to the HSA amendment. This assumption seems plausible. Despite other factors and policies may have differentially affected young Indian women, these do not vary by religion and are therefore controlled for by the cohort-state fixed effects.

Table A.5 reports the results of the first two strategies used to deal with the possible endogeneity of treatment. Panel A shows the estimated coefficients for the model in equation (1.1) when women who got married around the amendments in their state of residence are excluded from the analysis. Panel B includes the estimation results for the model in equation (A.1). Table A.6 reports the 2SLS estimates, obtained by instrumenting $\textit{HSAA Exposed}_{irsc}$ with $\textit{HSAA Eligible}_{rsc}$.[9] A comparison between the magnitude of the coefficients in tables A.5 and A.6 with those in table 1.1 suggests that OLS estimates over the entire sample are in general downward biased. The overall conclusion is confirmed, indicating that the endogeneity of time at marriage is not a major concern: HSAA exposure improves women's health outcomes. While the marginal effects on the probability of being underweight or mildly anaemic are not statistically significant, exposure to HSA amendments is found to increase women's BMI and to reduce their probability of being severely or moderately anaemic.

I address the limitations of linear probability models by estimating the baseline model over the restricted sample and equation (A.1) over the full sample by maximum-likelihood probit models. Moreover, I estimate a maximum-likelihood two-equation probit model specified as follows:

$$\begin{cases} \textit{HSAA Exposed}_{irsc} = \beta_1 \textit{HSAA Eligible}_{rsc} + X'_{irsc}\gamma_1 + \alpha_r + \alpha_c + \alpha_s + \alpha_{rs} + \alpha_{rc} + \alpha_s * (c - 1955) + \epsilon_{1,irsc} \\ y_{irsc} = \beta_2 \textit{HSAA Exposed}_{irsc} + X'_{irsc}\gamma_2 + \delta_r + \delta_c + \delta_s + \delta_{rs} + \delta_{rc} + \delta_s * (c - 1955) + \epsilon_{2,irsc} \end{cases} \tag{A.2}$$

[9] The correlation between the potentially endogenous treatment variable and the eligibility instrument is 0.92 and the very large F statistics in the first stage provide reassurance against a weak instrument problem.

Table A.5: HSA Amendments and Women's Health: Robustness Checks

	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
Panel A: Restricted Sample							
HSAA Exposed	0.392***	0.229**	-0.0183	-0.0240	-0.0127**	-0.0407***	-0.0210
	(0.136)	(0.100)	(0.176)	(0.0151)	(0.00548)	(0.0138)	(0.0175)
N	74,656	52,308	22,348	74,656	71,249	71,249	71249
Panel B: Unrestricted Sample							
HSAA Eligible	0.905***	0.544**	0.498	-0.00695	-0.0303***	-0.0692**	0.00759
	(0.336)	(0.227)	(0.491)	(0.0337)	(0.0106)	(0.0303)	(0.0444)
N	66,146	45,430	20,716	66,146	62,902	62,902	62,902
Mean Dependent Variable	21.42	19.24	26.69	0.2648	0.0154	0.1559	0.5298

$*p < 0.10$, $**p < 0.05$, $***p < 0.01$. 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, religion-cohort fixed effects, and individual controls. Individual controls include women's years of schooling, number of children under 5 in the household, a household wealth index, and indicator variables for 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. $HSAAExposed = \mathbb{I}$(Marriage Year $\leq$ HSAA Year) $\times$ $\mathbb{I}$ (Hindu, Sikh, Buddhist or Sikh). HSAA Eligible equal to the interaction between an indicator variable equal to one if a woman was 14 or younger at the HSAA implementation in her state and to zero if she was 23 or older and an indicator variable for being Hindu, Buddhist, Sikh or Jain. In Panel B the sample exclude women who got married on the year of the amendment, the year after or the year before.

For ease of estimation, I include state linear trend instead of state-cohort fixed effects. Table A.7 shows the marginal effects of being exposed to HSA amendments on binary health outcomes, while table A.8 presents the estimation results for the system (A.2). For simplicity, I report only the marginal effect of HSAA Exposed$_{irsc}$. The findings are overall confirmed.

Table A.6: HSA Amendments and Women's Health: Instrumental Variable

	Body Mass Index				Pr(Anaemia)		
	All Sample	$BMI \leq 23$	$BMI > 23$	Pr($BMI \leq 18.5$)	Severe	Moderate	Mild
	(1) 2SLS	(2) 2SLS	(3) 2SLS	(4) 2SLS	(5) 2SLS	(6) 2SLS	(7) 2SLS
HSAA Exposed	1.013***	0.593**	0.590	-0.00777	-0.0339***	-0.0775**	0.00849
	(0.374)	(0.246)	(0.570)	(0.0374)	(0.0118)	(0.0337)	(0.0494)
N	66,146	45,430	20,716	66,146	62,902	62,902	62,902
Kleibergen-Paap F-statistic	18,774.1	17,664.2	3,147.4	18,774.1	18,057.3	18,057.3	18,057.3
Mean Dependent Variable	21.42	19.24	26.69	0.2648	0.0154	0.1559	0.5298

*$p < 0.10$, **$p < 0.05$, ***$p < 0.01$. 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, religion-cohort fixed effects, and individual controls. Individual controls include women's years of schooling, number of children under 5 in the household, a household wealth index, and indicator variables for 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.

Table A.7: HSA Amendments and Women's Health: Robustness Checks (Probit)

	Body Mass Index	Pr(Anaemia)		
	Pr($BMI \leq 18.5$)	Severe	Moderate	Mild
	(1) Probit	(2) Probit	(3) Probit	(4) Probit
Panel A: Restricted Sample				
HSAA Exposed	-0.0406**	-0.00768	-0.0372***	-0.0211
	(0.0171)	(0.0490)	(0.0120)	(0.0179)
N	74239	51748	70975	71210
Panel B: Unrestricted Sample				
HSAA Eligible	-0.0619	-0.0201***	-0.0602***	0.00717
	(0.0394)	(0.00322)	(0.0232)	(0.0455)
N	65,762	45,361	62,737	62,900
Mean Dependent Variable	0.2648	0.0154	0.1559	0.5298

*$p < 0.10$, **$p < 0.05$, ***$p < 0.01$. 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, religion-cohort fixed effects, and individual controls. Individual controls include women's years of schooling, number of children under 5 in the household, a household wealth index, and indicator variables for 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.

Table A.8: HSA Amendments and Women's Health: Bivariate Probit

	Body Mass Index	Pr(Anaemia)		
	Pr($BMI \leq 18.5$)	Severe	Moderate	Mild
	(1) BiProbit	(2) BiProbit	(3) BiProbit	(4) BiProbit
HSAA Exposed	0.0023 (0.0193)	-0.0081*** (0.0014)	-0.0321** (0.0126)	-0.0185 (0.0225)
N	66,146	62,902	62,902	62,902
Mean Dependent Variable	0.2648	0.0154	0.1559	0.5298

*$p < 0.10$, **$p < 0.05$, ***$p < 0.01$. 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, religion-cohort fixed effects, state specific linear trends, and individual controls. Individual controls include women's years of schooling, number of children under 5 in the household, a household wealth index, and indicator variables for 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.

Falsification Tests. In order to assess the validity of my identification strategy, I perform two different falsification tests. First, I compare health outcomes of younger and older women, who could not be exposed to HSA amendments as they were too old at the time of implementation. In particular, I estimate the specification in equation (A.1) with the main regressor being the interaction between an indicator variable equal to one if a woman was 23 to 30 years old at the time of the HSA Amendment in her state and to zero if she was 30 or older, and an indicator variable for being Hindu, Buddhist, Sikh or Jain. If there was a difference in trends between the reformed and non-reformed states, I would find a statistically significant and positive coefficient on the newly defined interaction term. This could cast doubt on the validity of the identification strategy discussed above. Second, I estimate the effect of HSA Amendments on men's health outcome. I combine data on BMI, anemia level and demographic characteristics for adult males with information about women's eligibility within a household and I estimate the model in equation (A.1) with the dependent variable now being males' health outcomes. If there was a general increasing trend common to Hindu individuals, both men and women, in the reforming states, the estimated treatment effect would be positive and significant for men as well.[10]

Table A.9 contains the estimation results. For simplicity, the table reports only a subset of the health outcomes considered above. Columns 1 to 4 show the estimation results for the model in equation (A.1) with the main regressor now being the interaction between an indicator variable equal to one if a woman was 23 to 30 years old at the HSAA implementation in her state and to zero if she was 30 or older, and an indicator variable for being Hindu, Buddhist, Sikh or Jain. I find no significant effect on women's BMI, on their probability of being underweight nor on their probability of being anaemic, providing reassurance in favor of the validity of the empirical exercise. Columns 5 to 8 report the estimation results for the model in equation (A.1) when males' health outcomes are considered and the individual controls include males' schooling and an indicator variable for having worked in the past 12 months.[11] The estimated coefficients and marginal effects are not statistically different from zero, suggesting no violation of the identification assumptions.

[10]While an increase in men's health *could* be rationalized by a change in expenditure patterns induced by the increase in women's bargaining power - for example for food and medical expenses -, evidence of it may cast doubt on the validity of the identification strategy.

[11]HSAA Eligible still refers to eligibility of the woman. Males' health outcomes are taken from the NFHS-3 men surveys; each woman is then matched to a male respondent in her household. For simplicity, I retain information about only one male respondent in each household.

Table A.9: Falsification Tests

	Women				Men			
	BMI			Pr(Anaemia)	BMI			Pr(Anaemia)
	All Sample	$BMI \leq 23$	Pr($BMI < 18.5$)	Moderate	All Sample	$BMI \leq 23$	Pr($BMI < 18.5$)	Moderate
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
	OLS	OLS	OLS	OLS	OLS	OLS	OLS	OLS
$\mathbb{I}$(Control cohorts) ×	0.382	-0.325	0.0163	-0.0457				
$\mathbb{I}$(Hindu, Buddhist, Sikh, Jain)	(0.493)	(0.353)	(0.0376)	(0.0425)				
HSAA Eligible					-0.197	-0.0291	0.0413	-0.0193
					(0.329)	(0.236)	(0.0443)	(0.0277)
N	56,760	38,853	56,760	53,756	33,711	25,074	33,711	31,557

*$p < 0.10$, **$p < 0.05$, ***$p < 0.01$. 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, religion-cohort fixed effects, and individual controls. Individual controls include women's years of schooling, number of children under 5 in the household, a household wealth index, and indicator variables for 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. $\mathbb{I}$(Control cohorts) is equal to one is a woman was 23 to 30 at the HSAA in her state and to zero if she was 30 or older.

Expenditure Patterns. The improvement in women's ability to inherit property introduced by the HSA amendments, and the consequent increase in their lifetime unearned income, could have an effect both on women's bargaining power and on household expenditure behavior. In order to assess whether I am indeed capturing the effect of a change in bargaining power, rather than a change in income and wealth, I investigate how exposure to HSAA by the wife of the head of the household affects expenditure patterns. To this aim, I use expenditure data from the 2011-2012 NSS Consumer Expenditure Survey. Firstly, I analyze the relationship between HSAA exposure and total expenditure, and expenditures on durable and non-durable goods, respectively. If HSA amendments determined an increase in overall expenditure, claiming existence of a causal link between women's intra-household bargaining power and their health would be problematic. I then analyze the effect of exposure to HSA amendments on food expenditure shares. My motivation is twofold. On one hand, this type of analysis may provide additional evidence of the effect of exposure to the reforms on the household expenditure levels: if HSAA exposure increased household total expenditure, we would expect the budget share of food to decrease as a consequence of the standard view that, being a necessity, food has an income elasticity less than one. On the other hand, it allows to test whether modeling Indian household in a unitary fashion is a sensible choice. Rejection of the unitary model - which assumes that the household acts as a single decision unit maximizing a common utility function - would provide support to the decision of modeling Indian household in a collective fashion.[12] Exogenous changes in distribution factors, i.e., factors affecting bargaining power but not preferences, are frequently used in the literature to test for the pooling of resources by household members. In the spirit of Attanasio and Lechene (2002, 2010), I investigate the effect of HSAA exposure on food budget shares. Any significant effect would provide evidence against the unitary model.

The empirical specification is as in equation (A.1), where y_{irsc} is the logarithm of total household i

[12]There is a large consensus in the literature that a model should account for household members having different preferences and for the intra-household distribution of powers. However, evidence against the unitary model does not automatically imply evidence in favor of the collective model.

Table A.10: HSA Amendments and Expenditure Patterns

	Expenditure (log)			Food Budget
	Total	Non Durables	Durables	Share
	(1)	(2)	(3)	(4)
	OLS	OLS	OLS	OLS
HSAA Eligible	0.0400	0.0322	0.299	1.250*
	(0.0406)	(0.0394)	(0.199)	(0.722)
Total Expenditure (log)				-7.231***
				(0.220)
N	65,090	65,090	65,090	65,090

*$p < 0.10$, **$p < 0.05$, ***$p < 0.01$. NSS 68^{th} Round Consumer Expenditure data. 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, religion-cohort fixed effects, and household controls. Household controls include women's and men's highest education, household size, head of household's age and indicator variables for land ownership, for living in rural areas, for being part of Scheduled Castes, Scheduled Tribes or Other Backward Classes, and for the presence of a salary earner in the household. 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. Robust standard errors in parentheses. Standard errors clustered at the primary sampling unit (village) level (3,753). Sampling weights applied. Budget shares are multiplied by 100.

expenditure, of expenditure on non-durable goods, of expenditure on durable goods, or the budget share devoted to food or other selected commodities, respectively. Now, $HSAA\ Eligible_{rsc}$ is the interaction between an indicator variable equal to one if the wife of the head of the household was 14 or younger at the time of the HSA Amendment in her state and to zero if she was 23 or older, and an indicator variable for being Hindu, Buddhist, Sikh or Jain. X'_{irsc} is a vector of controls for household i, of religion r, living in state s, with the wife of the head of the household born in year c. As standard in the literature, when looking at the effect of HSAA exposure on food budget shares, I condition the model on the logarithm of total expenditure. Due to data availability, I only estimate an intent-to-treat effect. Standard errors are clustered at the first sampling unit level (2001 Census villages in rural areas and 2007-2012 Urban Frame Survey blocks in urban areas).

Table A.10 reports the estimation results of the effect of HSAA eligibility of the wife of the head of household on expenditure patterns. All specifications include individual controls and religion-state,

state-cohort and cohort-religion fixed effects. Column 1 to 3 show the estimated coefficients of the intent-to-treat effect on the logarithm of total expenditure, expenditure on non-durable goods and durable goods respectively. Expenditure data refer to the month prior to the survey interview. I find no statistically significant effect on expenditure levels. Columns 4 report the estimation results of the model in equation (A.1) where the outcome variable is the budget share on food. I find that HSAA eligible households devote more resources to food, which is in line with the findings in Attanasio and Lechene (2002, 2010) on the effect on PROGRESA transfers to women in Mexico. These results represent an empirical rejection of the unitary model of the household and of the income pooling hypothesis. In such a framework, HSAA eligibility could not significantly affects how expenditure is allocated to goods, as in unitary households “only household income matters for choice outcomes and not the source of the income" (Browning et al. (2014)).

A.4 A Fully Specified Collective Model of Indian Households

The empirical estimation of resource shares assumes Engel curves for the private assignable goods to be linear in $\ln y$ and requires resource shares to be independent of y. However, it does not invoke restrictions on other goods' demand functions. I here provide a fully specified household model that delivers linearity of Engel curves and resource shares that are independent of y. $h = (h^1, ..., h^K)$ 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 $t = f, m, c$.

Assume each household member of type t, $t = f, m, c$, has utility given by Muellbauer's Price Independent Generalized Logarithmic (Piglog) model. In this framework, the indirect utility functions can be written as follows:

$$v\left(\frac{y}{G_t(p)}\right) + F_t(p) = \ln\left[\ln\left(\frac{y}{G_t(p)}\right)\right] + F_t(p) \tag{A.3}$$

with $F_t(p) = p_t \exp(-a' \ln \tilde{p})$, for some constant vector a with elements a_k summing to one. Let $\tilde{p}$ contain the prices of all goods except the prices of the private assignable goods for adult females, adult males and children, p_f, p_m, p_c. The assumption of assigning the same utility function to members of the same type is data driven and constrained by the number of different assignable goods that can be observed. For simplicity, consider the transformation $\psi_t(v + F_t, \tilde{p}) = \exp(v + F_t)$, which not depending upon $\tilde{p}$ makes the individual Engel curves for all goods to be the same as those of the private assignable goods. While this assumption is not invoked in the estimation of the system of Engel curves in equation (1.6), it simplifies the derivation of the full model. Let $G_t(p) = G_t(p_t, \tilde{p})$, which makes explicit the assumption that the, for example, the price p_m of the good that is assignable to the father does not appear in the woman's utility function, and hence does not appear in $G_f(p_f, \tilde{p})$. The indirect utility function for individuals of type t is given by:

$$V_t = \ln\left(\frac{y}{G_t(p_t, \tilde{p})}\right) \exp\left(p_t \exp(-a' \ln \tilde{p})\right) \tag{A.4}$$

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

$$\tilde{U}(U_f, U_m, U_c, p/y) = \sum_{t \in \{f,m,c\}} \mu_t(p/y)[U_t + \xi_t(p/y)] \tag{A.5}$$

where $\mu_t(p/y)$ is the *Pareto weight* and $\xi_t(p/y)$ are *externality functions* capturing the extent to which one's consumption affect the other member of the household. μ_t and ξ_t must be homogeneous of degree zero, so that $\mu_t(p/y) = \mu_t(p)$ and $\xi_t(p/y) = \xi_t(p)$ (Dunbar et al. (2013)). Pareto weights are traditionally interpreted as measures of intra-household bargaining power. The larger the value of μ_t, the greater is the weight that type t members' preferences receive in the household program.

Let A be the matrix defining the extent of joint consumption in each household and characterizing the consumption technology function. In other words, the consumption technology function converts purchased quantities by the household, h, in *private good equivalents*, x. While the general framework allows for a non-diagonal A, I here assume it to be diagonal. Non-zero off diagonal elements of A may arise when the extent to which one good is shared depends upon other goods, e.g.,, if leisure time is a consumption good, then the degree to which a motorcycle use is shared may depend on the time involved. A *private* good is defined as a good with its corresponding diagonal element of A equal to 1, and all off-diagonal elements in that row or column are equal to 0. Each household solves the following program:

$$\max_{x_f,x_m,x_c,h} \xi(p) + \sum_{t\in\{f,m,c\}} \mu_t(p)U_t \quad \text{subject to} \quad h = A(Fx_f + Mx_m + Cx_c)$$
$$y = h'p \tag{A.6}$$

where $\xi(p) = \sum_{t\in\{f,m,c\}} \mu_t(p)\xi_t(p)$ and F, M and C are the total number of adult women, adult men and children in the household, respectively. Let λ_t the *resource share* for each individuals of type t, such that $\lambda_t = \sum_k \frac{A_k p_k x_{tk}}{y}$. This program can be decomposed in two steps: 1) the optimal allocation of resources across members and 2) the individual maximization of their own utility function. This results follows directly from the second welfare theorem, as any Pareto efficient allocation can be implemented as an equilibrium of this economy, possibly after lump sum transfers between members. Conditional on knowing λ_t, each household member chooses x_t as the bundle maximizing U_t subject to the Lindahl type shadow budget constraint $\sum_k A_k p^k x_t^k = \lambda_t y$. Substituting the indirect utility functions $V_t(A'p, \lambda_t y)$ in the program above simplifies the optimization problem to the choice of optimal resource shares subject to the constraint that total resources shares must sum to one. The household program can be therefore written as:

$$\max_{\lambda_f,\lambda_m,\lambda_c} \xi(p) + \sum_{t\in\{f,m,c\}} \mu_t(p)V_t(A'p, \lambda_t(p/y)y) \quad \text{subject to} \quad F\lambda_f(p/y) + M\lambda_m(p/y) + C\lambda_c(p/y) = 1 \tag{A.7}$$

and, given the functional form assumptions discussed above, as:

$$\max_{\lambda_f,\lambda_m,\lambda_c} \xi + \sum_{t\in\{f,m,c\}} \tilde{\mu}_t \left[\ln \frac{\lambda_t y}{G_t}\right] \quad \text{subject to} \quad \Lambda_f + \Lambda_m + \Lambda_c = 1 \tag{A.8}$$

with $\tilde{\mu}_t(p) = \mu_t(p)\exp(A_t p_t \exp(-a'(\ln\tilde{p} + \ln\tilde{A})))$. The first order conditions of the Lagrangean are as follows:

$$\begin{cases} \frac{\tilde{\mu}_f}{\Lambda_f} = \frac{\tilde{\mu}_m}{\Lambda_m} = \frac{\tilde{\mu}_c}{\Lambda_c} = \nu \\ \\ \Lambda_f + \Lambda_m + \Lambda_c = 1 \end{cases} \tag{A.9}$$

with ν being the Lagrange multiplier. Explicit formulas for resource shares can be derived from the solution of the system of equations above. For individuals of type $t = f, m, c$, the overall resource shares are defined as

$$\Lambda_t(p/y) = \Lambda_t(p) = \frac{\tilde{\mu}_t(p)}{\sum_{t\in\{f,m,c\}} \tilde{\mu}_t(p)} \tag{A.10}$$

Assuming equal sharing among individuals of the same type, individual resources shares for type t household members are defined as

$$\lambda_t(p) = \frac{\tilde{\mu}_t(p)/T}{\sum_{t\in\{f,m,c\}} \tilde{\mu}_t(p)} \tag{A.11}$$

with $T = F, M, C$. As required for identification, these explicit formulas for the resource shares do not depend on y.

Each household member chooses x_t as the bundle maximizing U_t subject to the constraint $\sum_k A_k p^k x_t^k = \lambda_t y$. The individuals' demand function for good k, $k = 1, ..., K$, can be obtained by Roy's identity to equation (A.4), with prices equal to Ap_k and income equal to $\lambda_t y$:

$$\begin{aligned} x_t^k(\lambda_t y, A'p) &= -\frac{\partial V_t(\lambda_t y, A'p)/\partial Ap_k}{\partial V_t(\lambda_t y, A'p)/\partial \lambda_t y} \\ &= \lambda_t y \left[\frac{\partial G_t}{\partial Ap_k} \frac{1}{G_t} - \frac{\partial (Ap_t \exp(-a' \ln \tilde{p}))}{\partial Ap_k} \left(\ln(\lambda_t y) - \ln G_t \right) \right] \\ &= \frac{\partial G_t}{\partial Ap_k} \frac{\lambda_t y}{G_t} - \frac{\partial (Ap_t \exp(-a' \ln \tilde{p}))}{\partial Ap_k} \left[\ln(\lambda_t y) - \ln G_t \right] \lambda_t y \end{aligned} \tag{A.12}$$

for each good k and any individual t. Using the consumption technology function, the household demand function for a generic good k, as fraction of the total expenditure, is then defined as:

$$\frac{h_k(p,y)}{y} = A_k \sum_{t\in\{f,m,c\}} \frac{T x_t^k(\lambda_t(p)y, A'p)}{y} \tag{A.13}$$

Equation (A.12) can be rewritten as:

$$x_t^k(\lambda_t y, A'p) = \lambda_t y \alpha_t^k + \lambda_t y \beta_t^k \ln(\lambda_t y) \tag{A.14}$$

The overall demand of good k from individuals of type t generated from the solution of the decentralized optimization problems is given by:

$$T x_t^k(\lambda_t(p)y, A'p) = \Lambda_t y \alpha_t^k + \Lambda_t y \beta_t^k (\ln y + \ln \Lambda_t - \ln T) \tag{A.15}$$

and the corresponding budget share is:

$$\frac{T x_t^k(\lambda_t(p)y, A'p)}{y} = \Lambda_t(p)\alpha_t^k(p) + \Lambda_t(p)\beta_t^k(p)(\ln y + \ln \Lambda_t(p) - \ln T) \tag{A.16}$$

where $T = F, M, C$, $t = f, m, c$, and $\Lambda_t = T\lambda_t$.

An Engel curve represents the relationship between a budget share and total expenditure, holding prices constant. The Engel curve for a generic good k, under the explicit assumptions made above, takes the following form:

$$\frac{h_k(y)}{y} = A_k \sum_{t\in\{f,m,c\}} [\Lambda_t \alpha_t^k + \Lambda_t \beta_t^k (\ln y + \ln \Lambda_t - \ln T)] \tag{A.17}$$

and for private assignable goods it simplifies to:

$$W_t(y) = \frac{h_t(y)}{y} = \Lambda_t \alpha_t + \Lambda_t \beta_t (\ln y + \ln \Lambda_t - \ln T) \tag{A.18}$$

A.5 Robustness Checks

Table A.11: Determinants of Women's Resource Shares: Expenditure On Non-Durables

	Adult Women's Resource Share		
	All Households		
	With & Without Children < 15	With Children < 15 Only	Withou Children < 15 Only
	(1) NLSUR	(2) NLSUR	(3) NLSUR
No. Adult Women	0.0406***	0.0304***	0.0574***
	(0.00415)	(0.00413)	(0.00965)
No. Adult Men	-0.0230***	-0.0195***	-0.0340***
	(0.00299)	(0.00310)	(0.00671)
No. Children	-0.00645***	0.000965	
	(0.00206)	(0.00194)	
Fraction of Female Children	0.0158***	0.0133***	
	(0.00531)	(0.00456)	
𝕀(Daughter in Law)	0.0101	0.00721	0.0276
	(0.00622)	(0.00593)	(0.0184)
𝕀(Unmarried Daughter above 15)	-0.00380	0.00723	0.00606
	(0.00670)	(0.00659)	(0.0167)
𝕀(Widow)	-0.0206***	-0.0337***	-0.0156
	(0.00780)	(0.00830)	(0.0184)
𝕀(HSAA Eligible)	0.00766**	0.00862**	0.0201**
	(0.00369)	(0.00430)	(0.00924)
𝕀(Hindu, Buddhist, Sikh, Jain)	-0.0393***	-0.0161**	-0.0179
	(0.00908)	(0.00684)	(0.0156)
𝕀(SC, ST, Other Backward Caste)	0.0554***	0.0603***	0.0514***
	(0.00757)	(0.00763)	(0.0126)
𝕀(Salary Earner)	-0.0229***	-0.0180***	-0.00902
	(0.00444)	(0.00409)	(0.0101)
𝕀(Land Ownership)	0.00805	0.0126*	-0.0333*
	(0.00814)	(0.00752)	(0.0195)
𝕀(Female Higher Education)	0.0336***	0.0272***	0.0333**
	(0.00809)	(0.00748)	(0.0162)
𝕀(Male Higher Education)	0.0285***	0.0401***	0.0785***
	(0.00539)	(0.00581)	(0.0127)
𝕀(Rural)	-0.0319***	-0.0277***	-0.0404***
	(0.00646)	(0.00567)	(0.0120)
Avg. Age Diff. (Men 15 to 79 - Women 15 to 79)	-0.0153	-0.109***	0.0209
	(0.0385)	(0.0405)	(0.0842)
Avg. Age Women 15 to 79	-0.677	-0.618	-1.886
	(0.565)	(0.646)	(1.256)
(Avg. Age Diff. (Men 15 to 79 - Women 15 to 79))^2	-0.207**	0.0918	-0.443**
	(0.104)	(0.111)	(0.206)
(Avg. Age Women 15 to 79)^2	1.218	1.686	3.739
	(1.356)	(1.613)	(3.017)
(Avg. Age Diff. (Men 15 to 79 - Women 15 to 79))^3	0.0655	0.507	-0.527
	(0.485)	(0.599)	(0.880)
(Avg. Age Women 15 to 79)^3	-0.579	-1.497	-2.210
	(1.046)	(1.307)	(2.334)
Avg. Age Children 0 to 14	-0.0196	-0.0274	
	(0.0451)	(0.0558)	
𝕀(North)	-0.0803***	-0.0869***	-0.0858***
	(0.0137)	(0.0146)	(0.0241)
𝕀(East)	-0.0190	-0.0240	-0.0405
	(0.0145)	(0.0154)	(0.0262)
𝕀(North-East)	0.00953	0.0633***	0.169***
	(0.0184)	(0.0224)	(0.0293)
𝕀(South)	-0.0208	-0.0193	-0.0427
	(0.0143)	(0.0157)	(0.0263)
Constant	0.402***	0.315***	0.757***
	(0.0812)	(0.0872)	(0.173)
N	73,709	47,198	26,546
LL	-584,345.9	-378,427.1	-190,331.4
No. Parameters	318	318	188

$p < 0.10$, $**p < 0.05$, $***p < 0.01$. 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.

Table A.12: Determinants of Women's Resource Shares: Food Engel Curve Excluded

	Adult Women's Resource Share		
	All Households		
	With & Without Children < 15	With Children < 15 Only	Without Children < 15 Only
	(1) NLSUR	(2) NLSUR	(3) NLSUR
No. Adult Women	0.0380***	0.0331***	0.0618***
	(0.00558)	(0.00461)	(0.00951)
No. Adult Men	-0.0245***	-0.0229***	-0.0331***
	(0.00395)	(0.00357)	(0.00644)
No. Children	-0.00235	0.00377	
	(0.00281)	(0.00238)	
Fraction of Female Children	0.0133*	0.0133**	
	(0.00731)	(0.00539)	
$\mathbb{I}$(Daughter in Law)	0.0123	0.00859	0.0220
	(0.00838)	(0.00700)	(0.0175)
$\mathbb{I}$(Unmarried Daughter above 15)	-0.000271	0.0101	-0.00609
	(0.00906)	(0.00782)	(0.0165)
$\mathbb{I}$(Widow)	-0.00533	-0.0362***	-0.0239
	(0.0104)	(0.00959)	(0.0177)
$\mathbb{I}$(HSAA Eligible)	0.0161***	0.0119**	0.0224**
	(0.00432)	(0.00505)	(0.00978)
$\mathbb{I}$(Hindu, Buddhist, Sikh, Jain)	-0.0174	-0.0186**	-0.0106
	(0.0109)	(0.00802)	(0.0150)
$\mathbb{I}$(SC, ST or Other Backward Caste)	0.0383***	0.0682***	0.0589***
	(0.00905)	(0.00851)	(0.0125)
$\mathbb{I}$(Salary Earner)	-0.0202***	-0.0192***	-0.0169*
	(0.00613)	(0.00489)	(0.00972)
$\mathbb{I}$(Land Ownership)	-0.00557	0.00916	-0.0191
	(0.0116)	(0.00919)	(0.0182)
$\mathbb{I}$(Female Higher Education)	0.0360***	0.0316***	0.0385**
	(0.0111)	(0.00880)	(0.0155)
$\mathbb{I}$(Male Higher Education)	0.0174**	0.0415***	0.0796***
	(0.00703)	(0.00651)	(0.0121)
$\mathbb{I}$(Rural)	-0.0381***	-0.0312***	-0.0398***
	(0.00916)	(0.00675)	(0.0115)
Avg. Age Diff. (Men 15 to 79 - Women 15 to 79)	0.0153	-0.127***	0.00661
	(0.0513)	(0.0471)	(0.0785)
Avg. Age Women 15 to 79	-0.441	-0.727	-2.089*
	(0.724)	(0.768)	(1.139)
(Avg. Age Diff. (Men 15 to 79 - Women 15 to 79))^2	-0.170	0.112	-0.455**
	(0.140)	(0.134)	(0.193)
(Avg. Age Women 15 to 79)^2	0.625	2.005	4.094
	(1.709)	(1.927)	(2.693)
(Avg. Age Diff. (Men 15 to 79 - Women 15 to 79))^3	-0.0328	0.576	-0.758
	(0.653)	(0.689)	(0.784)
(Avg. Age Women 15 to 79)^3	-0.0359	-1.796	-2.541
	(1.299)	(1.567)	(2.047)
Avg. Age Children 0 to 14	0.0260	-0.0392	
	(0.0636)	(0.0668)	
$\mathbb{I}$(North)	-0.0719***	-0.0972***	-0.0779***
	(0.0168)	(0.0167)	(0.0232)
$\mathbb{I}$(East)	-0.0121	-0.0226	-0.0170
	(0.0182)	(0.0184)	(0.0252)
$\mathbb{I}$(North-East)	0.00140	0.0588**	0.177***
	(0.0224)	(0.0247)	(0.0296)
$\mathbb{I}$(South)	-0.0225	-0.0302*	-0.0517**
	(0.0170)	(0.0181)	(0.0244)
Constant	0.381***	0.375***	0.751***
	(0.105)	(0.102)	(0.160)
N	73,642	47,159	26,518
LL	-31,1134.4	-19,5918.9	-85,244.5
No. Parameters	264	264	184

$p < 0.10$, $^{**}p < 0.05$, $^{***}p < 0.01$. 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 A.13: Determinants of Women's Resource Shares: Restricted Samples

	Adult Women's Resource Share			
	$F = M = 1, C = \{0, 1\}$			Married Couples
	With & Without Children < 15	With Children < 15 Only	Without Children < 15 Only	With Children < 15 Only
	(1) NLSUR	(2) NLSUR	(3) NLSUR	(4) NLSUR
𝕀(Child)	-0.0417			
	(0.0516)			
𝕀(Child is Female)	0.0367	0.0435**		
	(0.0339)	(0.0199)		
𝕀(Daughter in Law)	-0.0458	-0.0478	0.152	
	(0.225)	(0.153)	(0.140)	
𝕀(Unmarried Daughter above 15)	-0.0784	-0.0206	-0.0713	
	(0.144)	(0.184)	(0.0602)	
𝕀(Widow)	-0.00973	0.00999	-0.0138	
	(0.0624)	(0.0937)	(0.0266)	
𝕀(Two Children)				-0.00837
				(0.00736)
𝕀(Three Children)				0.00601
				(0.0100)
𝕀(Four Children)				0.00252
				(0.0146)
𝕀(Five Children)				0.0157
				(0.0234)
Fraction of Female Children				0.0151*
				(0.00798)
𝕀(HSAA Eligible)	0.0252*	0.0344*	0.000849	0.0292***
	(0.0130)	(0.0184)	(0.0136)	(0.00797)
𝕀(Hindu, Buddhist, Sikh, Jain)	0.000261	-0.0586*	0.0198	-0.00572
	(0.0297)	(0.0332)	(0.0152)	(0.0115)
𝕀(SC, ST or Other Backward Caste)	0.0419	0.0764**	-0.0367***	0.0839***
	(0.0321)	(0.0307)	(0.0139)	(0.0143)
𝕀(Salary Earner)	-0.00464	-0.0232	-0.00511	-0.0224***
	(0.0243)	(0.0214)	(0.0135)	(0.00737)
𝕀(Land Ownership)	0.0285	0.0120	0.0324*	-0.00435
	(0.0310)	(0.0259)	(0.0168)	(0.0116)
𝕀(Female Higher Education)	0.00935	-0.0189	-0.0324	0.0217
	(0.0424)	(0.0369)	(0.0209)	(0.0149)
𝕀(Male Higher Education)	-0.0103	0.0619**	0.0361**	0.0346***
	(0.0323)	(0.0291)	(0.0166)	(0.0108)
𝕀(Rural)	-0.0789**	-0.0687**	0.00408	-0.0348***
	(0.0314)	(0.0276)	(0.0114)	(0.00965)
Age Diff. (Man 15 to 79 - Woman 15 to 79)	0.0115	0.0987	-0.00672	0.0458
	(0.198)	(0.213)	(0.0910)	(0.0643)
Age Woman 15 to 79	0.0343	-0.861	-2.600**	-1.435
	(2.206)	(2.069)	(1.110)	(1.275)
(Age Diff. (Man 15 to 79 - Woman 15 to 79))^2	-0.0621	0.428	-0.198	-0.114
	(0.481)	(0.666)	(0.197)	(0.231)
(Age Woman 15 to 79)^2	0.125	2.569	5.039**	3.483
	(5.255)	(5.209)	(2.510)	(3.351)
(Age Diff. (Man 15 to 79 - Woman 15 to 79))^3	-0.162	-0.219	-0.675	0.844
	(1.383)	(1.795)	(0.564)	(1.247)
(Age Woman 15 to 79)^3	-0.262	-2.408	-3.413*	-2.474
	(4.010)	(4.185)	(1.795)	(2.837)
Age Child 0 to 14	0.0221	-0.0295		-0.227**
	(0.429)	(0.295)		(0.0970)
𝕀(North)	-0.0809*	-0.132**	0.113**	-0.0992***
	(0.0472)	(0.0545)	(0.0500)	(0.0226)
𝕀(East)	0.000901	0.000255	0.0258	-0.00730
	(0.0509)	(0.0604)	(0.0488)	(0.0237)
𝕀(North-East)	0.0133	0.0477	0.0474	0.0310
	(0.0625)	(0.0645)	(0.0678)	(0.0285)
𝕀(South)	-0.00921	-0.00950	-0.208***	0.0291
	(0.0468)	(0.0468)	(0.0468)	(0.0250)
Constant	0.359	0.519*	0.928***	0.500***
	(0.310)	(0.268)	(0.165)	(0.160)
N	11,138	4,168	6,975	32,511
LL	-92038.4	-32982.8	-47746.4	-274537.9
No. Parameters	294	294	172	

$p < 0.10$, $^{**}p < 0.05$, $^{***}p < 0.01$. 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.

A.6 Testing the Collective Framework

The theoretical framework presented in section 1.4 relies on households satisfying the assumptions regarding joint consumption and Pareto efficiency. Testing for these assumption in this context is equivalent to checking the validity of the Browning et al. (2013) (hereinafter, BCL) structure of household demand functions. BCL is a model of household demands, which are connected via the structural model to singles' demands. In my application, as in Dunbar et al. (2013), I do not impose the BCL assumptions regarding comparability of preferences of single versus couples. For this test, however, I consider imposing this additional assumption. I consider a *couple* a household with one adult man and one adult woman, independently on their actual marital status. I follow Dunbar et al. (2013) and I use additional data on singles to provide validation of these structural assumptions.[13] The BCL framework can be summarized as follows:

$$\begin{cases} W_t^{couple} = \lambda_t(\alpha_t + \beta_t \ln \lambda_t) + \lambda_t \beta_t \ln y \\ W_t^{alone} = a_t + b_t \ln y \end{cases} \tag{A.19}$$

for $t = f, m$. The restrictions imposing similarity of Engel curves required for identification constrain $b_m = b_f = \beta_m = \beta_f$. These restrictions imply two testable implications: 1) since λ_t cannot be negative, the slopes of men's and women's private assignable have the same sign; 2) the slopes of household demands must be proportional to those of singles' demands, with factors of proportionality that sum to 1.

To test implication 1, I compare the predicted slopes with respect to log-expenditure for men's and women's clothing obtained by estimating the model using observations for nuclear households without children only (see table A.13, column 3). Less than 1% of the predicted slopes are positive, both for women and for men. Moreover, the restriction that the slopes of men's and women's private assignable have the same sign is satisfied all the times. There are no households in which the slopes are of opposite

[13] For more details, see Dunbar et al. (2013) on-line Appendix, section A.5.1.

signs. To test implication 2, I combine a sample of 8,428 nuclear households without children with a sample of 5,918 singles from the NSS survey and estimate a linear SUR regressions of the men's and women's clothing budget share on the log of total expenditure and all demographic variables except those relating to other household members. I interact all regressors with a dummy for couples. The sum of the men's and women's ratios of slopes in couples versus single households is equal to 1.02 and not statistically different from 1.

That the BCL's assumptions on joint consumption and Pareto efficiency cannot be rejected provides an empirical validation of the theoretical framework of this paper.

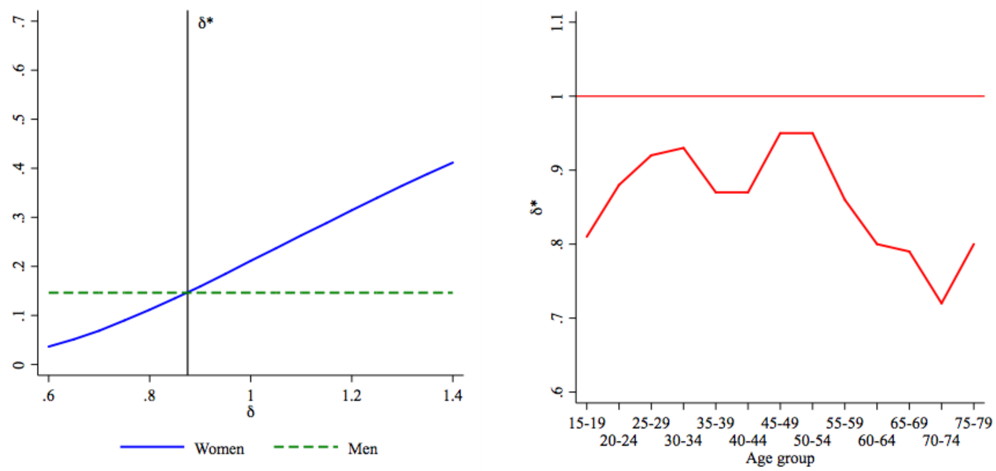
A.7 Simulation: Gender Specific Poverty Lines

It is standard among the development practitioners to assume poverty lines to be the same for men and women. While section 1.5.2 constrains poverty thresholds to be the same across genders, I here relax this assumption: while I fix the male poverty lines to the World Bank thresholds of $US 1.25/day and $US 2/day, I allow female poverty lines to be only proportional to that of males, with a factor of proportionality equal to δ. This exercise is equivalent to allowing women to have lower (or higher) basic expenditure needs relative to men. For $\delta = 1$, female and male poverty lines are the same and equal to the World Bank thresholds; for $\delta < 1$, the female poverty threshold is lower than that for males, while for $\delta > 1$ the poverty line for women is higher than for men.

I simulate gender and gender-age specific poverty rates in these alternative scenarios. In lights of the evidence of excess female poverty presented in section 1.5.2, the simulation exercise allows me to address concerns about that fact that women's needs need not be equal to men's and answer the following questions: How much lower should the poverty threshold for women be to obtain equal poverty rates across genders? Alternatively: How much higher should men's basic requirements be to rationalize excess female poverty? I define δ^* as the value of δ that equalizes poverty rates across genders, i.e., the fractions of women and men living in poverty.

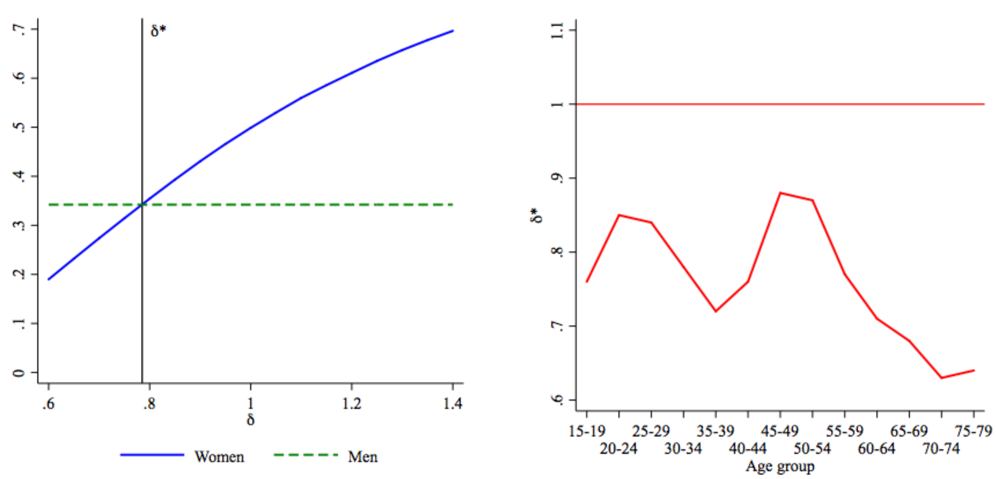
Figure A.11 shows the results for extreme poverty ($US 1.25/day), while figure A.12 presents the simulation results for average poverty ($US 2/day). When all age groups are pooled for the calculations (panels A), the simulation exercise indicates that female poverty lines equal to $US 1.08/day and $US 1.57/day would deliver extreme and average poverty rates that are equal across genders. The graphs displays the gender specific poverty estimates against the factor of proportionality δ. When gender-age specific poverty rates are taken into account (panels B), the simulation exercise indicate that at post-reproductive ages the female poverty threshold would need to be as low as less than 90 cents/day (70 percent of males) to equalize extreme poverty rates across gender and as low as $US 1.2/day to make the female and male average poverty rates the same. The graph shows the values δ^* equalizing female and male poverty rates in all thirteen 5-year age groups, from 15-19 to 75-79.

Investigating whether setting gender specific poverty lines makes sense from a policy perspective and whether men and women differ in their basic expenditure needs at different ages goes far beyond the scope of this paper. Future work, however, should focus on disentangling how much of excess female poverty is caused by discrimination and gender inequalities, and how much is driven by gender specific necessities and requirements, which may vary over the life cycle.



(A) Poverty Rates (B) δ^*

Figure A.11: Equalizing Male and Female Poverty (1.25 US$/day)



(A) Poverty Rates (B) δ^*

Figure A.12: Equalizing Male and Female Poverty (2 US$/day)

Appendix B

Women's Empowerment and Family Health: A Two-Step Approach

B.1 Additional Tables

Table B.1: Correlations: "NO say" Responses and Woman's Power ($\widehat{R}$)

Woman has NO say in ...	own health care	large hh pur-chases	visits to family	purchases to daily life	"NO say" Index	Woman's Power ($\widehat{R}$)
own health care	1.00					
large hh purchases	0.43	1.00				
visits to family	0.40	0.54	1.00			
purchases to daily life	0.45	0.51	0.49	1.00		
"NO say" Index	0.72	0.81	0.78	0.79	1.00	
Woman's Power ($\widehat{R}$)	-0.07	-0.08	-0.08	-0.12	-0.11	1.00

Table B.2: Women's Health, OLS Estimates

	Weight (kg)	Body Mass Index	Pr(BMI≤18.5)	Pr(Anemic)	Avg. Birth Space (months)
	(1)	(2)	(3)	(4)	(5)
Woman's Power ($\widehat{R}$)	-0.415	-0.811	0.0287	0.102^{**}	30.33^{***}
	(1.475)	(0.593)	(0.0611)	(0.0506)	(3.110)
Observations	16,513	16,498	16,498	15,525	14,003
Mean Dependent Variable	48.24	20.89	0.30	0.16	35.11

$^{*}p < 0.10$, $^{**}p < 0.05$, $^{***}p < 0.01$. Estimates are obtained using the NFHS-3 data. All specifications include state, cohort (woman), state-religion, and cohort-religion fixed effects, and state time trends. The sample includes married women of age 15 to 49 included in nuclear households with up to 4 children. All specifications include a religion dummy, equal to 1 if a woman is Hindu, Buddhist, Sikh or Jain, number of children below 15, the fraction of female children, woman's years of education, a household wealth index, indicator variables for owning a BPL card, for living in rural areas, and for being part of Scheduled Castes, Scheduled Tribes or Other Backward Classes, for having worked in the past 12 months. Anemia includes mild and moderate anemia. Bootstrap standard errors in parentheses (2,500 repetitions). Standard errors clustered at the primary sampling unit (village) level.

Table B.3: Men's Health, OLS Estimates

	Weight (kg)	Body Mass Index	Pr(BMI≤18.5)	Pr(Anemic)
	(1)	(2)	(3)	(4)
Woman's Power ($\ln \widehat{R}$)	1.188	0.818	0.00630	-0.0274
	(2.149)	(0.696)	(0.0802)	(0.0608)
Observations	8,536	8,533	8,533	7,789
Mean Dependent Variable	56.85	21.13	0.24	0.10

$^{*}p < 0.10$, $^{**}p < 0.05$, $^{***}p < 0.01$. Estimates are obtained using the NFHS-3 data. All specifications include state, cohort (woman), state-religion, and cohort-religion fixed effects, and state time trends. The sample includes married men of age 15 to 54 included in nuclear households with up to 4 children. All specifications include a religion dummy, equal to 1 if a woman is Hindu, Buddhist, Sikh or Jain, number of children below 15, the fraction of female children, man's age and years of education, a household wealth index, indicator variables for owning a BPL card, for living in rural areas, and for being part of Scheduled Castes, Scheduled Tribes or Other Backward Classes, for having worked in the past 12 months. Anemia includes mild and moderate anemia. Bootstrap standard errors in parentheses (2,500 repetitions). Standard errors clustered at the primary sampling unit (village) level.

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Table B.4: Children's Health, OLS Estimates

	Weight (kg)	Height (cm)	Pr(Any Vaccine)	Pr(BCG)	Pr(DPT)	Pr(Polio)
	(1)	(2)	(3)	(4)	(5)	(6)
Woman's Power ($\widehat{R}$)	-0.322	-1.422	-0.0443	-0.223***	-0.237***	-0.0541
	(0.404)	(1.535)	(0.0546)	(0.0746)	(0.0779)	(0.0559)
Observations	10,993	10,967	12,000	11,944	11,884	11,985
Mean Dependent Variable	10.67	83.94	0.91	0.78	0.75	0.91

*$p < 0.10$, **$p < 0.05$, ***$p < 0.01$. Estimates are obtained using the NFHS-3 data. All specifications include state, cohort (woman), state-religion, and cohort-religion fixed effects, and state time trends. The sample includes married women of age 15 to 49 included in nuclear households with up to 4 children. All specifications include a religion dummy, equal to 1 if a woman is Hindu, Buddhist, Sikh or Jain, number of children below 15, the fraction of female children, woman's years of education, a household wealth index, indicator variables for owning a BPL card, for living in rural areas, and for being part of Scheduled Castes, Scheduled Tribes or Other Backward Classes, for having a mother who worked in the past 12 months. Bootstrap standard errors in parentheses (5,000 repetitions). Standard errors clustered at the primary sampling unit (village) level.

Appendix C

Long-Term Effects of Access to Health Care: Medical Missions in Colonial India

C.1 Additional Figures and Tables

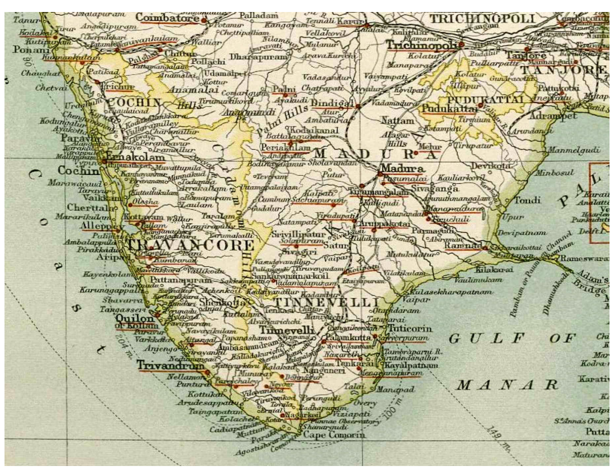
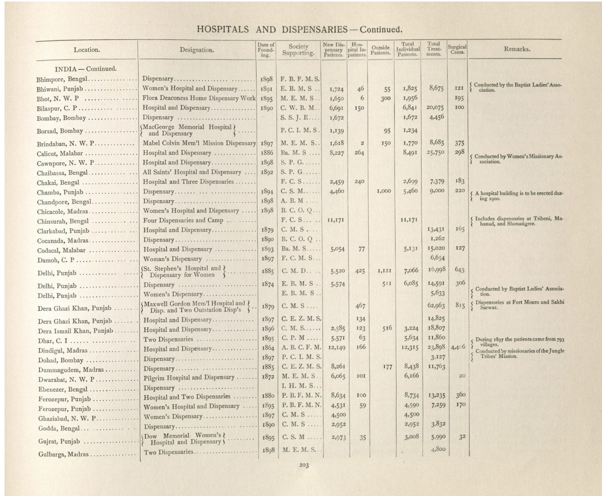


Figure C.1: Protestant missions locations as of 1910 (Statistical Atlas of Christian Missions)



HOSPITALS AND DISPENSARIES — Continued.

Location.	Designation.	Date of Founding.	Society Supporting.	New Dispensary Patients.	Hospital In-patients.	Outside Patients.	Total Individual Patients.	Total Treatments.	Surgical Cases.	Remarks.
INDIA — Continued.										
Bhimpore, Bengal	Dispensary	1898	F. B. F. M. S.							
Bhiwani, Punjab	Women's Hospital and Dispensary	1891	E. B. M. S	1,724	46	55	1,825	8,675	121	Conducted by the Baptist Ladies' Association.
Bhot, N. W. P	Flora Deaconess Home Dispensary Work	1895	M. E. M. S	1,650	6	300	1,956		195	
Bilaspur, C. P	Hospital and Dispensary	1890	C. W. B. M.	6,691	150		6,841	20,075	100	
Bombay, Bombay	Dispensary		S. S. J. E.	1,672			1,672	4,456		
Borsad, Bombay	MacGeorge Memorial Hospital and Dispensary		P. C. I. M. S.	1,139		95	1,234			
Brindaban, N. W. P	Mabel Colvin Mem'l Mission Dispensary	1897	M. E. M. S.	1,618	2	150	1,770	8,685	375	
Calicut, Malabar	Hospital and Dispensary	1886	Ba. M. S	8,227	264		8,491	25,750	298	
Cawnpore, N. W. P	Hospital and Dispensary	1898	S. P. G.							Conducted by Women's Missionary Association.
Chaibassa, Bengal	All Saints' Hospital and Dispensary	1892	S. P. G.							
Chakai, Bengal	Hospital and Three Dispensaries		F. C. S.	2,459	240		2,699	7,379	183	
Chamba, Punjab	Dispensary	1894	C. S. M.	4,460		1,000	5,460	9,000	220	A hospital building is to be erected during 1900.
Chandpore, Bengal	Dispensary	1898	A. B. M							
Chicacole, Madras	Women's Hospital and Dispensary	1898	B. C. O. Q.							
Chinsurah, Bengal	Four Dispensaries and Camp		F. C. S.	11,171			11,171			Includes dispensaries at Tribeni, Mahanad, and Shonatigree.
Clarkabad, Punjab	Hospital and Dispensary	1879	C. M. S					13,431	165	
Cocanada, Madras	Dispensary	1890	B. C. O. Q					1,262		
Codacal, Malabar	Hospital and Dispensary	1893	Ba. M. S.	5,054	77		5,131	15,020	127	
Damoh, C. P	Woman's Dispensary	1897	F. C. M. S.					6,654		
Delhi, Punjab	St. Stephen's Hospital and Dispensary for Women	1885	C. M. D.	5,520	425	1,121	7,066	16,998	643	
Delhi, Punjab	Dispensary	1874	E. B. M. S	5,574		511	6,085	14,591	306	
Delhi, Punjab	Women's Dispensary		E. B. M. S					5,633		Conducted by Baptist Ladies' Association.
Dera Ghazi Khan, Punjab	Maxwell Gordon Mem'l Hospital and Disp. and Two Outstation Disp's	1879	C. M. S		467			62,963	815	Dispensaries at Fort Monro and Sakhi Sarwar.
Dera Ghazi Khan, Punjab	Hospital and Dispensary	1897	C. E. Z. M. S.		134			14,825		
Dera Ismail Khan, Punjab	Hospital and Dispensary	1896	C. M. S.	2,585	123	516	3,224	18,807		
Dhar, C. I	Two Dispensaries	1895	C. P. M	5,571	63		5,634	11,860		During 1897 the patients came from 793 villages.
Dindigul, Madras	Hospital and Dispensary	1864	A. B. C. F. M.	12,149	166		12,315	23,898	4,406	
Dohad, Bombay	Dispensary	1897	P. C. I. M. S.					3,127		Conducted by missionaries of the Jungle Tribes' Mission.
Dummagudem, Madras	Dispensary	1885	C. E. Z. M. S.	8,261		177	8,438	11,763		
Dwarahat, N. W. P	Pilgrim Hospital and Dispensary	1872	M. E. M. S	6,065	101		6,166		20	
Ebenezer, Bengal	Dispensary		I. H. M. S.							
Ferozepur, Punjab	Hospital and Two Dispensaries	1880	P. B. F. M. N.	8,634	100		8,734	13,235	360	
Ferozepur, Punjab	Women's Hospital and Dispensary	1895	P. B. F. M. N.	4,531	59		4,590	7,259	170	
Ghaziabad, N. W. P	Women's Dispensary	1897	C. M. S	4,500			4,500			
Godda, Bengal	Dispensary	1890	C. M. S	2,952			2,952	3,832		
Gujrat, Punjab	Dow Memorial Women's Hospital and Dispensary	1895	C. S. M	2,973	35		3,008	5,990	32	
Gulbarga, Madras	Two Dispensaries	1898	M. E. M. S.					4,800		

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Figure C.2: Page from the Centennial Survey Survey of Foreign Missions

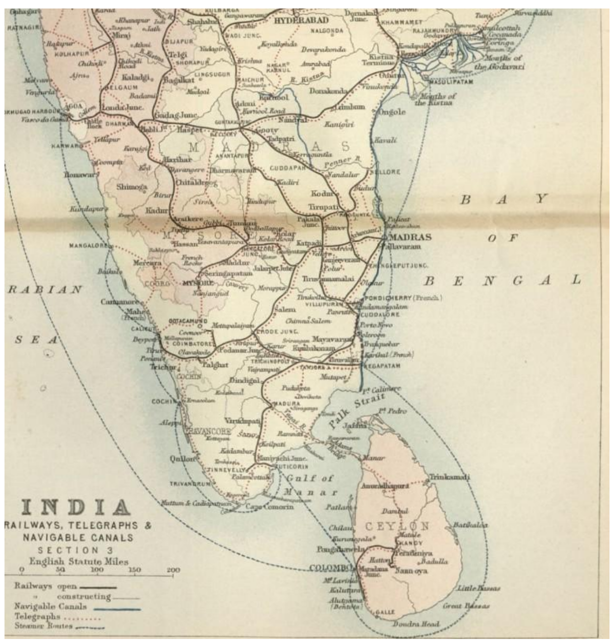

Figure C.3: Status of Indian railways network as of 1891 (Constable's Hand Atlas of India)

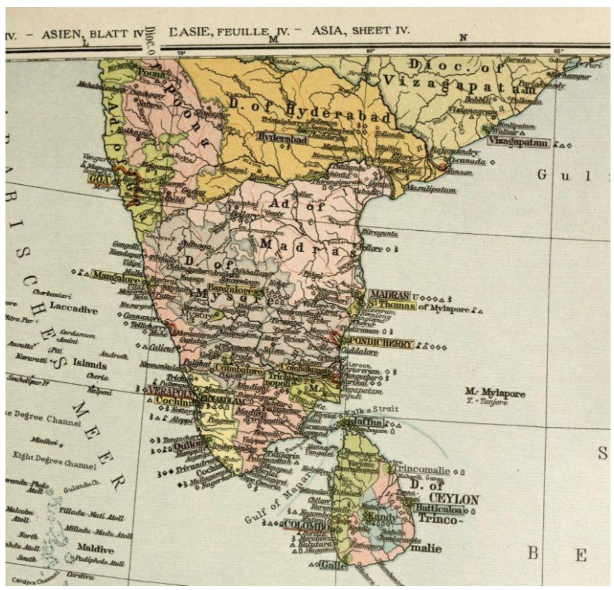

Figure C.4: Catholic missions locations as of 1911 (Atlas Hierarchicus)

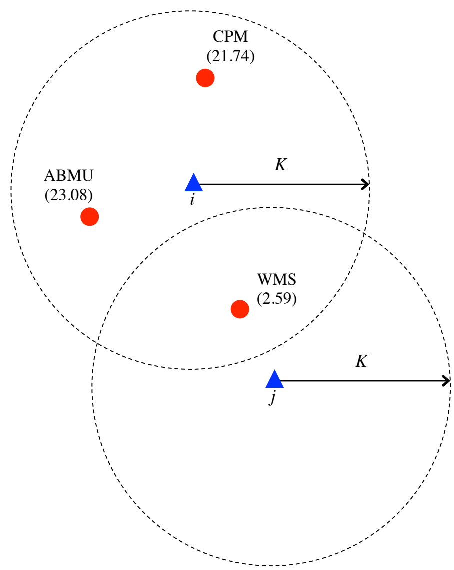

Figure C.5: Instrumental Variable: An Illustrative Example

Note: Consider the case of two individuals (i and j) and three Protestant medical missions associated with three different missionary societies: Canadian Presbyterian Mission (CPM, with *Society Hospital Share* = *21.74%*), American Baptist Missionary Union (ABMU, 23.08%), and Wesleyan Missionary Society (WMS, 2.59%). K is the buffer radius (in kilometers). For individual i, $\textit{Hospital share}_i^K = 21.74 + 23.08 + 2.59 = 47.41$. For individual j, $\textit{Hospital share}_j^K = 2.59$.

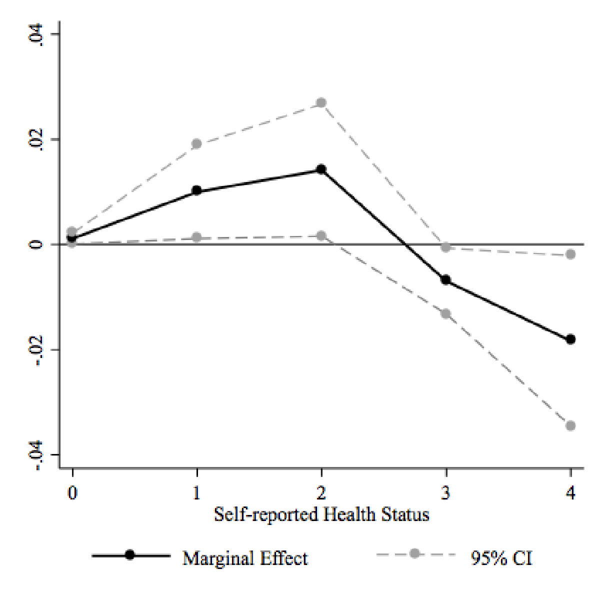

Figure C.6: Self-reported Health: Ordered Probit Marginal Effects

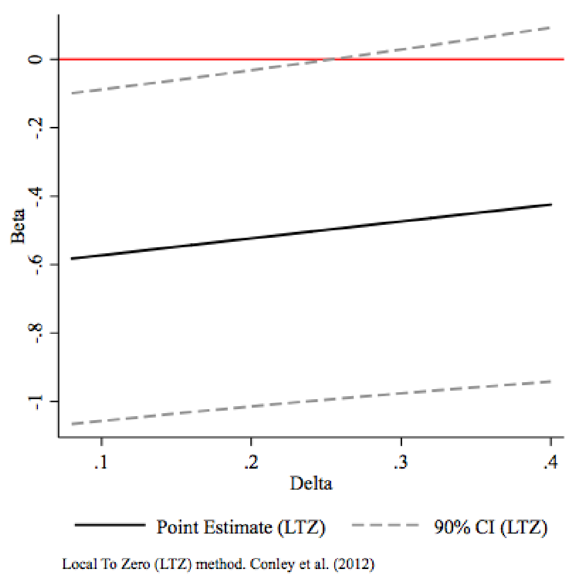

Figure C.7: Estimated β by Direct Effect of Instrument

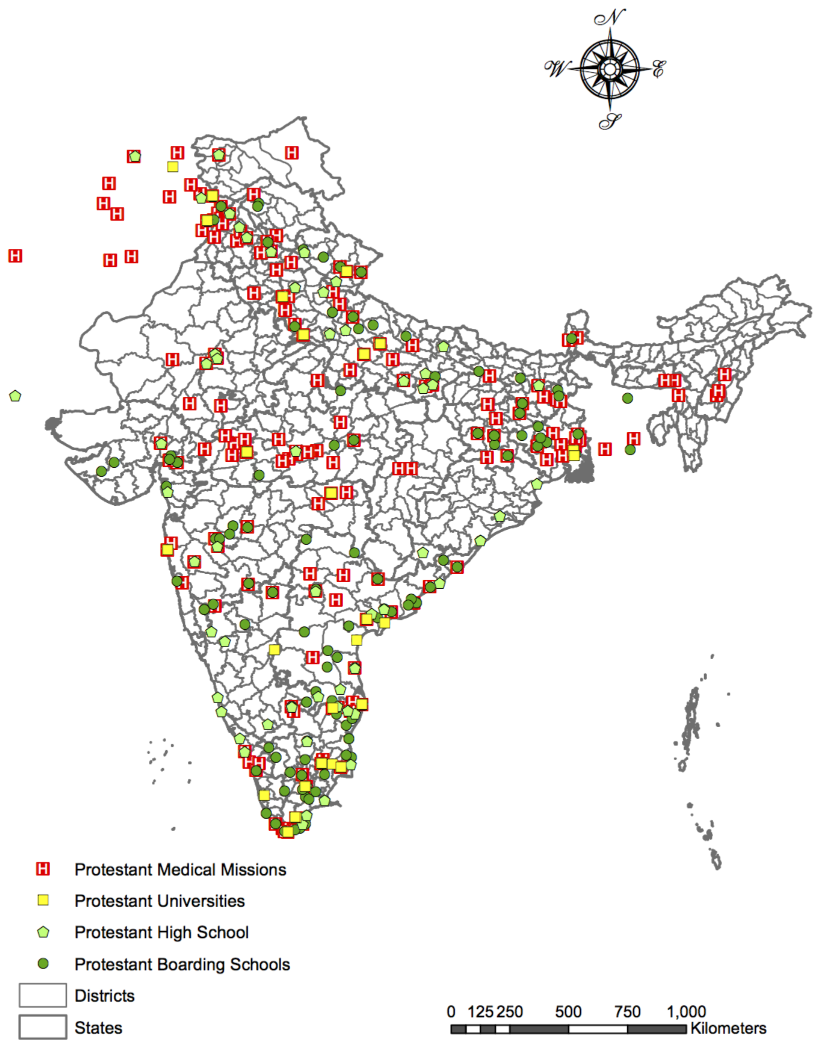

Figure C.8: Protestant Medical Missions and Schools

Table C.1: Missionary Societies and Hospital Shares

Society	Origin	*Society hospital share*
Canadian Presbyterian Mission	Canada	21.74%
Church of England Zenana Missionary Society	England	3.08%
Cambridge Mission to Delhi	England	0.00%
Church Missionary Society	England	8.87%
English Baptist Missionary Society	England	0.87%
English Presbyterian Church Mission	England	69.23%
Friends' Foreign Missions Association	England	22.22%
London Missionary Society	England	19.39%
Society for the Propagation of Gospel	England	1.60%
Society of St John the Evangelist	England	0.00%
Wesleyan Missionary Society	England	2.59%
Zenana, Bible and Medical Missions	England	0.00%
Basel Evangelical Missionary Society	Germany	7.14%
Independent	Independent	58.82%
Poona and Indian Village Mission	India	0.00%
Ranagat Medical Mission	India	0.00%
Bethel Santhal Mission	India	0.00%
Presbyterian Church of Ireland Missionary Society	Ireland	21.05%
Church of Scotland Mission	Scotland	21.05%
Edinburgh Medical Missionary Society	Scotland	66.67%
Free Church of Scotland	Scotland	26.83%
United Presbyterian Church of Scotland	Scotland	7.83%
American Board for Commissioners of Foreign Missions	United States	29.59%
American Baptist Missionary Union	United States	23.08%
Free Baptist Foreign Missionary Society	United States	0.00%
Methodist Episcopal Missionary Society	United States	18.66%
Presbyterian Board of Foreign Missions North	United States	42.34%
Woman's Union Missionary Society	United States	14.29%
Welsh Calvinist Methodist Foreign Missions Society	Wales	0.00%

Society hospital share is the ratio between each society's number of medical missions and total missions outside India. Only societies included in the estimation sample. Data from the Centennial Survey of Foreign Mission (1902).

Table C.2: OLS Estimates (Log-Log Model)

	BMI (log)				
	OLS (1)	OLS (2)	OLS (3)	OLS (4)	OLS (5)
Distance from Protestant medical mission (log)	-0.0119***	-0.0166***	-0.0110***	-0.0110***	-0.0110***
	(0.00240)	(0.00447)	(0.00416)	(0.00413)	(0.00416)
𝟙(Female)			-0.00653*	-0.00673*	-0.00673*
			(0.00358)	(0.00359)	(0.00387)
𝟙(Married)			0.00448	0.00467	0.00467
			(0.00472)	(0.00472)	(0.00436)
𝟙(Primary school completed)			0.0132***	0.0130***	0.0130***
			(0.00409)	(0.00410)	(0.00403)
𝟙(Major ethnic group)			-0.00561	-0.00446	-0.00446
			(0.00482)	(0.00482)	(0.00484)
Age			0.00723***	0.00720***	0.00720***
			(0.00102)	(0.00102)	(0.00103)
Age^2			-0.0000689***	-0.0000685***	-0.0000685***
			(0.0000129)	(0.0000129)	(0.0000130)
Household size			-0.00697**	-0.00703**	-0.00703**
			(0.00314)	(0.00314)	(0.00293)
Household size^2			0.000371*	0.000380*	0.000380**
			(0.000203)	(0.000203)	(0.000190)
𝟙(Wealth 2^{nd} quartile)			0.0134***	0.0133***	0.0133***
			(0.00443)	(0.00443)	(0.00440)
𝟙(Wealth 3^{rd} quartile)			0.0249***	0.0250***	0.0250***
			(0.00439)	(0.00437)	(0.00430)
𝟙(Wealth 4^{th} quartile)			0.0497***	0.0499***	0.0499***
			(0.00507)	(0.00510)	(0.00483)
Population in 2000 (000s)				0.000358	0.000358
				(0.000249)	(0.000307)
Altitude (meters)				0.0000359*	0.0000359*
				(0.0000203)	(0.0000204)
No. rivers/water sources				-0.00136	-0.00136
				(0.00168)	(0.00169)
Latitude				0.00890	0.00890
				(0.0100)	(0.0101)
Longitude				-0.0110	-0.0110*
				(0.00700)	(0.00639)
Population in 1900 (000s)				-0.00000119	-0.00000119
				(0.0000397)	(0.0000639)
Cropland in 1900 (km^2)				0.0000555	0.0000555
				(0.0000825)	(0.0000862)
No. colonial railways				-0.0142*	-0.0142*
				(0.00810)	(0.00749)
N	7,046	7,046	7,046	7,046	7,046
Mean Dependent Variable	3.05	3.05	3.05	3.05	3.05
District FE	No	Yes	Yes	Yes	Yes
Individual controls	No	No	Yes	Yes	Yes
Geographical controls	No	No	No	Yes	Yes
Historical controls	No	No	No	Yes	Yes
Spatial standard errors	No	No	No	No	Yes

Only individuals of age 20 to 60 and BMI 15 to 30 are included in the estimation sample. Robust standard errors in parentheses. * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$. Standard errors clustered at the village (ward) level in columns 1 to 4. Standard errors corrected for spatial correlation in column 5.

Table C.3: Robustness Check: Estimation Sample

	Age	BMI			
	15-65	18.5-30	15-35	15-45	25-45
	OLS (1)	OLS (2)	OLS (3)	OLS (4)	OLS (5)
Distance from Protestant medical mission (log)	-0.173* (0.105)	-0.229** (0.108)	-0.248* (0.129)	-0.239* (0.134)	0.325 (0.288)
N	7,883	4,966	7,155	7,174	699
Mean dependent variable	20.36	20.80	20.60	20.65	27.95
District FE	Yes	Yes	Yes	Yes	Yes
Individual controls	Yes	Yes	Yes	Yes	Yes
Geographical controls	Yes	Yes	Yes	Yes	Yes
Historical controls	Yes	Yes	Yes	Yes	Yes
Distance from Protestant mission (log)	Yes	Yes	Yes	Yes	Yes
Distance from Catholic mission (log)	Yes	Yes	Yes	Yes	Yes

In column 1, only individuals with BMI 15 to 30 are included in the estimation sample. In column 2 to 5, only individuals of age 20 to 60 are included in the estimation sample. Robust standard errors in parentheses. * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$. Standard errors clustered at the village (ward) level.

Table C.4: Robustness Check: Instrument

	Distance from Protestant medical mission (log)			BMI		
				10km (med. only)	30km (med. only)	50km (all miss.)
	OLS (1)	OLS (2)	OLS (3)	2SLS (4)	2SLS (5)	2SLS (6)
Hospital share 10 km (medical only)	-2.103*** (0.527)					
Hospital share 30 km (medical only)		-1.319*** (0.343)				
Hospital share 50 km (all missions)			-0.143*** (0.0430)			
Distance from Protestant medical mission (log)				-0.445* (0.265)	-0.584** (0.275)	-0.771* (0.406)
Kleibergen-Paap rk Wald F statistic	-	-	-	15.93	14.77	10.99
N	7,046	7,046	7,046	7,046	7,046	7,046
Mean dependent variable	4.02	4.02	4.02	20.42	20.42	20.42
District FE	Yes	Yes	Yes	Yes	Yes	Yes
Individual controls	Yes	Yes	Yes	Yes	Yes	Yes
Geographical controls	Yes	Yes	Yes	Yes	Yes	Yes
Historical controls	Yes	Yes	Yes	Yes	Yes	Yes

Only individuals of age 20 to 60 and BMI 15 to 30 are included in the estimation sample. Robust standard errors in parentheses. * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$. Standard errors clustered at the village (ward) level.

Table C.5: Robustness Check: Societies With and Without Medical Missionaries

	Non-Medical (1)	Medical (2)	Difference (3)
Date of Organization	1876	1857	19***
			(5.44)
No. Missionary Stations	44.50	220.34	-175.84***
			(-4.45)
Income from Home Sources per Station (US$)	1,195.42	1346.60	-151.18
			(-0.64)
Income from the Foreign Field per Station (US$)	224.18	181.44	42.74
			(0.43)
No. Missionaries per Station	1.65	1.58	0.07
			(0.28)
No. Native Christians per Station	412.14	117.19	294.95
			(0.85)
No. Churches per Station	0.52	1.02	-0.50
			(-1.34)
No. Sunday Schools per Station	0.52	1.49	-0.97
			(-1.67)
N	203	92	295

* $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$. The table compares societies operating worldwide in 1902 depending on whether they had medical missionaries. Societies compared are those also present in India. The number of missionary stations includes principals and all other substations.

Table C.6: Robustness Check: Test of Overidentifying Restrictions

	Generated Instr. Only (1)	Generated and Excluded Instr. (2)
Distance from Protestant medical mission (log)	-0.4838*	-0.5759***
	(0.268)	(0.214)
N	7046	7046
Hansen J Statistics	16.3	16.4
Degrees of Freedom	18	19
P-value	0.573	0.632

* $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$. All variables are centered to remove district fixed effects. 19 additional instruments are generated using the residuals from the first stage regression (Lewbel (2012)).

Table C.7: Additional Results: School Proximity

	BMI				
	OLS (1)	OLS (2)	OLS (3)	OLS (4)	OLS (5)
Distance from Protestant medical mission (log)	-0.227*	-0.193*	-0.204*		-0.216*
	(0.120)	(0.113)	(0.107)		(0.122)
Distance from Protestant boarding school (log)	0.0249			0.0280	0.0914
	(0.110)			(0.123)	(0.141)
Distance from Protestant high school (log)		-0.0748		-0.123	-0.0883
		(0.110)		(0.119)	(0.124)
Distance from Protestant university (log)			-0.0547	-0.0441	-0.0603
			(0.128)	(0.173)	(0.167)
Distance from Protestant mission (log)	0.0353	0.0405	0.0322	-0.00421	0.0408
	(0.111)	(0.112)	(0.110)	(0.113)	(0.114)
Distance from Catholic mission (log)	-0.134	-0.117	-0.124	-0.161	-0.120
	(0.117)	(0.115)	(0.117)	(0.114)	(0.116)
N	7,046	7,046	7,046	7,046	7,046
Mean Dependent Variable	20.42	20.42	20.42	20.42	20.42
District FE	Yes	Yes	Yes	Yes	Yes
Individual controls	Yes	Yes	Yes	Yes	Yes
Geographical controls	Yes	Yes	Yes	Yes	Yes
Historical controls	Yes	Yes	Yes	Yes	Yes
Distance from Protestant mission (log)	Yes	Yes	Yes	Yes	Yes
Distance from Catholic mission (log)	Yes	Yes	Yes	Yes	Yes

Only individuals of age 20 to 60 and BMI 15 to 30 are included in the estimation sample. Robust standard errors in parentheses. * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$. Standard errors clustered at the village (ward) level.

Table C.8: Additional Results: Medical Mission Activities

	BMI			
	OLS (1)	OLS (2)	OLS (3)	OLS (4)
$\mathbb{1}$(No. treatments and patients $\geq$ median)	0.265 (0.178)	0.900 (0.564)		
Distance from Protestant medical mission (log)		-0.194* (0.108)		
Distance from Protestant medical mission (log) $\times$ $\mathbb{1}$(No. treatments and patients $\geq$ median)		-0.147 (0.135)		
$\mathbb{1}$(Surgery at closest Protestant medical mission)			0.377** (0.179)	0.645 (0.582)
Distance from Protestant medical mission (log)				-0.251** (0.112)
Distance from Protestant medical mission (log) $\times$ $\mathbb{1}$(Surgery at closest Protestant medical mission)				-0.046 (0.143)
N	7,046	7,046	7,046	7,046
Mean Dependent Variable	20.42	20.42	20.42	20.42
District FE	Yes	Yes	Yes	Yes
Individual controls	Yes	Yes	Yes	Yes
Geographical controls	Yes	Yes	Yes	Yes
Historical controls	Yes	Yes	Yes	Yes
Distance from Protestant mission (log)	Yes	Yes	Yes	Yes
Distance from Catholic mission (log)	Yes	Yes	Yes	Yes

Only individuals of age 20 to 60 and BMI 15 to 30 are included in the estimation sample. Robust standard errors in parentheses. * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$. Standard errors clustered at the village (ward) level.

Table C.9: Descriptive Statistics (DLHS-3)

	Obs. (1)	Mean (2)	St. Dev. (3)
Panel A: Household Survey			
$\mathbb{1}$(Water Treatment)	485,219	0.30	0.46
$\mathbb{1}$(Open Defecation)	485,533	0.66	0.47
$\mathbb{1}$(Traditional Medicine)	458,569	0.04	0.19
Distance from Protestant medical mission	465,092	89.26	106.09
Distance from Protestant mission	485,540	49.93	100.41
Distance from Catholic mission	485,540	79.09	105.37
Household size	485,540	5.36	2.62
Age of head of household	485,537	46.43	14.39
$\mathbb{1}$(Hindu)	485,539	0.77	0.42
$\mathbb{1}$(Primary school completed by head of hh)	287,838	0.80	0.40
$\mathbb{1}$(Wealth 1^{st} quartile)	485,537	0.26	0.44
$\mathbb{1}$(Wealth 2^{nd} quartile)	485,537	0.27	0.44
$\mathbb{1}$(Wealth 3^{rd} quartile)	485,537	0.23	0.42
$\mathbb{1}$(Wealth 4^{th} quartile)	485,537	0.24	0.43
Village population (Census 2001)	485,540	2299.22	4272.09
Population in 1900	485,540	131.74	228.70
No. colonial railways	485,540	0.35	0.86
Cropland in 1900 (km^2)	485,540	35.91	32.91
Altitude (meters)	485,540	385.67	533.22
No. rivers/water sources	485,540	1.90	2.00
Latitude	465,092	24.24	5.16
Longitude	465,092	81.59	6.41
Panel B: Ever Married Women Survey			
Pneumonia Knowledge Index	441,168	0.22	0.24
$\mathbb{1}$(Iron supplements during last pregnancy)	441,168	0.81	0.39
$\mathbb{1}$(Full ANC during last pregnancy)	157,503	0.13	0.34
$\mathbb{1}$(Safe delivery)	157,488	0.42	0.49
$\mathbb{1}$(Child received check-up within 24 hrs from birth)	154,556	0.40	0.49
Distance from Protestant medical mission	422,332	87.85	100.11
Distance from Protestant mission	441,168	49.95	94.53
Distance from Catholic mission	441,168	78.05	99.11
Household size	441,168	6.36	3.12
Woman's age	441,168	31.45	8.62
$\mathbb{1}$(Hindu)	441,166	0.78	0.41
$\mathbb{1}$(Primary school completed)	441,123	0.38	0.49
$\mathbb{1}$(Wealth 1^{st} quartile)	441,168	0.28	0.45
$\mathbb{1}$(Wealth 2^{nd} quartile)	441,168	0.27	0.44
$\mathbb{1}$(Wealth 3^{rd} quartile)	441,168	0.21	0.40
$\mathbb{1}$(Wealth 4^{th} quartile)	441,168	0.24	0.43
Village population (Census 2001)	441,168	2307.91	4212.86
Population in 1900	441,168	137.47	233.89
No. colonial railways	441,168	0.37	0.88
Cropland in 1900 (km^2)	441,168	37.39	32.88
Altitude (meters)	441,168	357.25	487.81
No. rivers/water sources	441,168	1.87	1.97
Latitude	422,332	24.26	5.05
Longitude	422,332	81.29	6.20

DLHS-3 data. The wealth quartiles are based on a wealth index computed as the first principal component that combines information about a set of household assets: number of rooms in dwelling, and asset ownership, such as refrigerator, scooter, radio, and bike. Population, altitude, number of colonial railways and cropland data are extracted in 5 km-radius buffers around the DLHS village locations. Elevation data are available at the 30" × 30" grid cell; original population in year 1900 and cropland data are available at the 5' × 5' resolution. Latitude and longitude are in decimal degrees.

C.2 Long-Term Transmission and Health Production Function

To illustrate the three transmission channels, it may be useful to consider individual health as being determined by a set of inputs as follows:

$$H = \mathscr{H}(P, I, D) \tag{C.1}$$

where P is health potential, I are health inputs and D are exogenous environmental conditions. These may depend on local observable characteristics, i.e. $D = \mathscr{D}(W)$. We define A_{SR} as current access to health care and A_{LR} as long-term access to health care, i.e. access to health care of previous generations. Since access to health care and health status of one generation can leave their mark on the health potential of the following cohorts (Bhalotra and Rawlings (2011), Coneus and Spiess (2012)), we consider individual health potential to be determined by A_{LR} and by socio-economic individual characteristics, i.e. $P = \mathscr{P}(A_{LR}, X)$. Moreover, health inputs may depend on current access to health care, health culture C, such as hygiene and health awareness and health promoting practices, and socio-economic individual characteristics, i.e. $I = \mathscr{I}(A_{SR}, C, X)$. In particular, health culture may be shaped by current and past access to health care: $C = \mathscr{C}(A_{SR}, A_{LR})$. On one side, proximity to health facilities today may directly affect health culture and stimulate the diffusion of healthy practices. On the other side, health and hygiene awareness may have been bequeathed by the family or the social environment and therefore depend on the access to health care of previous generations. An individual's health production function is therefore a function of long-term access to health care, access to health care today, socio-economic and demographic individual characteristics and local environmental conditions. Assuming a simple Cobb-Douglas production function and conditioning on observable characteristics at the individual and at the village level, individual health is produced as follows:

$$H = \mathscr{H}(A_{LR}, A_{SR}) = \kappa(A_{LR})^{\beta_1}(A_{SR})^{\beta_2} \tag{C.2}$$

which we estimate empirically in section 3.7.1. We assume κ to be additively separable in district level and individual idiosyncratic components, i.e. $\kappa_{ivd} = \alpha_d + \epsilon_{ivd}$.

Bibliography

ACEMOGLU, D., S. JOHNSON, AND J. A. ROBINSON (2001): "The Colonial Origins of Comparative Development: An Empirical Investigation," *American Economic Review*, 91, 1369–1401. [79]

AGARWAL, B. (1995): "A Field of One's Own," *Cambridge Books*. [10], [62]

ALESINA, A., P. GIULIANO, AND N. NUNN (2013): "On the Origins of Gender Roles: Women and the Plough," *The Quarterly Journal of Economics*, 128, 469–530. [79]

ALMOND, D. AND B. MAZUMDER (2011): "Health Capital and the Prenatal Environment: The Effect of Ramadan Observance during Pregnancy," *American Economic Journal: Applied Economics*, 3, 56–85. [80]

ANDERSON, S. AND G. GENICOT (2015): "Suicide and Property Rights in India," *Journal of Development Economics*, 114, 64 – 78. [8]

ANDERSON, S. AND D. RAY (2010): "Missing Women: Age and Disease," *Review of Economic Studies*, 77, 1262–1300. [1], [2], [5], [7], [28], [38], [39], [40], [42], [117]

——— (2012): "The Age Distribution of Missing Women in India," *Economic and Political Weekly*, 47, 87–95. [2], [5], [7], [42]

——— (2015): "Missing Unmarried Women," Working Paper 21511, National Bureau of Economic Research. [5], [7], [25]

ANUKRITI, S., S. BHALOTRA, AND H. TAM (2015): "Missing Girls: Ultrasound Access, Fertility, and Gender Discrimination," *Unpublished Manuscript*. [7]

APPS, P. F. AND R. REES (1988): "Taxation and the Household," *Journal of Public Economics*, 35, 355 – 369. [9], [46]

ATTANASIO, O. AND V. LECHENE (2002): "Tests of Income Pooling in Household Decisions," *Review of Economic Dynamics*, 5, 720 – 748. [8], [44], [135], [137]

——— (2010): "Conditional Cash Transfers, Women and the Demand for Food," IFS Working Papers W10/17, Institute for Fiscal Studies. [13], [135], [137]

ATTANASIO, O. P. AND V. LECHENE (2014): "Efficient Responses to Targeted Cash Transfers," *Journal of Political Economy*, 122, 178 – 222. [16]

BAI, Y. AND J. K.-S. KUNG (2014): "Diffusing Knowledge While Spreading God's Message: Protestantism and Economic Prosperity in China, 1840–1920," *Journal of the European Economic Association*. [80]

BANERJEE, A., A. DEATON, AND E. DUFLO (2004): "Health, Health Care, and Economic Development: Wealth, Health, and Health Services in Rural Rajasthan," *The American economic review*, 94, 326. [79], [109]

BANERJEE, R. AND A. SACHDEVA (2014): "Pathways to Preventive Health, Evidence from India's Rural Road Program," *Unpublished Manuscript*. [84]

BARGAIN, O. AND O. DONNI (2012): "Expenditure on Children: A Rothbarth-type Method Consistent with Scale Economies and Parents' Bargaining," *European Economic Review*, 56, 792–813. [46]

BARRIENTOS, A., M. GORMAN, AND A. HESLOP (2003): "Old Age Poverty in Developing Countries: Contributions and Dependence in Later Life," *World Development*, 31, 555 – 570, chronic Poverty and Development Policy. [37]

BARRO, R. J. AND R. MCCLEARY (2003): "Religion and Economic Growth," NBER Working Papers 9682, National Bureau of Economic Research, Inc. [80]

BARRO, R. J. AND R. M. MCCLEARY (2005): "Which Countries Have State Religions?" *The Quarterly Journal of Economics*, 120, 1331–1370. [80]

BAUM, C. F. AND M. E. SCHAFFER (2012): "IVREG2H: Stata Module to Perform Instrumental Variables Estimation Using Heteroskedasticity-based Instruments," Statistical Software Components, Boston College Department of Economics. [105]

BECKER, S. O. AND L. WOESSMANN (2009): "Was Weber Wrong? A Human Capital Theory of Protestant Economic History," *The Quarterly Journal of Economics*, 124, 531–596. [80]

BHALOTRA, S. AND S. B. RAWLINGS (2011): "Intergenerational Persistence in Health in Developing Countries: The Penalty of Gender Inequality?" *Journal of Public Economics*, 95, 286 – 299, new Directions in the Economics of Welfare: Special Issue Celebrating Nobel Laureate Amartya Sen's 75th Birthday. [166]

BHALOTRA, S., C. VALENTE, AND A. VAN SOEST (2010): "The Puzzle of Muslim Advantage in Child Survival in India," *Journal of Health Economics*, 29, 191–204. [80]

BHALOTRA, S. R. AND T. COCHRANE (2010): "Where Have All the Young Girls Gone? Identification of Sex Selection in India," IZA Discussion Papers 5381, Institute for the Study of Labor (IZA). [7]

BHARADWAJ, P. AND L. K. LAKDAWALA (2013): "Discrimination Begins in the Womb: Evidence of Sex-Selective Prenatal Investments," *Journal of Human Resources*, 48, 71–113. [7]

BLACK, R. E., L. H. ALLEN, Z. A. BHUTTA, L. E. CAULFIELD, M. DE ONIS, M. EZZATI, C. MATHERS, J. RIVERA, MATERNAL, C. U. S. GROUP, ET AL. (2008): "Maternal and Child Undernutrition: Global and Regional Exposures and Health Consequences," *The lancet*, 371, 243–260. [127]

BLUNDELL, R., P. CHIAPPORI, AND C. MEGHIR (2005): "Collective labor supply with children," *Journal of Political Economy*, 113, 1277–1306. [46]

BONGAARTS, J. AND C. Z. GUILMOTO (2015): "How Many More Missing Women? Excess Female Mortality and Prenatal Sex Selection, 1970–2050," *Population and Development Review*, 41, 241–269. [7]

BOSE, N. AND S. DAS (2015): "Women's Inheritance Rights, Household Allocation and Gender Bias," *Unpublished Manuscript*. [8]

BRAINERD, E. AND N. MENON (2013): "Religion and Health in Early Childhood: Evidence from the Indian Subcontinent," Working Papers 65, Brandeis University, Department of Economics and International Business School. [80]

BROWNING, M., F. BOURGUIGNON, P.-A. CHIAPPORI, AND V. LECHENE (1994): "Income and outcomes: A structural model of intrahousehold allocation," *Journal of Political* ..., 102, 1067–1096. [9], [16], [46]

BROWNING, M. AND P. A. CHIAPPORI (1998): "Efficient Intra-Household Allocations: A General Characterization and Empirical Tests," *Econometrica*, 66, pp. 1241–1278. [9], [16], [46]

BROWNING, M., P.-A. CHIAPPORI, AND A. LEWBEL (2013): "Estimating Consumption Economies of Scale, Adult Equivalence Scales, and Household Bargaining Power," *Review of Economic Studies*, 80, 1267–1303. [3], [9], [16], [30], [46], [54], [146]

BROWNING, M., P.-A. CHIAPPORI, AND Y. WEISS (2014): *Economics of the Family*, Cambridge University Press. [14], [16], [137]

BURMASTER, D. E. AND E. A. CROUCH (1997): "Lognormal Distributions for Body Weight as a Function of Age for Males and Females in the United States, 1976-1980," *Risk Analysis*, 17, 499–505. [88]

BÜTIKOFER, A., A. LEWBEL, AND S. SEITZ (2010): "Health and Retirement Effects in a Collective Consumption Model of Older Households," Boston College Working Papers in Economics 767, Boston College Department of Economics. [9]

CAGÉ, J. AND V. RUEDA (2014): "The Long-term Effects of the Printing Press in Sub-Saharan Africa," *Available at SSRN 2292660*. [79], [80], [90], [94], [107] CALVI, R. (2016): "Why Are Older Women Missing in India? The Age Profile of Bargaining Power and Poverty," *Unpublished Manuscript*. [45], [58], [67]

CHAUDHURY, N. AND J. S. HAMMER (2004): "Ghost Doctors: Absenteeism in Rural Bangladeshi Health Facilities," *The World Bank Economic Review*, 18, 423–441. [79]

CHEN, M. A. AND J. DRÈZE (1995): "Widowhood and Well-being in Rural North India." . [7], [25]

CHERCHYE, L., B. DE ROCK, A. LEWBEL, AND F. VERMEULEN (2012a): "Sharing Rule Identification For General Collective Consumption Models," *IZA Working Paper, No. 6571*. [17]

CHERCHYE, L., B. DE ROCK, AND F. VERMEULEN (2012b): "Economic Well-being and Poverty Among the Elderly: An Analysis Based on a Collective Consumption Model," *European Economic Review*, 56, 985–1000. [9], [46]

CHIAPPORI, P.-A. (1988a): "Nash-Bargained Households Decisions: A Comment," *International Economic Review*, 29, pp. 791–796. [9], [16]

——— (1988b): "Rational Household Labor Supply," *Econometrica*, 56, pp. 63–90. [46]

——— (1992): "Collective Labor Supply and Welfare," *Journal of Political Economy*, 100, pp. 437–467. [9], [16], [46], [54]

CHIAPPORI, P. A. AND I. EKELAND (2006): "The Micro Economics of Group Behavior: General Characterization," *Journal of Economic Theory*, 130, 1–26. [46]

CHIAPPORI, P.-A., B. FORTIN, AND G. LACROIX (2002): "Marriage Market, Divorce Legislation, and Household Labor Supply," *Journal of Political Economy*, 110, 37–72. [24]

CLARKE, D. (2014): "PLAUSEXOG: Stata Module to Implement Conley et al.'s Plausibly Exogenous Bounds," Statistical Software Components, Boston College Department of Economics. [105]

COALE, A. J. (1991): "Excess Female Mortality and the Balance of the Sexes in the Population: An Estimate of the Number of "Missing Females"," *Population and Development Review*, 17, pp. 517–523. [1], [6]

CONDE-AGUDELO, A. AND J. M. BELIZÁN (2000): "Maternal Morbidity and Mortality Associated with Interpregnancy Interval: Cross Sectional Study," *Bmj*, 321, 1255–1259. [60], [73]

CONEUS, K. AND C. K. SPIESS (2012): "The Intergenerational Transmission of Health in Early Childhood-Evidence from the German Socio-Economic Panel Study," *Economics and Human Biology*, 10, 89 – 97. [166]

CONLEY, T. G. (2008): "Spatial Econometrics," in *The New Palgrave Dictionary of Economics*, ed. by S. N. Durlauf and L. E. Blume, Basingstoke: Palgrave Macmillan. [98]

CONLEY, T. G., C. B. HANSEN, AND P. E. ROSSI (2012): "Plausibly Exogenous," *Review of Economics and Statistics*, 94, 260–272. [104], [105]

CULLETON, B. F., B. J. MANNS, J. ZHANG, M. TONELLI, S. KLARENBACH, AND B. R. HEMMELGARN (2006): "Impact of Anemia on Hospitalization and Mortality in Older Adults," *Blood*, 107, 3841–3846. [127]

DAS, J. AND J. HAMMER (2007): "Money for Nothing: The Dire Straits of Medical Practice in Delhi, India," *Journal of Development Economics*, 83, 1–36. [109]

DASGUPTA, M. (2005): "Explaining Asia's "Missing Women": A New Look at the Data," *Population and Development Review*, 31, 529–535. [1], [7]

DE BENOIST, B., E. MCLEAN, I. EGLI, AND M. COGSWELL (2008): "Worldwide Prevalence of Anaemia 1993-2005: WHO Global Database on Anaemia," . [127]

DEATON, A. (2008): "Height, Health, and Inequality: The Distribution of Adult Heights in India," *American Economic Review*, 98, 468–74. [87]

DEATON, A. AND C. PAXSON (1995): "Measuring Poverty Among the Elderly," Tech. rep., National Bureau of Economic Research. [37]

DEININGER, K., A. GOYAL, AND H. NAGARAJAN (2013): "Women's Inheritance Rights and Intergenerational Transmission of Resources in India," *Journal of Human Resources*, 48, 114–141. [5], [8], [45]

DENNIS, J. S. (1902): *Centennial Survey of Foreign Mission: A Statistical Supplement to "Christian Missions and Social Progress", Being a Conspectus of the Achievements and Results of Evangelical Missions in all Lands at the close of the Nineteenth Century*, New York: Revell. [84], [94], [104]

DRÈZE, J. AND P. SRINIVASAN (1997): "Widowhood and Poverty in Rural India: Some Inferences from Household Survey Data," *Journal of Development Economics*, 54, 217–234. [7], [25]

DRÈZE, J. ET AL. (1990): "Widows in Rural India," *Discussion Paper-Development Economics Research Programme, Suntory-Toyota International Centre for Economics and Related Disciplines*. [7], [25]

DUFLO, E. (2001): "Schooling and Labor Market Consequences of School Construction in Indonesia: Evidence from an Unusual Policy Experiment," *American Economic Review*, 91, 795–813. [128]

——— (2003): "Grandmothers and Granddaughters: Old-Age Pensions and Intrahousehold Allocation in South Africa," *World Bank Economic Review*, 17, 1–25. [9], [44]

DUNBAR, G. R., A. LEWBEL, AND K. PENDAKUR (2013): "Children's Resources in Collective Households: Identification, Estimation, and an Application to Child Poverty in Malawi," *American Economic Review*, 103, 438–71. [3], [9], [14], [16], [17], [18], [25], [28], [34], [35], [46], [54], [57], [67], [70], [75], [139], [146]

FERNÁNDEZ-VAL, I. (2009): "Fixed Effects Estimation of Structural Parameters and Marginal Effects in Panel Probit Models," *Journal of Econometrics*, 150, 71–85. [12], [112]

FITZGERALD, R. (2001): "'Clinical Christianity': The Emergence of Medical Work as a Missionary Strategy in Colonial India," in *Health, Medicine and the Empire: New Perspectives on Colonial India*, ed. by B. Pati and M. Harrison, London: Sangam Books. [78], [82], [108], [111]

FOGEL, R. W. (1994): "Economic Growth, Population Theory, and Physiology: The Bearing of Long-Term Processes on the Making of Economic Policy," *American Economic Review*, 84, 369–95. [87], [98], [126]

——— (2004): *The Escape from Hunger and Premature Death, 1700-2100*, no. 9780521808781 in Cambridge Books, Cambridge University Press. [87], [126]

FULFORD, S. (2014): "The Puzzle of Marriage Migration in India," *Unpublished Manuscript*. [26]

GALLEGO, F. A. AND R. WOODBERRY (2009): "Christian Missionaries and Education in Former African Colonies: How Competition Mattered," Working Papers ClioLab 2, EH Clio Lab. Instituto de Economía. Pontificia Universidad Católica de Chile. [80]

GASPARINI, L., J. ALEJO, F. HAIMOVICH, S. OLIVIERI, AND L. TORNAROLLI (2007): "Poverty Among the Elderly in Latin America and the Caribbean," *Documentos de Trabajo del CEDLAS*. [37]

GLAESER, E. L. AND A. SHLEIFER (2002): "Legal Origins," *The Quarterly Journal of Economics*, 117, 1193–1229. [79]

GREENE, W., C. HAN, AND P. SCHMIDT (2002): "The Bias of the Fixed Effects Estimator in Nonlinear Models," *Unpublished Manuscript, Stern School of Business, NYU*. [12], [112]

HARARI, M. (2014): "Women's Inheritance Rights and Bargaining Power: Evidence from Kenya," *Unpublished Manuscript*. [8], [45]

HARDIMAN, D. (2008): *Missionaries and Their Medicine: A Christian Modernity for Tribal India*, Manchester University Press. [82]

HEATH, R. AND X. TAN (2014): "Intrahousehold Bargaining, Female Autonomy, and Labor Supply: Theory and Evidence from India," Unpublished Manuscript, Department of Economics, University of Washington. [4], [5], [8], [20], [26], [68], [128]

HECKMAN, J., S. URZUA, AND E. VYTLACIL (2006): "Understanding Instrumental Variables in Models with Essential Heterogeneity," *Review of Economics and Statistics*, 88, 389–432. [48]

HECKMAN, J. J. (2010): "Building Bridges between Structural and Program Evaluation Approaches to Evaluating Policy," *Journal of Economic Literature*, 48, 356–98. [48]

HECKMAN, J. J. AND E. J. VYTLACIL (2007): "Chapter 70 Econometric Evaluation of Social Programs, Part I: Causal Models, Structural Models and Econometric Policy Evaluation," Elsevier, vol. 6, Part B of *Handbook of Econometrics*, 4779 – 4874. [48]

HJELMBORG, J. V., C. FAGNANI, K. SILVENTOINEN, M. MCGUE, M. KORKEILA, K. CHRISTENSEN, A. RISSANEN, AND J. KAPRIO (2008): "Genetic Influences on Growth Traits of BMI: A Longitudinal Study of Adult Twins," *Obesity*, 16, 847–852. [88]

JAIN, T. (2014): "Where There Is a Will Fertility Behavior and Sex Bias in Large Families," *Journal of Human Resources*, 49, 393–423. [8]

JAYACHANDRAN, S. AND I. KUZIEMKO (2011): "Why Do Mothers Breastfeed Girls Less than Boys? Evidence and Implications for Child Health in India," *The Quarterly Journal of Economics*, 126, 1485–1538. [7]

JHA, P., R. KUMAR, P. VASA, N. DHINGRA, D. THIRUCHELVAM, AND R. MOINEDDIN (2006): "Low Male-to-female Sex Ratio of Children Born in India: National Survey of 1 Million Households," *The Lancet*, 367, 211–218. [7]

JOHNSON, H. B. (1967): "The Location of Christian Missions in Africa," *Geographical Review*, 57, pp. 168–202. [90]

KALAIVANI, K. ET AL. (2009): "Prevalence & Consequences of Anaemia in Pregnancy," *Indian J Med Res*, 130, 627–633. [127]

KAUR, K. (2014): "Anaemia 'A Silent Killer' Among Women in India: Present Scenario," *European Journal of Zoological Research*, 3, 32–36. [59], [126]

KLEIN GOLDEWIJK, K., A. BEUSEN, AND P. JANSSEN (2010): "Long-term Dynamic Modeling of Global Population and Built-up Area in a Spatially Explicit Way: HYDE 3.1," *The Holocene*. [86]

KLEIN GOLDEWIJK, K., A. BEUSEN, G. VAN DRECHT, AND M. DE VOS (2011): "The HYDE 3.1 Spatially Explicit Database of Human-induced Global Land-use Change Over the Past 12,000 Years," *Global Ecology and Biogeography*, 20, 73–86. [86]

KOCHAR, A. (1999): "Evaluating Familial Support for the Elderly: The Intrahousehold Allocation of Medical Expenditures in Rural Pakistan," *Economic Development and Cultural Change*, 47, pp. 621–656. [7]

La Ferrara, E. and A. Milazzo (2014): "Customary Norms, Inheritance, and Human Capital: Evidence from a Reform of the Matrilineal System in Ghana," CEPR Discussion Papers 10159, C.E.P.R. Discussion Papers. [8], [45]

Lancaster, G., P. Maitra, and R. Ray (2008): "Household Expenditure Patterns and Gender Bias: Evidence from Selected Indian States," *Oxford Development Studies*, 36, 133–157. [7]

Lewbel, A. (2012): "Using Heteroskedasticity to Identify and Estimate Mismeasured and Endogenous Regressor Models," *Journal of Business & Economic Statistics*, 30, 67–80. [104], [105], [162]

——— (2016): "The Identification Zoo - Meanings of Identification in Econometrics," *Forthcoming on Journal of Economic Literature*. [48]

Lewbel, A. and K. Pendakur (2008): "Estimation of Collective Household Models with Engel Curves," *Journal of Econometrics*, 147, 350–358. [3], [9], [16], [46]

Lise, J. and S. Seitz (2011): "Consumption Inequality and Intra-household Allocations," *The Review of Economic Studies*, 78, 328–355. [46]

Lundberg, S. J., R. A. Pollak, and T. J. Wales (1997): "Do Husbands and Wives Pool Their Resources? Evidence from the United Kingdom Child Benefit," *The Journal of Human Resources*, 32, 463–480. [44]

Mantovanelli, F. G. (2014): "The Protestant Legacy: Missions and Literacy in India," *Unpublished Manuscript*. [80], [107]

Mathew, E. (1999): "Growth of Literacy in Kerala: State Intervention, Missionary Initiatives and Social Movements," *Economic and Political Weekly*, 2811–2820. [78]

Mazzocco, M. (2007): "Household Intertemporal Behaviour: A Collective Characterization and a Test of Commitment," *The Review of Economic Studies*, 74, 857–895. [43]

Mazzocco, M., C. Ruiz, and S. Yamaguchi (2014): "Labor Supply and Household Dynamics," *American Economic Review*, 104, 354–59. [43]

Menon, M., K. Pendakur, and F. Perali (2012): "On the Expenditure-dependence of Children's Resource Shares," *Economics Letters*, 117, 739–742. [17]

Menon, N. and K. McQueeney (2015): "Christianity and Infant Health in India," *Unpublished Manuscript*. [80]

Milazzo, A. (2014): "Why Are Adult Women Missing? Son Preference and Maternal Survival in India," Policy Research Working Paper Series 6802, The World Bank. [5], [60], [73] NUNN, N. (2009): “The Importance of History for Economic Development,” *Annual Review of Economics*, 1, 65–92. [79]

——— (2010): “Religious Conversion in Colonial Africa,” *American Economic Review*, 100, 147–52. [90]

OSTER, E. (2009): “Does Increased Access Increase Equality? Gender and Child Health Investments in India,” *Journal of Development Economics*, 89, 62–76. [7]

PAL, P. AND J. GHOSH (2007): “Inequality in India: A Survey of Recent Trends,” Working Papers 45, United Nations, Department of Economics and Social Affairs. [77]

PATEL, K. V. (2008): “Epidemiology of Anemia in Older Adults,” *Seminars in Hematology*, 45, 210 – 217, anemia in the Elderly. [127]

PEARL, J. (2009): “Causal inference in statistics: An Overview,” *Statistics Surveys*, 3, 96–146. [48]

PLANNING COMMISSION (2014): “Report of the Expert Group to Review the Methodology for Measurement of Poverty,” *Government of India. New Delhi*. [35]

QUISUMBING, A. R. AND J. A. MALUCCIO (2003): “Resources at Marriage and Intrahousehold Allocation: Evidence from Bangladesh, Ethiopia, Indonesia, and South Africa*,” *Oxford Bulletin of Economics and Statistics*, 65, 283–327. [45]

REGGIO, I. (2011): “The Influence of the Mother’s Power on Her Child’s Labor in Mexico,” *Journal of Development Economics*, 96, 95 – 105. [45]

RICHTER, J. (1908): *A History of Missions in India*, F. H. Revell. [81]

ROSENBLUM, D. (2015): “Unintended Consequences of Women’s Inheritance Rights on Female Mortality in India,” *Economic Development and Cultural Change*, 63, pp. 223–248. [8]

ROY, K. AND A. CHAUDHURI (2008): “Influence of Socioeconomic Status, Wealth and Financial Empowerment on Gender Differences in Health and Healthcare Utilization in Later Life: Evidence from India,” *Social Science & Medicine*, 66, 1951 – 1962. [7]

ROY, S. (2008): “Female Empowerment Through Inheritance Rights: Evidence from India,” Unpublished Manuscript. [4], [5], [8], [26], [45], [68]

——— (2013): “Empowering Women: Inheritance Rights and Female Education in India,” Unpublished Manuscript. [5], [8], [45]

SCHULTZ, T. P. (2002): “Wage Gains Associated with Height as a Form of Health Human Capital,” *American Economic Review*, 92, 349–353. [78], [87], [98]

——— (2003): “Wage Rentals for Reproducible Human Capital: Evidence from Ghana and the Ivory Coast,” *Economics & Human Biology*, 1, 331–366. [78], [87], [98]

——— (2009): “Population and Health Policies,” Working Papers 974, Economic Growth Center, Yale University. [59], [80], [87], [126]

SEN, A. (1990): “More Than 100 Million Women Are Missing,” *New York Review of Books*, pp. 61–66. [1]

——— (1992): “Missing women.” *Bmj*, 304, 587–588. [1]

SHIWAKU, K., E. ANUURAD, B. ENKHMAA, K. KITAJIMA, AND Y. YAMANE (2004): “Appropriate Body-Mass Index For Asian Populations and Its Implications For Policy and Intervention Strategies,” *Lancet*. [59], [87], [126]

SRINIVISAN, R. (2010): “Health Care in India-Vision 2020,” *New Delhi, India: Government of India, Planning Commission of India*. [77]

STECKEL, R. H. (2008): “Heights and Human Welfare: Recent Developments and New Directions,” NBER Working Papers 14536, National Bureau of Economic Research, Inc. [87], [98]

SUBRAMANIAN, S. AND A. DEATON (1990): “Gender Effects In Indian Consumption Patterns,” Tech. rep. [7]

THOROGOOD, M., P. APPLEBY, T. KEY, AND J. MANN (2003): “Relation Between Body Mass Index and Mortality in an Unusually Slim Cohort,” *Journal of epidemiology and community health*, 57, 130–133. [10], [126]

TOMMASI, D. AND A. WOLF (2016): “Overcoming Weak Identification in the Estimation of Household Resource Shares,” ECARES working paper 2016-12. [58]

UDRY, C. (1996): “Gender, Agricultural Production, and the Theory of the Household,” *Journal of political Economy*, 1010–1046. [16]

VERMEULEN, F. (2002): “Collective Household Models: Principles and Main R‘esults,” *Journal of Economic Surveys*, 16, 533–564. [9], [16], [46]

VISSCHER, T. L. S., J. C. SEIDELL, A. MENOTTI, H. BLACKBURN, A. NISSINEN, E. J. M. FESKENS, D. KROMHOUT, AND FOR THE SEVEN COUNTRIES STUDY RESEARCH GROUP (2000): “Underweight and Overweight in Relation to Mortality Among Men Aged 40-59 and 50-69 Years: The Seven Countries Study,” *American Journal of Epidemiology*, 151, 660–666. [10], [126]

VOENA, A. (2015): “Yours, Mine, and Ours: Do Divorce Laws Affect the Intertemporal Behavior of Married Couples?” *American Economic Review*, 105, 2295–2332. [24], [43] VYTLACIL, E. (2002): "Independence, Monotonicity, and Latent Index Models: An Equivalence Result," *Econometrica*, 70, 331–341. [48]

WAILER, H. T. (1984): "Height, Weight and Mortality: The Norwegian Experience," *Acta Medica Scandinavica*, 215, 1–56. [87], [126]

WEBER, M. (1930): *The Protestant Ethic and the Spirit of Capitalism, 1700-2100*, London: G. Allen and Unwin. [80]

WEIL, D. N. (2005): "Accounting for the Effect of Health on Economic Growth," Tech. rep. [87], [98]

WOLPIN, K. I. (2013): *The Limits of Inference Without Theory*, MIT Press. [48]

WORBOYS, M. (2000): *Spreading Germs: Disease Theories and Medical Practice in Britain, 1865-1900*, Cambridge University Press. [111]

WORLD BANK (2014): "Women, Business and the Law 2014. Removing Restrictions to Enhance Gender Equality," *World Bank Report*. [10], [62]

WORLD HEALTH ORGANIZATION (1999): *World Health Report 1999 - Making a Difference*, World Health Organization. [4]

——— (2015): *World Health Statistics 2015*, World Health Organization. [4]

ZHENG, W., D. F. MCLERRAN, B. ROLLAND, X. ZHANG, M. INOUE, K. MATSUO, J. HE, P. C. GUPTA, K. RAMADAS, S. TSUGANE, ET AL. (2011): "Association Between Body-mass Index and Risk of Death in More than 1 million Asians," *New England Journal of Medicine*, 364, 719–729. [10], [126]

ZIMMERMANN, L. (2012): "Reconsidering Gender Bias in Intrahousehold Allocation in India," *Journal of Development Studies*, 48, 151–163. [7]