Why Are Adult Women Missing?
Son Preference and Maternal Survival in India*

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Very preliminary and incomplete. Comments are welcome.

Abstract

This paper explores the phenomenon of missing women of reproductive and adult ages in India. By using individual-level data, I compare the age structure and health indicators of women by the sex of their first-born and uncover several new findings. First, among mothers with at least one child, the share of those with first-born daughter decreases with women's age (observed at the time of the survey). Second, women with first-born daughter are significantly more likely to be anemic when young (especially under age 30) while they show better anthropometric indicators and no different incidence of anemia as they get older. Interestingly, differences in anemia prevalence arise only some time after the birth of the first child. Third, women with first-born daughter have higher fertility, tend to space births less and might thus be at higher risk of dying for complications during or after childbirth and maternal depletion. Larger differences are consistently found for women with two consecutive daughters. I argue that all these findings are consistent with a selection effect in which healthier women of higher socioeconomic status are more likely to survive the birth of daughters. To my knowledge, this is the first paper arguing that women's excess mortality can (at least partly) be an unintended consequence of son preference.

Keywords: gender bias, fertility, birth spacing, maternal mortality and morbidity
JEL classification: D63, J13, J16, I10

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1 Introduction

There is an extensive literature on parental preferences over a child's gender. Sen's 1990 article documented high sex ratios at birth (the ratio of males to females who are born) in many South and Southeast Asian countries: he argued that women might be missing because of sex-selective abortion, infanticide and neglect of female children. Sex ratios for children aged 0–6 have been steadily increasing since 1961 in India, and evidence of female foeticide has been widely documented (Jha et al, 2006).[1]. The diffusion of prenatal sex determination technologies starting from the end of the 1980s has been found to be associated with the increasing sex ratios and preferential prenatal treatment for boys (Bhalotra and Cochrane, 2010, and Bharadwaj and Nelson, 2012).[2] Other authors have studied fertility behavior and shown that subsequent fertility is higher for women who had girls among earlier-born children (son-preferring fertility stopping rules) (Chowdhury and Bairagi, 1990; Clark, 2000; Dreze and Murthi, 2001; Arnold et al., 2002). The 'try until you have a son' fertility rule may also 'passively' lead to lower amount of resources being allocated to female children within the household. This may happen either as a direct consequence of the fact that the weaning time for girls will be shorter than that for boys as parents want to try again for a son, or because girls tend be born in larger households (eg, see Jayachandran and Kuziemko, 2011, for gender disparities in breastfeeding duration, and Jensen, 2003, on the gender differential in education).[3]

Previous studies mostly focused on gender bias at birth or young ages in explaining the phenomenon of missing women. A notable exception is Anderson and Ray (2010), who provide a decomposition of the excess female deaths by age and causes of death using aggregate data for different parts of the world. While they still find evidence of severe bias against young (and unborn) female children, they show that the majority of missing women in India are of reproductive and adults ages.[4] Recently, Anderson and Ray

[1]According to decennial Census data, the number of girls per 1000 boys aged 0–6 was 976 in 1961 and reached 914 (its lowest level) in 2011.

[2]Abortion was legalized in 1971 in India. In 1994, the Government of India passed the Prenatal Diagnostic Techniques Regulations and Misuse Act (PNDT Act) to make the use of ultrasound or amniocentesis for the purpose of sex determination illegal. However, there is evidence that this Law is often ignored and not enforced (Arnold et al., 2002, George, 2002).

[3]Barcellos et al. (2012) find that, accounting for the effect of son-preferring stopping rules on family size, boys still receive more parental investment (including childcare time, breastfeeding, vaccinations, and vitamin supplementation) than girls.

[4]Anderson and Ray (2010) find that the sources of excess female mortality among

(2012) replicate this exercise for India. They confirm that most missing women are found in adulthood and add that there is 'significant state-wise variation in the distribution of missing women across the age groups' (p. 3). Moreover, they observe that 'if we compare maternal mortality rates and the percentage of the female population that is missing at reproductive ages across the states, we do indeed see a positive correlation.' (p. 15) The World Development Report (WDR) 2012 corroborates these findings and argues that the 'female disadvantage in mortality during the reproductive ages is in part driven by the risk of death in pregnancy and childbirth and associated long-term disabilities' (World Bank, 2011, p. 78).[5]

While previous recent studies have documented the excess of female mortality among adult women, this paper attempts to make a step forward in understanding the causes of excess female mortality among women of reproductive and adult ages. It makes at least two contributions to the existing literature. First, while Anderson and Ray (2010, 2012) use aggregate mortality data and find that, contrary to previously-held wisdom, many women are missing in adulthood, this paper uses individual-level data for India and explores whether and how fertility behavior and preferences can partly explain this pattern. Second, it compares health outcomes across different age groups to uncover a selection effect in which healthier women of higher socioeconomic status might be more likely to survive the birth of daughters. The main novelty of this paper is that it provides an original explanation for the phenomenon of missing adult women that essentially coincides with the one that has been brought forward to explain overt discrimination at younger ages, namely parental preferences for sons. The difference between the two phenomena though is that differential mortality of adult women with daughters arises unintentionally (being indirectly caused by son preference) while gender bias at younger ages can be both intentional or not.

I pool three rounds of the India National Family Health Survey (NFHS) and find several interesting new patterns. First, I provide non-parametric evidence that, among women with at least one child ever born, the share

women of reproductive ages (15–44) in India are maternal mortality and injuries. Among among infants and adolescents (age 0–14) respiratory and infectious diseases are the main causes of excess mortality, while among older women (age 45 and older) it is cardiovascular disease.

[5]One of every 140 women die from causes related to childbirth in India (estimates for 2008 from the WHO/UNICEF/UNFPA/World Bank, 2010). The adult lifetime risk of maternal death is defined as the probability of dying from a maternal cause during a womans reproductive lifespan.

of women with first-born daughter decreases with mother's age (reported at the time of the survey). To confirm this result, I also estimate an ordered logit model in which the dependent variable is a categorical variable for each age-group to show that the conditional age distribution shifts to the left for women with first-born girl (compared to first-born boys). Second, women with first-born daughter are 1.2 percentage points more likely to be anemic when young (especially under age 30) while they show significantly higher anthropometric indicators (weight-for-height and body mass index) and no different incidence of anemia as they get older. Interestingly, differences in anemia prevalence arise only some time after the birth of the first child. This suggests that women are similar at the time of first-birth, thus ruling out other interpretations based on biological differences among women with daughters or sons. This results are obtained after controlling for age at first birth, height (a proxy for health status before the first pregnancy), pregnancy and breastfeeding status, and other observable women's characteristics. Third, compared to women with first-born son, women with first-born daughter have 0.28 more children, are 7.3 percentage points less likely to use contraceptives (and in particular less likely of being sterilized), 1.3 more likely to have had a terminated pregnancy, and 5.1 percentage points more likely to desire more children. While the effects are stronger when considering the first two female births, I focus on the first-born to maintain a causal interpretation of the results. The identifying assumption is that the sex of the first-born is uncorrelated with the error term after conditioning for a set of observable characteristics. Lastly, using mother fixed effects to control for unobserved mother's heterogeneity, I estimate the effect of the sex of the child on birth spacing. I find that the birth of a girl increases the probability that the mother waits less than 24 (15) months by 2 (0.8) percentage points (equivalent to about a 6 and 9 percent effect, respectively).

My hypothesis is that the patterns identified in the data are consistent with selective mortality through son-preferring fertility stopping behavior. In particular, as it is shown in the empirical part, women with first-born daughter have higher fertility and closely spaced births. This reproductive behavior is associated with higher risk of death for complications during or after childbirth and maternal depletion. The link between son-preferring fertility behavior and maternal mortality is supported by medical evidence. Specifically, a high number of pregnancies is associated with higher lifetime risk of death due to pregnancy and short birth intervals are associated with poor child and maternal health outcomes. This is because close birth spacing does not allow a woman to regain her physical strength and nutrients required to have a successful following pregnancy. This fertility behavior is strongly associated with anemia, which in turn is an underlying cause of maternal mortality. I argue that a *selection effect* in which women of higher socioeconomic status are more likely to survive the birth of daughters can in part explain why adult women with first-born daughter are fewer and better-off in terms of nutritional indicators and other observable characteristics.

The remainder of the paper is organized as follows: Section 2 provides some background information on sex ratios in India, and on the prevalence of maternal mortality and anemia; Section 3 describes the data; Section 4 provides some descriptive evidence; Section 5 introduces the empirical methodology used to analyze fertility outcomes, birth spacing, health outcomes and discusses the results; Section 6 discusses three alternative interpretations of the results, and Section 7 concludes.

2 Fertility behavior and maternal health outcomes

A high number of pregnancies is associated with higher lifetime risk of death due to pregnancy (WHO, 2008). Short intervals between births have been found to be related to poor child health and maternal outcomes.[6] Medical research has shown that short birth intervals (<24 months) are associated with poor child and maternal health outcomes (Setty-Venugopal and Upadhyay, 2002).[7] For example, compared with mothers who give birth at 9- to 14-month intervals, women who have their babies at 27- to 32-month birth intervals are: 1.3 times more likely to avoid anemia; 1.7 times more likely to avoid third-trimester bleeding; and 2.5 times more likely to survive childbirth (Conde-Agudelo, and Belizan, 2000).[8]

[6] Birth spacing refers to the time interval (# months) between births.

[7] With reference to child health, compared with children born less than 2 years after a previous birth, children born 3 to 4 years after a previous birth are 1.5 times more likely to survive the first week of life, and 2.4 times more likely to survive to age five (DHS 2002). Also, there are effects on health of older children (e.g., breastfeeding duration, see Jayachandran, and Kuziemko, 2011). Palloni and Millman (1986) and Palloni and others (1994) discuss the channels through which short birth spacing and breastfeeding affect early childhood mortality. First of all, breast milk contains the nutritional requirements for infant growth and protects against the formation of bacteria and of malnutrition. Short preceding birth interval does not allow the mother the time to regain the strength between pregnancies, leading to the inability to adequately breastfeed, while short succeeding intervals reduces maternal care for older children.

[8] see Conde-Agudelo, Rosas-Bermdez, and Kafury-Goeta (2007) for a review of medical studies on the relationship between birth spacing and maternal health.

The World Bank (2011) observes that maternal mortality and morbidity related to childbirth are one of the two mechanisms driving excess mortality in the reproductive years.[9] The WHO estimates that 63000 maternal deaths occurred in India (out of 358000 worldwide) in 2008, the highest number in the world. India is making progress in the effort to improve maternal health (the fifth MDG) and the maternal mortality ratio went from 570 in 1990 to 230 in 2008.[10] Maternal mortality is difficult to measure and likely to go underestimated. It is defined by the WHO as the 'death of a woman while pregnant or within 42 days of termination of pregnancy, irrespective of the duration and site of the pregnancy, from any cause related to or aggravated by the pregnancy or its management but not from accidental or incidental causes.' (WHO, 2008). This definition only refers to complications occurring at the time of birth and within 42 days of delivery. However, maternal mortality 'implies not only death during childbirth but also concurrent morbidities brought on by the experience of pregnancy and childbirth.' (WDR, p. 128). These morbidities include, for example, anemia, malaria, hepatitis, tuberculosis, cardiovascular disease, and obstetric fistula.

Anemia is pervasive in India, affecting more than 55% of women age 15–49 (NFHS-3). It usually results from a nutritional deficiency of iron, folate, vitamin B12, or some other nutrients (iron-deficiency anaemia). It is more prevalent during pregnancy and among breastfeeding mothers, when the nutritional requirements increase substantially. Iron-deficiency anemia is an important cause of maternal morbidity and, when severe, mortality in India (World Bank, 1996, p.31).

In the WDR, the World Bank (2011) observes that 'maternal mortality is fundamentally different from excess female mortality at other ages in that, to reduce it, societies must focus on an intrinsically female condition and specifically on improving the maternal health care system.' (p. 128) In India, many women do not receive adequate maternal health care.[11]

3 Data

The data used in this paper are from the India National Family Health Survey (NFHS), which contains individual-level information on birth histories (including children who have died), birth intervals, health and anthropomet-

[9] The other mechanism, HIV/AIDS, is more pertinent for the Africa region.

[10] The maternal mortality ratio is the number of maternal deaths during a given time period per 100 000 live births during the same time-period (WHO, 2008).

[11] The NFHS-3 reports that only 39% of births took place in health facilities and 58% of women did not receive any postnatal check-up after their most recent birth.

ric indicators (for children and their mothers), and other variables. These are repeated cross-sectional survey data for a representative sample of households in India. I pool the three available survey rounds conducted in 1992/3, 1998/9, 2005/6. Pooling different survey waves allows to focus on age-effects rather that birth cohort-effects. The women covered in the NFHS are aged 15–49, and I consider the sample of all those who had at least one child ever born.

The total fertility rate in the NFHS-3 is 2.7, down to 2.9 in the NFHS-2. This means that, on average, an Indian woman will give birth to 2.7 children by the end of her childbearing years. Knowledge of contraception is nearly universal in India: 98% of women and 99% of men age 15-49 know one or more methods of contraception and contraceptive prevalence rate for currently married women in India is 56%, up from 48% in NFHS-2.[12]

Using the NFHS-3 birth history data, the sex ratio at birth (here defined as the fraction of male births) is 0.52 for all births.[13] Figure A.1 and figure A.2 show the sex ratios (at birth) by the birth year of the child in the pooled dataset and by survey year, respectively. Pooling different survey rounds allows to distinguish between age and cohort-specific effects. The figure shows that the decreasing trend in the sex ratio is common to all the survey rounds, independently of the cohort of birth of the child, and not specific to a particular period.[14] The most recent survey (NFHS-3) also shows an increasing sex ratio at birth from the end of the 1980s onward, which corresponds to the timing of the diffusion of ultrasound technologies and is consistent with evidence from Census data. Figure A.3 shows the sex ratio by mother's age: consistent with the previous graphs, older women tend to report more male births. Finally, it is important to note that the decreasing trend is specific to first births, while no similar pattern is found for higher

[12] Female sterilization, with a prevalence of 37%, accounts for 66% of all contraceptive use, down from 71% of all contraceptive use at the time of NFHS-2; The highest adoption rate of female sterilization, at 67%, is among women with three children who have two sons. The most common spacing methods are condoms and the rhythm method, each used by 5% of currently married women (India NFHS-3).

[13] Based on several data sources, Anderson and Ray (2010) report that sex ratios at birth are about 0.514 in developed countries (including the 'Established Market Economies' as defined by the World Bank: Western Europe, Canada, United States, Australia, New Zealand, and Japan), 0.518 in India, 0.539 in China, and 0.508 in Sub-Saharan Africa. *Overall* sex ratios in India and China are similar, both around 0.514. Several biological, environmental, and genetic factors can partly explain the variation across different areas.

[14] Using a different dataset for India (with retrospective fertility histories as in the NFHS), Rosenblum (2012) similarly finds higher sex ratios for births occurred in the past, but argues that the extent of the bias in the data is small. He also discusses about potential recall and survival bias in the data.

Table 1: Summary statistics, by sex of the first-born child

	first-born girl	first-born boy	diff	se	Obs
education years	3.482	3.438	0.044**	(0.022)	244512
husband eduyrs	6.036	6.014	0.022	(0.026)	243334
age	32.062	32.239	-0.177***	(0.042)	244803
age first birth	18.913	18.878	0.035**	(0.017)	244803
age first marriage	16.773	16.743	0.030*	(0.016)	244690
# children ever born	3.441	3.246	0.195***	(0.010)	244803
husband living in	0.929	0.93	-0.001	(0.001)	230290
female head	0.099	0.095	0.004**	(0.002)	244800
muslim	0.125	0.127	-0.003	(0.002)	244729
hindu	0.817	0.814	0.003	(0.002)	244729
sikh	0.018	0.018	-0.000	(0.000)	244729
sched caste	0.167	0.167	0.000	(0.002)	241080
sched tribe	0.088	0.086	0.002	(0.001)	241080
Health indicators					
anemic	0.548	0.538	0.011***	(0.003)	149779
weight-for-height	-1.362	-1.358	-0.004	(0.007)	154633
height	151.55	151.519	0.031	(0.037)	156402
BMI	20.595	20.586	0.009	(0.024)	156146

All women aged 15–49 with at least one child ever born. Using survey weights. ***, **, and * indicate significance at 1%, 5% and 10% levels. Anemia, height, weight-for-height, BMI only in waves 1998 & 2005.

birth-order births (figure A.4). This is consistent with the association between sex of the first-born, reproductive behavior, and maternal survival. In the next sections, I'll explore why there are systematically higher sex ratios for births (especially at birth order one) that occurred back in the past from the time of the survey.

Table 1 reports summary statistics for the pooled sample of women age 15–49 with at least one child ever born, by the sex of the first-born. The table shows that some characteristics significantly differ between the two subsamples of women. In particular, it shows that women with first-born girls are more educated, younger, married and had the first child later in life, had more children, are more likely to be the head of the household, and more likely to be anemic. While the higher number of children ever born is a direct consequence of the widespread use of son-preferring fertility rules in India, the differences in the other characteristics needs further investigation.

Table 2 reports the same summary statistics as in table 1 for the subsamples of women aged 15–30 and 31–49. The leftmost upper panel shows that younger women appear to be quite similar in terms of observable characteristics. Women with first-born girl aged 15–30 still have significantly more children (because of son-preferring fertility rules), but their education level and age are more similar to women with first-born boy. The rightmost upper panel shows a very different picture: the differences reported in table 1 are intensified for women aged 31–49. Older women with first-born girl seem to be 'better off' in terms of their own and husband's education, they married and had first child later, and are more likely to be head of the household.

The bottom panel shows the differences for the health indicators. The leftmost panel shows significant differences between younger women in the two groups: younger women with first-born daughter are significantly more likely to be anemic and are also characterized by lower weight-for-height, while the rightmost panel shows that these differences disappear and become positive for older women.

This information is consistent with a selection effect through son-preferring fertility behavior. In particular, women with first-born daughter have higher fertility, tend to space births less (as it will be shown later) and thus are at higher risk of dying because of maternal depletion. Younger women in their reproductive years might physically suffer from having children shortly spaced, and this is reflected in the incidence of anemia and anthropometric indicators. Among women with first-born girl, the ones who survive to older ages are probably the ones who are healthier and from a higher socio-economic background and the descriptive evidence above described seems to reflect this fact. More evidence in support of this channel will be shown in the next sections.

Birth spacing In the India NFHS pooled dataset, of all births reported by each woman with at least two children ever born, 10 percent are less than 15 months apart, 16 percent less than 18 months, 35 percent less than two years. The mean birth interval is 31.8 months, and the median is 27.[15] The birth interval increases with birth order (for given # of children as the preferences for family size affects the interval at each birth order). The sex of the first child affects the interval between the first and second child (as it will be investigated further in the following sections): it's 32.03 months after the birth of a girl, and 32.88 after a boy. The survival status of the child who opens the birth interval also matters: the median interval is 28 months if previous child is alive, 23 months if dead.

[15] To have at least one birth interval for each woman, it's necessary to consider the sample of women with at least two children ever born

Table 2: Summary statistics, by age group and sex of the the first-born

	Women age 15-30					Women age 31-49				
	firstG	firstB	diff	se	obs	firstG	firstB	diff	se	obs
education years	3.676	3.72	-0.044	(0.033)	110396	3.309	3.193	0.116***	(0.030)	134116
husband eduyrs	6.152	6.185	-0.033	(0.037)	109858	5.931	5.865	0.066*	(0.035)	133476
age	24.621	24.628	-0.007	(0.027)	110533	38.753	38.873	-0.120***	(0.036)	134270
age first birth	18.377	18.37	0.007	(0.021)	110533	19.394	19.32	0.074***	(0.025)	134270
age first marriage	16.532	16.53	0.002	(0.020)	110471	16.988	16.928	0.060**	(0.023)	134219
# children ever born	2.494	2.352	0.142***	(0.010)	110533	4.292	4.025	0.267***	(0.014)	134270
husband living in	0.912	0.914	-0.002	(0.002)	107346	0.944	0.944	0.000	(0.002)	122944
female head	0.083	0.081	0.002	(0.002)	110532	0.114	0.108	0.006***	(0.002)	134268
muslim	0.136	0.137	-0.001	(0.003)	110491	0.115	0.119	-0.004*	(0.002)	134238
hindu	0.815	0.813	0.002	(0.003)	110491	0.819	0.815	0.004	(0.003)	134238
sikh	0.015	0.015	0.000	(0.001)	110491	0.019	0.02	-0.001	(0.001)	134238
sched caste	0.176	0.172	0.004	(0.003)	108996	0.16	0.163	-0.003	(0.003)	132084
sched tribe	0.098	0.094	0.004*	(0.002)	108996	0.078	0.079	-0.001	(0.002)	132084
Health indicators										
anemic	0.573	0.554	0.019***	(0.005)	65400	0.528	0.524	0.004	(0.004)	84379
weight-for-height	-1.389	-1.368	-0.021**	(0.010)	67850	-1.338	-1.349	0.011	(0.011)	86783
height	151.51	151.495	0.015	(0.054)	68432	151.585	151.54	0.045	(0.050)	87970
BMI	19.779	19.807	-0.028	(0.029)	68305	21.289	21.236	0.053	(0.036)	87841

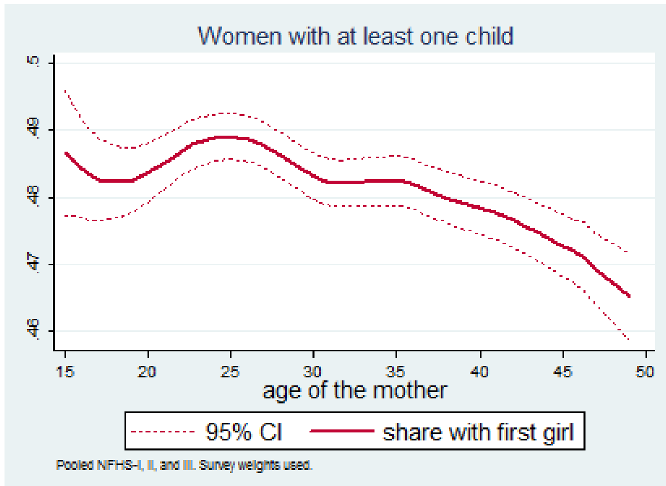
All women aged 15–49 with at least one child ever born. Using survey weights. ***, **, and * indicate significance at 1%, 5% and 10% levels. Anemia, height, weight-for-height, BMI only in waves 1998 & 2005.

4 Non-parametric evidence

Figure 1 reports non-parametric evidence (using Kernel-weighted local polynomial smoothing) of the share of women with first-born girl, by the age of the mother observed at the time of the survey. The confidence intervals at the 95% level of significance are also reported. The figure shows that the sex ratio for first births is in the biological range for women age 15–30, while the share of women with first-born girl (the inverse of the sex ratio, defined as the ratio of male to female children born) shows a steep decreasing trend for women older than 30 (below the biological range). This is equivalent to saying that, among women with at least one child ever born, the share of those who had a first-born girl (as opposed to a first-born boy) is lower among older women, especially older than 30. By pooling surveys for different years, it is possible to focus on age-effects rather than cohort of birth effects.

Appendix Figure A.5 shows that the decreasing sex ratio is found only among women with zero years of education, while the sex ratio is stable for women with at least one year of education. Also, figure A.6 shows that the decreasing trend is found for women in rural areas, while no clear pattern is visible for those in urban areas. Since education can be considered a proxy for socioeconomic status and that urban households have better access to

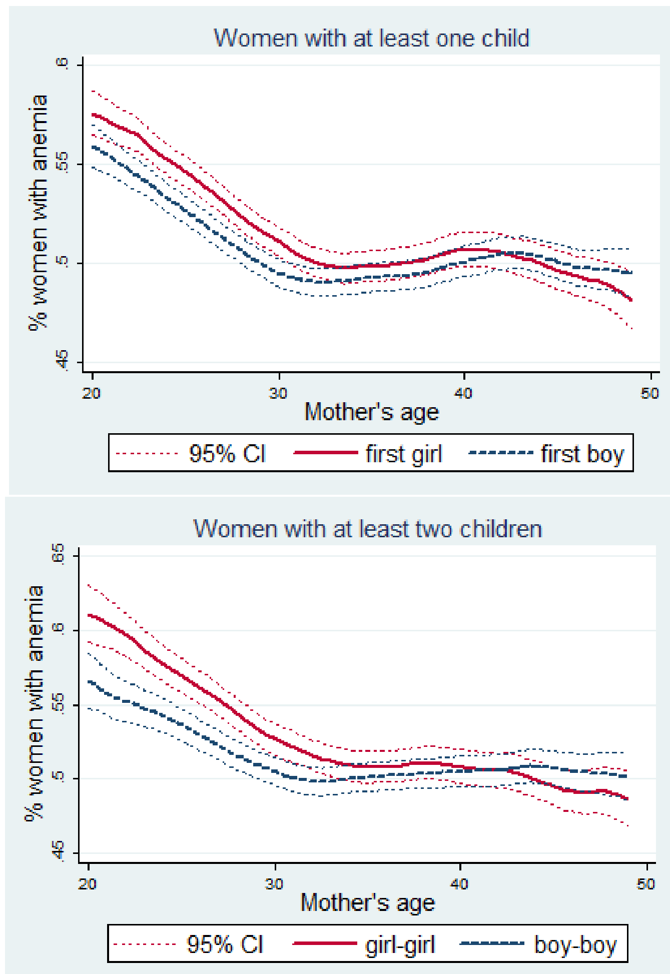
Figure 1: Share of women with first-born girl by mother's age



health facilities, this suggests an intuitive result: women who are worse-off and with poorer access to health services are the ones who are most likely to suffer from causes related to childbirth (aggravated by son-preferring fertility stopping rules) and thus most likely to die. Figure A.7 shows the non-parametric graph for the subsample of women with at least two children ever born. As expected, the share of women with two boys (girls) increases (decreases) with the mother's age, while the shares of women with mixed sex composition (boy-girl or girl-boy) lies always in the middle.

Next, I turn to the health indicators. The upper panel of figure 2 shows non-parametric evidence of the incidence of anemia (the % of women who are anemic) by the sex of the first-born and mother's age. It is first worth noting that the overall pattern reflects the common prevalence of anemia; higher during adolescence and during the reproductive years, decreasing and stabilizing afterwards. The incidence of anemia is significantly higher for women with first-born girl up to age 30, while there is no difference for older women. The bottom panel of figure 2 shows the non-parametric graph for the incidence of anemia among women with at least two children ever born. As it would be expected in a context where son-preferring fertility stopping rules are commonly practiced, the incidence of anemia for young women with two daughters is higher than that for women who had only one daughter as first-born (shown in the upper panel), and it is significantly higher than for women with two boys. The pattern by mother's age is similar

Figure 2: % anemic women, by age of the mother at the time of the survey



to that shown in figure 2. Intuitively, this results might be the consequence of the fact that women with girls among earlier-born children tend to be pregnant or breastfeeding more often during their reproductive years than women who had boys. I replicate the same non-parametric graphs using the subsamples of women who are not currently pregnant or breastfeeding. Appendix figure A.8 shows that the patterns are confirmed also when using these subsamples. This suggests that the differences found in the prevalence of anemia between women with first-born daughter and son might mostly be driven by birth spacing or other factors.

Figure 3 shows the non-parametric graph for weight-for-height.[16] Both the upper and bottom panels (for women with at least one or two children, respectively) show a similar pattern. Overall, weight-for-height increases for women older than 40, reflecting the fact that women tend to gain weight as they get older. There are no differences in weight-for-height for women below age 40 (even though the bottom panel shows a slightly higher weight-for-height for women with two sons). Among older women, those with daughters among first-born children show better anthropometric outcomes (similar results obtained for the body mass index in appendix figure A.9)

Lastly, figure 4 reports nonparametric evidence for height by the sex of the first-born and mother's age. Height reflects nutritional status in childhood and therefore should be affected by fertility behavior to a lower extent than other nutritional indicators.[17] As expected, the figure shows that women's height does not significantly differ by the sex of their first-born child.

5 Empirical strategy and results

To understand whether son preference may lead to selective maternal mortality through son-preferring fertility behavior, I test how the sex of the first-born affects reproductive behavior. In addition to confirming findings from the demographic literature documenting son-preferring reproductive behavior in India, I also provide estimates of the effects on child spacing that include mother fixed effects. This allows to control for unobserved mother's heterogeneity that may affect the estimates. To my knowledge, this approach has not been previously used. In this section I first describe the empirical strategy used to estimate the effects of the sex of the first-born

[16] Weight-for-Height standard deviations from the reference median based on the DHS reference standard.

[17] However, evidence suggests that weight and height can be affected if pregnancies occur before they have completed their adolescence growth spurt (World Bank, 1996).

Figure 3: Weight-for-height, by age of the mother at the time of the survey

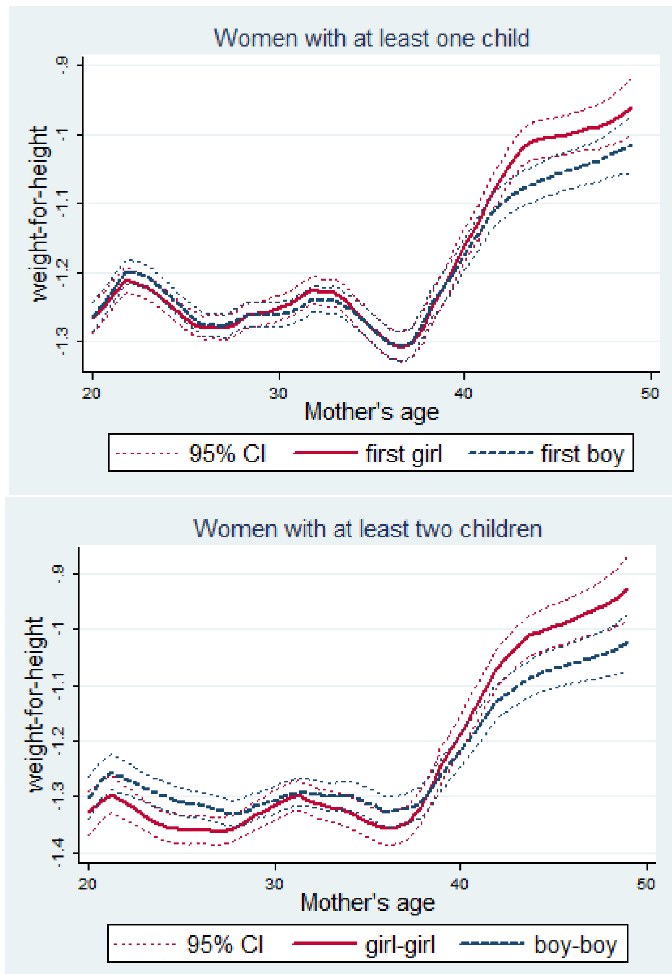
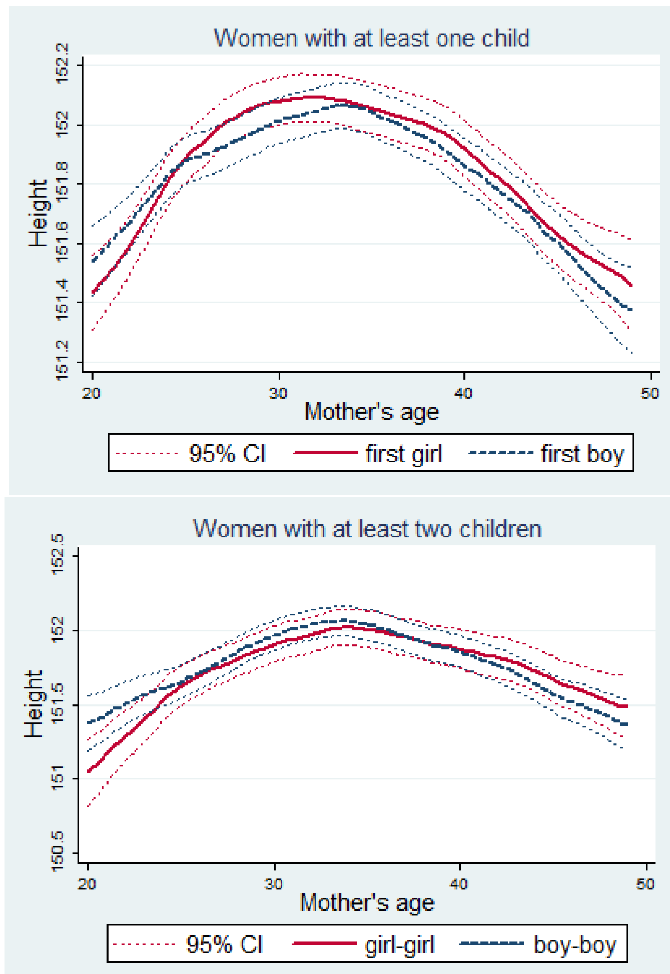


Figure 4: Height, by age of the mother at the time of the survey



on fertility behavior, preferences, and spacing and then discuss the results. Second, I confirm findings from the non-parametric graphs in section 4 by using econometric analysis.

5.1 Fertility regressions

To investigate the effect of son preference on fertility, I analyze whether the sex of the first-born affects realized fertility, by considering *all children ever born* to a woman[18]. Given that the sex of the first born can reasonably be considered as random, the empirical strategy should be straightforward: unbiased estimates should be obtained by simply regressing the number of children ever born on the sex of the first-born. However, as discussed in the previous section, there are differences in observables between women with a first-born boy or girl. These differences may have developed over time as a consequence of son-preferring fertility stopping rules (ie, women with a first-born daughter tend to have more children, and higher fertility may influence their health status and ultimately their survival), or may be due to selective recall bias. Even though these types of bias would lead to an underestimation of the effect of first-born girls on fertility, in all regressions I include a set of observable covariates in order to reduce this bias. If couples desire to have male children, women with first-born daughter should exhibit higher fertility than women with first-born son, after controlling for observable characteristics.

In order to understand if the sex of the first-born also affects the *desire* to have more children (which may differ from *realized* fertility), I construct a dummy variable based on the survey question: *'Would you like to have another child, or would you prefer not to have anymore children?'*. I also examine if the current use of contraceptives (and separately, of sterilization that is the most common contraceptive method in India), and the probability to have had a terminated pregnancy differs for women with a first-born daughter and first-born son.

I estimate the following regression (for women with one child or more):

$$y_{i,s,b,r} = \beta_1 (firstborngirl)_i + \gamma X_{i,r,b,s} + \alpha_s + \mu_b + \delta_s + \epsilon_{i,s,b,r} \qquad (1)$$

[18] The number of children ever born includes all births to each woman, including dead children and children living elsewhere. Compared to the number of surviving children, this is a more pertinent measure of fertility as it is not affected by differential mortality potentially *caused* by son preference. Clark (2000) uses the number of children ever born in studying fertility stopping rules in India.

with mother i, resident in state s, born in year b, surveyed in round r. $y_{i,s,b,r}$ is the dependent variable, which can alternatively be the # of children ever born, a dummy equal to one if the woman reports the desire to have more children, and a dummy for the current use of any contraceptive methods, of being sterilized, and of having had an abortion. $firstborngirl$ indicates whether the first child ever born is a girl (as opposed to a boy); $X_{i,s,b,r}$ is the set of covariates including: age of the mother (and age squared), age of the mother's partner, seven 5-year age groups, age at first marriage, age at first birth, number of years of education (of the mother and her partner), a wealth index, urban dummy, caste, and religion.[19,20] α_s, γ_b, δ_r are state, cohort of birth, and survey round fixed effects, respectively. Given that there is differential mortality across genders, in all regressions I control for the survival status of the child (a dummy for whether the first-born child is dead and its interaction with $firstborngirl$). Identification relies on the assumption that the sex of the first-born is exogenous (uncorrelated with the error term) after conditioning for the observable characteristics.

Regression 1 is estimated using OLS when the dependent variable is the # of children ever born, and using a probit model for the probability of desiring more children, of using a contraceptive method, of being sterilized, and of having had an abortion.

If women practice son-preferring fertility stopping behavior, I expect to find $\beta_1 > 0$ for the regressions for realized (and desired) fertility, and abortion, while $\beta_1 < 0$ for the use of contraceptives and sterilization. The analysis mainly focuses on the sex of the first-born to maintain a causal interpretation (Dahl and Moretti, 2008). Since fertility decisions and decisions related to marital status and living arrangements (as shown in section 6.4) are probably taken simultaneously, the sex of the first two, three, etc, children born cannot be considered as random anymore. However, the effects should be stronger for women who had more children and more daughters among earlier-born children. There is also a potential issue of selection bias when considering the subsamples of women with two, three or more children and comparing realized fertility depending on all children already born. For example, consider evaluating the effect of the first two girls on fertility. In

[19] The caste and religion variables are from the woman dataset. Caste include Scheduled Caste, and Scheduled Tribe (other backward classes (OBC) and other castes are the omitted category). Religion include Hindu, Sikh, Muslims, Christian (Buddhist, Jain, and other religions are the omitted category).

[20] The wealth index is a principal component index of services and durable goods owned by the household, including electricity, radio, television, refrigerator, bicycle, motorcycle, and car.

this case, the sample of women with two or more children would be used. While the sex of the first-born can be considered as random, following children represent a choice that may correlate with other household decisions. Moreover, if the choice to continue having children after a first-born girl or a first-born boy is correlated with unobserved characteristics, the estimated coefficients would be biased. Since it is possible to observe women that progress to next parity only if they have reached the previous one, those who have had children after daughters might be different than women who progressed after sons.[21]

Table 3 reports the results. Column (2) shows that women with first-born girl have 0.28 more children ever born (significant at 1% level), which implies a percent effect of 8.7 (with respect to women with first-born son). The probabilities of using contraceptives and of being sterilized decrease by 7.3 and 8.2 percentage points for women who had a first-born girl, respectively. Moreover, an abortion is 1.3 percent more likely for women with first-born girl. Finally, women with first-born girl are 5.1 percentage points more likely to report that they want more children. All these results suggest that the sex of the first-born strongly affects subsequent *realized* and *desired* fertility and are consistent with the idea that women intentionally continue bearing children until they have the desired number of sons. In particular, the effect for the use of contraceptives and abortion points to a conscious decision.

5.2 Birth spacing regressions

If there is pressure on women (or couples) to have male children, mothers who had daughters first might try to conceive another child sooner after the birth of a daughter (compared to those who had sons). To study the association between the length of birth intervals and the sex composition of earlier-born children, I use the sample of all children ever born to each woman (the unit of analysis is the child, and not the woman as in the previous section).[22] I analyze the average succeeding birth interval (expressed in # of months), which indicates how long the mother waits to have another child *after* the realization of the sex of the previous. Given that the NFHS includes information on the interval between each birth for all children, I can use mother fixed effects and exploit the variation in the length of the interval within the fertility history of each woman. This allows to control for all observables and unobservables characteristics that may correlate with

[21] This is a similar empirical issue to that of *dynamic selection bias*, exposed in Cameron and Heckman (1998), where they study schooling decisions over the life cycle.

[22] I exclude births of order higher than 10.

Table 3: Fertility outcomes by the sex of the first-born

dep. var.	# children ever born		contraceptives		sterilized		abortion		wants more ch	
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)
first-born girl	0.196***	0.283***	-0.048***	-0.073***	-0.052***	-0.082***	0.011***	0.013***	0.047***	0.051***
	(0.011)	(0.008)	(0.003)	(0.003)	(0.002)	(0.003)	(0.002)	(0.002)	(0.002)	(0.002)
first child dead		0.783***		-0.160***		-0.151***		0.029***		0.165***
		(0.019)		(0.005)		(0.004)		(0.005)		(0.006)
first child dead*		-0.289***		0.073***		0.094***		-0.007		-0.044***
first-born girl		(0.026)		(0.008)		(0.009)		(0.007)		(0.003)
Survey dummies	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
State Fes	No	Yes	No	Yes	No	Yes	No	Yes	No	Yes
Birth year Fes	No	Yes	No	Yes	No	Yes	No	Yes	No	Yes
Religion, Caste	No	Yes	No	Yes	No	Yes	No	Yes	No	Yes
Age group	No	Yes	No	Yes	No	Yes	No	Yes	No	Yes
Other controls	No	Yes	No	Yes	No	Yes	No	Yes	No	Yes
Observations	244803	224534	244803	224534	244803	224493	244783	224516	235206	224311
R2 (or Pseudo R2)	0.008	0.549	0.017	0.209	0.007	0.250	0.001	0.029	0.012	0.392
Percent effect	6.0	8.7	8.8	13.5	13.0	20.5	5.8	6.8	23.5	25.5
Mean first boy	3.25	3.25	0.54	0.54	0.40	0.40	0.19	0.19	0.20	0.20

OLS estimates in columns 1-2. Probit estimates in columns 3-10 (marginal effects reported). Robust standard errors adjusted for clustering at the household level in parentheses. All women aged 15–49. Using survey weights. Each regression also includes 7 five-year age groups dummies, age at first birth and marriage, mother and partner's age and education years, urban dummy, whether first child died (and its interaction with first born girl), and a durable index. ***, **, and * indicate significance at 1%, 5% and 10% levels.

the error term. I estimate the following regression:

$$y_{i,j,t} = \beta_1 \; girl_{i,j,t-1} + \sum_k \gamma_k birthorder_{i,j,k} + \alpha_j + \epsilon_{i,j,t} \qquad (2)$$

with child i, mother j, t: current child, $t-1$: preceding child. $y_{i,j,t}$ is the succeeding birth interval in months (the time between the birth of child $t-1$ and t), or a dummy =1 if interval <24 or 15 months. α_j mother fixed effects. $birthorder_{i,j,k}$ set of k dummies for birth order. Identification with FEs requires to consider mothers with at least 3 children ever born. I also control for the survival status of the previous child (the one who opens the interval) and its interaction with her/his sex (*girl*). The succeeding birth interval at each birth should be shorter ($\beta_1 < 0$) if a girl is born rather than a boy.

A potential source of bias is represented by sex-selective abortion. There is evidence that female fetuses might be selectively aborted and that this is more common among second or higher order children (Bhalotra and Cochrane, 2010). If this is the case, abortion would be more likely after the birth of a first-born girl rather than a boy. Since the NFHS does not contain detailed information on abortion (which would anyway be highly subject to under-reporting) it is not possible to directly control for this. However, it is important to note that this would lead to an underestimation of the (negative) effect of first-born girl on the length of the birth interval as more time will be needed to have a following completed pregnancy after an abortion.

The first two columns of table 4 show the estimated effect of having a daughter on the length of the birth interval (in # of months) without including the mother fixed effects. The coefficients are -1.03 and -0.72 (both highly statistically significant at 1% level) meaning that, compared to mothers who had a boy, those who have a girl wait on average 0.7-1 months less to have the next child. Column (3) includes the mother fixed effects and yelds a similar coefficient (-0.83 months). Columns (4)-(9) show the results for the probability that the interval is shorter than 24 (columns 4-6) or 15 months (columns 7-9), which are the intervals below which child's and mother's health can be negatively affected. Column (9) shows that having a girl increases the likelihood of the interval being shorter that 15 months by 0.8 percentage points (significant at 1% level), which is equivalent to a large 8.9 percent effect. These findings are consistent with son-preferring fertility stopping rules.

Table 4: Birth spacing by the sex of the first-born

dep. var.	# months			interval < 24 months			interval < 15 months		
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)
girl	-1.031***	-0.723***	-0.834***	0.022***	0.017***	0.020***	0.009***	0.007***	0.008***
	(0.058)	(0.058)	(0.089)	(0.002)	(0.002)	(0.003)	(0.001)	(0.001)	(0.001)
dead child	-6.540***	-4.735***	-4.692***	0.197***	0.165***	0.150***	0.141***	0.127***	0.111***
	(0.114)	(0.114)	(0.154)	(0.003)	(0.003)	(0.005)	(0.002)	(0.002)	(0.003)
girl*dead child	1.266***	0.891***	1.150***	-0.032***	-0.025***	-0.033***	-0.013***	-0.010***	-0.009*
	(0.146)	(0.146)	(0.189)	(0.005)	(0.005)	(0.006)	(0.004)	(0.004)	(0.005)
# children ever born		-1.776***			0.032***			0.014***	
		(0.018)			(0.000)			(0.000)	
Birth-order Fes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Mother Fes	No	No	Yes	No	No	Yes	No	No	Yes
Observations	552431	552431	552431	552431	552431	552431	552431	552431	552431
R-squared	0.017	0.044	0.505	0.021	0.033	0.401	0.030	0.036	0.402
% effect	3.2	2.2	2.6	6.5	5	5.9	10	7.8	8.9
Mean first-born boy	32.2	32.2	32.2	0.34	0.34	0.34	0.09	0.09	0.09

OLS estimates. Robust standard errors adjusted for clustering at the household level in parentheses. All women aged 15–49. Excludes births of birth order 10 or higher. Using survey weights. ***, **, and * indicate significance at 1%, 5% and 10% levels.

Table 5: Differences in the conditional age distribution between women with first-born girl or boy from ordered logit regression. Dep. variable: 5-year age groups.

women with:	at least one child first girl vs first boy		at least two children girl-girl vs boy-boy	
	change	95% CI	change	95% CI
	(1)	(2)	(3)	(4)
$Pr(y = 15 - 19 \vert x)$	0.0001	[0.0001, 0.0001]	0.0000	[0.0000, 0.0000]
$Pr(y = 20 - 24 \vert x)$	0.0022	[0.0019, 0.0026]	0.0018	[0.0016, 0.0021]
$Pr(y = 25 - 29 \vert x)$	0.0210	[0.0182, 0.0239]	0.0240	[0.0209, 0.0272]
$Pr(y = 30 - 34 \vert x)$	0.0022	[0.0015, 0.0030]	0.0346	[0.0305, 0.0388]
$Pr(y = 35 - 39 \vert x)$	-0.0218	[-0.0247, -0.0189]	-0.0468	[-0.0525, -0.0412]
$Pr(y = 40 - 44 \vert x)$	-0.0035	[-0.0039, -0.0030]	-0.0124	[-0.0140, -0.0109]
$Pr(y = 45 - 49 \vert x)$	-0.0003	[-0.0004, -0.0003]	-0.0013	[-0.0015, -0.0011]
Observations	224534		186992	
Pseudo R-sq	0.446		0.423	

Computed from the ordered logit estimates. The values reported are the changes in predicted probabilities when the 'first-born girl' changes from 0 to 1, with all the other independent variables set at their mean value. All regressions include individual controls as in table 3. Changes are reported with 95% confidence intervals by the delta method in parentheses. Robust standard errors adjusted for clustering at the household level in parentheses. All women aged 15–49. Excludes births of birth order 10 or higher. Using survey weights. ***, **, and * indicate significance at 1%, 5% and 10% levels.

5.3 Age structure

To confirm the non-parametric evidence shown in figure 1, I estimate an ordered logit model in which the dependent variable is a categorical variable with 7 categories for each of the 5-year age groups (from 15–19 to 45–49). The controls are similar to the ones included in regression 1. After estimating the ordered logit model, I compute changes in the outcomes predicted probabilities when the independent variable 'first-born girl' (or girl-girl when considering the subsample of women with at least two children) changes from 0 to 1, with all the other independent variables are set at their mean value. Changes are reported with 95% confidence intervals. Columns (1) and (3) in table 5 show the changes in the predicted probabilities for women with at least one or two children, respectively. The results confirm that the age distribution for women who had a first-born girl lies to the left of the distribution for women with first-born boy, conditional on the number of children and all other covariates (set at their mean).

5.4 Health outcomes

Next, I estimate the effect of the sex of the first-born on several health and anthropometric outcomes for women in different age groups. I estimate the following regression:

$$
\begin{aligned}
y_{i,s,b,r} = {} & \beta_1 (firstborngirl)_i + \beta_2 (firstborngirl)_i (ageover30)_i + \\
& + \gamma X_{i,r,b,s} + \alpha_s + \mu_b + \delta_s + \epsilon_{i,s,b,r}
\end{aligned} \tag{3}
$$

with mother i, resident in state s, born in year b, surveyed in round r. $y_{i,s,b,r}$ is the the incidence of anemia (the % of women who are anemic), weight-for-height, and body mass index (BMI). $firstborngirl$ indicates whether the first child ever born is a girl (as opposed to a boy), which is also interacted with a dummy equal to one if the woman is older than 30); $X_{i,s,b,r}$ is the same set of covariates as in regression 1. α_s, γ_b, δ_r are state, cohort of birth, and survey round fixed effects, respectively. In addition, I control for potential confounding factors such as pregnancy and breastfeeding status of mothers which may be correlated with the error term. Moreover, as a proxy for women's health status before birth, I control for height which reflects nutritional status in childhood. Identification relies on the assumption that the sex of the first-born is exogenous (uncorrelated with the error term) after conditioning for the observable characteristics. Regression 3 is estimated using OLS.

Columns (1) to (4) of table 6 shows the results for the incidence of anemia. Consistent with the non-parametric evidence shown in figure 2, women under the age of 30 with a first-born girl are 1.5 percentage points more likely to be anemic, while this difference disappears for women older than 30 (column 1). Columns (2) to (4) show that this result is robust to the inclusion of the number of children ever born, dummies for pregnancy and breastfeeding status and woman's height among the regressors. Consistent with the medical evidence, women with many children, pregnant or breastfeeding are more likely to be anemic. Columns (5) to (12) of table 6 show the results for weight-for-height and BMI. First, both these indicators of women's nutritional status are negatively correlated with the number of children ever born (controlling for all other covariates, including proxies for socioeconomic status). The estimates show that, among women age 30–49, those with first-born girl show significantly higher weight-for-height and BMI.

Table 6: Health outcomes and sex of first-born, by age group

dep. var.	% anemic				weight-for-height				BMI			
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)	(11)	(12)
first girl	0.015***	0.012***	0.012***	0.012***	-0.008	-0.000	0.005	0.005	0.011	0.003	0.018	0.018
	(0.004)	(0.004)	(0.004)	(0.004)	(0.011)	(0.011)	(0.011)	(0.011)	(0.032)	(0.032)	(0.032)	(0.032)
first girl*	-0.012**	-0.011*	-0.012**	-0.012**	0.029**	0.025*	0.033***	0.033***	0.084**	0.091**	0.119***	0.118***
age≥30	(0.006)	(0.006)	(0.006)	(0.006)	(0.013)	(0.013)	(0.013)	(0.013)	(0.041)	(0.041)	(0.041)	(0.041)
currently		0.042***	0.039***	0.039***		-0.071***	-0.038***	-0.037***		-0.278***	-0.171***	-0.178***
breastfeeding		(0.005)	(0.005)	(0.005)		(0.008)	(0.008)	(0.008)		(0.027)	(0.027)	(0.027)
pregnant		0.017**	0.016**	0.017**		-0.267***	-0.262***	-0.262***		1.096***	1.112***	1.112***
		(0.008)	(0.008)	(0.008)		(0.013)	(0.013)	(0.013)		(0.045)	(0.045)	(0.044)
# children			0.004***	0.004***			-0.037***	-0.037***			-0.118***	-0.117***
			(0.001)	(0.001)			(0.003)	(0.003)			(0.008)	(0.008)
height				-0.001***				0.005***				-0.036***
				(0.000)				(0.001)				(0.002)
Observations	136419	136419	136419	136216	140895	140895	140895	140895	142201	142201	142201	142200
R-squared	0.047	0.048	0.048	0.048	0.232	0.235	0.236	0.237	0.248	0.253	0.255	0.257

OLS estimates. Robust standard errors adjusted for clustering at the household level in parentheses. All women aged 15–49. Using survey weights. Includes age at first birth and marriage, mother and partner's age and education years, urban dummy, whether first child died, and a durable index, birth year dummies, wave dummies, state FEs, religion and caste. ***, **, and * indicate significance at 1%, 5% and 10% levels.

6 Alternative interpretations

I next turn to discussing and testing the plausibility of three alternative interpretations of the results. First, I check whether the results might be driven by biological differences. There is evidence that a female birth is more likely if the mother is malnourished (because male fetuses are more susceptible to death in utero than female) (Andersson and Bergstrom, 1998, Almond and Mazumder, 2011). If this is the case, this interpretation would not be compatible with a pattern in which differences in the nutritional indicators for women with first-born daughter or son arise only *after* the time of first birth. In order to check for this possibility, I plot the prevalence of anemia by the age and sex of the first-born (using Kernel-weighted local polynomial smoothing as in Section 4) to understand when these differences arise. The upper panel of figure 5 shows the % of women (with at least one child) with anemia by the age of the first child and the sex of the first-born. The figure shows that the incidence of anemia is the same for women with first-born daughter or son at the time of birth of the first child (ie, when the first child is aged 0) and that difference develop over time. Specifically, women with first-born daughter start being more likely to be anemic approximately one year after the birth of their daughter, and differences increase over time. The bottom panel of figure 5 focuses on women with at least two children and shows the prevalence of anemia by the age of the second child for women who had at least two daughters or two sons among their earlier-born children. The figure shows that, at the time of birth of the second child, women who already had a first-born daughter are more anemic (but not significantly so) and, again, differences in the incidence of anemia increase over the years. This evidence is not consistent with the explanation based on biological differences. Instead, it probably reflects son-preferring fertility behavior in which women with daughters want to 'try again for a son' by having more children (and shortly spaced). This behavior is reflected in higher anemia prevalence, which is an indicator that responds quickly to reproductive behavior and it is associated with maternal mortality. Finally, it is important to note that the results for the anemia prevalence in table 6 are unchanged when mother's height is included among the controls, thus suggesting that the results are not driven by the nutritional status before birth.

The second check deals with selective recall bias. With son preference, women might under-report female births, especially those those occurred back in the past. In this regard, it is important to note that several studies find that son preference is not decreasing over time in India (if anything,

Figure 5: Share of women with anemia, by age of the first (or first two) child (children) and the sex composition

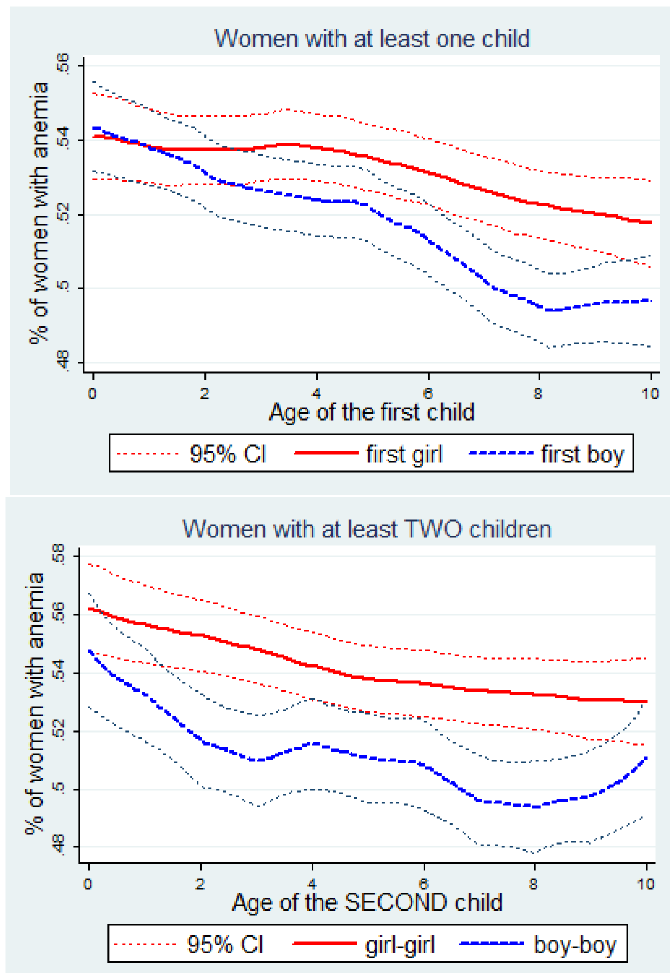


Table 7: Dep. variable: Mother's age

	(1)
first-born girl	-0.287***
	(0.031)
first-born girl*wave2	0.026
	(0.044)
first-born girl*wave3	0.080*
	(0.044)
wave2	0.273***
	(0.036)
wave3	0.926***
	(0.039)
Obs	224534
R-squared	0.830

OLS estimates. Robust standard errors adjusted for clustering at the household level in parentheses. All women aged 15–49. Using survey weights. Includes age at first birth and marriage, mother and partner's age and education years, urban dummy, whether first child died, state dummies, # children ever born, and a durable index. ***, **, and * indicate significance at 1%, 5% and 10% levels.

Census data report increasing male-biased sex-ratios at birth). Consistent with nondecreasing son preference, if the results of this paper are driven by selective recall bias (motivated by son preference), then I should not find evidence of decreasing maternal mortality among women with first-born girl in most recent surveys. To check for this possibility, I regress the mother's age on the variable 'first-born girl' and its interaction with the survey round dummies. Table 7 shows that women with first-born girl are younger than those with first-born boy especially among women in the first wave of the NFHS. Interestingly, this effect is attenuated for women in the last two survey rounds. This pattern would not be consistent with sex-selective recall bias but instead with the reduction in maternal mortality that has been taking place in India over the last years.

Finally, one concern may be that the practice of sex-selective abortion may be driving the results. One could argue that older and uneducated women might be the ones who selectively aborted female children, and this might explain why they are fewer as shown in figure A.5 (as they do not report aborted girls). However, the evidence on sex-selective abortion in India suggests that it is more prevalent among more educated women (who have access to the technologies for the detection of the sex of the fetus) and for higher-order children (*after* the first-born). Also, ultrasound technologies became increasingly available only after the end of the 1980s. Therefore, sex-selective abortion should be more frequent among younger women and not the ones that in figure A.5 appear to be missing (older and uneducated women). As a further check, I estimate that the probability that the second-born child is a male on the first-born being a girl. This probability is positive and significant only for younger women (especially if educated and in urban areas): women below 30 are 1.2 percentage points more likely to have a second-born boy if the first is a girl, while no difference for women 30 or older are found. Hence, sex-selective abortion should not affect the result obtained for maternal mortality, that is mostly found for older women.

7 Conclusions

This paper has shown some evidence that differential mortality for mothers with daughters (rather than sons) among earlier-born children might partly explain the phenomenon of missing women at reproductive and older ages in India recently uncovered by Anderson and Ray (2010, 2012). It proposes an original explanation in which maternal mortality can (at least partly) an unintended consequence of son preference through son-preferring fertility behavior.

A Appendix

Figure A.1: Sex ratio (% male births), by birth year of the child

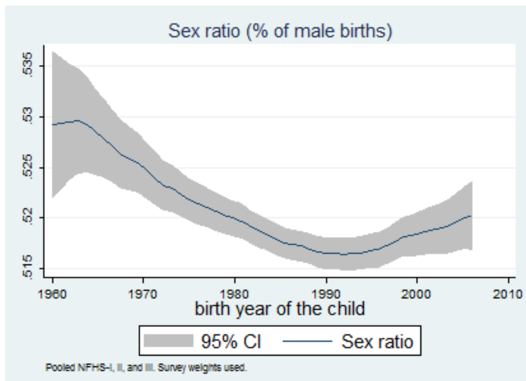


Figure A.2: Sex ratio (% male births), by birth year of the child and DHS wave

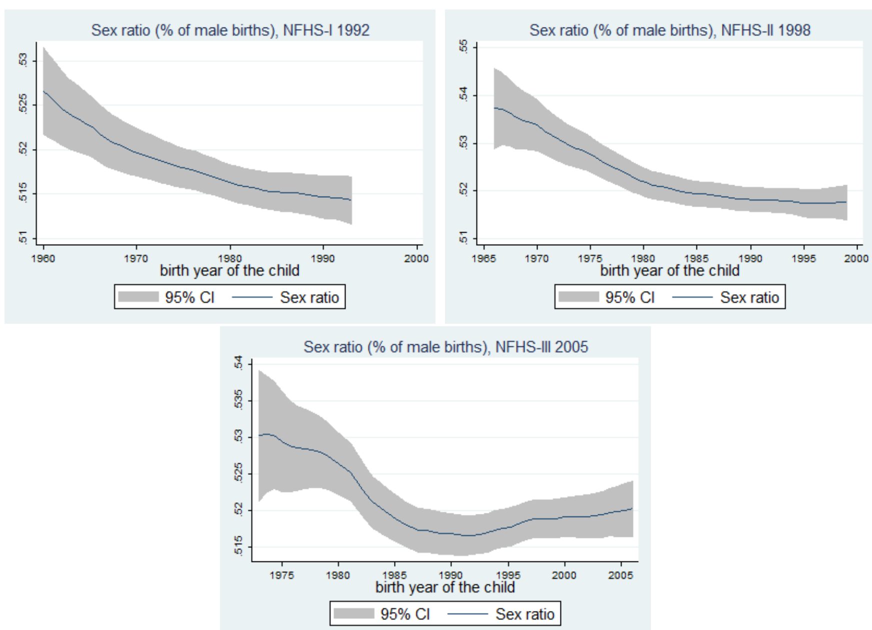


Figure A.3: Sex ratio (% male births), by age of the mother

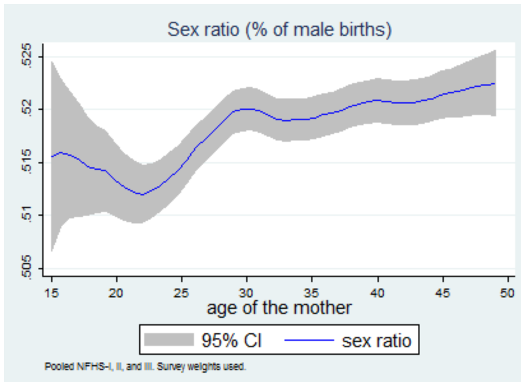


Figure A.4: Sex ratio (% male births), by birth year of the child and birth order

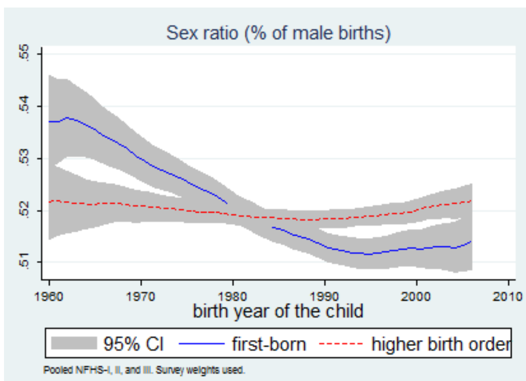


Figure A.5: Share of women with first-born girl by mother's age and education

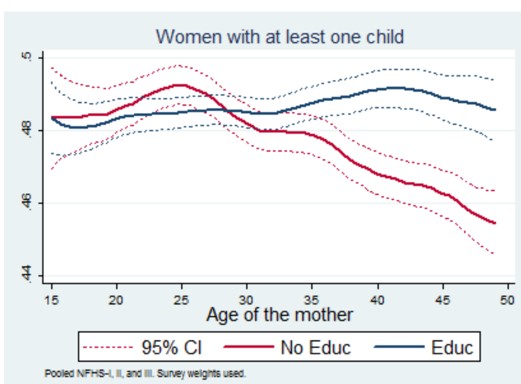


Figure A.6: Share of women with first-born girl by mother's age and urban/rural location

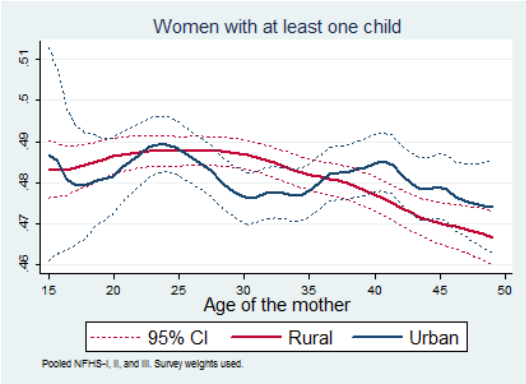


Figure A.7: Share of women by the sex of the first two children mother's age

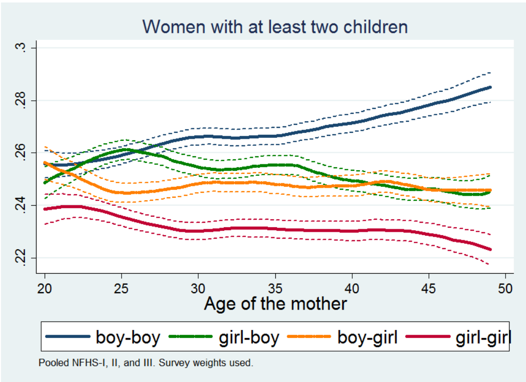


Figure A.8: % anemic women, by age of the mother at the time of the survey. Subsamples of women not currently breastfeeding or pregnant

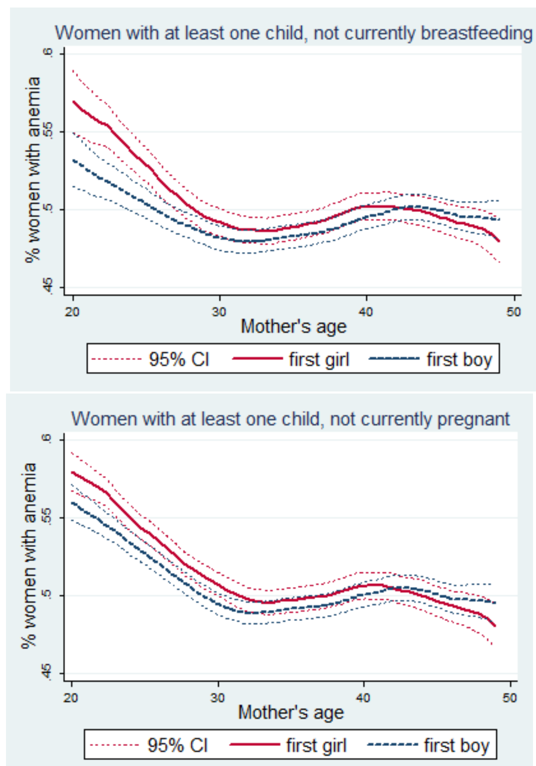


Figure A.9: Body mass index, by sex of the first-born and age of the mother at the time of the survey

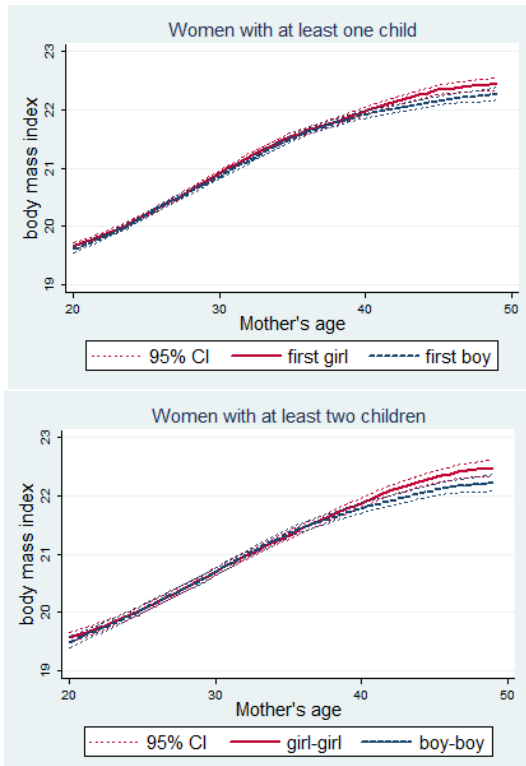
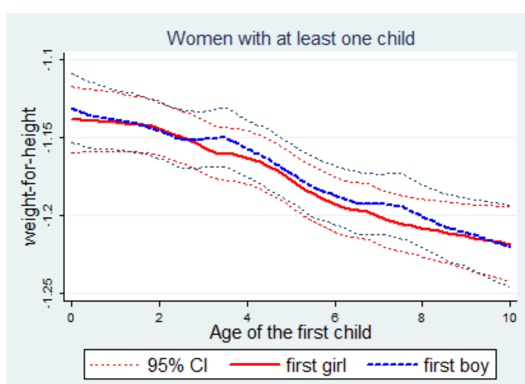


Figure A.10: Weight-for-height, by age of the FIRST child



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