POLICY RESEARCH WORKING PAPER 6802

Why Are Adult Women Missing?

Son Preference and Maternal Survival in India

Annamaria Milazzo

Abstract

This paper is the first to show that excess mortality among adult women can be partly explained by strong preference for male children, the same cultural norm widely known to cause excess mortality before birth or at young ages. Using pooled individual-level data for India, the paper compares the age structure and anemia status of women by the sex of their first-born and uncovers several new findings. First, the share of living women with a first-born girl is a decreasing function of the women's age at the time of the survey. Second, while there are no systematic differences at the time of birth, women with a first-born girl are significantly more likely to develop anemia when young (under the age of 30) and these differences disappear for older women. Moreover, among those in the older age group, they appear to be significantly better off in terms of various predetermined characteristics. These findings are consistent with a selection effect in which maternal and adult mortality is higher for women with first-born girls, especially the poor and uneducated with limited access to health care and prenatal sex diagnostic technologies. To ensure the desired sex composition of children, these women resort to a fertility behavior medically known to increase their risk of death. The observed sex ratios for first births imply that 2.2–8.4 percent of women with first-born girls are 'missing' because of son preference between the ages of 30 and 49.

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

ANNAMARIA MILAZZO[†]

---

1   Introduction

Research on missing women has documented severe gender bias at birth and at young ages due to pervasive son preference in India and other Asian countries (Das, 1987; Das Gupta, 1987; Sen, 1990; Arnold, Kishor and Roy, 2002; Jha et al., 2006, 2011). Recent studies confirm these previous findings but also demonstrate that most women are instead missing at adult ages (Anderson and Ray, 2010, 2012; World Bank, 2011).[1,2] Anderson and Ray (2012) find a large state-wise variation in the distribution of excess mortality across age groups in India and argue that the traditional explanation of strong preference for male children can only explain the phenomenon of missing women in states with abnormally male-biased sex ratios *at birth*. They contend that the explanation for excess mortality in adulthood must reside in other, unspecified, factors.[3] This paper provides evidence that, on the contrary, strong son preference is also partly responsible for high mortality rates among adult women.

India has the world's largest number of maternal deaths. 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). It is estimated that for every maternal death there are twenty other women who suffer pregnancy-related illness, infection, injury, and other physical and psychological long-term consequences (Nanda, Switlick, and Lule, 2005; Filippi et al., 2006; UNICEF, 2008). Anemia, one of the most common indirect causes of maternal mortality in Asia, is pervasive among Indian women. Medical research has shown that repeated and closely spaced pregnancies are associated with increased risk of maternal mortality and morbidity. This behavior is common among women who have only daughters and want to conceive again for a son. The diffusion of prenatal sex diagnostic technologies in the late 1980s has been found to be associated with increasing sex ratios at birth (the fraction of male births) in India which is strongly suggestive of sex selection (Arnold, Kishor, and Roy, 2002; Bhalotra and Cochrane, 2010). Research

---

[1]Missing women, or excess female mortality, is based on a counterfactual in which there is no discrimination based on gender (Sen, 1990). As a counterfactual, Anderson and Ray (2010) use developed countries, while in their study on India, Anderson and Ray (2012) also use Kerala as an alternative benchmark.

[2]Klasen and Vollmer (2013) challenge the methodology developed in Anderson and Ray (2010) and argue that the data used in their computations are largely unreliable (especially those on adult mortality for sub-Saharan Africa) and use a reference benchmark that does not account for differences in the disease and mortality environment. Anderson and Ray (2012) replicate their decomposition exercise for India and find that their results are robust to using Kerala as a benchmark (rather than developed countries), though of smaller magnitude. It is important to note that the issues at stake in this debate do not alter the motivation for this paper, which is to analyze the factors contributing to the high number of maternal deaths in India. High maternal mortality is certainly partly due to generally poor health conditions and services and, as argued in this paper, it is *aggravated* by strong son preference.

[3]Jayaraman, Ray, and Wang (2013) find evidence that women seek treatment for eye care later than men, while they do not seem to receive less surgical care in India.

has shown that sex-selective abortion is more common among educated women and *after* the first child is born (Das Gupta, 1987; Jha et al., 2011).

The central idea of this paper is that maternal and adult mortality is higher among women with a first-born daughter. As a result of strong son preference, such women are more likely to engage in fertility behavior that adversely affects their health. This effect should and is found to be stronger among women with generally more vulnerable health, lower access to maternal care and to technologies for fetal sex determination, typically the poor and uneducated.

I pool three rounds of the India National Family Health Survey (NFHS) and find several significant findings. First, there are no systematic differences between women whose first birth is male and women whose first birth is female close to the time they give birth. Significant differences appear in the sample of older women (i.e., aged 30 to 49), both in terms of variables related to fertility behavior, and predetermined characteristics. The latter is suggestive of selection among women with a first-born girl as a result of higher mortality. The empirical strategy rests on the identifying assumption of conditional exogeneity of the sex of the first-born. Second, compared to women with first-born sons, women with first-born daughters are 5.1 percentage points more likely to desire more children, 6.8 percentage points less likely to be sterilized, 2.5 percentage points more likely to have undergone an ultrasound test, 1.1 percentage points more likely to have had a terminated pregnancy, and to have given birth to 0.3 additional children. These results reinforce findings from the existing literature on son-preferring stopping behavior and sex-selective abortion. I investigate heterogeneous effects by women's education level (with education as a proxy for socio-economic status) and find that, as expected, educated women with a first-born girl are more likely to have had a terminated pregnancy and to exhibit an attenuated effect on fertility (compared to uneducated women). Moreover, a first-born girl increases the probability of a short birth interval by 0.5 percentage points (the effect is 0.8 percentage points for women with a second-born girl conditional on a first-born girl). This result is robust to the inclusion of mother fixed effects and is larger for uneducated women. Although the effects of two or more successive daughters on fertility outcomes are consistently larger (reflecting increasing pressure for having a son), I focus on the sex of the first-born to maintain a causal interpretation. Third, among women under the age of 30, those with a first-born girl develop anemia at a higher rate: they are in fact 1.4 percentage points more likely to be anemic (compared to those with a first-born boy), even after controlling for height, current pregnancy and breastfeeding status. Fourth, differences in anemia prevalence disappear for women older than 30. Fifth, the share of living women whose first birth is female decreases with mother's age at the time of the survey, especially among the uneducated. This suggests that maternal and adult mortality is higher among uneducated women with a first-born girl. State-wise variation confirms that this decreasing pattern is found in states with the highest proportion of missing women of adult ages (Assam and Bihar), and not in states with the most

male-biased sex ratios at birth (Punjab and Haryana).

I address potential alternative interpretations of the findings. First, based on evidence that poor maternal nutrition might be associated with the sex of the child (Mathews, Johnson, and Neil, 2008; Almond and Mazumder, 2011), I estimate the timing of the emergence of anemia differences. If predetermined health affects the sex of the child, differential incidence of anemia between women with a first-born girl or boy should be present at the time of birth (and even *before*, but this cannot be verified with the data at hand). The results indicate no differential incidence of anemia close to the time of birth. Women with a first-born girl start developing anemia approximately one and a half years after birth. Therefore, biological differences do not seem to explain the anemia pattern, which is rather consistent with behavioral factors. Second, I test whether the estimates are affected by selective reporting bias according to which women might under-report female births, and find no evidence supporting it. Third, I show nonparametric evidence that the decreasing share of women as a function of their age (found for the uneducated) cannot be affected by the diffusion of the practice of sex-selective abortion.[4] Finally, I check whether women suffer other forms of overt discrimination as direct 'punishment' for having had a first-born daughter. I do not find evidence that they are punished through withholding of food (or receive less iron supplements), while they appear to be more likely to have been the victims of severe domestic violence inflicted by their husbands. The effect on violence is specific to the period close to the time of birth and to uneducated women. Therefore, increased violence is possibly contributing to higher mortality among women with daughters as the result of strong preference for male children. All these findings are consistent with a model in which son preference induces women whose first birth is female into a fertility behavior that is medically known to be associated with higher risk of maternal mortality and morbidity. This pattern is especially found among uneducated women who have poorer access to maternal care. These women develop anemia at a higher rate and appear to suffer a higher risk of death.

This paper is related to three strands of literature. The first is the recent literature on the age distribution of missing women (Anderson and Ray, 2010, 2012; World Bank, 2011). While previous studies use aggregate data on mortality and find that most women are missing in adulthood, this paper uses individual-level data to examine the possible channels generating this phenomenon.

Second, it is related to the literature on the effects of son preference on children's health. Son-preferring fertility stopping rules can in fact lead to less resources being allocated to female children within the household even *unintentionally* on the part of parents. Jayachandran and Kuziemko

---

[4]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, Kishor, and Roy, 2002; Bhalotra and Cochrane, 2010).

(2011) show that gender disparities in breastfeeding duration arise because parents want to try again for a son and therefore girls are weaned sooner than boys as a result. Jensen (2003) finds that the gender differential in education is the consequence of the fact that girls tend to be born in larger households. Barcellos, Carvalho, and Lleras-Muney (2014) develop a methodology to account for the effect of son-preferring stopping rules on family size and find that boys still receive more parental inputs (including childcare time, breastfeeding, vaccinations, and vitamin supplementation) than girls in India. Recently, Jayachandran and Pande (2013) find that the cultural norm of eldest son preference can in part explain the child height gap between India and Africa. This paper contributes to this literature by suggesting another channel through which strong son preference can affect children's health, namely higher mortality of mothers (which often results in mortality of the newborn or increased health risks for older children).

Third, it refers to the literature on the consequences of the diffusion of prenatal sex diagnostic technologies and sex-selective abortion. Research has shown evidence of massive female foeticide in India, especially from the late 1980s with the diffusion of ultrasound testing (Arnold, Kishor, and Roy, 2002; Jha et al., 2006; Bhalotra and Cochrane, 2010; Jha et al., 2011).[5] The ability to detect the sex of the fetus has also been found to be associated with preferential prenatal treatment for boys (Bharadwaj and Lakdawala, 2013). Other authors have argued that sex selective abortion implies a substitution in discrimination against female children from the postnatal to the prenatal period (Goodkind, 1999). Hu and Schlosser (2011) find that the practice of selective abortion is associated with lower malnutrition among girls likely due to differential reductions in family size (larger for girls than boys). This paper provides evidence that, in addition to its general health benefits, easier access to maternal antenatal care, and in particular to prenatal sex determination technologies, can also reduce the risk of maternal mortality as women tend to rely less on a fertility behavior that entails major risks to their health. Importantly, lower maternal mortality does not necessitate easier access to sex-selective abortion, but is instead the outcome of improved maternal care for all women, especially those practicing son-preferring fertility behavior due to increased access to treatment of the complications deriving from it.

The remainder of the paper is organized as follows: section 2 provides some background information on the prevalence of maternal mortality and anemia in India; section 3 describes the data; section 4 presents nonparametric evidence; section 5 introduces the empirical methodology used to analyze fertility outcomes, birth spacing, health outcomes, and describes the results; section 6 discusses alternative interpretations; section 7 reports some calculations of the implied fraction of missing women due to son preference; and section 8 concludes.

---

[5]Akbulut-Yuksel, and Rosenblum (2012) find that, after the initial introduction of ultrasound, a recent rapid spread in ultrasound use in the 2000s in some parts of India is not associated with further increase in sex selection.

2 Son preference, fertility behavior and maternal health outcomes

2.1 Maternal mortality, morbidity, and anemia prevalence in India

Maternal mortality and morbidity related to childbirth are among the main drivers of excess female mortality during the reproductive years in India (Anderson and Ray, 2010; World Bank, 2011).[6]

The WHO estimates that 117,000 maternal deaths occurred in India in 2005 (out of 536,000 worldwide), more than in any other country (WHO, 2007). These figures are considered to underestimate the real burden of maternal mortality due to problems with its standard definition, measurement issues, and lack of adequate data (AbouZahr, 2003; Koblinsky et al., 2012). A maternal death is defined 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, 2010, p. 10). The dominant direct causes of maternal mortality are severe bleeding, hypertensive diseases, and infections, while indirect causes include anemia, HIV/AIDS and malaria (Ronsmans and Graham, 2006). Many direct causes, like severe bleeding, are themselves symptoms of underlying indirect causes such as anemia (Center for Reproductive Rights, 2008). Kahn et al. (2006) conduct a systematic review of the geographical variation in the distribution of the causes of maternal deaths. They find that anemia is the second most common cause after hemorrhage in Asia (causing 12 and 30 percent of maternal deaths, respectively).[7,8] Unsafe abortion is another direct cause, probably resulting in a significant number of deaths, although precise estimates do not exist due to underreporting and misclassification. Despite abortion having been legalized with the 1971 Medical Termination of Pregnancy Act, many Indian women continue to have illegal abortions performed by unqualified practitioners often in unsanitary conditions. Arnold, Kishor, and Roy (2002) argue that there are several reasons why women do not seek safe and legal abortions. These include that they might want to abort for reasons not allowed by the Act (i.e., abortion for sex selection or due to contraceptive failure); they cannot afford the cost of an abortion performed by a registered medical practitioner at a government-approved facility; or because of the unavailability of such facilities. Even though estimates are likely to be imprecise, Kahn et al. (2006) estimate that abortion contributes to about 6 percent of maternal deaths in Asia.

Although not accounted for in standard estimates, maternal mortality 'implies not only death

---

[6]The other main cause of excess mortality in the reproductive years (between the ages of 15–44) is death due to injuries, while cardiovascular disease is the most common cause among older women (Anderson and Ray, 2010).

[7]While hemorrhage is found to be the most important cause overall in developing countries, other causes are region-specific, such as sepsis and HIV in Africa, anemia in Asia, and abortion in Latin America.

[8]Other studies provide higher estimates. For example, Gragnolati et al. (2005) and Kalaivani (2009) state that anemia is responsible for almost one quarter and 40 percent of maternal deaths, respectively.

during childbirth but also concurrent morbidities brought on by the experience of pregnancy and childbirth.' (World Bank, 2011, p. 128). The global estimate of maternal deaths is considered to be the 'tip of the iceberg' since an estimated 10 to 20 million women suffer pregnancy-related physical or mental disabilities every year (Filippi et al., 2006, p.2).[9] Maternal morbidities can have long-lasting consequences –potentially affecting women's survival– that can also lead to impoverishment, violence, stigmatization, and marital instability (Koblinsky et al., 2012). Also not encompassed in the standard definition, other causes of maternal deaths include injuries, accidents, suicide, murder, and domestic violence while women are pregnant (AbouZahr, 2003; Ronsmans and Graham, 2006).

Iron-deficiency anemia is the most common nutritional deficiency worldwide (Mukuria, Aboulafia, and Themme, 2005). It is pervasive in India where it affects 56 percent of ever-married women aged 15–49 (using NFHS-III data). It commonly results from a nutritional deficiency of iron, folate, or vitamin B12. It is more prevalent during adolescence, pregnancy and among breastfeeding mothers when nutritional requirements increase substantially (Sharman, 2000; Kalaivani, 2009).[10] In addition to indirectly causing maternal mortality, anemia is an important risk factor for other adverse health outcomes including impaired cognitive performance, higher incidence of infection, prenatal and perinatal infant loss, and reduced work capacity (World Bank, 1996; Sharman, 2000; Gragnolati et al., 2005; Ackerson, and Subramanian, 2008).

Progress has been made in improving maternal health in India (Bhat 2002a; 2002b).[11] The maternal mortality ratio declined from 570 in 1990 to 230 in 2008 (WHO, 2010).[12] World Bank (2011) states 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). Notwithstanding better health conditions in recent times, many women are still not receiving adequate maternal health care in India.[13] The decrease in maternal mortality has not been paralleled by improvements in anemia prevalence. Instead, the incidence of anemia among ever-married women has worsened over time: from 52 percent in 1998 to 56 percent in 2005. Given its magnitude, anemia is an extremely

---

[9]Permanent and chronic conditions include obstetric fistula, urinary or faecal incontinence, scarred uterus, pelvic inflammatory disease, palsy. Other main sequelae are severe anemia, eclampsia, and infertility. Psychological obstetric morbidity includes puerperal psychoses, post-partum depression, suicide, and strong fear of pregnancy and childbirth (AbouZahr, 2003).

[10]Iron needs increase during the second and especially the third trimester of pregnancy. During pregnancy, the required amount of iron is higher than what can be absorbed from the diet, therefore iron supplementation is necessary (Sharman, 2000).

[11]Safe motherhood is one of the objectives of the 1996 Reproductive and Child Health Programme and the National Population Policy adopted by the Government of India in 2000.

[12]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, 2010).

[13]For instance, NFHS-III data indicate that only 39 percent of births take place in health facilities and 58 percent of women did not receive any postnatal check-up after their most recent birth (IIPS and Macro International, 2007).

serious public health issue in India (Gragnolati et al., 2005).

2.2 Fertility behavior and maternal health outcomes

It has been argued that in contexts where the preference for male children is pervasive, parents alter their fertility choices and can exhibit overt discrimination against their unborn and living female children (Das Gupta, 1987; Das, 1987; Clark, 2000; Arnold, Kishor, and Roy, 2002; Jha et al., 2006). Parents can ensure the desired number of sons in two main ways. First, they can avoid the birth of unwanted daughters by eliminating female fetuses through sex-selective abortion. Several studies have noted that increasing sex ratios at birth in India match the timing of the diffusion of prenatal sex determination technologies in the late 1980s (Arnold, Kishor, and Roy, 2002; Bhalotra and Cochrane, 2010). These and other studies (Das Gupta, 1987, and Jha et al., 2011) have also documented that sex selection is more common among educated women and richer families. Another key finding of this literature is that there is no evidence of sex selection at first birth (Das Gupta, 1987; Bhat and Zavier, 2007; Jha et al., 2011). Second, women can continue childbearing until the desired sex composition of children is reached, commonly known as son-preferring stopping behavior (as women with daughters only are more likely to keep having children) (Chowdhury and Bairagi, 1990; Clark, 2000; Dreze and Murthi, 2001; Arnold, Kishor, and Roy, 2002; Filmer, Friedman and Schady, 2009). Other authors have also shown that a female birth is associated with shorter birth intervals (Trussel et al., 1995; Arnold, Choe, and Roy, 1998), and shorter breastfeeding (Jayachandran and Kuziemko, 2011). In this paper, I use 'son-preferring fertility behavior' to refer to both stopping as well as reduced birth interval.

The adverse health consequences of a high number of births and short spacing between births are well documented in the medical literature. A high number of pregnancies is associated with higher lifetime risk of death due to pregnancy.[14] Medical research has shown that birth intervals shorter than 24 months are associated with poor child and maternal health outcomes (Setty-Venugopal and Upadhyay, 2002; King, 2003).[15] Closely spaced pregnancies increase the risk of maternal nutrient depletion as the mother is not allowed sufficient time to restore the adequate supply of nutrients needed to have a successive successful pregnancy (Sharman, 2000; King, 2003). Conde-Agudelo, and Belizan (2000) use a large cross-sectional dataset covering 456,889 women from 18 Latin American countries and find that, compared to women giving birth at 27 to 32-month intervals, women

---

[14]The adult lifetime risk of maternal mortality is defined as "the probability that a 15-year-old female will die eventually from a maternal cause if she experiences throughout her lifetime the risks of maternal death and the overall levels of fertility and mortality that are observed for a given population" (WHO, 2010, pp. 14). It is related to both the mortality associated with a single pregnancy and the number of pregnancies a woman has in her reproductive years (UNICEF, 2008).

[15]Birth interval (or birth spacing) refers to the time interval (number of months) between births (i.e., the length of time from birth to next conception plus nine months).

with intervals shorter than 15 months (i.e., less than six months from birth to conception of the next child) are: 1.3 times less likely to avoid anemia; 1.7 times less likely to avoid third-trimester bleeding; and 2.5 times less likely to survive childbirth.[16] Therefore, short spacing is found to affect maternal mortality both directly and indirectly through increased risk of anemia.

3   Data

The data used in this paper are from the India National Family Health Survey (NFHS) which contains detailed information on women's entire birth histories, birth intervals (at the child level), health indicators (for children and their mothers), domestic violence, and other demographic and socioeconomic characteristics of women and the households they live in. The NFHSs are repeated cross-sectional surveys for a representative sample of women in India.[17] The women sampled for individual interviews in the NFHS are aged 15–49, and I consider the sample of all those who had at least one child ever born. Moreover, I exclude women with 9 or more children ever born to reduce heterogeneity due to women with extraordinarily high fertility (the 98th percentile of the pooled sample of women).[18] I pool the three available surveys, conducted in 1992/93 (NFHS-I), 1998/99 (NFHS-II), and 2005/06 (NFHS-III).

Using the pooled NFHS birth history data, the sex ratio at birth (here defined as the fraction of male births) is 0.519 for first births and 0.521 for all births. Among women who had their first birth during the ten years before the survey the sex ratios for first and all births are 0.513 and 0.518, respectively.[19,20] Figures 1 and 2 show the sex ratio at birth (for *all* births) by birth year in the pooled dataset and separately by survey year, respectively. Pooling different survey rounds allows me to distinguish between age and cohort-specific effects. The figures show a decreasing trend of

---

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

[17]While NFHS-I and II only interviewed ever-married women, NFHS-III includes those never married. To ensure comparability, I exclude never married women from the analysis. Since I only consider women with at least one child ever born (the vast majority of whom are or have been married), the fraction of excluded women is quite small: 72 women, equivalent to 0.02% of all women interviewed in NFHS-III.

[18]Total fertility rate (TFR) in the NFHS-III is 2.7, down from 3.4 in NFHS-I. Therefore, on average, an Indian woman in NFHS-III will give birth to 2.7 children by the end of her childbearing years. The TFR is based on age-specific fertility rates during the three years preceding the survey (IIPS and Macro International, 2007).

[19]The biologically "normal" sex ratio at birth ranges from 1.03 to 1.06 (male to female births), corresponding to a fraction of male births approximately in the range of 0.507 – 0.515 (Arnold, Kishor, and Roy, 2002).

[20]Sex ratios have been found to be affected by several biological, environmental, and genetic factors (other than behavioral factors). These include improvements in maternal health status (Bhat 2002a), mother's malnutrition (Mathews, Johnson, and Neil, 2008; Almond and Mazumder, 2011; Andersson and Bergstrom, 1998), prevalence of Hepatitis B (Oster, 2005), birth order, timing of conception, hormonal conditions (James, 1987). Some of these biological factors and how they might affect the findings of this paper are discussed in more detail in section 6.1.

the sex ratio which is common to all the survey rounds, independently of the cohort of birth of the child, and not specific to a particular period.[21] The most recent survey (NFHS-III) also shows an increasing sex ratio at birth starting from the end of the 1980s onward, which corresponds to the timing of the diffusion of ultrasound technologies (also consistent with the sex ratio in the Indian Census). Figure 3 shows that the decreasing trend is specific to first births, while the sex ratio for higher-order births is somewhat stable throughout the whole survey period and significantly higher than that for first births in 1980 and after. This is further indicative of the use of sex selective abortion, known to be much more common among higher order births. Indeed, pooled NFHS data on parental gender preferences indicate that, albeit more than one third of ever-married women want to have more sons than daughters (compared to only 2.5 percent who want more daughters than sons), more than 80 percent of them want *at least* one daughter and about 86 percent want *at least* one son.[22] Moreover, 42 (14.5) percent desire at least two sons (daughters).

Table 1 reports summary statistics for the pooled sample of ever-married women aged 15–49 by the sex of their first-born (refer to the data appendix for a description and availability of these variables in the NFHSs).[23] Panel (a) reports socio-demographic and fertility-related variables. Most characteristics differ significantly between the sample of women with a first-born girl and a first-born boy. In particular, the table shows that women with first-born girls are significantly more likely to be educated, to have married and had their first child later in life, to have (and desire to have) more children, to use contraception (and have undergone permanent sterilization)[24], to have performed a prenatal ultrasound test, and to have had a shorter spacing between the first

---

[21]Rosenblum (2013) uses the 2002-2004 Reproductive and Child Health Survey (RCH-II) and finds a similar pattern: higher sex ratios for births occurred in the past, especially for lower birth orders. He tests for the exogeneity of the sex of the first born, discusses potential recall or survival bias in the data (without attempting to test or disentangle the two), and argues that the extent of the bias is small.

[22]These figures are based on the NFHS question about the number of children women would like to have if they could go back to the time before they started childbearing. NFHS-III indicates that average ideal family size among ever-married women is 2.5 children (down from 2.9 in NFHS-I). Interviewed women were also asked the ideal sex composition for these children: in NFHS-I they indicated 1.6 sons, 1.1 daughters, and 0.2 of either sex, while in NFHS-III 1.2 sons, 0.9 daughters, and 0.4 of either sex.

[23]All the estimates shown in this paper have been computed using sample weights provided by the NFHS. Given that three surveys were conducted in different years, the sample weights used in the pooled sample have been multiplied by the fraction of the total female population aged 15–49 in India (obtained from the United Nations Population Division World Population Prospects: 2012 Revision, $http://esa.un.org/wpp/Excel-Data/population.htm$) and the number of women aged 15-49 interviewed in each survey round.

[24]Knowledge of contraception is nearly universal among ever-married Indian women in NFHS-III: 99.2% of women know one or more methods of contraception and current use is 55% (up from 98.9% and 45.3% in NFHS-II, respectively). Female sterilization, with a prevalence of 37%, is a permanent contraception method (to *limit* the number of births) that accounts for 67.4% of all contraceptives used. The highest adoption rate of female sterilization, 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 women (NFHS-III data).

10

and second birth. The differences found in the variables related to fertility behavior are a direct consequence of pervasive son preference in India, while those in the education and demographic variables (arguably predetermined to first birth) are indicative of selection among women with a first-born girl. Panel (b) of table 1 reports summary statistics for health indicators and shows that, compared to women with a first-born boy, those with a first-born girl are more likely to be anemic. Panels (c) and (d) show that there are no significant differences between women by the sex of their first-born child in terms of food consumption or violence perpetrated by husbands against them.

Table 2 reports selected summary statistics for three subsamples of women (by the sex of their first-born child): those who had their first birth less than a year before the time of the survey, and the 15–29, 30–49 age groups. The idea is to understand the timing of the emergence of the differences highlighted in table 1. If the sex of the first-born is random at birth but subsequently affects women's fertility behavior (as has been documented in places with son preference), there should be no observed differences among women with first-born girls and boys close to the time of their first birth. Differences should instead emerge and widen over time as women engage in son-preferring fertility behavior.[25] As conjectured, panel (a) shows that, except for the desire to have more children (expected to be immediately affected by the sex of the first-born), all the other variables are balanced between the two samples. Panel (b) considers women aged 15–29 (who are in their most active reproductive years) and shows that, while there are no significant differences in terms of predetermined variables (education, age at first birth and marriage), fertility behavior and anemia status are significantly affected by the gender of the first-born. Panel (c) shows that most of the observed differences in table 1 emerge among women aged 30–49. Women with a first-born girl in this age group also have better educated husbands and show no differential incidence of anemia.

This evidence confirms that women are similar at the time of first birth and develop differences over time as a likely result of son-preferring fertility behavior. In fact, they appear to respond to a first female birth by having substantially more children and shorter spacing, using contraceptives less often, and exhibiting a higher probability of having had an ultrasound test and a terminated pregnancy. This behavior increases their risk of mortality and morbidity (especially after repeated pregnancies and unsafe abortions). It is therefore likely that women who engage most in son-preferring fertility behavior and have lower access to safe abortion –typically women of lower socioeconomic status– are more likely to die. This selection effect might explain the observed differences (especially in the predetermined variables) among living women by the sex of their first child as well as the decreasing sex ratio for first births. More evidence in support of this channel will be shown in the next sections.

---

[25]This approach is conceptually similar to Barcellos, Carvalho, and Lleras-Muney (2014) who study gender differences in parental inputs in India by restricting the analysis to families with children who are still young to account for the effect of child gender on family size.

4 Nonparametric evidence

This section examines the age structure and health status of women by the sex of the first-born by using kernel-weighted local polynomial regressions. The resulting smoothed values are displayed in graphs with 95% confidence intervals. The age range of women considered in this section is 20–49.[26]

4.1 Share of women with a first-born girl

The nonparametric graphs in this section plot the share of women with a first-born girl by their age at the time of the survey. This share is simply the inverse of the sex ratio for first births. Therefore, rather than plotting the sex ratio by birth year of the child (done in section 3), the following graphs show the age distribution of women with a first-born girl (due to pooling, these are independent of cohort effects, except for non-overlapping years of the NFHSs).

The left panel of figure 4 shows that the sex ratio for first births is in the biological range for women below the age of 30 (including the confidence bands). However, the share of women with first-born girls decreases almost monotonically for women older than 30, where it is 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 is lower among older women, particularly among those older than 30. The right panel shows that the same pattern is found in each round of the NFHSs (though with larger confidence bands and more fluctuation likely due to smaller sample sizes when using each survey separately). Interestingly, the upper graph in figure 5 shows that the decreasing pattern is found only among women who did not complete any years of schooling, while the share is rather stable for women with at least one year of education at all ages.[27] Similar patterns are found when using each NFHS round separately (see bottom two graphs in figure 5).[28]

Considering education as a plausible proxy for socioeconomic status, this suggests that poorer women suffer higher maternal and adult mortality. Having lower access to health services (and in particular to prenatal sex diagnostic techniques in more recent decades), uneducated women may rely more heavily on son-preferring fertility behavior to achieve the desired sex composition of children and thus suffer the adverse consequences on their health and survival.

State-wise variation. There is substantial state-wise variation in sex ratios at birth and excess mortality in adulthood within India. Anderson and Ray (2012) find a distinctive pattern: gender

---

[26]Women aged 15–19 are dropped because of very large confidence bands due to a small fraction of them who already has had a birth –these represent 4.5% of the entire pooled sample.

[27]In Milazzo (2012), I find similar patterns when using DHS data for Nigeria. Compared to India, this is a context in which women are also found to engage in son-preferring fertility behavior (though not as strongly), and where the lifetime risk of maternal mortality is higher.

[28]Appendix figure A.1 also shows a decreasing pattern for women in rural areas (not found in urban areas).

bias at birth is highest in the North-West where excess female mortality among adult women is in turn relatively low (the "Group 1" states), while North-Eastern states have low gender bias at birth together with the highest excess female mortality in adulthood (the "Group 2" states). Southern states, like Kerala, exhibit the lowest excess female mortality in India (for this reason, Kerala was used as an alternative benchmark in their analysis). This evidence is confirmed by official statistics on sex ratios (2011 Indian Census) and maternal mortality (Office of the Registrar General India, 2011).[29] Two North-Western states, Punjab and Haryana, have the most male-biased sex ratio at birth of all Indian states. At the same time, they have below-national-average maternal mortality ratios. On the other hand, Assam and Bihar in the North-East have below-national-average (i.e., less male-biased) sex ratios at birth and are among the states with the highest maternal mortality ratios in India.

Figure 6 shows the share of women with first-born girls by age in Punjab, Haryana, Assam, Bihar, and Kerala. Panel (a) shows that Punjab and Haryana exhibit a steady share of women with first-born girls at all ages. Assam and Bihar, shown in panel (b), exhibit a strikingly different pattern: the share of living women with a first-born girl is steeply decreasing with women's age. This is consistent with high prevalence of sex-selective abortions in the North-West (reflected in skewed sex ratios at birth) and higher maternal mortality in the North-East. Finally, as expected, Kerala shows a rather flat pattern, consistent with low excess female mortality across all age groups.

While more research is warranted, this paper offers a possible explanation for the apparent puzzle singled out in Anderson and Ray (2012). This is based on differential prevalence of sex selective abortion and son-preferring fertility behavior (as the two main ways through which parents can achieve the desired sex composition of children) across the states of India.

4.2 Health indicators

Turning to health indicators (available only in NFHS-II and III), figure 7 shows the incidence of anemia by the sex of the first-born and women's age at the time of the survey.[30] The general age pattern reflects the fact that anemia is typically higher during adolescence and the reproductive years due to higher physiological demand for iron while pregnant and lactating (Sharman, 2000; Mukuria, Aboulafia, and Themme, 2005). Notably, the figure shows that the incidence of anemia is significantly higher among women with a first-born girl (compared to women with a first-born boy) approximately up to age 30, while there is no differential incidence for older women.

Intuitively, this result might be the consequence of son preference: women with first-born girls tend to be pregnant (or breastfeeding) more often during their reproductive years than women who have first-born boys. Although higher iron needs during pregnancy are accounted for in the NFHS

---

[29]The sex ratio from Census data refers to all children aged 0 to 6.

[30]For a description of how hemoglobin testing is conducted in the NFHS and DHS surveys refer to Sharman (2000).

definition of the hemoglobin cut-off value for anemic women[31], I replicate the same graph using the subsamples of women who are not currently pregnant or breastfeeding (cut-off values are not different for lactating women). Appendix figure A.2 shows that the patterns are found also in these subsamples. This suggests that differential incidence of anemia between women with a first-born girl or boy may be driven by short birth spacing (or past pregnancies).

Figure 8 shows a graph for BMI. Overall, BMI increases with women's age (within the range of 19 to 22).[32] There are no apparent differences in BMI by the sex of the first-born for women below age 40. It is also worth noting that the BMI index excludes women who are pregnant at the time of the survey and who gave birth less than three months before the survey (IIPS and Macro International, 2007). Among older women, those with a girl as first-born child show a slightly higher nutritional status.[33] Lastly, figure 9 reports a graph for height. Height reflects nutritional status in childhood and adolescence and therefore should be affected by fertility behavior to a lower extent than other nutritional indicators that are not predetermined to first birth. Yet, mortality selection might also entail higher survival possibilities for taller women who have been better nourished during childhood (similarly to Bozzoli, Deaton, and Quintana-Domeque, 2009). Moreover, taller women are more likely to be from a higher socioeconomic status and thus enjoy better access to prenatal diagnostics. The figure shows that, as expected, there are no height differences among younger women with a first-born girl or boy, while there is some weak evidence that older women with first-born girls are taller (especially above age 45).

5 Empirical strategy and results

This section first discusses the identification assumption, then explores how the sex of the first-born affects fertility behavior, and lastly examines the age structure and anemia status of women using parametric analysis.

5.1 Identification assumption

As shown in table 2, women with a first-born girl or boy do not exhibit significant differences in terms of observable predetermined characteristics close to the time of first birth or when they are

---

[31]A woman is defined as being anemic if the measured hemoglobin count is of less than 12 grams per deciliter (cut-off for all women) and less than 11 grams for pregnant women.

[32]The BMI is defined as weight (kg) divided by height (m^2). A cut-off point of 18.5 is used to define chronic energy deficiency and a BMI of 25 or above indicates overweight (IIPS and Macro International, 2007).

[33]While the association between fertility behavior and anemia status is widely documented in the medical literature, BMI is not among the outcomes mentioned in a review of the literature on the effects of short spacing (Conde-Agudelo, Rosas-Bermdez, and Kafury-Goeta, 2007). Evidence that mother's malnutrition at conception might affect the sex ratio is discussed in Section 6.1.

relatively young. The randomness of the sex of the first-born is in fact widely supported. First, as previously noted, several studies document that sex-selective abortion is much more common among higher-order births. Second, the vast majority of Indian parents want at least one daughter. Third, data from the NFHS indicate that the sex ratio for first births during ten years before the time of each survey is in the normal range. The sex ratio for *first* births after 1990 (timing of the diffusion of ultrasound testing) is 0.514, which is in the biological range.[34] It is in the normal range even for educated women (known to practice more sex-selection): among educated (uneducated) women the sex ratio for first births is 0.515 (0.513). Section 6.3 provides additional evidence that sex-selection is not practiced at first-birth.

Given possible selective mortality (most visible in the subsample of older women in panel (c) of table 2), the empirical strategy relies on the assumption of *conditional* exogeneity of the sex of the first-born child. All the regressions for women's fertility behavior control for a vast set of observable characteristics. The analysis of women's health status additionally focuses on women who recently had their first birth and are therefore not yet selected.

It is important to note that selective mortality would lead to underestimate the effect of a first-born girl on fertility. This is because women who did not survive (and therefore do not appear in the data) are probably the ones who suffered the most the consequences of an 'extreme' fertility behavior due to son preference (both because of poorer health and access to maternal care).

5.2   Mechanism: Fertility behavior and sex of the first-born child

5.2.1   Fertility regressions

Although most Indian women want more than two children, the sex of the first-born still affects the fraction of male children they can have through their lifetime. In fact, as previously noted, more than a third of women want more sons than daughters and more than 40 percent want at least two sons. To examine desired and realized fertility as well as other outcomes that are likely to be affected in contexts with son preference, I estimate the following regression under the assumption of conditional exogeneity of the sex of the first-born.

$$y_{i,s,b,r} = \beta_1 \ fg_i + \beta_2 \ fg_i \ ^* \ edu_i + \beta_3 \ edu_i + \gamma X_{i,s,b,r} + \alpha_s + \mu_b + \delta_r + \epsilon_{i,s,b,r} \qquad (1)$$

with woman i, resident in state s, born in year b, surveyed in round r. $y_{i,s,b,r}$ is the dependent variable, alternatively a dummy for whether the woman currently reports the desire to have more children, a dummy for whether she has been sterilized, whether she has undergone ultrasound

---

[34]The sex ratio for *all* births occurred after 1990, including those of higher-order is 0.519.

testing, whether she has ever had an abortion, and the number of children ever born.[35] fg_i indicates whether the first child ever born is a girl, edu_i is a dummy variable equal to one if the woman completed at least one year of education, zero if she has completed none, $fg_i^*edu_i$ tests whether the effect is heterogeneous between educated and uneducated women, $X_{i,s,b,r}$ is the set of covariates including: age of the woman and her partner (and age squared), seven five-year age groups, age at first marriage, age at first birth, whether the woman's partner completed one or more years of education, woman's religion and caste, a wealth index, and a urban dummy.[36,37] α_s, γ_b, δ_r are state, birth year, and survey round fixed effects, respectively. All the regressions (except that for the number of children ever born) also include the number of live births. Since most of the younger women in the sample have not completed their fertility, I restrict the sample to women who have been sterilized or who report they do not want more children when estimating the regression for the number of children ever born. Regression 1 is estimated using a probit model for binary outcomes, and a linear model for the other dependent variable.

If women's fertility behavior is affected by the sex of the first-born and this is due to son preference, β_1 is expected to be positive in the regressions for desired (and realized) fertility, ultrasound testing, and abortion, while β_1 should be negative for the use of sterilization. In terms of heterogeneous effects, predictions are as follows.[38] Women's desire for having more children after the birth of a first girl should not differ much between educated or not as they both have strong incentives to try to conceive again for a boy. All women are expected to seek sterilization when the desired sex composition of children is reached. The use of sterilization is in fact widespread among all women irrespective of their education (often women in rural areas have access to 'sterilization camps'). The use of ultrasound testing is instead part of the routine antenatal care for most educated women (especially those with high education: 74 percent of women with more than 10 years of education in NFHS-III performed the test), but is only rarely undergone among the uneducated. Given their high access to prenatal maternal care, educated women are not expected to resort more to ultrasound testing in response to a first-born girl than the uneducated (since most educated do

---

[35]The number of children ever born includes all live 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 affected by son preference.

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

[37]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.

[38]Average values of the independent variables for educated and uneducated women are as follows: 21 and 24 percent for the desire to want more children, 38 and 35 percent for female sterilization, 40 and 10 percent for ultrasound testing, 21 and 18 percent for terminated pregnancy, and 2.9 and 4.2 children ever born (the latter calculated on the sample of women who have been sterilized or do not want to have more children), respectively.

it irrespective of the sex of the child).[39] In turn, the educated are expected to be more likely to selectively abort after a first-born girl, consistent with existing evidence. The uneducated may still practice abortion but, being generally poorer, it is plausible to think that they might seek illegal abortion, probably underreported. The sign of the differential effect on the number of children is a priori ambiguous. The uneducated have higher ideal and realized family size so, given the number of children, they have higher chances to have a son, and the educated will have more children too if they do not practice sex selection at low birth-order. At the same time, the uneducated can keep having children if they have no sons, or have an abortion performed.

Table 3 reports the results. The first column shows the unconditional estimates for each of the outcomes studied. Column (1) shows that, compared to women with a first-born boy, women with a first-born girl are 4.9 percentage points more likely to desire more children (the effect is significant at 1% level). Column (2) shows that the effect is slightly higher at 5.1 percentage points (equivalent to a 25 percent effect relative to the average for women with a first-born boy) after conditioning for all the observables and fixed effects. Column (3) explores heterogeneous effects: women with at least one year of education are on average less likely to say they want more children, but the effect of a first-born girl is not statistically different than for uneducated women (who have a higher desired number of children to start with). A first-born girl substantially decreases the probability that a woman is sterilized (columns 4–6). The effect is of 6.8 percentage points (column (5), equivalent to a 18 percent effect). The effect on sterilization does not appear to differ significantly for women with education. Column (8) shows that women with a first-born girl are 2.5 percentage points more likely to have had an ultrasound test for their most recent births. This effect is significant at 1% level and equivalent to a 9.5 percent effect. Educated women are more likely to have undergone ultrasound testing as part of their antenatal care. This reflects their higher access to health care relative to poorer uneducated women. As expected, the increase in the probability of an ultrasound test after a first-born girl does not differ between educated and uneducated women. As can be seen in column (11), an abortion is 1.1 percentage points (significant at 1% level) more likely for women with first-born girls (equivalent to a 5.8% percent effect). Column (12) shows that education is positively associated with the probability of having ever had a terminated pregnancy, and that this effect is stronger for women with a girl as first-born child (the estimate for the interaction term is 1.4 percentage points, significant at 1% level). Although uneducated women are also more likely –thought to a smaller extent– to have had an abortion after a first-born girl, this result confirms findings from the existing literature documenting that sex-selective abortion is more common among the educated. A nonparametric plot of the likelihood to have ever had a terminated pregnancy by sex of the first-born and age reveals that, among the uneducated, a first-born girl has an effect only

---

[39]The use of ultrasound testing seems to be motivated by the early detection of abnormalities for first-born children, and by sex-selection for higher-order births (Bath and Zavier, 2007).

after age 30 (while among the educated an effect is visible also among younger women) (appendix figure A.3). This is consistent with poorer women selectively aborting a female fetus later in life, probably when the costs of having an additional daughter become unsustainable (i.e., dowry costs).[40] However, as the variable for terminated pregnancy does not distinguish among the reasons for termination (e.g., miscarriage, abortion, or stillbirth), it is important to consider this evidence as only indicative of higher recourse to induced abortion for sex selection.[41] Column (13) shows that women with first-born girls unconditionally have 0.301 more children ever born, 0.295 after including all the controls (both significant at 1% level). The latter estimate implies a 8.7 percent effect. As expected, the effect is statistically different for educated women for whom the effect of a first-born girl on fertility is attenuated. Appendix table A.1 additionally estimates the effect of the sex of the first two or three children on fertility outcomes of women who had at least two or three live births. Consistent with increasing pressure for giving birth to a male child, additional daughters after the first substantially increase the effects on all the examined outcomes.

Altogether, these results suggest that the sex of the first-born strongly affects women's fertility behavior, and that educated women tend to rely more on sex-selective abortion, while uneducated women rely on son-preferring fertility behavior.

5.2.2 Birth spacing regressions

The NFHS reports the entire fertility history of each woman. I exploit information on the length of the interval between successive births. The idea is to test whether women with a first-born girl try to conceive the following child sooner than women with a first-born boy.

In a first step, I estimate whether the spacing between first and second child is affected by the sex of the first child, as well as whether the interval between second and third child is affected by the sex of the second child (conditional on the first child being a girl). To do this, I use a regression similar to equation 1, in which the dependent variable is either the number of months or the probability of an interval shorter than 15 months. To evaluate the effect of the sex of the first-born, I first restrict the sample to women with at least two children ever born (who report at least one birth interval), and then to the sample of women with at least three children ever born for the effect of the sex of the second child.

The availability of birth histories allows a second empirical strategy that exploits the variation in the length of the interval and the sex of the child *within* the reported history of each woman. This is done with mother fixed effects. However, this methodology does not allow the estimation

---

[40]It is also consistent with a higher probability of miscarriage for women engaging in son-preferring behavior.

[41]Only NFHS-II provides some more details on whether the abortion was spontaneous or induced. Evidence suggests that at least some of the abortions reported as spontaneous are instead induced (due to misreporting) (Arnold, Kishor and Roy, 2002).

of the effect of the sex of a child at a specific birth order and identification requires restricting the sample to all women with at least three children ever born. I estimate the following regression:

$$y_{i,j,(k-1,k)} = \beta_1 \; girl_{i,j,k-1} + \sum_k \gamma_k \; b_{i,j,k} + \chi_j + \epsilon_{i,j,k-1} \tag{2}$$

with child i, mother j, birth order k. $y_{i,j,(k-1,k)}$ is the time interval (between birth of child $k-1$ and k), expressed in number of months or as a dummy variable equal to one if the interval is shorter than 15 months. $girl_{i,j,k-1}$ is a dummy equal to one if the sex of the child $k-1$ is female. $b_{i,j,k}$ are k dummies for each birth order. χ_j denote mother fixed effects. I exclude births of birth order four and higher to evaluate the effect of the sex of earlier-born girls on spacing. This regression is estimated using the entire sample, and the subsamples of uneducated and educated women. If there is son preference, the birth interval is expected to be shorter ($\beta_1 < 0$) after a girl is born compared to a boy. This regression is estimated using a linear model for the number of months, and a probit model for the probability of having an interval shorter than 15 months.

A potential source of bias in the regressions for birth spacing is represented by the practice of sex-selective abortion. This is because female fetuses are especially likely to be selectively aborted after a first-born girl. It is important to note that this would generate downward biased estimates of the (negative) effect of a first-born girl on the length of the birth interval since more time will be needed to have a completed pregnancy following an abortion. This is especially the case for educated women who are much more likely to have had a terminated pregnancy after a first-born girl than the uneducated.[42]

Table 4 shows the results of the regression without mother fixed effects (number of months in columns (1)–(2), and probability of a short interval in columns (3)–(4)). Panel (a) shows the effect of a first-born girl on the time to the next birth. Conditional on the covariates, women with a first-born girl wait 0.8 months less compared to women with a first-born boy (significant at 1% level, and equivalent to a 2.4 percent effect). Column (2) explores heterogeneous effects. A first-born girl reduces the interval particularly for uneducated women, while among the educated the effect is significantly attenuated. This is consistent with educated women with a first-born girl aborting if the next fetus is detected to be a female. Columns (3)–(4) show that, while education is on average significantly and negatively associated with the probability of a short birth intervals, a first girl increases the probability of a short interval by 0.5 percentage points (significant at 1% level, and equal to a large 5.7 percent effect), and this effect is not different between educated and

---

[42]As a check, I added the variable for having ever had a terminated pregnancy as a control in the regressions (without mother fixed effects) and the results for birth spacing are consistently stronger. However, as previously noted, this variable probably underestimates the prevalence of induced abortions, due to under-reporting especially if performed illegally.

uneducated women. The effect of a second-born girl on the birth interval conditional on having a first-born girl is shown in panel (b). Again, the time interval is significantly reduced after the birth of a second girl (column (1)), and now the effect is only found among uneducated women (column (2)). Moreover, the birth of two consecutive girls increases the probability of a short interval by 0.8 percentage points (significant at 1% level) (column (3)). This effect implies a 8.7 percent effect. Similar to the result in column (2), the effect on short spacing is almost doubled in magnitude (1.5 percentage points and significant at 1% level) for uneducated women (column (4)). The net effect is instead negative for women with at least one year of education.

Table 5 shows the estimates including fixed effects at the mother-level. The first column shows the estimated effect of having a girl at first or second order on the length of the interval to next birth. The coefficient is -0.60 months (significant at 1% level) for all women, -0.63 and -0.50 for uneducated and educated women, respectively (columns (2) and (3)). Columns (4)–(6) show that the birth of a girl increases the likelihood of a short interval by 0.5 percentage points (significant at 1% level), which is equal to the estimate obtained without mother fixed effects (column (3) of table 4). As expected and consistent with estimates in table 4, an effect (equal to 0.6 percentage points) is found only in the subsample of uneducated women.

This evidence is again consistent with educated women being more likely to turn to sex-selective abortion while uneducated ones to son-preferring fertility behavior with associated health risks.

5.3 Outcomes: Age structure and women's health outcomes

5.3.1 Age structure

To confirm the nonparametric evidence shown in figure 4, I estimate an ordered logit model in which the dependent variable is a categorical variable with seven five-year age groups (from 15–19 to 45–49). The controls are the same as in regression 1. After estimating the ordered logit model, I compute changes in the predicted outcome probabilities when the independent variable 'first-born girl' changes from 0 to 1, with all the other independent variables set at their mean value. Changes are reported with 95% confidence intervals computed with the delta method. Column (1) and (2) in table 6 show the changes in the predicted probabilities and the 95% confidence bands, 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 a first-born boy, conditional on the number of children and all other covariates (set at their mean).

5.3.2 Health outcomes

Next, I estimate the effect of the sex of the first-born on anemia status and other health indicators for women in different age groups with the following regression:

$$y_{i,s,b,r} = \beta_1 \, fg_i + \beta_2 \, fg_i \, ^* \, over30_i \, + \gamma X_{i,s,b,r} + \alpha_s + \mu_b + \delta_r + \epsilon_{i,s,b,r} \qquad (3)$$

with woman i, resident in state s, born in year b, surveyed in round r. $y_{i,s,b,r}$ is a dummy variable equal to one if the woman is anemic, zero otherwise. fg_i indicates whether the first child ever born is a girl, which is also interacted with a dummy equal to one if the woman is aged 30 or older ($over30_i$). The cutoff at age 30 is chosen based on the nonparametric evidence previously presented (section 4) and the age at which the conditional age distribution for women with first-born girls shifts to the left of the distribution of those with a first-born boy (table 6). α_s, μ_b, δ_r are state, cohort of birth, and survey round fixed effects, respectively. In addition to the set of covariates and fixed effects as in regression 1, $X_{i,s,b,r}$ also includes dummies for whether the woman is currently pregnant or breastfeeding. These women are in fact more likely to be anemic and their status might be correlated with the error term. Another potential concern is whether past health (predetermined to first-birth) affects both the probability of having a first-born girl and current anemia status (further discussed in the next section). For this reason, I also control for height, a proxy for women's health status during childhood. Regression 3 is estimated using a probit model. Again, 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.

Table 7 shows the results. Column (1) shows the unconditional estimates. Women with a first-born girl are 1.1 percentage points more likely to be anemic (significant at 1% level). Columns (2)-(3) include most covariates as well as the controls for women's current pregnancy and breastfeeding status and the coefficient of a first-born girl is slighly reduced (as expected). Height is included in column (4) and does not affect the estimated coefficient for a first-born girl. This allows me to rule out the possibility that past health conditions are correlated with the sex ratio and with current anemia status. Column (5) controls for the number of children ever born and the coefficient of interest is slightly reduced. The patterns shown in figure 7 are consistent with the result in column (6) indicating that a first-born girl significantly increases the incidence of anemia for women below age 30 (the increase is 1.4 percentage points or a 2.5 percent effect), while this effect is significantly reduced for older women. A further decrease in the incidence of anemia for women aged 45 and older is suggested by estimates in column (7), though it is not significant. Columns (8)-(9) split the sample into women by education and shows that the effect is found only among women with no education. This is consistent with increased health risks associated with risky son-preferring fertility behavior that is most common among them. The last two columns check whether the age pattern of anemia is purely reflecting higher fertility. The estimates of the interactions between the number of children ever born and the older age groups show that, as expected, a higher number of live births significanlty increases the prevalence of anemia among younger women. However, as

21

opposed to the effect of a first-born girl, this effect is not offset among those aged 30 and older. This suggests that anemia persists among women with more children (irrespective of their gender) over the years, while the impact of a first-born girl is more complex in that it also leads to 'riskier' pregnancies that further increase their risk of mortality (and selection).

To check how the sex of the first-born affects the *intensity* of anemia, I estimate an ordered logit model in which the dependent variable is a categorical variable with four ordered categories: severely anemic, moderately anemic, mildly anemic, and not anemic. Table 8 shows the changes in predicted probabilities computed using the estimates for the entire sample of women, and the subsample of younger (aged 15–29) and older women (aged 30–49). It can be seen that, conditional on the covariates, a first-born girl is associated with a significant increase in the probability of suffering from any level of anemia, and a decrease in the probability of not being anemic (implied percent effects indicate that women aged 15–29 are 4.3, 3.3, and 1.1 more likely to be severely, moderately and mildly anemic, respectively).

Finally, table 9 considers other indicators of women's nutritional status, including BMI (and undernourishment), and height (and shortness) as the dependent variables of regression 3. The estimates for BMI and chronic energy deficiency show a similar pattern: women above the age of 30 exhibit significantly better nutritional conditions, controlling for height. The results for height are not significant, but in the same direction as the other indicators (women 45 and older with a first-born girl appear slightly taller).

In sum, these estimates support the hypothesis that a first-born girl drives women into a fertility behavior that significantly jeopardizes their health and can have fatal effects during their reproductive years and cause long-term disabilities. They also support the hypothesis of selective mortality in light of better health conditions of living older women with first-born girls.

6   Alternative interpretations and robustness

Next, I turn to the plausibility of four alternative interpretations of the results.

6.1   Biological differences

There is evidence that maternal diet and poor nutrition around the time of conception are associated with a female birth (Mathews, Johnson, and Neil, 2008; Almond and Mazumder, 2011). This is because male fetuses are more fragile (Waldron, 1983; Kraemer, 2000) and susceptible to death *in utero* than females. Therefore selective mortality may be the result of biological and genetic factors rather than behavioral. To rule out this possibility, I perform several checks.

First, as previously noted, all the regressions for health outcomes control for height, an indicator of nutritional status in childhood. The estimates with and without the inclusion of height are

similar, indicating that past health does not seem to be associated with the sex of the first-born.[43]

Second, this interpretation would not be compatible with a pattern in which differential incidence of anemia between women with a first-born girl or boy is not present at the time of birth and arises only some time *after* first birth. To understand the timing of the emergence of differences in anemia status, I plot the prevalence of anemia among women by the sex of their first-born child and the time since their first birth (using Kernel-weighted local polynomial smoothing). Figure 10 shows that there is no differential incidence of anemia at the time of first birth, and that women with a first-born girl start developing anemia approximately 16 months after the birth of their daughter and that differences increase over time.[44] I also estimate a probit regression for the probability of being anemic on 'first-born girl' and its interaction with the interval of time since birth (controlling for height and other covariates). Column (1) of table 10 shows that from the time of birth until the 11th month there is no differential anemia status, while differences seem to develop gradually over time, though not significantly. As expected, differential incidence of anemia develops much faster among uneducated women, for which the increase becomes significant about three years after the birth of the first child. No similar pattern is found for educated women (columns (2)–(3)).

A third check explores whether there is differential incidence of anemia at the time of *each* birth from birth order 1 to 3. This is to make sure that anemia status does not affect the sex of the child at any birth. I run a probit regression for the probability of a female birth at birth order 1 to 3 on whether the mother is anemic (and the different levels of anemia severity) restricting the dataset to women who gave birth less than a year before the time of the survey. Table 11 shows no association between anemia status and the sex of the child at any birth order.[45]

None of these results is consistent with explanations based on biological differences.

6.2   Reporting bias

Another concern is that, as they are based on self-reported retrospective information, birth histories in the NFHS might suffer from misreporting and recall bias especially in relation to births (and deaths) occurred many years before the time of the survey (Smith, 1994; Byass et al., 2007).

---

[43]The results of the regressions for fertility behavior are robust to the inclusion of height (including height implies restricting the sample to second and third surveys in which measurements were taken). Results shown in table A.2.

[44]The timing of emergence of anemia is consistent with higher iron requirements particularly in the third trimester of pregnancy. Iron stores are especially depleted for women who had a short interval. Therefore, women with a first-born girl (found to be more likely to have a short interval) appear to develop anemia towards the end of their second pregnancy and these anemia differences seem to persist over time.

[45]No association (from birth order 1 to 3) is found even when using other nutritional indicators such as BMI (and the two cutoffs for underweight and obesity) and shortness. This is in contrast with Andersson and Bergstrom (1998) who find that short maternal stature and obesity are independently related to low sex ratio at birth using data for 684 women from the Central African Republic.

However, to affect the findings, recall bias would have to be systematically gender-biased and specific to first-births. In societies with strong son preference women might deliberately under-report female births (even though there is no reason they should misreport those that occurred at first birth only).[46] If this is the case, age at first birth should be negatively associated with a first female birth. This is because women who do not report their first-born daughters and instead report a later born boy will appear as if they had a first-born boy, but later. I test for this possibility by estimating a probit model for the probability of a female at first birth regressed on age at first birth and other controls. Results instead indicate that age at first birth *positively* and significantly predicts a first female birth (marginal effect of 0.001, with a standard error of 0.000, equally for all women and those aged 30 and older separately).[47] This result is not consistent with gender-biased recall bias but instead supports selective mortality.

As a further check, it is useful to note that gender-biased recall bias (arguably motivated by strong preferences for male children), if it exists, should not be decreasing over time in line with nondecreasing son preference in India (if anything, Census data report *increasing* male-biased sex-ratios at birth). Since both selective mortality and gender-biased recall bias imply that women with a first-born girl appear younger on average (older women might have died or be more likely to under-report a first-born girl, respectively), I check the age distribution of women with a first-born girl across the three NFHS rounds. Specifically, I regress women's age at the time of the survey on the variable 'first-born girl' and its interaction with the survey round dummies. The dependent variable is expressed in terms of seven age groups (in this case the regression is an ordered logit model), or as age in years (estimated with a linear model). As can be seen in column (1) table 12, women with a first-born girl are on average younger than those with a first-born boy in the first wave of the NFHS and this effect is attenuated for women in the last two survey rounds. Column (2) shows a similar pattern, although it is not significant. This pattern is not consistent with sex-selective recall bias but instead with the reduction in maternal mortality that has been taking place over the last decades in India.

6.3   Selective mortality or selective abortion?

One could argue that figure 5 can be interpreted as evidence of selective abortion of female children (at first pregnancy) among older and uneducated women. However, as discussed in the previous sections, there is vast evidence in the literature that sex selection is instead common at higher-

---

[46]Gender-biased recall bias is not compatible with the state-wise patterns presented in section 4.1. If women do not want to report a first female birth because of son preference, then the decreasing share of women with a first-born girl as a function of women's age should also be found in states with the highest sex ratios, i.e. Punjab and Haryana.

[47]Covariates include state fixed effects, urban dummy, birth year fixed effects, 7 five-year age groups dummies, survey round fixed effects, whether she and her partner have completed one or more years of education, a durables index, religion and caste. Result available upon request.

order births and among women with higher education and from more prosperous families (who have access to the technologies for sex detection of the fetus).

As an additional check, figure 11 plots the sex ratio for first and higher-order births by birth year of the child and women's education. Panel (a) shows that, for educated women, the sex ratio for first births is constant over time (and in the normal range). As expected, the sex ratio for higher-order births shows a steep increase after 1990 most likely due to the spread of prenatal sex diagnostic technologies. This evidence clearly suggests that sex selective abortion is not practiced at first-birth among the educated. Among the uneducated (panel b), it is the sex ratio for first births only that is decreasing over time (while it is decreasing only slightly for higher order births).[48] This pattern cannot be the result of sex-selective abortion as the period in which the sex ratio for first birth is found to be abnormally high is prior to the diffusion of ultrasound testing. It is instead consistent with selective mortality due to the association between sex of the first-born and risky fertility behavior or unsafe abortion.

6.4 Direct punishment (mistreatment)

There is evidence that domestic violence is associated with malnutrition (after controlling for socioeconomic variables) potentially through withholding of food or psychological stress (Ackerson and Subramanian, 2008). This section tests whether women with a first-born girl are more likely to suffer overt discrimination in the allocation of food (or nutritional supplements) or through domestic violence and discusses their possible contribution to selective mortality.

Food consumption and iron supplementation. To test whether women are 'punished' for having given birth to a girl through withholding of food, I estimate a probit model for the probability of consuming certain types of food at least once a week on the sex of the first-born. The dependent variables are dummies based on whether women self-report consuming specific food items daily or weekly. These include milk, pulses, green vegetables, fruits, and eggs or meat or fish. The independent variables of interest are 'first-born girl' and its interaction with the dummy for education level. All the regressions include height, current pregnancy or breastfeeding status as well as the standard set of covariates. The regression is estimated using three samples: women who had their first birth less than a year before the time of the survey, women aged 15–29, and 30–49.

To explain the anemia pattern, lower consumption of nutritious food (in particular eggs/meat/fish and green vegetables, that are rich in iron) should be found among younger women with a first-born girl, especially among the uneducated. Table 13 shows that there is no differential consumption of

---

[48]The pattern of the sex ratio for higher order births among the uneducated might be explained partly by selective mortality of women with female children at low birth order back in time, and to the practice of sex selective abortion in the period after 1990 (although much less common than among the educated– compare panel (a) and (b)).

25

food between younger women with a first-born girl or boy (also by education level) (panel (a) and (b)). Among older women, those with a first-born girl appear to have a slightly higher consumption of milk, green vegetables, and fruits (panel (b)). This can be partly explained by selective mortality.

Table 14 further checks whether women are differentially likely to receive iron and folic acid (IFA) supplements during pregnancy. It can be seen that a first-born girl is not associated with having received more supplements during the first (columns (1)-(2)) or second (columns (3)-(4)) pregnancies. The results also indicate that consumption of food or iron supplementation does not increase after the birth of a girl to compensate for higher nutritional needs.

Domestic violence. Next, I test whether self-reported experience of physical, emotional or sexual violence perpetrated by the husband (collected in NFHS-III) differs between women by the sex of their first-born. Table 15 presents the results. Panel (a) considers the subsample of women who had their first birth less than a year before the time of the survey. As can be seen in columns (1)-(2), women with a first-born girl do not appear to be more likely to have suffered *any* forms of physical violence. It is only when the variable for physical violence is disaggregated between severe and less severe that an effect is visible. Columns (4) in fact shows that among uneducated women, those with a first-born girl are 3.6 percentage points more likely to have been the victims of severe violence (including strangling or burning, or threatening with a gun or knife). This effect is in accord with Bloch and Rao (2002) who find that the number of male children reduces the probability of violence consistent with increased husband's satisfaction with the marriage in a male-dominated society that greatly values males. The effect on severe violence is specific to the period immediately after the birth of the child (11 months or less) seemingly as a form of immediate punishment inflicted on women (panel (b) and (c) indicate no effect on women aged 15–29 and 30–49). It is also specific to uneducated women (who are generally more likely to experience spousal violence). No effect is found for the other forms of violence.

Domestic violence thus appears to be another channel (complementary to or potentially exacerbating fertility behavior) that contributes to explaining excess mortality of women with first-born girls around the time of the first-birth. As for extreme fertility behavior, this is also motivated by strong son preference. However, even though it is consistent with selective mortality (the weakest are more likely die), the fact that it is only specific to the time of birth does not reflect the differential age patterns of anemia status (higher among women with a first girl during their entire reproductive years or below age 30, as shown in figure 7). The anemia patterns are instead more consistent with the explanation based on son-preferring fertility behavior.

7   Implied effect on women's excess mortality

Next, I turn to calculating the implied impact of son preference on mortality among adult women from the observed sex ratios for first births in the data.[49] The observed sex ratio (s_o) can be expressed as: $s_o = \frac{s_n}{1-m(1-s_n)}$; where s_n is the biologically normal sex ratio, and m is the mortality rate among women with first-born girls due to son preference (i.e., with no son preference: $m = 0$ and $s_o = s_n$). Therefore, $m = \frac{s_o - s_n}{s_o(1-s_n)}$. Table 16 reports the observed sex ratios by age group in the entire sample and the subsample of uneducated women. Figure 12 shows the implied values of m (for each age group) among women with first-born girls assuming a normal sex ratio of 0.514. As expected, the inferred mortality rates increase with age, ranging from 2.2 to 8.4% for women aged 30 to 49 (3.7 to 13% among uneducated women).

8   Conclusions

This paper examines the phenomenon of excess mortality among women of reproductive and adult ages in India. It is the first to show that strong preferences for male children are partly driving female mortality in adulthood. In settings with son preference, there are two main ways for parents to ensure the desired sex composition of children. They can have repeated and closely spaced pregnancies until the ideal number of sons has been reached, a behavior medically known to increase the risk of maternal mortality and morbidity. Parents can also selectively abort female fetuses. In India, the latter appears to be more common among well-off, educated women, and the former among the poor and uneducated. Therefore, differential access to prenatal care as well as differences in initial health status between women from different socioeconomic strata can lead to differential mortality due to son preference.

I explore this issue by using three rounds of the National Family Health Survey for India. The main results can be summarized as follows. First, there are no systematic differences among women with a first-born girl or boy close to the time of birth. Second, women with first-born daughters are more likely to desire more children, to have performed an ultrasound test, to have had a terminated pregnancy, to have given birth to more children, and less likely to be sterilized. The effect on abortion is significantly larger for educated women (who also exhibit an attenuated effect on fertility), which confirms existing evidence that sex selection is more common among them. Third, women with a first-born girl are more likely to develop anemia more than a year *after* birth. Fourth, differential incidence of anemia disappears in the subsample of older women. This age pattern of anemia is stronger for women with no education. Fifth, the share of living women with

---

[49]This approach is adapted from Abrevaya (2009) who estimates the prevalence of sex-selective abortions from the observed sex ratios.

a first-born girl decreases with the women's age at the time of the survey, and this is again specific to uneducated women. These results are consistent with a model in which son preference induces women into a fertility behavior that increases their risk of death and morbidity. Women with first-born daughters, especially the poor and uneducated among them, are in fact more likely to develop anemia and appear to suffer higher maternal and adult mortality. Higher mortality can also be partly attributed to increased domestic violence close to the time of birth for these women. The findings also indicate that neither women's food consumption nor iron supplementation is affected by the sex of their first-born. This can inform targeted interventions aiming to compensate higher nutritional needs of women with first-born girls.

Altogether, the findings of this paper vividly emphasize how vulnerable women are in certain contexts and help to understand the factors exacerbating the burden of maternal mortality. Although progress in reducing maternal mortality has been made over the last decades, many women continue not to receive adequate maternal care, especially among the poorest. To ensure further progress, it is crucial to guarantee access and availability of maternal care services to all strata of the population. This is of foremost importance in the context of the fifth Millennium Development Goal that aims to reduce by three-quarters the maternal mortality ratio and achieve universal access to reproductive health care by 2015.

Finally, the paper highlights once again how deeply rooted son preference is in Indian society and its potential for hindering safe motherhood and better child health.

Demographic, fertility and education variables. "*eduyrs*" is the number of completed years of education. "*eduyrs>0*" is a dummy equal to one if the woman completed at least one year of education, zero otherwise. "*wants more children*" is a dummy equal to one if the woman wants more children at the time of the survey, zero otherwise. "*sterilized*" is a dummy equal to one if the woman has undergone female sterilization, zero otherwise (or if she is using another method of contraception). "*ultrasound*" is a dummy equal to one if the woman has undergone an ultrasound test (zero otherwise), asked only with reference to the two most recent births in the last three years preceding the survey (NFHS 1998), and to any pregnancy in the last five years preceding the survey (NFHS 2005) (not asked in NFHS 1992). "*ever had terminated pregnancy*" is a dummy equal to one if the woman ever had a pregnancy that terminated in a miscarriage, abortion, or still birth, zero otherwise. "*spacing*" refers to the time interval between first and second birth and it is therefore available only for women who had at least two births.

Health indicators. "*anemic*" is a dummy equal to one if the measured hemoglobin count is less than 12 grams per deciliter (less than 11 grams for pregnant women). "*mild*" anemia refers to hemoglobin count between 10.0 and 11.9 grams per deciliter (10.0 and 10.9 for pregnant women). "*moderate*" anemia refers to hemoglobin count between 7.0 and 9.9 grams per deciliter. "*severe*" anemia refers to hemoglobin count less than 7.0 grams per deciliter. Appropriate adjustments in these cutoff points were made for women living at altitudes above 1,000 meters and women who smoke (NFHS data). Hemoglobin measurement is available for all women in NFHS 1998 and 2005 (not available in NFHS 1992). The body mass index (BMI) is calculated as weight in kilograms divided by the square of height in meters and it is not available for currently pregnant women and those who gave birth in the two months preceding the survey. A value of the BMI lower than 18.5 is used to define chronic energy deficiency. Height measurement is in centimeters and women whose height is below 145 cm are considered to be of short stature. BMI and height are available only in NFHS 1998 and 2005.

Spousal violence. Spousal violence variables are based on a module of NFHS 2005 that collected detailed information on physical, sexual, and emotional violence perpetrated by husbands against their wives (based on women's self-report). "*any severe violence*" includes whether the spouse ever tried to strangle or burn, or ever threatened her with knife or gun. "*any less severe violence*" includes whether the spouse ever pushed, slapped, punched, kicked or dragged her. "*any emotional violence*" includes whether the husband ever said something to humiliate her in front of others, or threatened to hurt someone close to her, or insulted her. "*any sexual violence*" includes whether the husband forced her to perform any unwanted sexual acts. Domestic violence variables available for all married women in NFHS 2005.

Food consumption. Food consumption variables are based on NFHS questions on frequency of consumption of certain food groups. The variables are dummies equal to one if the woman consumes the indicated type of food at least once a week. Food consumption is available for all women in NFHS 1998 and 2005 (not available in NFHS 1992).

References

[1] Abrevaya, Jason. 2009. "Are There Missing Girls in the United States? Evidence from Birth Data" *American Economic Journal: Applied Economics*, Vol. 1, No. 2, pp. 1-34.

[2] AbouZahr, Carla. 2003. "Global burden of maternal death and disability." *British Medical Bulletin*, Vol. 67, pp. 1-11.

[3] Ackerson, L. K., and S. V. Subramanian. 2008. "Domestic Violence and Chronic Malnutrition among Women and Children in India" *American Journal of Epidemiology*, Vol. 167, No. 10, pp. 1188–1196.

[4] Akbulut-Yuksel, Mevlude, and Daniel Rosenblum. 2012. "The Indian Ultrasound Paradox." IZA Discussion Paper No. 6273.

[5] Almond, Douglas, and Bhashkar Mazumder. 2011. "Health Capital and the Prenatal Environment: The Effect of Ramadan Observance During Pregnancy." *American Economic Journal: Applied Economics*, Vol. 3, No. 4, pp. 56–85.

[6] Anderson, Siwan and Debraj Ray. 2010. "Missing Women: Age and Disease." *Review of Economic Studies*, Vol. 77, No. 4, pp. 1262-1300.

[7] —. 2012. "The Age Distribution of Missing Women in India." *Economic & Political Weekly*, Vol. 47, No. 47-48, pp. 87–95.

[8] Andersson, Roland, and Staffan Bergstrom. 1998. "Is Maternal Malnutrition Associated with a Low Sex Ratio at Birth?" *Human Biology*, Vol. 70, No. 6, pp. 1101-1106.

[9] Arnold, F., M. K. Choe, and T. Roy. 1998. "Son Preference, the Family-Building Process and Child Mortality in India." *Population Studies*, Vol. 52, pp. 301–315.

[10] Arnold, Fred, Sunita Kishor, and T. K. Roy. 2002. "Sex-selective Abortions in India." *Population and Development Review*, Vol. 28, No. 4, pp. 759–785.

[11] Barcellos, Silvia H., Leandro Carvalho, and Adriana Lleras-Muney. 2014. "Child Gender And Parental Investments in India: Are Boys And Girls Treated Differently?" *American Economic Journal: Applied Economics*, Vol. 6, No. 1, pp. 157-189.

[12] Bhalotra, Sonia, and Tom Cochrane. 2010. "Where Have All the Young Girls Gone? Identification of Sex Selection in India." IZA Discussion Paper No. 5381.

[13] Bharadwaj, Prashant and Leah K. Lakdawala. 2013. "Discrimination Begins in the Womb: Evidence of Sex Selective Prenatal Investments." *Journal of Human Resources*, Vol. 48, No. 1, pp. 71–113.

[14] Bhat, Mari. 2002a. "On the Trail of 'Missing' Indian Females: II: Illusion and Reality." *Economic and Political Weekly*, Vol. 37, No. 52, pp. 5244–5263.

[15] —. 2002b. "Maternal Mortality in India: An Update." *Studies in Family Planning*, Vol. 33, No.3, pp. 227–236.

[16] Bhat, Mari, and Francis Zavier. 2007. "Factors Influencing the Use of Prenatal Diagnostic Techniques and the Sex Ratio at Birth in India." *Economic and Political Weekly*, Vol. 42, No. 24.

[17] Bloch, Francis, and Vijayendra Rao. 2002. "Terror as a Bargaining Instrument: A Case Study of Dowry Violence in Rural India." *American Economic Review*, Vol. 92, No. 4, pp. 1029–1043.

[18] Bozzoli, Carlos, Angus Deaton, and Climent Quintana-Domeque. 2009. "Adult Height and Childhood Disease." *Demography*, Vol. 46, No. 4, pp. 647–669.

[19] Center for Reproductive Rights (CPR). 2008. "Maternal Mortality in India." Center for Reproductive Rights, New York.

[20] Chowdhury, Mridul K. and Radheshyam Bairagi. 1990. "Son Preference and Fertility in Bangladesh." *Population and Development Review*, Vol. 16, No. 4, pp. 749–757.

[21] Clark, Shelley. 2000. "Son Preference and Sex Composition of Children: Evidence from India." *Demography*, Vol. 37, No. 1, pp. 95–108.

[22] Conde-Agudelo, A., Anyeli Rosas-Bermdez, and Ana C. Kafury-Goeta. 2007. "Effects of birth spacing on maternal health: a systematic review" *American Journal of Obstetrics and Gynecology*, Vol. 196, No. 4, pp. 297–308.

[23] Conde-Agudelo, A., and Belizan, J.M. 2000. "Maternal morbidity and mortality associated with inter-pregnancy interval: Cross sectional study" *British Medical Journal*, Vol. 321, No. 7271, pp. 1255–1259.

[24] Das, Narayan. 1987. "Sex Preference and Fertility Behavior: A Study of Recent Indian Data." *Demography*, Vol. 24, No. 4, pp. 517–530.

[25] Das Gupta, M. (1987): "Selective Discrimination Against Female Children in Rural Punjab, India." *Population and Development Review*, Vol. 13, No. 1, pp. 77–100.

[26] —. 2005. "Explaining Asia's 'Missing Women': A New Look at the Data." *Population and Development Review*, Vol. 31, No. 3, pp. 529–535.

[27] Das Gupta, Monica, Jiang Zhenghua, Li Bohua, Xie Zhenming, Woojin Chung, Bae Hwa-Ok. 2003. "Why is Son preference so persistent in East and South Asia? a cross-country study of China, India and the Republic of Korea." *Journal of Development Studies*, Vol. 40, No. 2, pp. 153–187.

[28] Dreze, Jean and Mamta Murthi. 2001. "Fertility, Education, and Development: Evidence from India." *Population and Development Review*, Vol. 27, No. 1, pp. 33–63.

[29] Filippi, Veronique, Carine Ronsmans, Oona M R Campbell, Wendy J Graham, Anne Mills, Jo Borghi, Marjorie Koblinsky, and David Osrin. 2006. "Maternal health in poor countries: the broader context and a call for action." *Lancet*, Vol. 368, No. 9546, pp. 1535-41.

[30] Filmer, Deon, Jed Friedman, and Norbert Schady. 2009. "Development, Modernization, and Childbearing: The Role of Family Sex Composition." *The World Bank Economic Review*, Vol. 23, No. 3, pp. 371–398.

[31] George, Sabu M. 2002. "Sex Selection/Determination in India: Contemporary Developments." *Reproductive Health Matters*, Vol. 10, No. 19, pp. 190–192.

[32] Goodkind, Daniel. 1999. "Should prenatal sex selection be restricted? Ethical questions and their implications for research and policy." *Population Studies*, Vol. 53, No. 1, pp. 49–61.

[33] Gragnolati, Michele, Meera Shekar, Monica Das Gupta, Caryn Bredenkamp and Yi-Kyoung Lee. 2005. "India's Undernourished Children: A Call for Reform and Action." Health, Nutrition and Population (HNP) Discussion Paper, World Bank.

[34] Hu, Luojia, and Analia Schlosser. 2011. "Prenatal Sex Selection and Girls Well-Being: Evidence from India." IZA Discussion Paper No. 5562.

[35] International Institute for Population Sciences (IIPS) and Macro International. 2007. "National Family Health Survey (NFHS-3), 2005-06: India: Volume I". Mumbai: IIPS.

[36] Jayachandran, Seema and Ilyana Kuziemko. 2011. "Why Do Mothers Breastfeed Girls Less Than Boys: Evidence and Implications for Child Health in India." *The Quarterly Journal of Economics*, Vol. 126, No. 3, pp. 1485–1538.

[37] Jayachandran, Seema, and Rohini Pande. 2013. "Choice Not Genes. Probable Cause for the India-Africa Child Height Gap." *Economic & Political Weekly*, Vol. 48, No. 34.

[38] Jayaraman, Rajshri, Debraj Ray, and Shing-Yi Wang. 2013. "Gender Differentials in Eye Care: Access and Treatment." working paper.

[39] James, William H. 2009. "Variation of the Probability of a Male Birth Within and Between Sibships." *Human Biology*, Vol. 81, No. 1, pp. 13–22.

[40] Jensen, Robert. 2003. "Equal Treatment, Unequal Outcomes? Generating Sex Inequality Through Fertility Behavior" Mimeo, Harvard University.

[41] Jha, Prabhat, Rajesh Kumar, Priya Vasa, Neeraj Dhingra, Deva Thiruchelvam, and Rahim Moineddin. 2006. "Low Male-to-Female Sex Ratio of Children Born in India: National Survey of 1.1 Million Households." *Lancet*, Vol. 367, No. 9506, pp. 211–218.

[42] Jha, Prabhat, Maya A Kesler, Rajesh Kumar, Faujdar Ram, Usha Ram, Lukasz Aleksandrowicz, Diego G Bassani, Shailaja Chandra, and Jayant K Banthia. 2011. "Trends in Selective Abortions of Girls in India: Analysis of Nationally Representative Birth Histories from 1990 to 2005 and Census Data from 1991 to 2011." *Lancet*, Vol. 377, No. 9781, Pages 1921–1928.

[43] Khan, Khalid S, Daniel Wojdyla, Lale Say, A Metin Glmezoglu, and Paul F A Van Look. 2006. "WHO analysis of causes of maternal death: a systematic review." *Lancet*, Vol. 367, pp. 1066-74.

[44] Kalaivani K. 2009. "Prevalence & Consequences of Anaemia in Pregnancy." *Indian Journal of Medical Research*, Vol. 130, No. 5, pp. 627–633.

[45] King, Janet C. 2003. "The Risk of Maternal Nutritional Depletion and Poor Outcomes Increases in Early or Closely Spaced Pregnancies." *Journal of Nutrition*, Vol. 133, No. 5, pp. 1732–1736.

[46] Klasen, Stephan. 1996. "Nutrition, Health, and Mortality in Sub Saharan Africa: Is There a Gender Bias?" *Journal of Development Studies*, Vol. 32, No. 6, pp. 913-933.

[47] Klasen, Stephan, and Sebastian Vollmer. 2013. "Missing Women: Age and Disease: A Correction." University of Göttingen, Discussion Paper No. 133.

[48] Koblinsky, Marge, Mahbub Elahi Chowdhury, Allisyn Moran, and Carine Ronsmans. 2012. "Maternal Morbidity and Disability and Their Consequences: Neglected Agenda in Maternal Health." *Journal of Health, Population, and Nutrition*, Vol. 30, No. 2, pp. 124–130.

[49] Kraemer, Sebastian. 2000. "The fragile male." *British Medical Journal*, Vol. 321, pp. 1609-12.

[50] Mathews, Fiona, Paul J. Johnson, and Andrew Neil. 2008. "You are what your mother eats: evidence for maternal preconception diet in uencing foetal sex in humans." *Proceedings of the Royal Society*, Vol. 275, No. 1643, pp. 1661–1668.

[51] Milazzo, Annamaria. 2012. "Son Preference, Fertility, and Family Structure. Evidence from Reproductive Behavior among Nigerian Women" Bocconi University, working paper.

[52] Mukuria, Altrena, Casey Aboulafia, and Albert Themme. 2005. "The Context of Womens Health: Results from the Demographic and Health Surveys, 1994-2001." DHS Comparative Reports No. 11, ORC Macro, Calverton, Maryland, USA.

[53] Nanda, Geeta, Kimberly Switlick, and Elizabeth Lule. 2005. "Accelerating Progress towards Achieving the MDG to Improve Maternal Health: A Collection of Promising Approaches." Health, Nutrition and Population (HNP) Discussion Paper, World Bank.

[54] Office of the Registrar General India. 2011. "Special Bulletin on Maternal Mortality in India 2007-09." Sample Registration System, Ministry of Home Affairs, Government of India.

[55] Oster, Emily. 2005. "Hepatitis B and the Case of the Missing Women." *Journal of Political Economy*, 2005, vol. 113, No. 6, pp. 1163–1216.

[56] Pallikadavath, Saseendran, and R. William Stones. 2006. "Maternal and Social Factors Associated with Abortion In India: A Population-Based Study." *International Family Planning Perspectives*, Vol. 32, No. 3, pp. 120-125.

[57] Ronsmans, Carine, and Wendy J Graham. 2006. "Maternal mortality: who, when, where, and why." *Lancet*, Vol. 368, pp. 1189-200.

[58] Rosenblum, Daniel. 2013. "The Effect of Fertility Decisions on Excess Female Mortality in India." *Journal of Population Economics*, Vol. 26, No. 1, pp. 147-180.

[59] Sen, Amartya. 1990. "More Than 100 Million Women Are Missing." *The New York Review of Books*, Vol. 37, No.20.

[60] Setty-Venugopal, V and U.D. Upadhyay. 2002. "Birth Spacing: Three to Five Saves Lives." *Population Reports*, Series L, No. 13. Baltimore: Johns Hopkins Bloomberg School of Public Health, Population Information Program.

[61] Sharman, Almaz. 2000. "Anemia Testing in Population-Based Surveys General Information and Guidelines for Country Monitors and Program Managers." ORC Macro, Calverton, Maryland, USA.

[62] Smith, Herbert L. 1994. "Nonreporting of Births or Nonreporting of Pregnancies? Some Evidence from Four Rural Counties in North China." *Demography*, Vol. 31, No. 3, pp. 481–486.

[63] Trussell, James Linda G. Martin, Robert Feldman, James A. Palmore, Mercedes Concepcion and Datin Noor Laily Bt. Dato' Abu Bakar. 1985. "Determinants of Birth-Interval Length in the Philippines, Malaysia, and Indonesia: A Hazard-Model Analysis." *Demography*, Vol. 22, No. 2, pp. 145–168.

[64] UNICEF. 2008. "The State of the World's Children 2009. Maternal and Newborn Health." UNICEF, New York.

[65] Waldron, Ingrid. 1983. "Sex Differences in Human Mortality: The Role of Genetic Factors." *Social Science & Medicine.*, Vol. 17, No. 6, pp. 321-333.

[66] World Bank. 1996. "Improving Women's Health in India." Development in Practice, World Bank.

[67] —. 2011. "World Development Report 2012. Gender Equality and Development." World Bank.

[68] World Health Organization (WHO). 2008. "Worldwide Prevalence of Anaemia 1993-2005. WHO Global Database on Anaemia." World Health Organization.

[69] —. 2010. "Trends in Maternal Mortality: 1990 to 2008. Estimates developed by WHO, UNICEF, UNFPA and The World Bank." World Health Organization, Geneva.

[70] —. 2007. "Maternal mortality in 2005: estimates developed by WHO, UNICEF, UNFPA, and the World Bank." World Health Organization, Geneva.

[71] —. 2011. "World Health Statistics." World Health Organization, Geneva.

Figures

Figure 1: Sex ratio (% male births), by cohort of birth
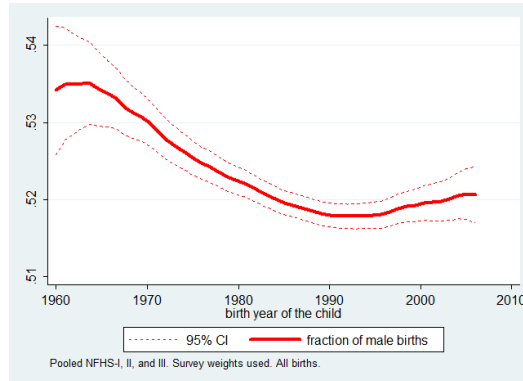
Figure 2: Sex ratio (% male births), by cohort of birth and NFHS round
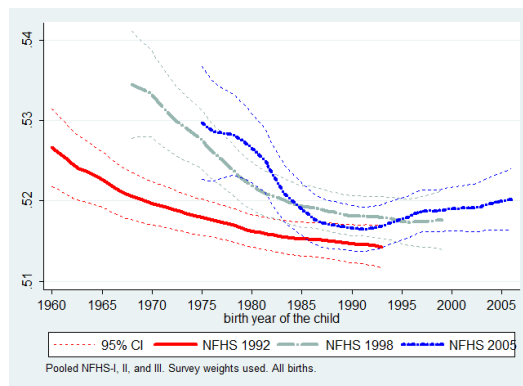
Figure 3: Sex ratio (% male births), by cohort of birth and birth order
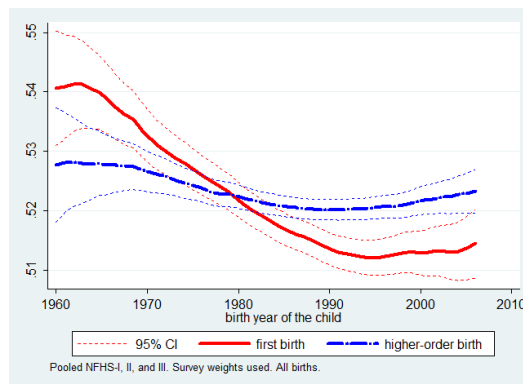
Figure 4: SHARE OF WOMEN WITH A FIRST-BORN GIRL BY AGE AND NFHS ROUND
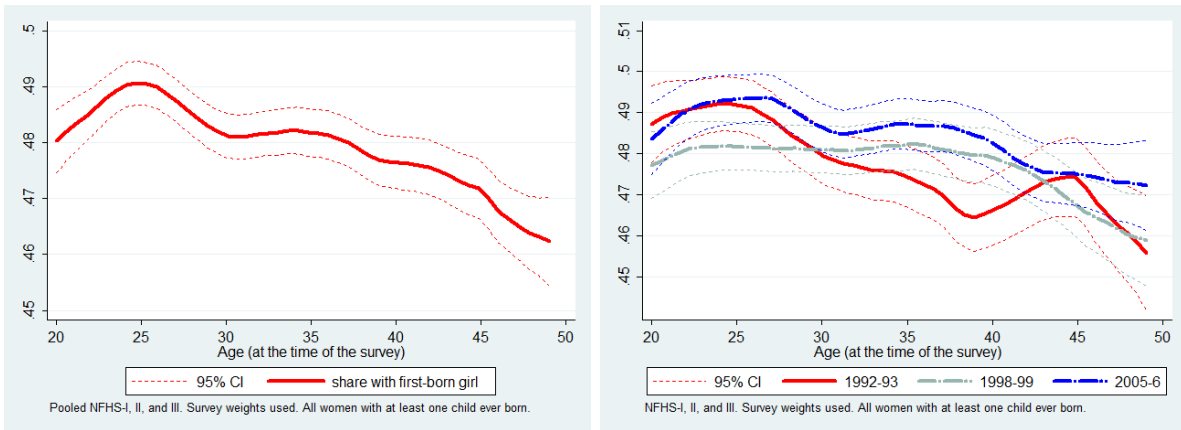
Figure 5: SHARE OF WOMEN WITH A FIRST-BORN GIRL BY AGE AND EDUCATION
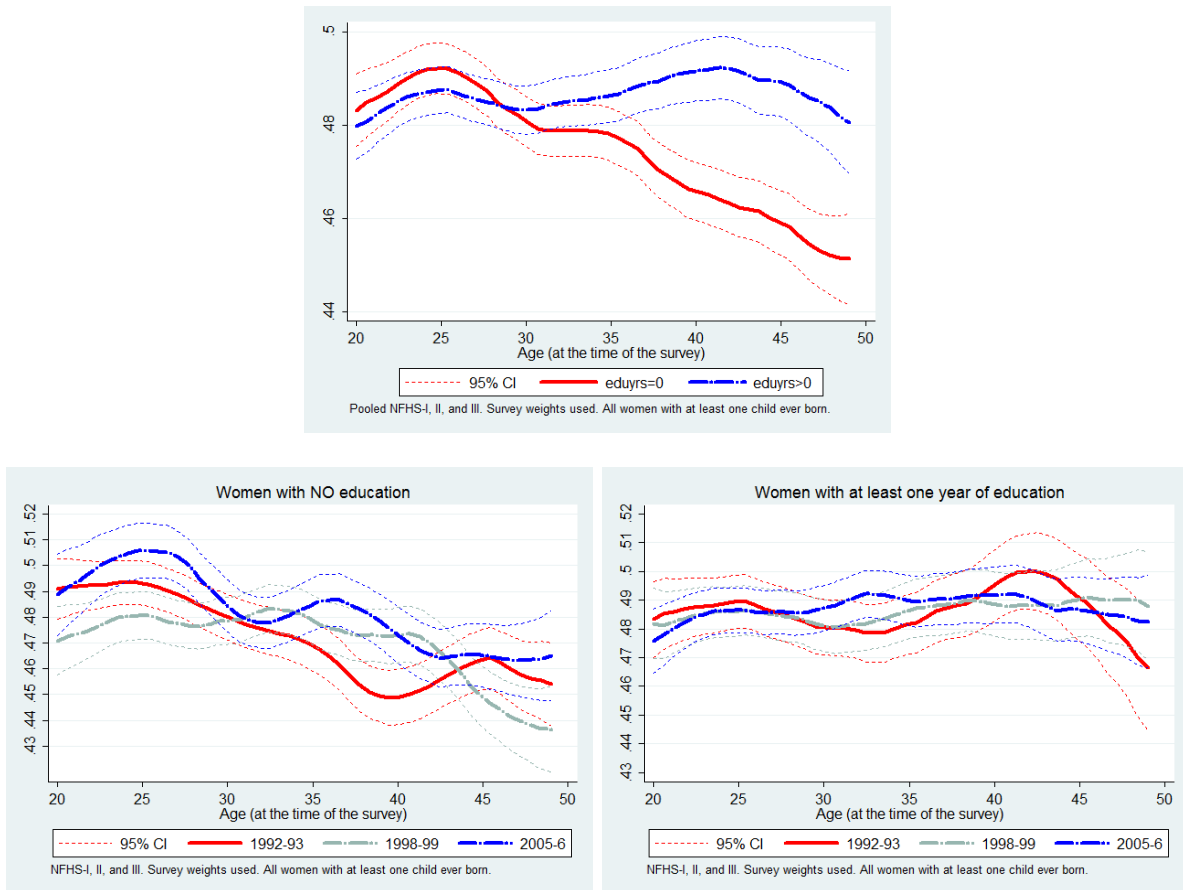
Figure 6: SHARE OF WOMEN WITH A FIRST-BORN GIRL BY AGE IN SELECTED INDIAN STATES

a) group 1 states: highest proportion missing at birth
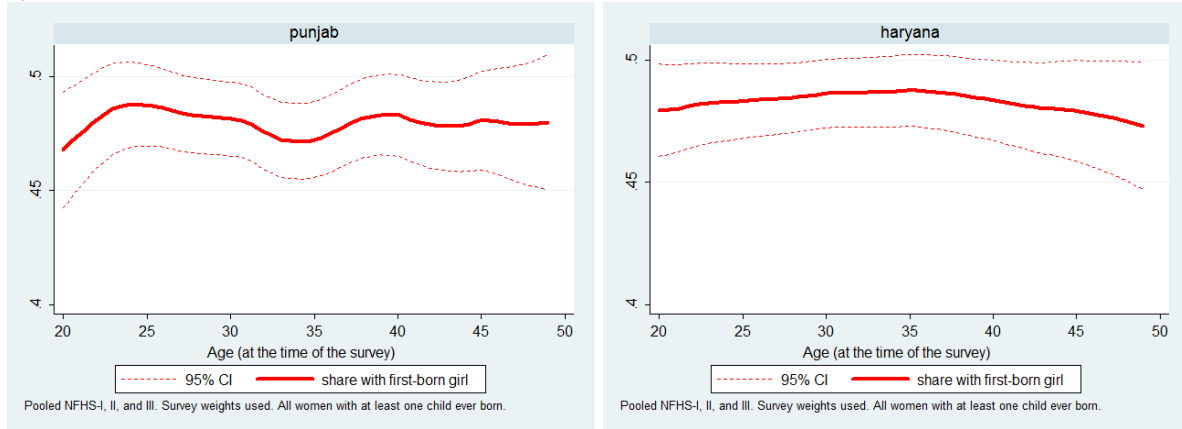
b) group 2 states: highest proportion of missing women in India & the majority missing at adult ages (15+)
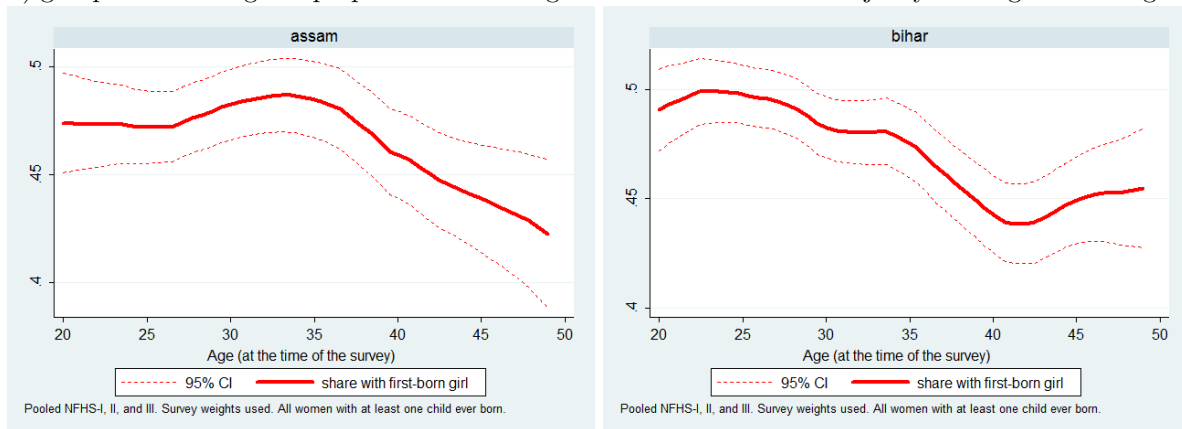
c) benchmark: lowest proportion of missing women in India & lowest female excess mortality at all ages
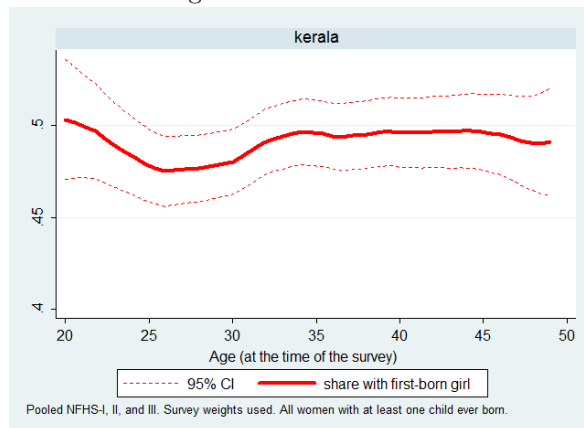
Figure 7: Share of anemic women by age and sex of the first-born child
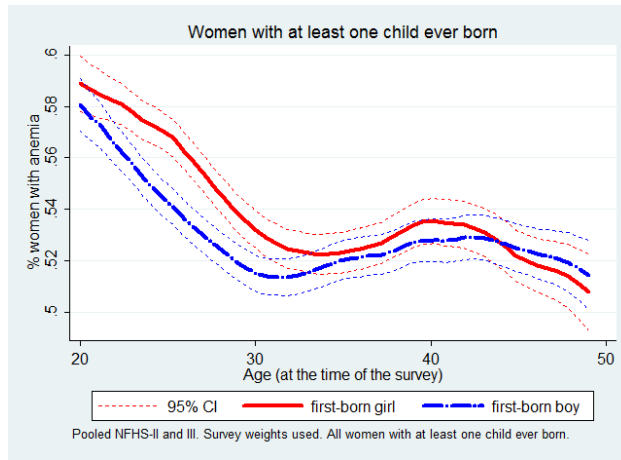
Figure 8: Body mass index (BMI) by age and sex of the first-born child
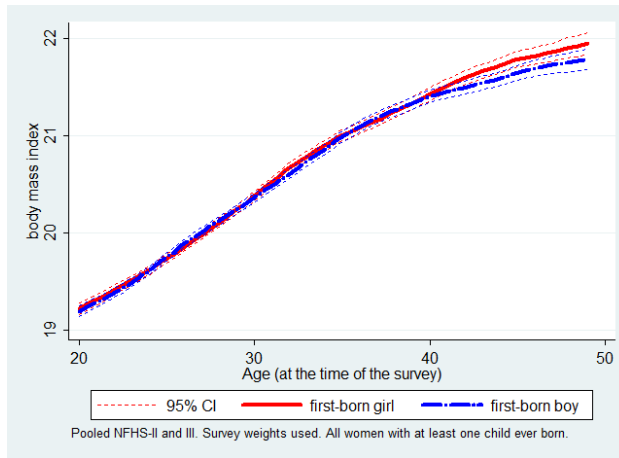
Notes: The BMI index excludes women who were pregnant at the time of the survey or women who had given birth during the two months preceding the survey (NFHS).

Figure 9: Height by age and sex of the first-born child

37
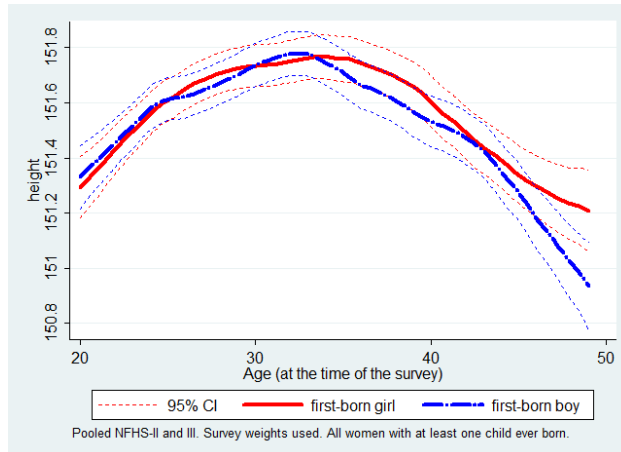
Figure 10: SHARE OF ANEMIC WOMEN BY LENGTH OF TIME SINCE FIRST BIRTH
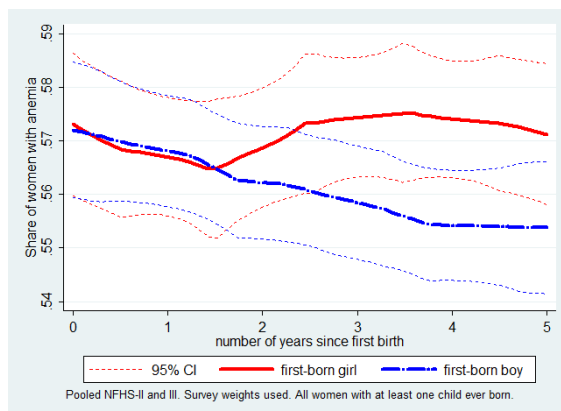
Figure 11: SEX RATIO (% MALE BIRTHS), BY COHORT OF BIRTH, BIRTH ORDER, AND EDUCATION LEVEL
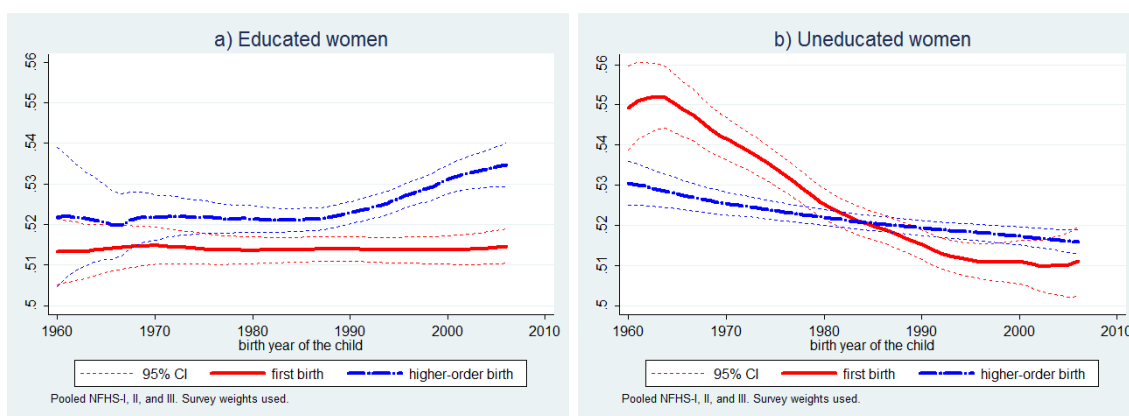
Figure 12: INFERRED PERCENT OF MISSING WOMEN WITH FIRST-BORN GIRLS BY AGE GROUP
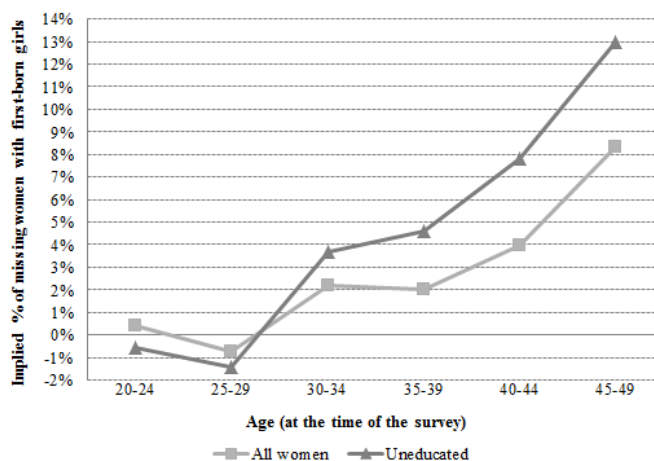
Tables

Table 1: Summary statistics for all ever-married women and by sex of the first-born child

	all women	first-born girl	first-born boy	diff	se	Obs
a) Demographic, fertility and education variables						
eduyrs>0	0.453	0.458	0.449	0.009***	(0.002)	239826
husband eduyrs>0 (dummy)	0.698	0.699	0.697	0.002	(0.002)	238745
eduyrs	3.512	3.538	3.489	0.049**	(0.022)	239826
husband eduyrs	6.087	6.102	6.073	0.029	(0.026)	238745
age	31.930	31.831	32.021	-0.190***	(0.042)	240112
age at first birth	18.923	18.945	18.902	0.043**	(0.017)	240112
age at first marriage	16.790	16.807	16.774	0.033**	(0.016)	240072
wants more children	0.227	0.252	0.204	0.048***	(0.002)	230795
# children ever born	3.203	3.300	3.114	0.186***	(0.009)	240112
sterilized	0.361	0.336	0.385	-0.049***	(0.002)	240112
currently use any contraceptives	0.520	0.497	0.543	-0.046***	(0.003)	240112
ultrasound	0.268	0.275	0.262	0.013***	(0.004)	60112
ever had terminated pregnancy	0.194	0.199	0.189	0.010***	(0.002)	240093
spacing (bw first and second birth) (# months)	32.636	32.193	33.051	-0.858***	(0.105)	198208
short spacing (bw first and second birth) (<15 months)	0.090	0.093	0.088	0.005***	(0.002)	198208
currently pregnant	0.056	0.061	0.051	0.010***	(0.001)	240112
currently breastfeeding	0.301	0.319	0.285	0.034***	(0.002)	240112
b) Health indicators						
anemic	0.541	0.547	0.536	0.011***	(0.003)	147427
mildly anemic	0.372	0.375	0.370	0.005*	(0.003)	147427
moderately anemic	0.151	0.154	0.148	0.006***	(0.002)	147427
severely anemic	0.017	0.017	0.017	-0.000	(0.001)	147427
height	151.527	151.546	151.508	0.038	(0.037)	153864
height<145cm	0.123	0.123	0.123	-0.000	(0.002)	153864
body mass index (BMI)	20.581	20.593	20.570	0.023	(0.025)	144590
BMI<18.5	0.346	0.346	0.346	-0.000	(0.003)	144590
c) Food consumption (at least once a week)						
consumes milk	0.550	0.549	0.552	-0.002	(0.003)	162661
consumes pulses	0.887	0.887	0.886	0.000	(0.002)	162655
consumes green vegetables	0.891	0.892	0.891	0.001	(0.002)	162642
consumes fruits	0.354	0.355	0.354	0.001	(0.003)	162573
consumes eggs or meat or fish	0.413	0.412	0.413	-0.001	(0.003)	162628
d) Spousal violence						
any severe physical violence	0.124	0.126	0.124	0.002	(0.004)	62564
any less severe physical violence	0.365	0.365	0.366	-0.001	(0.005)	62564
any emotional violence	0.162	0.164	0.159	0.005	(0.004)	62569
any sexual violence	0.100	0.102	0.098	0.004	(0.003)	62569

Demographic, fertility and education variables. "eduyrs" is the number of completed years of education. "eduyrs>0" is a dummy equal to one if the woman completed at least one year of education, zero otherwise. "wants more children" is a dummy equal to one if the woman wants more children at the time of the survey, zero otherwise. "sterilized" is a dummy equal to one if the mother has undergone female sterilization, zero otherwise (or if she is using another method of contraception). "ultrasound" is a dummy equal to one if the woman has undergone ultrasound test (zero otherwise), asked only with reference to the two most recent births in the last three years preceding the survey (NFHS 1998), and to any pregnancy in the last five years preceding the survey (NFHS 2005) (not asked in NFHS 1992). "ever had terminated pregnancy" is a dummy equal to one if the woman ever had a pregnancy that terminated in a miscarriage, abortion, or still birth, zero otherwise. Spacing refers to the time interval between first and second birth and it is therefore available only for women who had at least two births. **Health indicators**. "anemic" is a dummy equal to one if the measured hemoglobin count is less than 12 grams per deciliter (less than 11 grams for pregnant women). "mild" anemia refers to hemoglobin count between 10.0 and 11.9 grams per deciliter (10.0 and 10.9 for pregnant women). "moderate" anemia refers to hemoglobin count between 7.0 and 9.9 grams per deciliter. "severe" anemia refers to hemoglobin count less than 7.0 grams per deciliter. Appropriate adjustments in these cutoff points were made for women living at altitudes above 1,000 meters and women who smoke (NFHS data). Hemoglobin measurement is available for all women in NFHS 1998 and 2005 (not available in NFHS 1992). The body mass index (BMI) is calculated as weight in kilograms divided by the square of height in meters and it is not available for currently pregnant women and those who gave birth in the two months preceding the survey. A value of the BMI lower than 18.5 is used to define chronic energy deficiency. Height measurement is in centimeters and women whose height is below 145 cm are considered to be of short stature. BMI and height are available only in NFHS 1998 and 2005. **Spousal violence**. Spousal violence variables are based on a module of NFHS 2005 that collected detailed information on physical, sexual, and emotional violence perpetrated by husbands against their wives (based on women's self-report). "any severe violence" includes whether the spouse ever tried to strangle or burn, or ever threatened her with knife or gun. "any less severe violence" includes whether the spouse ever pushed, slapped, punched, kicked or dragged her. "any emotional violence" includes whether the husband ever said something to humiliate her in front of others, or threatened to hurt someone close to her, or insulted her. "any sexual violence" includes whether the husband forced her to perform any unwanted sexual acts. Domestic violence variables available for all married women in NFHS 2005. **Food consumption.** Food consumption variables are based on NFHS questions on frequency of consumption of certain food groups. The variables are dummies equal to one if the woman consumes the indicated type of food at least once a week. Food consumption is available for all women in NFHS 1998 and 2005 (not available in NFHS 1992). Using survey weights. ***, **, and * indicate significance at 1%, 5% and 10% levels, respectively.

Table 2: Selected summary statistics by sex of the first-born child, time since first birth, and age group

a) Less than a year after first birth

	first-born girl	first-born boy	diff	se	Obs
eduyrs>0	0.604	0.593	0.011	(0.012)	10212
husband eduyrs>0	0.785	0.791	-0.006	(0.010)	10169
eduyrs	5.186	5.133	0.053	(0.122)	10212
husband eduyrs	7.419	7.382	0.037	(0.124)	10169
age	20.694	20.683	0.011	(0.083)	10217
age at first birth	20.231	20.215	0.016	(0.083)	10217
age at first marriage	18.014	18.016	-0.002	(0.080)	10217
wants more children	0.894	0.861	0.033***	(0.008)	10126
# children ever born	1.008	1.009	-0.001	(0.002)	10217
sterilized	0.001	0.001	0.000	(0.001)	10217
ultrasound	0.37	0.356	0.014	(0.015)	5979
ever had terminated pregnancy	0.101	0.093	0.008	(0.007)	10216
anemic	0.584	0.58	0.004	(0.016)	5958
bmi	19.576	19.644	-0.068	(0.103)	4747
height	151.578	151.429	0.149	(0.187)	6252

b) Women aged 15–29

	first-born girl	first-born boy	diff	se	Obs
eduyrs>0	0.476	0.48	-0.004	(0.004)	97834
husband eduyrs>0	0.706	0.708	-0.002	(0.004)	97389
eduyrs	3.672	3.721	-0.049	(0.034)	97834
husband eduyrs	6.167	6.205	-0.038	(0.039)	97389
age	24.014	23.998	0.016	(0.026)	97958
age at first birth	18.287	18.276	0.011	(0.022)	97958
age at first marriage	16.48	16.485	-0.005	(0.021)	97947
wants more children	0.476	0.404	0.072***	(0.004)	96053
# children ever born	2.382	2.255	0.127***	(0.010)	97958
sterilized	0.195	0.252	-0.057***	(0.003)	97958
ultrasound	0.28	0.27	0.010*	(0.005)	43087
ever had terminated pregnancy	0.166	0.158	0.008***	(0.003)	97951
anemic	0.575	0.556	0.019***	(0.005)	57747
bmi	19.604	19.627	-0.023	(0.032)	52790
height	151.444	151.437	0.007	(0.057)	60368

c) Women aged 30–49

	first-born girl	first-born boy	diff	se	Obs
eduyrs>0	0.444	0.426	0.018***	(0.003)	141992
husband eduyrs>0	0.693	0.688	0.005*	(0.003)	141356
eduyrs	3.435	3.316	0.119***	(0.029)	141992
husband eduyrs	6.053	5.976	0.077**	(0.034)	141356
age	37.876	37.991	-0.115***	(0.037)	142154
age at first birth	19.456	19.369	0.087***	(0.025)	142154
age at first marriage	17.061	16.989	0.072***	(0.022)	142125
wants more children	0.072	0.05	0.022***	(0.002)	134742
# children ever born	4.008	3.752	0.256***	(0.012)	142154
sterilized	0.445	0.484	-0.039***	(0.003)	142154
ultrasound	0.262	0.236	0.026***	(0.008)	17025
ever had terminated pregnancy	0.225	0.213	0.012***	(0.003)	142142
anemic	0.526	0.521	0.005	(0.004)	89680
bmi	21.229	21.172	0.057	(0.035)	91800
height	151.622	151.559	0.063	(0.048)	93496

Demographic, fertility and education variables. "eduyrs" is the number of completed years of education. "eduyrs>0" is a dummy equal to one if the woman completed at least one year of education, zero otherwise. "wants more children" is a dummy equal to one if the woman wants more children at the time of the survey, zero otherwise. "sterilized" is a dummy equal to one if the mother has undergone female sterilization, zero otherwise (or if she is using another method of contraception). "ultrasound" is a dummy equal to one if the mother has undergone ultrasound test (zero otherwise), asked only with reference to the two most recent births in the last three years preceding the survey (NFHS 1998), and to any pregnancy in the last five years preceding the survey (NFHS 2005) (not asked in NFHS 1992). "ever had terminated pregnancy" is a dummy equal to one if the mother ever had a pregnancy that terminated in a miscarriage, abortion, or still birth, zero otherwise. **Health indicators**. "anemic" is a dummy equal to one if the measured hemoglobin count is less than 12 grams per deciliter (less than 11 grams for pregnant women). Appropriate adjustments in these cutoff points were made for women living at altitudes above 1,000 meters and women who smoke (NFHS data). Hemoglobin measurement is available for all women in NFHS 1998 and 2005 (not available in NFHS 1992). The body mass index (BMI) is calculated as weight in kilograms divided by the square of height in meters and it is not available for currently pregnant women and those who gave birth in the two months preceding the survey. Height measurement is in centimeters. BMI and height are available only in NFHS 1998 and 2005. Using survey weights. ***, **, and * indicate significance at 1%, 5% and 10% levels, respectively.

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

dep. var:	=1 if wants more children			=1 if sterilized			=1 if undergone ultrasound			=1 if ever had terminated pregnancy			number of children ever born		
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)	(11)	(12)	(13)	(14)	(15)
first-born girl	0.049***	0.051***	0.053***	-0.052***	-0.068***	-0.065***	0.012***	0.025***	0.025***	0.010***	0.011***	0.005*	0.301***	0.295***	0.324***
	(0.002)	(0.002)	(0.003)	(0.002)	(0.002)	(0.004)	(0.004)	(0.004)	(0.008)	(0.002)	(0.002)	(0.003)	(0.010)	(0.008)	(0.013)
first-born girl*eduyrs>0			-0.005			-0.008			0.000			0.014***			-0.058***
			(0.004)			(0.005)			(0.010)			(0.004)			(0.015)
eduyrs>0		-0.046***	-0.043***		0.025***	0.028***		0.086***	0.085***		0.017***	0.010***		-0.334***	-0.307***
		(0.002)	(0.002)		(0.003)	(0.004)		(0.005)	(0.008)		(0.003)	(0.003)		(0.011)	(0.013)
urban		-0.018***	-0.018***		-0.024***	-0.024***		0.099***	0.099***		0.023***	0.023***		-0.098***	-0.098***
		(0.002)	(0.002)		(0.004)	(0.004)		(0.006)	(0.006)		(0.003)	(0.003)		(0.010)	(0.010)
husband eduyrs>0		-0.021***	-0.021***		0.041***	0.041***		0.067***	0.067***		0.009***	0.009***		-0.111***	-0.111***
		(0.002)	(0.003)		(0.004)	(0.004)		(0.005)	(0.005)		(0.003)	(0.003)		(0.011)	(0.011)
age first birth		0.023***	0.023***		-0.034***	-0.034***		0.018***	0.018***		0.007***	0.007***		-0.153***	-0.153***
		(0.000)	(0.000)		(0.001)	(0.001)		(0.001)	(0.001)		(0.001)	(0.001)		(0.002)	(0.002)
age first marriage		-0.004***	-0.004***		0.001	0.001		0.001	0.001		-0.009***	-0.009***		-0.006**	-0.006**
		(0.000)	(0.000)		(0.001)	(0.001)		(0.001)	(0.001)		(0.001)	(0.001)		(0.003)	(0.003)
durables index		-0.010***	-0.010***		0.001	0.001		0.061***	0.061***		0.002**	0.002**		-0.134***	-0.134***
		(0.001)	(0.001)		(0.001)	(0.001)		(0.002)	(0.002)		(0.001)	(0.001)		(0.003)	(0.003)
# children ever born		-0.040***	-0.040***		0.017***	0.017***		-0.008**	-0.008**		-0.002**	-0.002**			
		(0.001)	(0.001)		(0.001)	(0.001)		(0.003)	(0.003)		(0.001)	(0.001)			
Observations	220087	220087	220087	220303	220234	220234	56669	56657	56657	220285	220275	220275	152291	152291	152291
Pseudo R2 or R2	0.003	0.393	0.393	0.003	0.225	0.225	0.000	0.307	0.307	0.000	0.0278	0.0278	0.008	0.480	0.480

Probit estimates in columns 1–12 (marginal effects reported). OLS estimates in columns 13–15. Robust standard errors adjusted for clustering at the household level in parentheses. "wants more children" is a dummy equal to one if the woman wants more children at the time of the survey, zero otherwise. "sterilized" is a dummy equal to one if the woman has undergone female sterilization, zero otherwise (or if she is using another method of contraception). "ultrasound" is a dummy equal to one if the woman has undergone ultrasound test (zero otherwise), asked only with reference to the two most recent births in the last three years preceding the survey (NFHS 1998), and to any pregnancy in the last five years preceding the survey (NFHS 2005) (not asked in NFHS 1992). "had terminated pregnancy" is a dummy equal to one if the woman ever had a pregnancy that terminated in a miscarriage, abortion, or still birth, zero otherwise. The regression for the number of children ever born is estimated on the subsample of women who have been sterilized or who do not want more children. Covariates include state fixed effects, urban dummy, birth year fixed effects, 7 five-year age groups dummies, survey round fixed effects, age at first birth and marriage, her and partner's current age (and square), whether she and her partner have completed one or more years of education, a durables index, religion and caste. The number of children ever born is included in columns 1–12. Using survey weights. ***, **, and * indicate significance at 1%, 5% and 10% levels, respectively.

Table 4: Birth spacing by sex of the child

a) **Time interval between first and second birth**

dep. var:	number of months		interval < 15 months	
	(1)	(2)	(3)	(4)
first-born girl	-0.778***	-0.960***	0.005***	0.004**
	(0.102)	(0.133)	(0.002)	(0.002)
first-born girl*eduyrs>0		0.422**		0.001
		(0.199)		(0.003)
eduyrs>0	0.143	-0.062	-0.005***	-0.006**
	(0.141)	(0.177)	(0.002)	(0.003)
Observations	182395	182395	182395	182395
Pseudo R2 or R2	0.026	0.026	0.010	0.010

b) **Time interval between second and third birth**
(conditional on first-born girl)

dep. var:	number of months		interval < 15 months	
	(1)	(2)	(3)	(4)
second-born girl	-0.523***	-0.944***	0.008***	0.015***
	(0.158)	(0.211)	(0.002)	(0.003)
second-born girl*eduyrs>0		1.149***		-0.019***
		(0.386)		(0.005)
eduyrs>0	0.483**	-0.102	-0.012***	-0.002
	(0.218)	(0.275)	(0.004)	(0.004)
Observations	63365	63365	63363	63363
Pseudo R2 or R2	0.024	0.024	0.010	0.010

OLS estimates in columns 1–2. Probit estimates in columns 3–4 (marginal effects reported). Robust standard errors adjusted for clustering at the household level in parentheses. Spacing is the length of time between births and it's expressed in number of months (columns 1–2), or as a dummy variable equal to one if the interval is shorter than 15 months (or less than six months from birth to conception of the next child). Covariates include state fixed effects, urban dummy, birth year fixed effects, 7 five-year age groups dummies, survey round fixed effects, age at first birth and marriage, her and partner's current age (and square), whether she and her partner have completed one or more years of education, a durables index, religion and caste, and number of children ever born. Using survey weights. ***, **, and * indicate significance at 1%, 5% and 10% levels, respectively.

Table 5: Birth spacing by sex of the child, with mother fixed effects

dep. var:	number of months			interval < 15 months		
Subsample:		eduyrs=0	eduyrs>0		eduyrs=0	eduyrs>0
	(1)	(2)	(3)	(4)	(5)	(6)
girl	-0.597***	-0.632***	-0.496**	0.005***	0.006***	0.002
	(0.113)	(0.129)	(0.229)	(0.002)	(0.002)	(0.004)
Observations	420703	243660	176488	420703	243660	176488
R-squared	0.593	0.556	0.658	0.506	0.484	0.548

OLS estimates. Robust standard errors adjusted for clustering at the household level in parentheses. The sample exclude births of birth order 4 and higher. Includes birth order fixed effects. Using survey weights. ***, **, and * indicate significance at 1%, 5% and 10% levels, respectively.

Table 6: Differences in the conditional age distribution by sex of the first-born child from ordered logit regression. Dep. variable: 5-year age groups.

women with:	at least one child ever born	
	first-born girl vs first-born boy	
	change	95% CI
	(1)	(2)
$Pr(y = 15 - 19 \| x)$	0.0000	[0.0000, 0.0001]
$Pr(y = 20 - 24 \| x)$	0.0026	[0.0022, 0.0029]
$Pr(y = 25 - 29 \| x)$	0.0283	[0.0247, 0.0318]
$Pr(y = 30 - 34 \| x)$	-0.0093	[-0.0106, -0.0081]
$Pr(y = 35 - 39 \| x)$	-0.0185	[-0.0208, -0.0161]
$Pr(y = 40 - 44 \| x)$	-0.0027	[-0.0031, -0.0024]
$Pr(y = 45 - 49 \| x)$	-0.0004	[-0.0004, -0.0003]
Observations	220303	
Pseudo R2	0.456	

Computed from the ordered logit estimates in which the dependent variable is a categorical variable for seven 5-year age groups and the independent variable of interest is "first-born girl" (column 1), and "girl-girl" (column 2). Covariates include state fixed effects, urban dummy, survey round fixed effects, age at first birth and marriage, partner's age (and square), whether she and her partner have completed one or more years of education, a durables index, religion and caste, and number of children ever born. Column 2 also includes dummies for whether the first two born children are boy-girl or girl-boy. The values reported are the changes in predicted probabilities when "first-born girl" or "girl-girl" change from 0 to 1, with all the other independent variables set at their mean value (for the regression in column 2, "girl-boy" and "boy-girl" are set at zero). 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. Using survey weights. ***, **, and * indicate significance at 1%, 5% and 10% levels, respectively.

Table 7: Anemia status by sex of the first-born child (or number of children ever born)

dep. var:	=1 if anemic										
Subsample								eduyrs=0	eduyrs>0		
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)	(11)
first-born girl	0.011***	0.010***	0.009***	0.009***	0.008**	0.014***	0.014***	0.020***	0.009		
	(0.003)	(0.003)	(0.003)	(0.003)	(0.003)	(0.005)	(0.005)	(0.007)	(0.007)		
first-born girl*age≥30						-0.011**	-0.010*	-0.020**	-0.002		
						(0.006)	(0.006)	(0.009)	(0.010)		
first-born girl*age≥45							-0.008	-0.006	-0.011		
							(0.013)	(0.016)	(0.018)		
currently pregnant			0.016*	0.016*	0.015*	0.014*	0.014*	0.008	0.023**	0.016*	0.015*
			(0.008)	(0.008)	(0.008)	(0.008)	(0.008)	(0.012)	(0.011)	(0.008)	(0.008)
currently breastfeeding			0.042***	0.042***	0.036***	0.036***	0.036***	0.038***	0.033***	0.036***	0.035***
			(0.005)	(0.005)	(0.005)	(0.005)	(0.005)	(0.007)	(0.007)	(0.005)	(0.005)
height				-0.001***	-0.001***	-0.001***	-0.001***	-0.002***	-0.001*	-0.001***	-0.001***
				(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)
eduyrs>0		-0.015***	-0.013***	-0.013***	-0.011**	-0.011**	-0.011**			-0.011**	-0.011**
		(0.005)	(0.005)	(0.005)	(0.005)	(0.005)	(0.005)			(0.004)	(0.005)
husband eduyrs>0		-0.013***	-0.012***	-0.012***	-0.011**	-0.011**	-0.011**	-0.011**	-0.008	-0.011**	-0.011**
		(0.005)	(0.005)	(0.005)	(0.005)	(0.005)	(0.005)	(0.005)	(0.010)	(0.005)	(0.005)
durables index		-0.021***	-0.020***	-0.020***	-0.019***	-0.019***	-0.019***	-0.022***	-0.017***	-0.019***	-0.019***
		(0.002)	(0.002)	(0.002)	(0.002)	(0.002)	(0.002)	(0.003)	(0.002)	(0.002)	(0.002)
urban		-0.008*	-0.006	-0.007*	-0.007	-0.006	-0.006	-0.005	-0.006	-0.007*	-0.007*
		(0.004)	(0.004)	(0.004)	(0.004)	(0.004)	(0.004)	(0.006)	(0.006)	(0.004)	(0.004)
# children ever born					0.007***	0.007***	0.007***	0.003*	0.016***	0.011***	0.011***
					(0.001)	(0.001)	(0.001)	(0.002)	(0.003)	(0.003)	(0.003)
# children ever born*age≥30										-0.004	-0.004
										(0.003)	(0.003)
# children ever born*age≥45											-0.001
											(0.004)
Covariates	N	Y	Y	Y	Y	Y	Y	Y	Y	Y	Y
Observations	134132	134132	134132	134132	134132	134132	134132	59448	74681	134132	134132
Pseudo R2	0.000	0.0352	0.0358	0.0360	0.0361	0.0362	0.0362	0.0310	0.0354	0.036	0.036

Probit estimates (marginal effects reported). Robust standard errors adjusted for clustering at the household level in parentheses. "anemic" is a dummy equal to one if the measured hemoglobin count is less than 12 grams per deciliter (less than 11 grams for pregnant women). Appropriate adjustments in these cutoff points were made for women living at altitudes above 1,000 meters and women who smoke (NFHS data). Hemoglobin measurement is available for all women in NFHS 1998 and 2005 (not available in NFHS 1992). Covariates include state fixed effects, urban dummy, birth year fixed effects, 7 five-year age groups dummies, NFHS 2005 survey round fixed effect, age at first birth and marriage, her and partner's current age (and square), whether she and her partner have completed one or more years of education, a durables index, religion and caste. Using survey weights. ***, **, and * indicate significance at 1%, 5% and 10% levels, respectively.

Table 8: Levels of severity of anemia by sex of the first-born child - Changes in predicted probabilities computed from ordered logit regression

	all women		women age 15–29		women age 30–49	
	change	95% CI	change	95% CI	change	95% CI
	(1)	(2)	(3)	(4)	(5)	(6)
$Pr(y =$severely anemic$|x)$	0.0004	[0.0001, 0.0007]	0.0006	[0.0001, 0.0010]	0.0003	[-0.0002, 0.0007]
$Pr(y =$moderately anemic$|x)$	0.0032	[0.0005, 0.0060]	0.0049	[0.0005, 0.0093]	0.0020	[-0.0014, 0.0055]
$Pr(y =$mildly anemic$|x)$	0.0037	[0.0005, 0.0068]	0.0043	[0.0004, 0.0083]	0.0026	[-0.0019, 0.0072]
$Pr(y =$not anemic$|x)$	-0.0073	[-0.0136, -0.0010]	-0.0098	[-0.0186, -0.0010]	-0.0049	[-0.0133, 0.0035]
Observations		134060		54328		79732
Pseudo R-sq		0.023		0.021		0.024

Computed from the ordered logit estimates in which the dependent variable is a categorical variable with four categories for the levels of anemia severity. "mild" anemia refers to hemoglobin count between 10.0 and 11.9 grams per deciliter (10.0 and 10.9 for pregnant women). "moderate" anemia refers to hemoglobin count between 7.0 and 9.9 grams per deciliter. "severe" anemia refers to hemoglobin count less than 7.0 grams per deciliter. Appropriate adjustments in these cutoff points were made for women living at altitudes above 1,000 meters and women who smoke (NFHS data). Hemoglobin measurement is available for all women in NFHS 1998 and 2005 (not available in NFHS 1992). The independent variable of interest is "first-born girl". The values reported are the changes in predicted probabilities when "first-born girl" change from 0 to 1, with all the other independent variables set at their mean value. Changes are reported with 95% confidence intervals by the delta method in parentheses. Covariates include state fixed effects, urban dummy, 2005 NFHS survey round fixed effect, age at first birth and marriage, partner's age (and square), whether she and her partner have completed one or more years of education, a durables index, religion and caste, number of children ever born, whether currently pregnant or breastfeeding, height. Column 1 also includes 7 five-year age groups dummies. Robust standard errors adjusted for clustering at the household level in parentheses. All women aged 15–49. Using survey weights. ***, **, and * indicate significance at 1%, 5% and 10% levels, respectively.

Table 9: Other health outcomes by sex of the first-born child

	BMI			BMI< 18.5			height			height<145cm		
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)	(11)	(12)
first-born girl	0.090***	0.018	0.018	-0.007**	0.000	0.000	0.053	0.042	0.042	-0.000	0.001	0.001
	(0.023)	(0.033)	(0.033)	(0.003)	(0.005)	(0.005)	(0.036)	(0.053)	(0.053)	(0.002)	(0.003)	(0.003)
first-born girl*age≥30		0.120***	0.102**		-0.013**	-0.011*		0.018	-0.009		-0.001	0.000
		(0.044)	(0.047)		(0.006)	(0.006)		(0.064)	(0.068)		(0.004)	(0.004)
first-born girl*age≥45			0.113			-0.014			0.171			-0.008
			(0.098)			(0.012)			(0.125)			(0.007)
Observations	131262	131262	131262	131262	131262	131262	140196	140196	140196	140196	140196	140196
R-squared	0.258	0.258	0.258	0.089	0.089	0.089	0.079	0.079	0.079	0.037	0.037	0.037

OLS estimates (columns 1–3 and 7–9). Probit estimates (marginal effects reported) (columns 4–6 and 10–12). Robust standard errors adjusted for clustering at the household level in parentheses. The body mass index (BMI) is calculated as weight in kilograms divided by the square of height in meters and it is not available for currently pregnant women and those who gave birth in the two months preceding the survey. A value of the BMI lower than 18.5 is used to define chronic energy deficiency. Height measurement is in centimeters and women whose height is below 145 cm are considered to be of short stature. BMI and height are available only in NFHS 1998 and 2005. Covariates include state fixed effects, urban dummy, birth year fixed effects, 7 five-year age groups dummies, NFHS 2005 survey round fixed effect, age at first birth and marriage, her and partner's current age (and square), whether she and her partner have completed one or more years of education, a durables index, religion and caste, number of children ever born, whether currently pregnant or breastfeeding. Using survey weights. ***, **, and * indicate significance at 1%, 5% and 10% levels, respectively.

Table 10: Anemia status by time since first birth

Sample:	all	eduyrs=0	eduyrs>0
	(1)	(2)	(3)
first-born girl	-0.005	-0.023	0.006
	(0.017)	(0.031)	(0.021)
first-born girl*12–23 months after first birth	-0.003	0.038	-0.025
	(0.025)	(0.040)	(0.029)
first-born girl*24–35 months after first birth	0.012	0.043	-0.001
	(0.022)	(0.042)	(0.026)
first-born girl*36–47 months after first birth	0.014	0.066*	-0.019
	(0.024)	(0.039)	(0.030)
Observations	22200	6584	15604
Pseudo R2	0.050	0.050	0.052

Probit estimates (marginal effects reported). The sample is of all women who had their first birth less than 48 months before the time of the survey. Robust standard errors adjusted for clustering at the household level in parentheses. "anemic" is a dummy equal to one if the measured hemoglobin count is less than 12 grams per deciliter (less than 11 grams for pregnant women). Appropriate adjustments in these cutoff points were made for women living at altitudes above 1,000 meters and women who smoke (NFHS data). Hemoglobin measurement is available for all women in NFHS 1998 and 2005 (not available in NFHS 1992). Covariates include state fixed effects, urban dummy, birth year fixed effects, years since first birth fixed effects, 7 five-year age groups dummies, NFHS 2005 survey round fixed effect, age at first birth and marriage, her and partner's current age (and square), whether she and her partner have completed one or more years of education, a durables index, religion and caste, number of children ever born, whether currently pregnant or breastfeeding, height. Using survey weights. ***, **, and * indicate significance at 1%, 5% and 10% levels, respectively.

Table 11: Probability of female child and anemia status less than a year after birth, by birth order

dep. var:	=1 if child birth order i is a female											
birth order:	1	2	3	1	2	3	1	2	3	1	2	3
anemic (any level)	-0.007	0.005	-0.019									
	(0.018)	(0.018)	(0.024)									
mild anemic				-0.006	0.000	-0.002						
				(0.018)	(0.018)	(0.023)						
moderately anemic							-0.004	0.007	-0.030			
							(0.023)	(0.026)	(0.027)			
severely anemic										0.029	-0.007	0.036
										(0.080)	(0.064)	(0.070)
Observations	5691	4938	2946	5691	4938	2946	5691	4938	2946	5691	4938	2946
Pseudo R2	0.0136	0.0162	0.0351	0.0136	0.0161	0.0349	0.0136	0.0162	0.0352	0.0136	0.0161	0.0349

Probit estimates (marginal effects reported). Robust standard errors adjusted for clustering at the household level in parentheses. "anemic" is a dummy equal to one if the measured hemoglobin count is less than 12 grams per deciliter (less than 11 grams for pregnant women). Appropriate adjustments in these cutoff points were made for women living at altitudes above 1,000 meters and women who smoke (NFHS data). Hemoglobin measurement is available for all women in NFHS 1998 and 2005 (not available in NFHS 1992). Covariates include state fixed effects, urban dummy, birth year fixed effects, 7 five-year age groups dummies, NFHS 2005 survey round fixed effect, age at first birth and marriage, her and partner's current age (and square), whether she and her partner have completed one or more years of education, a durables index, religion and caste, number of children ever born, whether currently pregnant or breastfeeding, height. Using survey weights. ***, **, and * indicate significance at 1%, 5% and 10% levels, respectively.

Table 12: Women's age at the time of the survey by sex of the first-born child and NFHS survey round

dep. var:	seven 5-year age groups	age
	ordered logit	OLS
	(1)	(2)
first-born girl	-0.180***	-0.287***
	(0.018)	(0.035)
first-born girl*wave 1998	0.033	0.013
	(0.024)	(0.046)
first-born girl*wave 2005	0.049*	0.068
	(0.027)	(0.042)
wave 1998	0.117***	0.229***
	(0.018)	(0.031)
wave 2005	0.447***	0.791***
	(0.017)	(0.029)
Observations	220303	220303
Pseudo R2 or R2	0.456	0.834

Ordered logit estimates in column (1), and OLS estimates in column (2). The dependent variable in column (1) is a categorical variable for 7 five-year age groups, and in column (2) it's the woman's age. Covariates include state fixed effects, urban dummy, survey round fixed effects, age at first birth and marriage, partner's age (and square), whether she and her partner have completed one or more years of education, a durables index, religion and caste, and number of children ever born. Robust standard errors adjusted for clustering at the household level in parentheses. All women aged 15–49. Using survey weights. ***, **, and * indicate significance at 1%, 5% and 10% levels, respectively.

Table 13: Food consumption by sex of the first-born child (daily or weekly consumption of specific food items)

a) Less than a year since first birth

	milk		pulses		green vegetables		fruits		eggs/meat/fish	
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)
first-born girl	-0.020	0.005	0.004	-0.014	0.002	-0.004	-0.001	0.024	-0.024	-0.027
	(0.016)	(0.028)	(0.007)	(0.013)	(0.009)	(0.014)	(0.018)	(0.028)	(0.017)	(0.031)
first-born girl*eduyrs>0		-0.039		0.029**		0.010		-0.035		0.004
		(0.038)		(0.015)		(0.018)		(0.036)		(0.040)
eduyrs>0	0.109***	0.129***	0.022*	0.006	0.045***	0.040**	0.140***	0.156***	0.033	0.031
	(0.026)	(0.032)	(0.012)	(0.014)	(0.013)	(0.016)	(0.023)	(0.032)	(0.025)	(0.032)
Observations	5972	5972	5961	5961	5959	5959	5973	5973	5977	5977
Pseudo R2	0.173	0.173	0.123	0.125	0.120	0.120	0.184	0.185	0.300	0.300

b) Women age 15–29

	milk		pulses		green vegetables		fruits		eggs/meat/fish	
first-born girl	-0.001	-0.003	-0.002	-0.005	0.000	-0.002	0.003	0.008	-0.003	-0.002
	(0.006)	(0.009)	(0.003)	(0.004)	(0.002)	(0.004)	(0.005)	(0.008)	(0.006)	(0.009)
first-born girl*eduyrs>0		0.004		0.006		0.005		-0.008		-0.002
		(0.011)		(0.005)		(0.005)		(0.010)		(0.011)
eduyrs>0	0.097***	0.095***	0.023***	0.019***	0.017***	0.014***	0.097***	0.101***	0.050***	0.051***
	(0.007)	(0.009)	(0.003)	(0.005)	(0.004)	(0.005)	(0.006)	(0.006)	(0.007)	(0.009)
Observations	56940	56940	56940	56940	56937	56937	56921	56921	56930	56930
Pseudo R2	0.168	0.168	0.097	0.097	0.105	0.105	0.177	0.177	0.283	0.283

c) Women age 30–49

	milk		pulses		green vegetables		fruits		eggs/meat/fish	
first-born girl	0.008*	0.005	0.003	0.002	0.004*	0.005	0.008*	0.005	0.004	0.004
	(0.005)	(0.007)	(0.002)	(0.003)	(0.002)	(0.003)	(0.004)	(0.006)	(0.005)	(0.007)
first-born girl*eduyrs>0		0.005		0.003		-0.000		0.007		-0.001
		(0.011)		(0.004)		(0.005)		(0.009)		(0.009)
eduyrs>0	0.081***	0.078***	0.017***	0.015***	0.017***	0.017***	0.096***	0.093***	0.010	0.010
	(0.006)	(0.008)	(0.003)	(0.004)	(0.003)	(0.004)	(0.006)	(0.007)	(0.006)	(0.008)
Observations	83236	83236	83231	83231	83224	83224	83187	83187	83213	83213
Pseudo R2	0.171	0.171	0.109	0.109	0.125	0.125	0.202	0.202	0.293	0.293

Probit estimates (marginal effects reported). Robust standard errors adjusted for clustering at the household level in parentheses. Food consumption variables are based on NFHS questions on frequency of consumption of certain food groups. The variables are dummies equal to one if the woman consumes the indicated type of food at least once a week. Food consumption is available for all women in NFHS 1998 and 2005 (not available in NFHS 1992). Covariates include state fixed effects, urban dummy, birth year fixed effects, 7 five-year age groups dummies, NFHS 2005 survey round fixed effect, age at first birth and marriage, her and partner's current age (and square), whether she and her partner have completed one or more years of education, a durables index, religion and caste, number of children ever born, whether currently pregnant or breastfeeding, height. Using survey weights. ***, **, and * indicate significance at 1%, 5% and 10% levels, respectively.

Table 14: Iron supplementation by sex of the first-born child

dep. var:	=1 if taken iron suppl. during:			
	first pregnancy		second pregnancy	
	(1)	(2)	(3)	(4)
first-born girl	-0.008	-0.001	0.004	0.001
	(0.006)	(0.009)	(0.007)	(0.009)
first-born girl*eduyrs>0		-0.015		0.005
		(0.013)		(0.014)
eduyrs>0	0.137***	0.144***	0.150***	0.147***
	(0.009)	(0.012)	(0.010)	(0.012)
currently pregnant	-0.000	-0.001	-0.014	-0.014
	(0.010)	(0.010)	(0.015)	(0.015)
currently breastfeeding	0.005	0.005	0.004	0.004
	(0.008)	(0.008)	(0.009)	(0.009)
# children ever born	-0.012	-0.012	-0.054***	-0.054***
	(0.010)	(0.010)	(0.010)	(0.009)
husband eduyrs>0	0.056***	0.056***	0.048***	0.048***
	(0.009)	(0.009)	(0.009)	(0.009)
durables index	0.043***	0.043***	0.043***	0.043***
	(0.003)	(0.003)	(0.003)	(0.003)
urban	0.033***	0.033***	0.015	0.015
	(0.009)	(0.009)	(0.009)	(0.009)
Observations	32050	32050	30099	30099
Pseudo R2	0.231	0.231	0.212	0.212

Probit estimates (marginal effects reported). Robust standard errors adjusted for clustering at the household level in parentheses. NFHS 1992 asked women whether they received iron and folic acid (IFA) tablets or syrup during pregnancy for each birth during the *four* years preceding the survey. NFHS 1998 asked the same question in relation to each birth during the last *three* years preceding the survey. NFHS 2005 asked the same question in relation to the most recent birth during the *five* years preceding the survey. Covariates include state fixed effects, urban dummy, birth year fixed effects, 7 five-year age groups dummies, survey round fixed effects, age at first birth and marriage, her and partner's current age (and square), whether she and her partner have completed one or more years of education, a durables index, religion and caste, number of children ever born, whether currently pregnant or breastfeeding. Using survey weights. ***, **, and * indicate significance at 1%, 5% and 10% levels, respectively.

Table 15: Spousal violence by sex of the first-born child

a) Less than a year since first birth

| | =1 if suffered physical violence | | | | | | =1 if suffered any emotional violence | | =1 if suffered any sexual violence | |
| | any | | severe | | less severe | | | | | |
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)
first-born girl	0.022	0.033	0.012	0.036**	0.024	0.037	0.012	-0.010	0.007	0.016
	(0.017)	(0.038)	(0.009)	(0.017)	(0.018)	(0.038)	(0.014)	(0.023)	(0.013)	(0.023)
first-born girl*eduyrs>0		-0.017		-0.035**		-0.020		0.038		-0.014
		(0.050)		(0.017)		(0.050)		(0.029)		(0.027)
eduyrs>0	-0.036	-0.026	-0.003	0.017	-0.031	-0.020	-0.021	-0.043*	-0.011	-0.002
	(0.033)	(0.040)	(0.011)	(0.015)	(0.031)	(0.039)	(0.019)	(0.023)	(0.016)	(0.019)
Observations	1867	1867	1492	1492	1867	1867	1843	1843	1677	1677
Pseudo R2	0.127	0.127	0.154	0.161	0.130	0.130	0.110	0.112	0.199	0.200

b) Women age 15–29

| | =1 if suffered physical violence | | | | | | =1 if suffered any emotional violence | | =1 if suffered any sexual violence | |
	any		severe		less severe					
first-born girl	-0.009	0.001	-0.008	-0.007	-0.007	0.003	0.001	0.004	0.001	0.008
	(0.010)	(0.016)	(0.016)	(0.009)	(0.010)	(0.016)	(0.005)	(0.008)	(0.006)	(0.008)
first-born girl*eduyrs>0		-0.018		-0.002		-0.020		-0.007		-0.014*
		(0.019)		(0.011)		(0.019)		(0.012)		(0.008)
eduyrs>0	-0.039***	-0.031**	-0.017**	-0.016**	-0.038***	-0.029*	-0.013	-0.009	-0.005	0.002
	(0.010)	(0.015)	(0.007)	(0.008)	(0.010)	(0.015)	(0.008)	(0.009)	(0.006)	(0.007)
Observations	21554	21554	21554	21554	21552	21552	21561	21561	21561	21561
Pseudo R2	0.091	0.091	0.067	0.067	0.091	0.092	0.042	0.042	0.099	0.099

c) Women age 30–49

| | =1 if suffered any physical violence | | | | | | =1 if suffered any emotional violence | | =1 if suffered any sexual violence | |
	any		severe		less severe					
first-born girl	-0.013*	-0.013	0.005	0.004	-0.013*	-0.012	0.005	0.007	0.003	0.005
	(0.014)	(0.010)	(0.004)	(0.005)	(0.007)	(0.009)	(0.005)	(0.007)	(0.003)	(0.005)
first-born girl*eduyrs>0		0.001		0.002		-0.002		-0.003		-0.003
		(0.013)		(0.008)		(0.013)		(0.009)		(0.007)
eduyrs>0	-0.053***	-0.053***	-0.024***	-0.025***	-0.052***	-0.051***	-0.013**	-0.011*	-0.008*	-0.007
	(0.009)	(0.009)	(0.007)	(0.008)	(0.009)	(0.009)	(0.006)	(0.006)	(0.005)	(0.006)
Observations	32633	32633	32633	32633	32633	32633	32634	32634	32634	32634
Pseudo R2	0.087	0.087	0.069	0.069	0.087	0.087	0.043	0.043	0.099	0.099

Probit estimates (marginal effects reported). Robust standard errors adjusted for clustering at the household level in parentheses. The spousal violence variables are based on a module of NFHS 2005 that collected detailed information on physical, sexual, and emotional violence perpetrated by husbands against their wives (based on women's self-report). "any severe violence" includes whether the spouse ever tried to strangle or burn, or ever threatened her with knife or gun. "any less severe violence" includes whether the spouse ever pushed, slapped, punched, kicked or dragged her. "any emotional violence" includes whether the husband ever said something to humiliate her in front of others, or threatened to hurt someone close to her, or insulted her. "any sexual violence" includes whether the husband forced her to perform any unwanted sexual acts. Domestic violence variables available for all married women in NFHS 2005. Covariates include state fixed effects, urban dummy, birth year fixed effects, 7 five-year age groups dummies, NFHS 2005 survey round fixed effect, age at first birth and marriage, her and partner's current age (and square), whether she and her partner have completed one or more years of education, a durables index, religion and caste, number of children ever born, whether currently pregnant or breastfeeding, height. Using survey weights. ***, **, and * indicate significance at 1%, 5% and 10% levels, respectively.

Table 16: Sex ratios for first births by age group

| | all women | | | women with no education | | |
age group	sex ratio	CI	n	sex ratio	CI	n
20-24	0.515	[0.510; 0.520]	38544	0.513	[0.505; 0.520]	18050
25-29	0.512	[0.508; 0.517]	50199	0.510	[0.504; 0.517]	22835
30-34	0.520	[0.515; 0.524]	45878	0.523	[0.517; 0.530]	21637
35-39	0.519	[0.514; 0.524]	40137	0.526	[0.519; 0.533]	19837
40-44	0.524	[0.518; 0.530]	31674	0.534	[0.527; 0.542]	16419
45-49	0.536	[0.529; 0.542]	24303	0.549	[0.540; 0.557]	13461

95% confidence intervals reported in square brackets. Using survey weights.
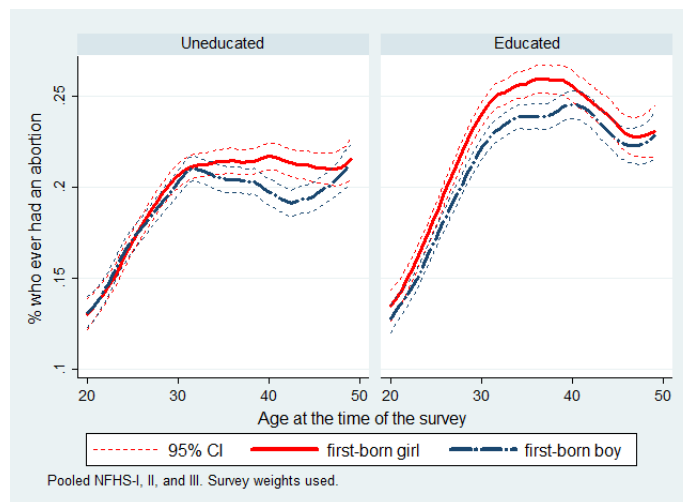
A   Appendix

Figure A.1: Share of women with first-born girl by age and urban/rural location
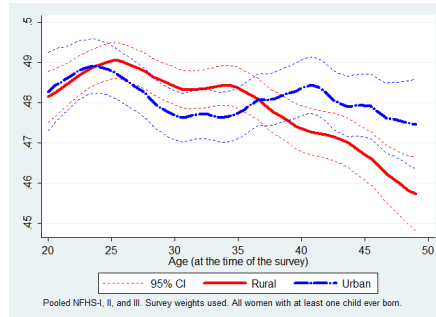
Figure A.2: Share of anemic women, by age (Subsamples)

Figure A.3: % of women who ever had a terminated pregnancy, by age and sex of the first-born child

Table A.1: Fertility outcomes by the sex of the earlier-born children

dep. var.	=1 if wants more children		=1 if sterilized		=1 if undergone ultrasound		=1 if ever had terminated pregnancy		number of children ever born	
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)
G,G	0.100***		-0.158***		0.049***		0.029***		0.639***	
	(0.003)		(0.004)		(0.008)		(0.004)		(0.011)	
G,B	0.009***		-0.056***		0.022***		0.013***		0.147***	
	(0.002)		(0.004)		(0.007)		(0.003)		(0.010)	
B,G	0.015***		-0.054***		0.005		0.015***		0.155***	
	(0.002)		(0.004)		(0.008)		(0.003)		(0.010)	
G,G,G		0.107***		-0.189***		0.067***		0.047***		0.756***
		(0.005)		(0.007)		(0.013)		(0.006)		(0.018)
G,G,B		0.016***		-0.077***		0.046***		0.017***		0.207***
		(0.002)		(0.007)		(0.012)		(0.005)		(0.016)
G,B,B		-0.011***		0.019***		0.026**		0.001		-0.076***
		(0.002)		(0.006)		(0.012)		(0.005)		(0.017)
G,B,G		0.021***		-0.052***		0.023**		0.024***		0.236***
		(0.003)		(0.007)		(0.010)		(0.005)		(0.017)
B,B,G		-0.003		0.002		0.023**		0.003		-0.046***
		(0.002)		(0.007)		(0.011)		(0.006)		(0.016)
B,G,B		-0.008***		0.003		0.010		0.009		-0.067***
		(0.002)		(0.007)		(0.012)		(0.006)		(0.016)
B,G,G		0.028***		-0.070***		0.007		0.026***		0.264***
		(0.003)		(0.007)		(0.011)		(0.006)		(0.017)
Observations	182582	126426	182752	126550	40236	23171	182724	126523	142863	103120
Pseudo R2	0.325	0.296	0.184	0.180	0.291	0.250	0.0262	0.0255	0.455	0.346

Probit estimates in columns 1–8 (marginal effects reported). OLS estimates in columns 9–10. Robust standard errors adjusted for clustering at the household level in parentheses. "wants more children" is a dummy equal to one if the woman wants more children at the time of the survey, zero otherwise. "sterilized" is a dummy equal to one if the woman has undergone female sterilization, zero otherwise (or if she is using another method of contraception). "ultrasound" is a dummy equal to one if the woman has undergone ultrasound test (zero otherwise), asked only with reference to the two most recent births in the last three years preceding the survey (NFHS 1998), and to any pregnancy in the last five years preceding the survey (NFHS 2005) (not asked in NFHS 1992). "had terminated pregnancy" is a dummy equal to one if the woman ever had a pregnancy that terminated in a miscarriage, abortion, or still birth, zero otherwise. The regression for the number of children ever born is estimated on the subsample of women who have been sterilized or who do not want more children. Covariates include state fixed effects, urban dummy, birth year fixed effects, 7 five-year age groups dummies, survey round fixed effects, age at first birth and marriage, her and partner's current age (and square), whether she and her partner have completed one or more years of education, a durables index, religion and caste, and number of children ever born. Height measurement (in centimeters) is available only in NFHS 1998 and 2005. The number of children ever born is included in columns 1–8. Using survey weights. ***, **, and * indicate significance at 1%, 5% and 10% levels, respectively.

Table A.2: Fertility outcomes by the sex of the first-born, robustness check (inclusion of height)

dep. var.	=1 if wants more children		=1 if sterilized		=1 if undergone ultrasound		=1 if ever had terminated pregnancy		number of children ever born	
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)
first-born girl	0.051***	0.051***	-0.072***	-0.072***	0.025***	0.025***	0.004	0.004	0.344***	0.344***
	(0.003)	(0.003)	(0.005)	(0.005)	(0.009)	(0.009)	(0.004)	(0.004)	(0.014)	(0.014)
first-born girl*eduyrs>0	-0.009**	-0.009**	-0.009	-0.009	-0.001	-0.001	0.020***	0.020***	-0.063***	-0.063***
	(0.004)	(0.004)	(0.007)	(0.007)	(0.011)	(0.011)	(0.005)	(0.005)	(0.017)	(0.017)
eduyrs>0	-0.037***	-0.037***	0.037***	0.037***	0.074***	0.074***	0.011***	0.011***	-0.307***	-0.305***
	(0.003)	(0.003)	(0.005)	(0.005)	(0.009)	(0.009)	(0.004)	(0.004)	(0.015)	(0.015)
height		-0.000**		0.000		0.000		-0.001***		-0.005***
		(0.000)		(0.000)		(0.000)		(0.000)		(0.001)
Observations	140062	140062	140196	140196	54100	54100	140189	140189	104999	104999
Pseudo R2 or R2	0.398	0.398	0.234	0.234	0.310	0.310	0.0302	0.0305	0.468	0.469

Probit estimates in columns 1–8 (marginal effects reported). OLS estimates in columns 9–10. Robust standard errors adjusted for clustering at the household level in parentheses. "wants more children" is a dummy equal to one if the woman wants more children at the time of the survey, zero otherwise. "sterilized" is a dummy equal to one if the woman has undergone female sterilization, zero otherwise (or if she is using another method of contraception). "ultrasound" is a dummy equal to one if the woman has undergone ultrasound test (zero otherwise), asked only with reference to the two most recent births in the last three years preceding the survey (NFHS 1998), and to any pregnancy in the last five years preceding the survey (NFHS 2005) (not asked in NFHS 1992). "had terminated pregnancy" is a dummy equal to one if the woman ever had a pregnancy that terminated in a miscarriage, abortion, or still birth, zero otherwise. The regression for the number of children ever born is estimated on the subsample of women who have been sterilized or who do not want more children. Covariates include state fixed effects, urban dummy, birth year fixed effects, 7 five-year age groups dummies, survey round fixed effects, age at first birth and marriage, her and partner's current age (and square), whether she and her partner have completed one or more years of education, a durables index, religion and caste, and number of children ever born. Height measurement (in centimeters) is available only in NFHS 1998 and 2005. The number of children ever born is included in columns 1–8. Using survey weights. ***, **, and * indicate significance at 1%, 5% and 10% levels, respectively.