Son Preference, Fertility and Family Structure

Evidence from Reproductive Behavior among Nigerian Women

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Abstract

Strong boy-bias and its consequences for young and unborn girls have been widely documented for Asia. This paper considers a country in Sub-Saharan Africa and finds that parental gender preferences do affect fertility behavior and shape traditional social institutions with negative effects on adult women's health and well-being. Using individual-level data for Nigeria, the paper shows that, compared to women with first-born sons, women with first-born daughters have (and desire) more children and are less likely to use contraceptives. Women with

daughters among earlier-born children are also more likely to have shorter birth intervals, a behavior medically known to increase the risk of child and maternal mortality. Moreover, they are more likely to end up in a polygynous union, to be divorced, and to be head of the household. The preference for sons is also supported by child fostering patterns in which daughters are substitutes for foster girls, while the same does not hold for sons and foster boys. These results can partly explain excess female mortality among adult women in Sub-Saharan Africa.

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Son Preference, Fertility and Family Structure. Evidence from Reproductive Behavior among Nigerian Women*

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1 Introduction

There is an extensive literature on parental preferences over a child's gender.¹ The preference for male children, or son preference, has been widely documented in South and East Asia where gender bias is severe especially before birth and at young ages. Sen (1990) documents abnormally high sex ratios at birth (the share of male births) in large parts of Asia: this evidence is suggestive of sex-selective abortion, infanticide and neglect of female children. Sub-Saharan Africa is typically regarded as a region with low or absent gender preferences.² Although it is well recognized that sons are strongly valued in traditional patrilineal societies in Sub-Saharan Africa, no evidence has been found of sex ratios favoring boys at birth (possibly due to lack of accessible technologies for fetal sex determination). Recent studies find that many women are indeed missing in Sub-Saharan Africa, especially during their reproductive years (Anderson and Ray, 2010; World Bank, 2011).³ This paper revisits the issue of son preference in Sub-Saharan Africa by examining patterns of fertility behavior and family structure in Nigeria. It shows that son preference is prevalent and that its deleterious effects are reflected not in missing young or unborn girl children but rather in the health and well-being of adult women. Increasing sex ratios at birth may be envisaged following a future diffusion of techniques for sex selection as parents seek ways to meet their demand for sons.

The reasons given for son preference are various and include cultural, kinship, economic and institutional factors that typically differ across different societies (Das Gupta et al., 2003; Almond, Edlund, and Milligan, 2009). In the Sub-Saharan African context, anthropological and demographic evidence emphasizes the dominant role of males in patrilineal societies in which descent and inheritance are transmitted through the male line (Goody, 1973, 1976; Isiugo-Abanihe, 1994). Furthermore, male children strengthen the relationship between the wife and her husband's kin (i.e., by guaranteeing the continuation of his lineage) and secure (the mother's) access to residence and inheritance upon the husband's death.

This paper makes two contributions to the existing literature. Anderson and Ray (2010) use

¹For a cross-cultural review of literature on parental preferences see, among others, Williamson (1976).

²Two notable exceptions are Garg and Morduch (1998) who show that children with relatively more sisters than brothers benefit in the intra-household allocation of resources in resource-constrained households in Ghana; and Klasen (1996) who uses population and mortality data for Sub-Saharan Africa to construct an adjusted aggregate reference ratio and finds a small and increasing anti-female bias.

³Anderson and Ray (2010) and World Bank (2011) find that the sources of excess female mortality among adult women in Sub-Saharan Africa are primarily HIV/AIDS and maternal mortality. Klasen and Vollmer (2013) use different reference standards than Anderson and Ray (2010) and confirm their findings for excess mortality among young female adults in Africa (though of a smaller magnitude).

aggregate mortality data and find evidence of excess mortality among adult women in Sub-Saharan Africa. In contrast, this paper uses individual-level data and analyzes how son preference affects women’s fertility preferences and choices in a context where sex ratios at birth are in the normal range and where little is known about parental gender preferences. In addition, it explores how parental gender preferences can affect the development of traditional social institutions common in the West African context, namely child fostering and polygyny, as well as other marital outcomes and arrangements, including divorce and female headship. These different sets of outcomes are analyzed separately and, taken together, can shed some light on the complex mechanisms through which the desire for own biological sons can be satisfied in Nigeria.

More specifically, the paper puts forward two channels through which adult women might be missing in Sub-Saharan Africa and in other contexts where fertility choices and family structure are found to be strongly affected by son preference. The first is son-preferring fertility behavior (or son-preferring stopping rules) in which women with girls among earlier-born children are more likely to continue having children and to have shorter birth intervals. This behavior is medically known to adversely affect women’s health (as well as children’s health) and increase the risk of maternal mortality and morbidity due to maternal depletion. Maternal mortality is in fact among the highest in the world in Nigeria (second only to India in absolute number of deaths), which alone accounts for 14 percent of all maternal deaths worldwide (WHO, 2010). Women of lower socioeconomic status are more likely to be affected due to their poorer access to maternal care and more vulnerable health. The second channel refers to the association between parental gender preferences and specific marital outcomes and living arrangements in which women are often found to be particularly vulnerable and disadvantaged (i.e., polygyny and female headship).

I use the 2008 Nigeria Demographic and Health Survey (NDHS) and find evidence of son preference in several dimensions. A comparison of women’s characteristics by the sex of the first-born reveals that while younger women (younger than 30) are very similar in terms of observables, significant differences appear in the subsample of older women, which is suggestive of selective mortality.⁴ I check for the exogeneity of the sex of the first-born child and discuss other potential sources of bias in the data. The empirical strategy relies on the identifying assumption that the sex of the first-born is uncorrelated with the error term after conditioning for a set of observable characteristics. Results are as follows. First, compared to women who had a first-born boy, women with a first-born girl exhibit a 2 percent increase in the number of children ever born. While the

⁴Selective mortality is also supported by nonparametric evidence of a lower fraction of women with a first-born girl among women in the older age group.

effect on fertility is stronger when considering the first two female births, I focus on the sex of the first-born to maintain a causal interpretation of the results. Second, I find that women’s desired fertility and use of contraceptives are affected by the sex of the first-born. In particular, women with a first-born girl are 2.3 percentage points more likely to report that they desire another child, and 1.1 percentage points less likely to use contraceptives. Third, I investigate whether the pace at which women have children is influenced by the sex of earlier-born children. I exploit the variation in birth spacing for children of the same mother by estimating regressions with mother fixed effects. I find that, among women who had daughters as the first two born children, those who have a third or fourth daughter are 3.6 (1.8) percentage points (significantly) more likely to wait less than 24 (15) months.

Second, the preference for male children is found to be significantly associated with the social institution of polygyny (the practice in which a man has several wives). Women with first-born daughters are in fact 1.2 percentage points more likely to end up in a polygynous union (i.e., the husband takes another wife). As expected, this effect is specific to first-rank wives (while no effect is found for higher-rank wives), as they are the ones who entered the current union as monogamous. Among women aged 30–49, those with first-born daughters are also 1.4 percentage points more likely to have a nonresident husband, 1.2 percentage points more likely to be ever divorced or separated, and 1.1 percentage points more likely to be the head of the household.

Further evidence is drawn from an examination of the institution of child fostering, in which biological children are temporarily sent to live with other families. I focus on the child-labor hypothesis as one of the motivations for fostering according to which children of each gender and age group have specific roles within the household. Households with an imbalance of biological girls or boys might decide to send or receive a child in order to achieve a balanced gender structure, thus maximizing household productivity (Akresh, 2009). If non-biological girls (boys) are substitutes for biological daughters (sons), households should respond symmetrically to the imbalance of daughters or sons in fostering decisions. Instead, I find that households with an excess of sons are one percentage point more likely to foster-in a girl, while those with an excess of daughters are not more likely to foster-in boys. This asymmetric response may suggest that, motivated by the desire to have their own biological son, households with an excess of daughters might continue bearing children instead of fostering-in outside boys. In contrast, households with an excess of sons are significantly more likely to receive a girl, as girls are often needed for domestic chores. This evidence is consistent with the idea that foster girls are considered substitutes for own biological daughters in fostering-in decisions, while foster boys are not substitutes for sons. Interestingly, results do not

differ between urban and rural households (or landed and landless households), suggesting that gender preferences may play a role in shaping fostering decisions beyond considerations related to the production system.

This paper is related to several strands of the literature. First, it relates to the literature documenting son preference in different areas of the world. Evidence of son-preferring fertility behavior is found especially for women in Asian countries (Chowdhury and Bairagi, 1990; Clark, 2000; Dreze and Murthi, 2001; Filmer, Friedman, and Schady, 2009). For the United States, in addition to the well-documented preference for sex-balance (Ben-Porath and Welch, 1976; Angrist and Evans, 1998), recent research has documented unusually high sex ratios among Asian mothers: this pattern is found to be related to the increasing availability of technologies to determine the sex of the fetus (Dahl and Moretti, 2008; Abrevaya, 2009). Similar evidence has been found among Asian immigrants in Canada (Almond, Edlund, and Milligan, 2009). In this paper, I quantitatively investigate if fertility decisions are gender-biased in a context where the sex ratio at birth is in the normal range.

Second, it speaks to the literature examining the effects of son preference in generating gender disparities in health and educational outcomes. In the context of India, the diffusion of prenatal sex determination technologies has not only been found to be associated with increasing sex ratios, but also with preferential *prenatal* treatment for boys (Bharadwaj and Lakdawala, 2013). The ‘try until you have a son’ fertility rule may also ‘passively’ lead to a lower amount of resources being allocated to female children within the household. This may happen either as a consequence of shorter breastfeeding for girls because parents with daughters want to ‘try again’ for a son (Jayachandran and Kuziemko, 2011), or because girls tend to be born in larger households (Jensen, 2003). Recently, Barcellos, Carvalho, and Lleras-Muney (2014) find that, accounting for the effect of son-preferring stopping rules on family size, boys still receive more parental investment (including childcare time, breastfeeding, vaccinations, and vitamin supplementations) than girls in India. While this literature has mainly focused on young (or not yet born) children, this work represents one of the first attempts to understand the effects of the preference for male children on adult women’s health and well-being.

Third, it relates to the economic and sociological literature on the effects of a child’s gender on marital stability, divorce, and parental involvement (see Lundberg, 2005, for a review). Dahl and Moretti (2008) analyze Census and other data sources for the United States and find that women with a first-born daughter are less likely to marry, more likely to be divorced, and that –after

divorce— fathers are more likely to obtain custody of sons compared to daughters. In a previous version, Dahl and Moretti (2004) also use Census data for Kenya and find an association between having a girl and the probability of being in a polygynous union. This paper exploits the entire women’s fertility histories (available in the NDHS but not in the Census, along with other variables such as a wife’s rank) and studies the effect of son preference on the probability that the husband becomes polygynous, that he does not co-reside with his wife, that the woman is ever divorced or separated, and that she is the head of the household.

Lastly, this paper contributes to the literature on child fostering (Isiugo-Abanbie, 1985; Bledsoe, 1990; Ainsworth, 1996; Akresh, 2009) by showing evidence that fostering patterns respond asymmetrically to a gender imbalance of daughters or sons, thereby suggesting a new motivation for household fostering decisions that is complementary to the ones already proposed in the literature.

The remainder of the paper is organized as follows. Section 2 provides some insights on the context and cultural background. Section 3 describes the data and some suggestive patterns in the variables of interest. Section 4 introduces the empirical methodology used to analyze fertility outcomes and birth spacing and further explores the identified patterns. Section 5 presents the results. Section 6 focuses on the interactions between son preference, marital outcomes, and child fostering decisions. Section 7 concludes.

2 Context

As in many parts of Sub-Saharan Africa, society is organized around the extended family which still represents the most basic unit of social organization in Nigeria. Family ties are strong and play an important role in shaping individual behavior, even though there are signs that the extended family system is weakening for some ethnic groups (Wusu and Isiugo-Abanihe, 2006). There is an extraordinary ethnic diversity in Nigeria, but all ethnic groups are predominantly characterized by patrilineality and patrilocality.⁵ Large progenies are strongly valued because they strengthen a man’s family status (Caldwell and Caldwell, 1987). The interactions between the principles of social organization, fertility preferences and family structure are documented in several demographic studies, such as Isiugo-Abanihe (1994):

‘Childlessness is the most dreaded tragedy for a man or a woman to experience in Nigeria’s

⁵The three largest ethnic groups are geographically concentrated in different areas of Nigeria. They are the Hausa-Fulani (North), Yoruba (South-West), and Igbo (South-East). They represent 28%, 18%, 16% of the population, respectively (NPC and ICF Macro, 2009).

patrilineal society. (...) The majority of the respondents felt that a man without a child, particularly without a son, will not be remembered in the family; his branch of the family will come to an end. For the same reason, a man who has only daughters may acquire a second wife to enhance the chance of having a son. Clearly, in such a patriarchal system, the perpetuation of the family line is a strong motivation for children.' Isiugo-Abanihe (1994, p. 154, from a survey of 3,073 Nigerian couples)

Indeed, the desire for many children, especially males, is motivated by descent transmission and the urge to provide continuation to the family lineage. More importantly, these societies are characterized by the fact that only males have the control of family landed property, as documented in Wusu and Isiugo-Abanihe (2006).⁶

'In most cultures only male children are allowed to share in family land holdings within the context of the extended family structure and communal ownership of land. Since farming is central to economic life, the most economically rewarding reproductive goal a couple could pursue is a large family size, ideally with many male children. Against this backdrop couples dread barrenness, and until a 'good' number of male children are born, extended family members exert pressure, which may culminate in the man marrying another wife' Wusu and Isiugo-Abanihe (2006, pp. 141–142)

In this context, sons are highly valued by women as well because they represent the only way through which they can inherit part of their deceased husband's property (Fapohunda and Todaro, 1988; Ewelukwa, 2002). This consequently generates pressure for bearing male children as a way to protect themselves in the state of widowhood. It may also lead to competition among co-wives as described in Bledsoe, Banja, and Hill (1998).

'In their husbands' compounds, women seek to establish their security and to gain a competitive edge over present and future co-wives and sisters-in-law by bearing a number of children, especially sons, who will retain rights of residence and inheritance in the compound and will eventually take over its leadership roles.' Bledsoe, Banja, and Hill (1998, p. 23)

⁶There are several sources of law in Nigeria: Common law (the *default* Law, applied only when other sources of law are not applicable), Customary Law, and Islamic law (Oba, 2011). Even though there is some variation in inheritance rules across states and ethnic groups, in most cases only male children are allowed to inherit the father's property or larger part of it. According to Islamic law, female children are entitled to inherit half of the share of their male siblings.

A peculiarity of the West-African context is the widespread fostering of children who are temporarily sent away from their biological parents to be raised in another family.⁷ The reasons for child fostering have been studied by several researchers (Isiugo-Abanbie, 1985; Bledsoe, 1990; Ainsworth, 1996; Akresh, 2009; Beck et al., 2011). These include insurance against idiosyncratic shocks to family income, education (i.e., children in rural households are sent to live with their urban kin members where there are schools), kinship fostering (i.e., as part of the obligations between members of the same extended family), and child labor. As argued by some demographers, high fertility is also made possible by the social institution of fostering, through which the costs of childrearing are partly shared with the extended family (Isiugo-Abanbie, 1985; Bledsoe, 1990). Fapohunda and Todaro (1988) argue that ‘*the presence of parental surrogates in the extended family alleviates problems of incompatibility between child care and work and, thereby, lowers the opportunity cost of children.*’ (p. 572). According to Smith (2004), the traditional high value associated with having many children (that allow larger kinship networks) might partly motivate the slow transition to lower fertility in Nigeria, even though families start lamenting the increasing costs of raising children due to modernization.

3 Data and descriptive evidence

I use the 2008 Nigeria Demographic and Health Survey (NDHS), a representative cross-sectional dataset of women aged 15 to 49 in Nigeria. It contains individual-level information on women’s birth histories, household composition, biological children living in the household or elsewhere, birth intervals, marital status, and other individual and household characteristics.

Summary statistics for the sample of women aged 15 to 49 with at least one child ever born are reported in table 1.

The mean number of children ever born is 4.35.⁸ This does not reflect completed fertility especially for younger women in the sample: the total fertility rate is in fact 5.7 births per woman (NPC and ICF Macro, 2009). As noted in the previous section, high fertility is strongly desired. This is reflected both in the high number of births and desired fertility: 63 percent of women express

⁷Child fostering is a widely accepted and practiced social institution in many parts of Sub-Saharan Africa. Lloyd and Desai (1992) analyze data from 16 DHS survey around the world and find that the percentages of children aged 10–14 living away from their mothers ranges from 0.9 in Tunisia, 4.2 in Sri Lanka, 5.7 in Brazil, 24 in Senegal, 29.4 in Ghana, and 40.9 in Liberia. Isiugo-Abanbie (1985) argues that ‘nowhere is it as institutionalized as in parts of West Africa’ (pp. 56).

⁸The number of children ever born includes children who have died.

Table 1: Summary statistics

	mean	median	st. dev.	min	max	N
# children ever born	4.35	4	2.76	1	18	23751
# daughters ever born	2.12	2	1.72	0	12	23751
# sons ever born	2.23	2	1.76	0	12	23751
# children alive	3.53	3	2.15	0	15	23751
# daughters alive	1.74	1	1.42	0	10	23751
# sons alive	1.79	2	1.45	0	11	23751
first child dead	0.18	0	0.39	0	1	23751
first daughter dead	0.08	0	0.27	0	1	23751
first son dead	0.10	0	0.31	0	1	23751
first-born girl	0.48	0	0.50	0	1	23751
first-born boy	0.52	1	0.50	0	1	23751
currently married	0.90	1	0.30	0	1	23750
currently divorced or sep	0.03	0	0.16	0	1	23750
ever divorced or sep.	0.12	0	0.32	0	1	23751
currently widowed	0.03	0	0.18	0	1	23750
ever widowed	0.05	0	0.22	0	1	23751
living with husband	0.90	1	0.30	0	1	21652
first union	0.86	1	0.34	0	1	23050
polygynous husband	0.34	0	0.47	0	1	21690
live in female headed hh	0.14	0	0.34	0	1	23751
female head	0.10	0	0.31	0	1	23751
age	32.13	31	8.61	15	49	23751
age at first marriage	17.69	17	4.63	7	45	23153
age at first birth	19.36	19	4.44	9	44	23751
husband's age	42.49	40	11.29	15	96	21457
contraceptive use	0.16	0	0.36	0	1	23751
wants another child	0.63	1	0.48	0	1	23598
wants no more children	0.22	0	0.42	0	1	23598
urban	0.32	0	0.47	0	1	23751
woman's eduyrs	4.91	4	5.26	0	22	23735
husband's eduyrs	6.16	6	5.79	0	21	22657
catholic	0.10	0	0.30	0	1	23575
other christian	0.37	0	0.48	0	1	23575
muslim	0.51	1	0.50	0	1	23575
hausa-fulani	0.34	0	0.47	0	1	23617
igbo	0.12	0	0.33	0	1	23617
yoruba	0.17	0	0.37	0	1	23617
weight-for-height	-0.55	-0.68	1.21	-4	5.92	22986
body mass index	23.07	22.14	4.66	12.1	59.81	23115

Notes. 2008 NDHS. Using sample weights.

the desire for another child. Although knowledge of any contraceptive method is widespread, current use among currently married women is much lower but increasing over time: from 6% in 1990, 13% in 2003, to 15% in 2008.⁹ Despite relatively low levels of contraceptive use, most pregnancies are ‘wanted’ by the vast majority of women: 88 percent of women with births in the five years

⁹68% of all married women and 90% of all married men know of at least one method of contraception, and 29% (45%) of all married women (men) reported ever using a method of contraception. Among women currently using contraception, about two-thirds are using a modern method (mostly male condom, injectables and pills), while one-third use traditional methods (rhythm method and withdrawal are the most common).

preceding the survey states that the births were wanted then, 7 percent were wanted later, and only 5 percent were unwanted. Marriage is almost universal for women in Nigeria: 90 percent are currently married, 3 percent are divorced or separated, 3 percent are widowed; and remarriage is not uncommon. About 34 percent of women have a polygynous husband, 10 percent are the head of the household, and 90 percent of women live with their husband.

Birth spacing Medical research has shown that short birth intervals are associated with poor child and maternal health outcomes (Setty-Venugopal and Upadhyay, 2002).¹⁰ With respect to maternal health, relative to women who give birth at 9- to 14-month intervals, those who have their children at 27- to 32-month birth intervals are: 1.3 times more likely to avoid anemia; 1.7 times more likely to avoid third-trimester bleeding; and 2.5 times more likely to survive childbirth (Conde-Agudelo, and Belizan, 2000).¹¹ Several studies also find that short birth intervals and reduced breastfeeding time have deleterious effects on child health and mortality (Palloni and Millman, 1986; Palloni, Pinto Aguirre, and Lastiri, 1994; and Jayachandran, and Kuziemko, 2011). Palloni and Millman (1986) observe that close birth spacing does not allow a woman to regain her physical strength and the nutrients required to have a successful new pregnancy. The effects go through retardation of fetal growth as well as the mother's inability to properly breastfeed the child, all leading to increased risk of death of the child. Moreover, competition for resources among siblings born in close succession may aggravate these negative effects.¹²

In the 2008 NDHS, of all births reported by each woman, 8 percent are less than 15 months apart, and 33 percent less than two years. The mean birth interval is 32.6 months, and the median is 27. The survival status of the child who opens the birth interval matters: the median interval is 29 months if the previous child is alive, 24 months if dead.

Given the association between birth spacing and child and maternal health outcomes, it is important to understand whether the desire for sons affects the pace at which mothers try to conceive and give birth.

¹⁰Birth spacing refers to the time interval (number of months) between births.

¹¹See Conde-Agudelo et al. (2007) for a review of medical studies on the relationship between birth spacing and maternal health.

¹²There is a striking negative relationship between the length of the birth interval and child mortality. For example, 252 children (out of 1000 live births) die under the age of 5 if the preceding birth interval is shorter than 24 months, while the number of children who die if the interval is 4 years or longer is reduced to 92 (NPC and ICF Macro, 2009).

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

Sex ratio at birth Using the 2008 NDHS birth history data, the sex ratio at birth (the fraction of male births) is 0.51 for *all* reported births.^{13,14} The left panel of figure 1 plots the sex ratio for all births reported in the women’s fertility histories, by the year of each birth using Kernel-weighted local polynomial smoothing. It shows a clear negative relationship between the sex ratio at birth and the birth year for births before 1995, and a rather stable (and biologically ‘normal’) sex ratio for more recent births. The availability of previous rounds of the NDHS allows me to plot the sex ratio for each survey (conducted in 1990, 1999, 2003, and 2008). The right panel shows that the decreasing pattern is visible in each survey round, thus it is not specific to the most recent one.

A possible explanation for this pattern is selective maternal mortality due to son preference (Milazzo, 2014).¹⁵ Specifically, women with first-born daughters tend to have repeated and shortly spaced pregnancies, a behavior medically known to increase their risk of mortality and morbidity. I use the pooled NDHSs and plot the share of women with a first-born girl by their age at the time of the survey. Figure 2 shows that the share of women with a first-born girl is lower among older women, in particular those older than 30. Figure 3 shows that the declining trend is specific to first births, but is not present for births of higher order. This is suggestive of an association between a first-born girl and subsequent realized fertility and spacing. Women of lower socioeconomic status

¹³Based on several data sources, Anderson and Ray (2010) report that the sex ratio at birth is about 0.514 in developed countries, 0.518 in India, 0.539 in China, and 0.508 in Sub-Saharan Africa. Several biological, environmental, and genetic factors can partly explain the variation across different areas.

¹⁴Garenne (2002) analyzes sex ratios at birth in African countries using the NDHS surveys. He finds that there is substantial variation within the African region (with Nigeria and Ethiopia being the countries with the highest sex ratio, equal to 0.517) and that the average sex ratio is 0.508. As for Nigeria, Garenne’s study does not include the most recent DHS surveys, in which the sex ratio is 0.512 (2008) and 0.515 (2003).

¹⁵In Milazzo (2014), I use the India National Family Health Surveys (NFHSs) and find a similar decreasing pattern for the sex ratio in India. I compare women’s age structure and health indicators (especially anemia) by the sex of their first born and show that son preference is associated with selective (i.e., higher among poorer women with a first-born girl) maternal mortality through son-preferring fertility behavior.

Figure 2: Share of women with a first-born girl, by woman's age at the time of the survey

Figure 3: Share of women with a first or higher-order girl, by woman's age at the time of the survey

are the most likely to suffer these adverse consequences on their health and mortality, probably due to worse access to maternal care and weaker health status. To see this, I split the sample of women with first-born girls by their education (those who have completed zero or at least one year of education, respectively). Figure 4 reveals a striking pattern: the share of women with first-born girls is declining with age only among uneducated women, consistent with higher risk of death among them.¹⁶ Selective mortality leads to an underestimation of the effect of daughters on fertility because (surviving) women reporting a first-born girl have fewer children on average.

Therefore, son preference might partly contribute to the high number of deaths due to maternal reasons in Nigeria. In Nigeria, a woman faces a one in 23 probability of dying due to complications related to pregnancy or childbirth during her reproductive lifespan (WHO, 2010).¹⁷ Wall (1998)

¹⁶In the absence of info on anemia in NDHS 2008, I compare weight-for-height and body mass index and find that women above the age of 30 with a first-born girl exhibit better nutritional conditions than those with a first-born boy (see summary statistics and t-tests in table 2). Similar results are found for India (Milazzo, 2014).

¹⁷The adult lifetime risk of maternal death is defined as the probability of dying from a maternal cause during a woman's reproductive lifespan. For comparison, the WHO estimates for broad world regions are: 1 in 32 in Africa; 1

Figure 4: Share of women with a first-born girl, by woman’s education and age at the time of the survey

observes that a series of factors contribute to high maternal mortality and morbidity in Nigeria, including low female education, marriage and first pregnancy often occurring at an early age, inadequate health facilities, and cultural and religious practices that restrict women’s access to adequate care.

Potential alternative explanations for the patterns found in figure 2, including factors related to biological conditions and misreporting, are discussed in appendix section A.1.

Women’s characteristics by the sex of their first-born child. Table 2 compares selected characteristics of women by the sex of their first-born child. Panel (a) shows that women with first-born girls are significantly more likely to be the head of the household, and less likely to live with their husbands. In terms of predetermined characteristics (i.e., that should not be affected by later fertility behavior), women with a first-born daughter have the first child later in life, are more educated, and have better educated husbands. These differences are statistically significant. The number of children ever born does not differ between the two groups: this can be partly explained by selective maternal mortality and differential infant mortality, i.e., male children are biologically weaker than females and more likely to die at birth and in infancy (Waldron, 1983).¹⁸ As mothers tend to replace a dead child by having more children, those with a first-born boy will thus have more children than they would have had if there were no biological gender differences in infant mortality.

in 150 in South-East Asia (incl. India); 1 in 84 in the Eastern Mediterranean Region; 1 in 2900 in Europe; 1 in 670 in the Americas; and 1 in 1100 in the Western Pacific Region (incl. China)

¹⁸Summary statistics in table 1 report that the fraction of women whose first-born child is dead is 18 percent: 10 percent with a first-born son, and 8 percent with a first-born daughter. When considering only children who died before reaching the age of two, these fractions are 8 and 6 percent for women with a first-born son and daughter, respectively. Higher mortality rates for boys in Nigeria are confirmed by WHO estimates: infant mortality rates (the probability of dying before reaching age 1) for *all births* are 92 and 80 deaths per 1,000 live births for male and female children, respectively, in 2009 (WHO, 2011).

Table 2: Selected women’s characteristics, by the sex of the first-born and age group

a) All women					
	first-born girl	first-born boy	diff	st. error diff	n
# children ever born	4.344	4.369	-0.025	(0.038)	23751
wants another child	0.637	0.629	0.008	(0.007)	23598
use contraceptives	0.157	0.157	-0.000	(0.005)	23751
age at first birth	19.423	19.295	0.128**	(0.064)	23751
polygynous husband	0.333	0.336	-0.003	(0.007)	21690
female headed hh	0.144	0.133	0.011**	(0.005)	23751
living with husband	0.891	0.903	-0.012**	(0.005)	21652
first union	0.862	0.863	-0.001	(0.005)	23050
woman’s eduyrs	5.026	4.811	0.215***	(0.076)	23735
husband’s eduyrs	6.329	5.999	0.330***	(0.083)	22657
weight-for-height	-0.536	-0.561	0.025	(0.018)	22986
body mass index	23.116	23.031	0.085	(0.068)	23115

b) Women aged 15–29					
	first-born girl	first-born boy	diff	st. error diff	n
# children ever born	2.541	2.545	-0.004	(0.032)	10048
wants another child	0.867	0.852	0.015**	(0.007)	9988
use contraceptives	0.132	0.132	-0.000	(0.008)	10048
age at first birth	18.345	18.236	0.109	(0.074)	10048
polygynous husband	0.267	0.275	-0.008	(0.010)	9173
female headed hh	0.107	0.109	-0.002	(0.007)	10048
living with husband	0.900	0.899	0.001	(0.007)	9157
first union	0.915	0.913	0.002	(0.006)	9492
woman’s eduyrs	4.889	4.861	0.028	(0.111)	10042
husband’s eduyrs	6.359	6.203	0.156	(0.125)	9308
weight-for-height	-0.571	-0.567	-0.004	(0.023)	9711
body mass index	22.114	22.136	-0.022	(0.089)	9775

c) Women aged 30–49					
	first-born girl	first-born boy	diff	st. error diff	n
# children ever born	5.649	5.639	0.010	(0.050)	13703
wants another child	0.471	0.475	-0.004	(0.009)	13610
use contraceptives	0.174	0.174	0.000	(0.007)	13703
age at first birth	20.202	20.032	0.170*	(0.094)	13703
polygynous husband	0.381	0.379	0.002	(0.009)	12517
female headed hh	0.170	0.149	0.021***	(0.007)	13703
living with husband	0.884	0.905	-0.021***	(0.006)	12495
first union	0.826	0.83	-0.004	(0.007)	13558
woman’s eduyrs	5.125	4.776	0.349***	(0.103)	13693
husband’s eduyrs	6.309	5.864	0.445***	(0.111)	13349
weight-for-height	-0.511	-0.557	0.046*	(0.025)	13275
body mass index	23.839	23.654	0.185*	(0.096)	13340

Notes. Robust standard errors adjusted for clustering at the household level in parentheses. Using survey weights. ***, **, and * indicate significance at 1%, 5% and 10% levels, respectively.

If women’s reproductive behavior is affected by the gender of their first-born child, differences in characteristics may develop among women over time. I split the sample into subsamples of women aged 15–29, and 30–49. Panel (b) shows that the differences are much smaller for younger women (other than the desire for more children, which is expected to be affected among women of

reproductive ages). Among older women (panel (c)), those with a first-born daughter are ‘better-off’ in terms of their own and their husband’s education, first gave birth at an older age, are more likely to be household heads, and are also characterized by higher weight-for-height and body mass index (BMI).¹⁹ These findings, especially those related to differences in predetermined characteristics among older women, are again consistent with selective mortality.

4 Empirical strategy

4.1 Fertility regressions

Given that the sex of the first-born can reasonably be considered as random, the empirical strategy should be straightforward: unbiased estimates should be obtained by simply regressing the number of children ever born on the sex of the first-born. However, as discussed in the previous section, there are differences in observable characteristics between women with a first-born daughter or son that may have developed over time as a consequence of son-preferring fertility behavior. This would generate a negative correlation between a first-born girl and fertility because surviving women have fewer children on average. In all regressions I include a set of observable covariates in order to reduce the potential (downward) bias. Conditional on the controls, if there is preference for male children, women with first-born daughters should exhibit higher fertility than women with first-born sons.

In order to understand if the sex of the first-born also affects the *desire* to have more children, I construct a dummy variable based on the NDHS survey question: ‘*Would you like to have another child, or would you prefer not to have anymore children?*’. I also consider the reported use of contraceptives as an outcome.²⁰

I estimate the following regression:

$$y_{i,t,r} = \beta_1(\text{firstborngirl})_i + \gamma X_{i,t,r} + \alpha_r + \gamma_t + \delta_r t + \chi_e t + \epsilon_{i,t,r} \quad (1)$$

with woman i , born in year t , and resident in region r . $y_{i,t,r}$ is the dependent variable, either the number of children ever born, the desire for more children, and contraceptive use. firstborngirl indicates whether the first child ever born is a girl; $X_{i,t,r}$ is a set of covariates including: age (of the woman and her husband) and age squared, age at first marriage, age at first birth, number of years

¹⁹The Body Mass Index (BMI) is a measure of women’s nutritional status. It is the ratio of the weight in kilograms to the square of the height in meters. Weight-for-height is another indicator of women’s nutritional status and it’s defined in terms of standard deviations from the reference median based on the DHS reference standard.

²⁰The variable for contraceptive use is defined as equal to one if the woman reports she is currently using *any* contraceptive method, zero otherwise.

of education (of the woman and her husband), a wealth index²¹, a dummy for urban areas, and woman’s ethnicity and religion.²² Controlling for age at first marriage and husband’s characteristics entails considering the sample of all currently married women. Given that marital status can also be affected by the sex of the first-born (this is analyzed in section 6.1), women with daughters might be more likely to be ever divorced or separated and, as a result, typically have lower fertility than those continuously married. To account for this possible (downward) bias, I first estimate regression 1 using the entire sample of women, and then restrict it to continuously married women (for whom the effect on fertility should be stronger as found in Dahl and Moretti, 2008).²³ α_r , γ_t are region and cohort of birth fixed effects, respectively. I also control for region and ethnic-specific trends, δ_{rt} and χ_{et} , respectively, to capture region or ethnic-specific trends that may be correlated with the error term. Given that boys are biologically weaker than girls and are more likely to die in infancy, I additionally control for the survival status of the child (a dummy for whether the first-born child died and its interaction with *firstborngirl*).²⁴ Results with and without controlling for child mortality are shown. As previously noted, since mothers tend to replace a lost child by having more children, omitting to control for differential mortality would lead to a downward bias of the effect of a first-born daughter on fertility. After controlling for the mortality of the first-born child and its interaction with *firstborngirl*, β_1 is the estimated effect of a first-born *surviving* daughter on the outcome of interest. Identification relies on the assumption that the sex of the first-born is exogenous (uncorrelated with the error term) after conditioning on observable characteristics. Regression 1 is estimated using OLS and a model for count data (accounting for over-dispersed outcome variables) when the dependent variable is the number of children ever born, and using a probit model for the probability of desiring more children and using contraceptives.

If women practice son-preferring fertility stopping behavior, I expect to find $\beta_1 > 0$ for the regressions for realized (and desired) fertility, and $\beta_1 < 0$ for the use of contraceptives. Although the effect on fertility should (and is found to) be larger for women who had more subsequent daughters among earlier-born children, the analysis focuses on the sex of the first-born, thus allowing a causal interpretation of the results (Dahl and Moretti, 2008). In fact, while the sex of the first-born can

²¹The DHS wealth index is a measure of a household’s cumulative living standard. It’s generated using principal component analysis based on the households ownership of consumer goods; dwelling characteristics; type of drinking water source; toilet facilities; and other characteristics that are related to a households socio-economic status.

²²Ethnicity include Hausa-Fulani, Igbo, Yoruba, and other minor ethnicities as the omitted category. Religion include Catholic, other Christians, Muslims, and Traditionalists (omitted).

²³The rest of the analysis for fertility behavior and birth spacing uses the sample of continuously married women (so that the fertility history at the time of the survey is relevant for the current union).

²⁴I also constructed two alternative variables for child death before reaching the age of 1 or 2 years, and results are qualitatively similar to the ones obtained with child death at any age. This is probably because most deaths happen in the first years of life in this context.

be considered as random, the decision to have more children after the first represents a choice that may correlate with other household decisions.²⁵

4.2 Birth spacing regressions

If there is social pressure for having a son, women who have had daughters among earlier-born children might try to conceive again sooner than women with sons. To study the association between the birth interval and the sex composition of earlier-born children, I use the sample of all children ever born to each woman (the unit of analysis is now the child). I exclude women with twin births and those with more than 12 children to reduce heterogeneity among mothers. I analyze whether the average birth interval is affected by the previous child's sex.²⁶ Given information on the interval between each birth for all children, I can use mother fixed effects and exploit the variation in the length of the interval *within* the fertility history of each woman. This allows me to control for all observable and unobservable characteristics that may correlate with the error term. I estimate the following regression:

$$y_{i,j,t} = \beta_1 \text{girl}_{i,j,t-1} + \sum_k \gamma_k \text{birthorder}_{i,j,t-1,k} + \alpha_j + \epsilon_{i,j,t-1} \quad (2)$$

with child i , mother j , current child t , and preceding child $t - 1$. $y_{i,j,t}$ is the birth interval in months (the time between the birth of child $t - 1$ and t), or dummy variables equal to one if the birth interval is shorter than 24 or 15 months (the critical intervals below which child and maternal health are adversely affected, respectively). α_j are the mother fixed effects, $\text{birthorder}_{i,j,t-1,k}$ is a set of k dummies for birth order, and $\text{girl}_{i,j,t-1}$ is a dummy equal to one if the previous child is a girl. I also add variables for the survival status of the preceding child (and its interaction with *girl*) to control for the fact that boys are more likely to die than girls in this context.

Given the high fertility context, I expect to find larger reductions in birth intervals for women with several consecutive daughters. Therefore, I estimate regression 2 on the following subsamples: the subsample of women who had a first-born daughter (and, separately, those with a first-born boy), the subsample of women who had a first and second-born daughter (and those with mixed

²⁵As it is possible to observe women that progress to the next parity only if they have reached the previous one, those who have had children after daughters might be different to those who progress after sons. This is a similar empirical issue to that of *dynamic selection bias*, exposed in Cameron and Heckman (1998), where they study schooling decisions over the life cycle.

²⁶Fayehun et al. (2011) find weak evidence of shorter spacing after the birth of *each* girl in Nigeria. Compared to them, I estimate a model with mother fixed effects to control for unobserved heterogeneity among mothers. Moreover, I consider the subsamples of women with one or more consecutive daughters to see if the increasing pressure for having sons leads to further shortening of the birth intervals.

sex composition), and the subsample of women with three girls (and, separately, those with three boys, two boys and one girl, and two girls and one boy) as earlier-born children. As a further check, I also estimate regressions using the full sample of women and including interaction terms between *girl* and the sex of earlier-born children.

Identification with fixed effects requires that women with at least three, four, or five children ever born are considered.²⁷ The prediction with son preference is that the birth intervals succeeding the birth of girls should be significantly shorter ($\beta_1 < 0$) than for boys. This effect should be larger the more girls there are among earlier-born children, because of mounting pressure for a son.

5 Main results

5.1 Fertility results

Table 3 shows the estimates of the effect of a first-born girl on the number of children ever born. In column (1), the coefficient on *firstborngirl* is positive but not significant. In column (2), age at first birth is added among the controls and the coefficient is now larger and significant at the 10% level. The number of years of education is included in column (3) and the coefficient of interest is very similar to column (2), probably because of the positive correlation between age at first birth and education. Column (4) additionally controls for husband’s characteristics, therefore restricting the sample to married women (with information on their husband’s age and education) and yields similar results. As expected, adding the controls for first-born child mortality in column (5) and restricting the sample to continuously married women (column 6) leads to a further increase in the coefficient of interest. Specifically, compared to women with a first-born son, women with a first-born (surviving) daughter have 0.07 more children (column 6) (this effect is significant at the 1% level), which correspond to a 1.7 percent increase of the number of children ever born (considering that the average number of children born to women with a first-born son in this sample is 4.32).²⁸ The effect is 2 percent in the subsample of women age 30–49 (column 8), who have had more children and are closer to completed fertility.²⁹

²⁷The percentages of women with at least three, four, or five children ever born in this sample are the following: 70%, 55%, and 42%, respectively.

²⁸The estimated effect obtained using a negative binomial regression model for over-dispersed count data is shown in appendix table A.1 and is consistent with the OLS estimates (region and ethnic trends are not included due to failed convergence of the model with all the controls). A poisson regression yields similar results (with all controls as in column 6 of table 3 including region and ethnic trends), but it is less appropriate than the negative binomial regression due to over-dispersion of the dependent variable (the Pearson test for goodness of fit of the poisson model rejects that the conditional variance of the variable ‘children ever born’ is equal to its conditional mean).

²⁹For comparison, Dahl and Moretti (2008) use Census data for the United States and find that women aged 18–40 with a first-born girl have 0.3 percent more children than women with a first-born boy. Using data for India, I find that women aged 15–49 with a first-born girl have 0.295 more births (equivalent to a 9 percent increase).

Table 3: The effect of a first-born girl on the number of children ever born

dep var.	# children ever born							
	Age 15-49						Age 30-49	
	all women			currently married women			all married	continuously married
Sample				all married	continuously married			
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
first-born girl	0.025 (0.027)	0.040* (0.022)	0.041* (0.022)	0.045* (0.023)	0.056** (0.025)	0.073*** (0.025)	0.090** (0.041)	0.115*** (0.043)
age	0.289*** (0.025)	0.488*** (0.020)	0.491*** (0.020)	0.466*** (0.023)	0.451*** (0.022)	0.441*** (0.023)	0.576*** (0.117)	0.529*** (0.126)
age sq.	-0.002*** (0.000)	-0.004*** (0.000)	-0.004*** (0.000)	-0.004*** (0.000)	-0.003*** (0.000)	-0.003*** (0.000)	-0.005*** (0.002)	-0.004*** (0.002)
age at first birth		-0.256*** (0.003)	-0.251*** (0.003)	-0.255*** (0.005)	-0.249*** (0.005)	-0.246*** (0.005)	-0.238*** (0.006)	-0.233*** (0.007)
eduyrs			-0.027*** (0.003)	-0.023*** (0.004)	-0.023*** (0.004)	-0.023*** (0.004)	-0.028*** (0.006)	-0.029*** (0.006)
husband age				0.038*** (0.008)	0.037*** (0.008)	0.035*** (0.008)	0.042** (0.017)	0.042** (0.017)
husband age sq.				-0.000*** (0.000)	-0.000*** (0.000)	-0.000*** (0.000)	-0.001*** (0.000)	-0.001*** (0.000)
husband eduyrs				-0.001 (0.003)	0.001 (0.003)	-0.000 (0.003)	-0.001 (0.005)	-0.002 (0.005)
age at first marriage				-0.006 (0.005)	-0.008 (0.005)	-0.012** (0.006)	-0.003 (0.006)	-0.007 (0.007)
first-born dead					0.531*** (0.045)	0.595*** (0.048)	0.619*** (0.067)	0.715*** (0.073)
first-born girl x first-born dead					0.010 (0.067)	-0.062 (0.069)	0.021 (0.099)	-0.079 (0.106)
Observations	23440	23440	23440	20479	20479	17587	11897	9775
R-squared	0.529	0.664	0.665	0.675	0.681	0.705	0.510	0.543

Notes. OLS estimates. Robust standard errors adjusted for clustering at the household level in parentheses. Sample of women indicated at the top of the table. Using survey weights. Includes six region dummies, urban dummy, birth year fixed effects, ethnicity and religion, ethnic and region-specific time trends, and a wealth index. ***, **, and * indicate significance at 1%, 5% and 10% levels, respectively.

The rising coefficient of *firstborngirl* across specifications in table 3 confirms that it is important to control for observable characteristics. In particular, surviving women with first-born daughters tend to be more educated and to have the first child later. Since education and age at first birth are negatively correlated with fertility, ignoring these controls would lead to underestimating the effect of a first-born girl on fertility. Similarly, not taking into account that boys are biologically weaker than girls at young ages and are therefore more likely to die would generate a downward bias since mortality is associated with higher fertility.

The results so far suggest that the sex composition of earlier-born children affects subsequent *realized* fertility. The next question is whether it has an effect on mothers' fertility preferences. Table 4 shows that women with a first-born girl are 1.8 percentage points (significant at the 5% level) more likely to report that they want another child than those with a first-born son (column

Table 4: The effect of a first girl on desired fertility and the use of contraception

dep. var:	=1 wants another child		=1 using contraceptives	
	(1)	(2)	(3)	(4)
first-born girl	0.018** (0.009)	0.023** (0.010)	-0.008 (0.005)	-0.011** (0.005)
age	0.004 (0.014)	0.003 (0.014)	0.010 (0.013)	0.011 (0.013)
age sq.	-0.000** (0.000)	-0.000** (0.000)	-0.000 (0.000)	-0.000 (0.000)
age at first birth	0.018*** (0.002)	0.019*** (0.002)	0.002 (0.001)	0.001 (0.001)
eduyrs	-0.003** (0.001)	-0.003** (0.001)	0.005*** (0.001)	0.005*** (0.001)
husband age	0.003 (0.003)	0.003 (0.003)	0.002 (0.002)	0.002 (0.002)
husband age sq.	-0.000 (0.000)	-0.000 (0.000)	-0.000 (0.000)	-0.000 (0.000)
husband eduyrs	0.004*** (0.001)	0.004*** (0.001)	0.001 (0.001)	0.001 (0.001)
age at first marriage	0.004*** (0.002)	0.004** (0.002)	-0.002* (0.001)	-0.002 (0.001)
first-born dead		0.109*** (0.014)		-0.032*** (0.009)
first-born girl x first-born dead		-0.014 (0.023)		0.015 (0.017)
Observations	17460	17460	17587	17587
Pseudo R2	0.277	0.281	0.222	0.223

Notes. Probit estimates, marginal effects reported. Robust standard errors adjusted for clustering at the household level in parentheses. Using survey weights. Includes six region dummies, urban dummy, birth year fixed effects, ethnicity and religion, ethnic and region-specific time trends, # children ever born, and a wealth index. ***, **, and * indicate significance at 1%, 5% and 10% levels.

1).³⁰ As expected, the effect is larger for women with a surviving first-born daughter (column 2), and equivalent to a 3.6 percent effect (compared to the mean for women with a first-born son).

Columns (3–4) of table 4 show the results for the use of contraception. Women with a first-born girl are 0.8 percentage points less likely to use any method of contraception, although the coefficient is not statistically significant. Column (2) estimates the same effect for women with a surviving daughter and, as expected, the effect on the use of contraception is larger (−1.1 percentage points, equivalent to 6.6 percent effect), now significant at the 5% level. This is an important effect considering that only a small fraction of women are currently using contraceptives.

Table 5 shows the estimated effects of the birth of two consecutive daughters on fertility (realized and desired) and contraceptive use for the subsample of women with at least two children ever born. As expected, the percent effects are larger: compared to women with two first born sons, women with two first born daughters have 2.6 percent more children, and are 8.6 percent more (less) likely to desire to have more children (to use contraceptives).

³⁰The regressions for desired fertility and contraceptive use also control for the number of children ever born.

Table 5: Sex of the first two-born children, realized (and desired) fertility and the use of contraception

dep. var:	# children ever	=1 wants another child	=1 currently using contrac.
	(1)	(2)	(3)
girl,girl	0.127*** (0.039)	0.051*** (0.015)	-0.015** (0.008)
one boy one girl	0.038 (0.032)	-0.003 (0.012)	-0.002 (0.007)
Observations	15065	14975	15027
R-squared (or Pseudo)	0.655	0.262	0.230

Notes. OLS estimates in column (1). Probit estimates in columns (2)-(3), marginal effects reported. Robust standard errors adjusted for clustering at the household level in parentheses. Using survey weights. Includes all the controls as in column (6) of table 3 as well as a dummy for whether the second child is dead and its interaction with second-born girl. The # children ever born is included in columns (2)-(3). ***, **, and * indicate significance at 1%, 5% and 10% levels, respectively.

Altogether, these results suggest that women with daughters *want* to have more children to a greater degree than women with sons, and they seem to intentionally continue bearing children until they have the desired number of sons.

5.2 Birth spacing results

I turn now to the impact of the sex composition of earlier-born children on the length of the interval between births. Columns (1) to (3) of table 6 show that the birth of a girl has no effect on the birth interval when considering all births. This is probably due to the fact that in a high fertility context the effect on spacing might become negative (and significant) only after the birth of several girls. The death of a child has a negative and significant effect on birth spacing as mothers probably attempt to conceive again to replace the lost child. It is interesting to note that the omission of the indicators for mortality leads to an overestimation of the effect of the birth of a girl. Consistent with the results found for fertility, women with male children appear to wait less on average than those with daughters due to the higher mortality of sons and the urge to replace the dead child. Columns (4) to (6) show that the birth of each successive girl implies a reduction in the interval for women with a first-born girl. However, this effect is not significant. Columns (7) to (9) focus on the subsample of women who had at least four children, among which the first two are girls. Column (8) shows that, on average, mothers wait 1.32 months less after the birth of a girl than a boy (this effect is significant at the 5% level). The estimates obtained using OLS are similar (column 9). Finally, columns (10) to (12) consider the subsample of women with 5 or more children. Interestingly, for women with girls as her first three born, the length of the interval after the birth of a girl is reduced relative to that of a boy by 3.25 months and this effect is significant at the 1% level.

Table 7 shows additional results for the subsamples of women with a first-born boy (column 2), first and second-born boys (column 4), three first-born boys (column 7), and mixed gender

Table 6: Length of the birth interval (# months), conditional on the gender of earlier-born children

Women with:	3+ children			3+ children			4+ children			5+ children		
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)	(11)	(12)
					first-born girl			girl			girl, girl, girl	
girl	0.300 (0.224)	0.139 (0.258)	0.015 (0.209)	-0.022 (0.377)	-0.171 (0.427)	-0.257 (0.332)	-1.021* (0.562)	-1.316** (0.616)	-1.682*** (0.507)	-2.724*** (0.874)	-3.252*** (0.980)	-3.232*** (0.842)
girl* dead child		0.530 (0.491)	0.706* (0.396)		0.550 (0.689)	0.276 (0.564)		1.198 (1.034)	0.619 (0.883)		2.120 (1.568)	0.936 (1.317)
dead child		-4.928*** (0.360)	-7.459*** (0.276)		-4.993*** (0.579)	-7.450*** (0.452)		-5.650*** (0.938)	-7.514*** (0.768)		-6.694*** (1.440)	-7.204*** (1.202)
Birth order dummies	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Mother FEs	Yes	Yes	No	Yes	Yes	No	Yes	Yes	No	Yes	Yes	No
Observations	51081	51081	51081	24082	24082	24082	11086	11086	11086	4808	4808	4808
R-squared	0.313	0.320	0.025	0.308	0.315	0.026	0.271	0.280	0.028	0.219	0.232	0.035

Notes. OLS estimates. Robust standard errors adjusted for clustering at the household level in parentheses. Using survey weights. ***, **, and * indicate significance at 1%, 5% and 10% levels, respectively.

Table 7: Length of the birth interval (# months), conditional on the gender of earlier-born children

	firstG	firstB	G,G	B,B	mixed	G,G,G	B,B,B	two boys one girl	two girls one boy
Women with:	3+ children		4+ children			5+ children			
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)
girl	-0.171 (0.427)	-0.039 (0.380)	-1.316** (0.616)	0.437 (0.568)	0.167 (0.364)	-3.252*** (0.980)	1.015 (0.904)	0.267 (0.425)	-0.391 (0.438)
girl*dead child	0.550 (0.689)	0.531 (0.733)	1.198 (1.034)	-1.612 (1.001)	1.036 (0.716)	2.120 (1.568)	-1.919 (1.557)	0.675 (0.861)	0.840 (0.824)
dead child	-4.993*** (0.579)	-4.849*** (0.458)	-5.650*** (0.938)	-4.348*** (0.574)	-4.995*** (0.520)	-6.694*** (1.440)	-3.870*** (0.762)	-4.958*** (0.551)	-5.328*** (0.715)
B.order dummies	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Mother FEs	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Observations	24082	26999	11086	12920	21685	4808	5748	14750	13278
R-squared	0.315	0.324	0.280	0.277	0.268	0.232	0.250	0.220	0.237

Notes. OLS estimates. Robust standard errors adjusted for clustering at the household level in parentheses. Using survey weights. ***, **, and * indicate significance at 1%, 5% and 10% levels, respectively.

composition (columns 5, 8, and 9). As expected, the birth of a girl has no effect on spacing for women who had a first-born boy or mixed gender offspring among earlier-born children. As an additional check, appendix table A.2 shows the results when using the full sample of women with at least three, four, and five children, respectively, and including the interaction terms between the sex composition of earlier-born children with the birth of a girl. Again, the results confirm that mothers tend to significantly reduce spacing between births if they have several daughters: these women are in fact those for whom the pressure for having a son is greater.

Next, I estimate the probability that the birth interval is shorter than 24 and 15 months. The results are shown in table 8: columns (1) and (3) show that, after the birth of any girl, mothers who had two girls as the first two born children are 2.1 percentage points more likely to wait less than 24 months, though not significantly so, and 1.5 percentage points more likely to wait less than 15 months (significant at the 10 percent level). In column (2), the sample is restricted to the first 4 children, to make sure that the effect on spacing is not driven by the behavior of women with many children. The results suggest that the pace at which mothers have children tends to accelerate especially among earlier female births (the third and/or the fourth girl born) rather than later born children. Considering that about 32% of women wait less than 24 month in this subsample, the effect is important and equal to a 11.2% increase. Column (4) shows that the likelihood that the next birth is spaced less than 15 months apart increases by 1.8 percentage points after the birth of a girl for mothers who already had two girls. This is equivalent to a 22% increase given that on average 8% of births are spaced less than 15 months apart in this subsample.

Figure 5 plots the distribution of the length of the interval for the subsample of women with two first born girls (and at least three children ever born) by the sex of the third and/or fourth

Table 8: Probability of a short birth interval (shorter than 24 or 15 months), conditional on the first two born girls

Subsample	<24 months		<15 months	
	All children	first 4 children	All children	first 4 children
	(1)	(2)	(3)	(4)
girl	0.021 (0.015)	0.036* (0.022)	0.015* (0.008)	0.018* (0.011)
girl*dead child	-0.046 (0.033)	-0.069 (0.048)	-0.015 (0.023)	-0.015 (0.032)
dead child	0.184*** (0.028)	0.193*** (0.044)	0.117*** (0.020)	0.109*** (0.029)
Birth order dummies	Yes	Yes	Yes	Yes
Mother FEs	Yes	Yes	Yes	Yes
Observations	11086	8149	11086	8149
R-squared	0.244	0.319	0.276	0.373

Notes. OLS estimates, linear probability model. Robust standard errors adjusted for clustering at the household level in parentheses. Using survey weights. ***, **, and * indicate significance at 1%, 5% and 10% levels, respectively.

child. The continuous (dotted) line represents the distribution after the birth of a girl (boy). Given that the length of the interval increases with birth order, I consider the first four births for each woman.³¹ to avoid confounding effects. As expected, the distribution of the length of the interval after the birth of a girl lies to the left of the distribution of the interval after a boy. The two distributions are statistically different from each other, as confirmed by the rejection of the null hypothesis of equality of distributions (Kolmogorov-Smirnov test, with a p-value of 0.004). Appendix figure A.1 shows that the distributions in the subsample of women who had two boys or mixed gender composition are not significantly different from each other. Consistently, figure 6 uses the subsample of women with three girls as the first three born children and shows that the whole distribution of the interval after each successive girl lies to the left of the distribution after each boy (and the two distributions are statistically different from each other as shown by the Kolmogorov-Smirnov test with p-value 0.047). This evidence confirms the idea that the pressure to have a son increases with the number of girls born, and leads to a significant reduction in spacing between births, with implications for the health status of mothers and their children.

6 Son preference and family structure

6.1 Marital outcomes

The anthropological and demographic evidence introduced in section 2 suggests that infertility and sex composition of earlier-born children are among the factors affecting the husband's decision to

³¹This also allows me to exclude the hypothesis that the effect only comes from women who had many children.

Figure 5: Distribution of the length of the birth interval after the birth of each girl or boy

Notes. Subsample of women with at least three children ever born. Excludes births of birth order five or higher. Kolmogorov-Smirnov test for the equality of distribution calculated for range of interval 0-60 months.

Figure 6: Distribution of the length of the birth interval after the birth of each girl or boy

Notes. Subsample of women with at least four children ever born. Excludes births of birth order six or higher. Kolmogorov-Smirnov test for the equality of distribution calculated for range of interval 0-60 months.

marry another woman.^{32,33} Being in a polygynous union has important consequences on women's and children's welfare. In particular, even though there are contrasting views in the literature on the

³²This could be due to the fact that women who only had girls are discriminated by their husband's kin (and the society in general) or because husbands might perceive that the probability of giving birth to boys at next births is lower. There is contrasting evidence in the biology literature that the sex realization of successive births is influenced by the sex of previous children (Ben-Porath and Welch, 1976; James, 2009; Stansfield and Carlton, 2007). Another interpretation could be that taking an additional wife, possibly younger, maximizes the probability of having many sons, given that the fertile period of the first wife is reduced (as it is reduced the possible number of male children that she can possibly bear during her lifetime).

³³Using Census data for Kenya, Dahl and Moretti (2004) find that women with daughters among earlier born children are more likely to be in a polygamous marriage. In the empirical analysis, they consider *resident* children (cohabitation is a very imperfect proxy given the flexibility of household structure in the African context). Moreover, they do not distinguish between wives according to their rank (likely because of lack of this information in the data). They do not find any effect of having a girl on divorce or separation, but again they do not consider the full fertility history of women and do not consider remarried women.

desirability of being in a polygynous union by women, several papers have documented that women and children in polygynous households exhibit poorer health outcomes because of competition for resources (see Bove and Vallengia, 2009, for a review of the empirical evidence on the relationship between polygyny and women’s health in Sub-Saharan Africa).³⁴ Female headship has often, but not always, been found to be associated with poverty and insecurity.³⁵ Horrell and Krishnan (2007) find that female-headed households in rural Zimbabwe are no different from male-headed ones in terms of income poverty, but they lack assets for agricultural production and are thus constrained in their capacity to improve productivity.

6.1.1 Data and results

I use the sample of all married (or ever-married) women to explore whether polygyny and other marital outcomes are associated with the sex of their first-born child. Among polygynous married women (with at least one child ever born), 43 percent are ranked first, 47 percent second, 8 percent third, and 1 percent higher rank.³⁶ The probability of being in their first union is lower for higher-rank wives (89 percent for rank one wives, 69 percent for rank two, 56 percent of rank three, while it is 91 percent for women in monogamous marriages).³⁷

I analyze the effect of a first-born daughter on four separate outcomes: that the husband takes another wife (the husband is polygynous), that he lives with his wife, and that the woman is ever divorced or separated or she is the head of the household. The empirical strategy is similar to that employed in the analysis of fertility outcomes illustrated in section 4.1.³⁸ Again, the identification assumption is that the sex of the first-born is conditionally exogenous. Each regression is run separately for all married (or ever-married women), and on the subsample of older women aged 30–49, who are more likely to have experienced the outcomes under consideration.

I exploit the information on the rank of each wife in a polygynous union and run a regression in which the dependent variable is equal to one if the woman is the first-rank wife, and zero if she is monogamously married (thus excluding wives of higher order rank). This is because the outcome of interest is whether the husband marries another woman as a consequence of the fact that the first wife had daughters instead of sons. Thus, I compare monogamous women to women who are

³⁴Boserup (1970) and Goody (1976) argue that the prevalence of polygyny is positively associated with the degree of female involvement in agriculture. Jacoby (1995) finds a relationship between polygyny and women’s productivity in agriculture in Côte d’Ivoire.

³⁵see Buvinic and Rao Gupta, 1997, for a review of the literature, and van de Walle, 2013, for a study of the welfare consequences of being a widowed female head in Mali.

³⁶The rank is typically based on the duration of marriage, with women ranked first as the ones married the longest.

³⁷Levirate marriage (traditional practice in which the wife of a deceased man is obliged to marry the husband’s brother) is not uncommon in Nigeria.

³⁸Additional controls include the number of children ever born and household size.

currently first-rank wives and so entered the union with their current husband as monogamous. Table 9 shows the results. Column (1) shows that women with a first-born surviving daughter are 1.2 percentage points (significant at 5% level) more likely to be first-rank wives (as opposed to monogamous) compared to women with a first-born son. As a placebo test, I run a regression in which I compare monogamous women and last-rank wives (of rank higher than the first). If husbands choose to marry another woman because the first had a daughter, the effect should be stronger for first-rank wives (compared to monogamous), while no effect (or a smaller effect, since the husband can always marry more women) should be found for higher rank wives. As expected, the probability of being in a polygynous union is not higher for last-rank wives who have had a first-born daughter rather than a first-born boy (the coefficient of first-born girl is 0.005 with a standard error equal to 0.006). The effect of a first-born girl on the likelihood of having a polygynous husband is larger for the older group of women (column 2). Another interesting result is that child mortality is strongly associated with polygyny, and that this effect is attenuated (but still positive) if the first-born deceased child is a girl. In other words, the death of a first-born son appears to affect the husband's decision to marry another woman more than the death of a first-born daughter. It is unclear why this is the case and further research is needed to explain this finding.

Finally, column (4) shows that the husband of women with first-born daughters is 1.4 percentage points less likely to live with his wife (considering all marriages, polygynous and not). Columns (6) and (8) show that ever-married women with first-born girls are also 1.2 percentage points more likely to be ever divorced or separated, and 1.1 percentage points more likely to be the head of the household, respectively. This evidence clearly suggests that the sex of children affects marriage outcomes for Nigerian women. Compared to women with a first-born son, women with first-born daughters appear to be more at risk of falling into a category of women who are typically more disadvantaged. This disadvantage stems from the fact that women often have limited property rights in traditional patrilineal societies and need husbands or other male relatives to access resources.

6.2 Child fostering

According to the child labor hypothesis for child fostering, children participate in household production by performing domestic tasks and children of each gender and age group have their specific roles (Akresh, 2009). Therefore, households with an excess of female or male children decide to send or receive a child in order to achieve a balanced gender structure and thus maximize household productivity. I build on this idea and investigate whether households' fostering decisions differ depending on the sex composition of children. If non-biological girls (boys) are substitutes for

Table 9: Sex of the first-born child, marital outcomes and living arrangements

Dep. var.:	=1 polygyn husband		=1 living with husband		=1 ever divorced or sep.		=1 is the female head	
Sample:	currently married women				ever-married women			
AGE	all	30-49	all	30-49	all	30-49	all	30-49
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
first-born girl	0.012** (0.006)	0.028*** (0.010)	-0.009** (0.004)	-0.014** (0.006)	0.003 (0.005)	0.012* (0.007)	0.003 (0.002)	0.011*** (0.004)
age	0.004 (0.013)	-0.012 (0.025)	0.014* (0.007)	0.016 (0.014)	0.029*** (0.010)	0.036* (0.019)	0.013*** (0.005)	-0.001 (0.009)
age sq.	0.000 (0.000)	0.000 (0.000)	-0.000** (0.000)	-0.000 (0.000)	-0.000** (0.000)	-0.000* (0.000)	-0.000*** (0.000)	0.000 (0.000)
age at first birth	-0.004*** (0.001)	-0.005*** (0.001)	0.001 (0.001)	0.000 (0.001)	-0.002*** (0.001)	-0.002* (0.001)	-0.000 (0.000)	-0.000 (0.001)
eduyrs	-0.007*** (0.001)	-0.008*** (0.002)	-0.002*** (0.001)	-0.002** (0.001)	-0.002*** (0.001)	-0.003** (0.001)	0.001*** (0.000)	0.001*** (0.000)
husband age	0.007*** (0.002)	0.002 (0.004)	-0.005*** (0.001)	-0.006*** (0.002)				
husband age sq.	-0.000*** (0.000)	-0.000 (0.000)	0.000*** (0.000)	0.000*** (0.000)				
husband eduyrs	0.002*** (0.001)	0.003** (0.001)	-0.005*** (0.001)	-0.005*** (0.001)				
age at first marriage	-0.004*** (0.001)	-0.004*** (0.001)	0.001 (0.001)	0.001 (0.001)	-0.006*** (0.001)	-0.006*** (0.001)	-0.000 (0.000)	-0.000 (0.000)
first-born dead	0.055*** (0.011)	0.070*** (0.016)	-0.002 (0.008)	0.001 (0.009)	0.042*** (0.008)	0.060*** (0.012)	-0.007** (0.003)	-0.006 (0.006)
first-born girl x first-born dead	-0.023** (0.011)	-0.043** (0.017)	0.002 (0.011)	0.001 (0.013)	0.006 (0.010)	-0.009 (0.014)	-0.007 (0.005)	-0.015** (0.006)
Observations	16343	9395	20310	11802	22112	12745	22072	12745
Pseudo R2	0.287	0.295	0.088	0.121	0.184	0.167	0.234	0.247

Notes. Probit estimates, marginal effects reported. Robust standard errors adjusted for clustering at the household level in parentheses. Using survey weights. ***, **, and * indicate significance at 1%, 5% and 10% levels, respectively. In Columns (1) and (2) the sample is restricted to currently married women who are monogamous or rank 1 (excluding women of rank 2 or higher); columns (3) and (4) include all currently married women; columns (5)-(8) include all ever-married women and exclude widows. Other controls include region fixed effects, urban dummy, birth year fixed effects, ethnicity and region-specific time trends, religion and ethnicity, a wealth index, # of children ever born, household size. Columns 5-8 include dummies for whether women are currently married.

biological daughters (sons), I should find symmetric responses for the imbalance with more girls or boys.³⁹ If instead there is preference for *own* biological sons but not for *own* daughters, responses should not be symmetric. It is important to note that this analysis is not meant to be a comprehensive analysis of the motivations for child fostering, but is instead a preliminary test of asymmetric patterns of fostering in/out that may be driven by gender preferences.

6.2.1 Household fostering decisions, data and empirical strategy

The analysis of fostering decisions is done at the household level. The number of biological children is based on fertility histories of the spouse/s of the head or of the female head, and includes all daughters and sons alive *before* fostering (summing up resident children and children living else-

³⁹Specifically, if household fostering decisions are influenced by this motivation with no gender bias, the following patterns should be observed. For households with an excess of daughters: (+) foster-in boys; (+) foster-out daughters, and for households with an excess of sons: (+) foster-in girls; (+) foster-out sons.

where).⁴⁰ The focus is on living children aged 6 through 14, for whom fostering is most common.⁴¹

First, I empirically examine whether the probability of sending or receiving a child is correlated with household demographic variables, such as the number of biological children (and separately, the number of boys and girls). Second, I construct two variables for the household gender imbalance: ‘*more sons 6–14*’, and ‘*more daughters 6–14*’ (similar to Akresh (2009)).⁴² These are dummies equal to one if the number of biological sons aged 6–14 (daughters) exceeds the number of daughters (sons) in the same age range, respectively. The omitted category is a dummy equal to one for households in which the number of sons equals the number of daughters. The latter category also includes households with no children in the relevant age group, which are the ones that typically receive more children in Nigeria.⁴³ After estimating the regressions on the full sample of households, I also examine households with at least one biological child aged 6–14 (for the fostering-in regressions) and separately households with at least one child, daughter or son (for the corresponding fostering-out regression). This allows a study of fostering decisions for households that have a gender imbalance versus those that are actually balanced because they have the same (strictly positive) number of boys and girls. I focus on the child-labor hypothesis, controlling for a set of additional household-level observable characteristics.⁴⁴

A foster child is defined as being a resident child whose parents are alive and live elsewhere. Accordingly, dummies for whether the household hosts at least one foster child, girl or boy are created (for the fostering-in regressions). In the NDHS, mothers also report whether each biological child is living in the household or elsewhere. Based on this information, I construct three dummies which are equal to one if at least one biological child, son or daughter of the household head lives elsewhere (for the fostering-out regressions).⁴⁵

Table A.3 shows some summary statistics. The sample excludes households in which there is no

⁴⁰If the household head has more than one wife, the sum of children over all spouses is considered.

⁴¹In the child fostering literature, it is common to consider children in this age range. The percentage of households sending out (receiving) at least one child 6–14 is 17% (6%), while it is 3.8% and 1.8% for children 0–5 (2008 NDHS).

⁴²The issues of simultaneity of household fostering and fertility decisions as well as selection into fostering are not dealt with in this analysis. However, I control for many observable household characteristics and use the variables for the gender imbalance (in addition to the number of children) which is a more exogenous proxy for demographics.

⁴³In the sample of households considered, the percentage of households that has at least one foster child is 6%. This percentage is 8% in the subsample of households with no children in that age group.

⁴⁴The set of controls includes: region fixed effects, urban dummy, wife (or woman head if female headed household) religion and ethnicity, years of education and age of the male head and his wife (or female head), a wealth index, household land ownership, wife and husband working in agriculture, a dummy for polygynous household head, a dummy for female headed household, number of children ever born, number of daughters and sons alive age 0–5, number of household members older than 15 (male and female).

⁴⁵There is no data on the survival status of the father or whether the child is living with the father in a separate compound for non-resident children. Therefore, the variables for fostered-out children are only based on mothers’ survival status (i.e., it overestimates the number of fostered-out children, as it includes paternal orphans and children living with father only).

woman eligible for the woman questionnaire (i.e., not in the 15–49 age group): for these households the number and sex of biological children is not available.⁴⁶ 17 percent of households have at least one child aged 6–14 who lives elsewhere, and 6 percent of households host at least one child aged 6–14.⁴⁷ Moreover, there are more female foster children than male. 24 percent of households have an excess of daughters or sons, while 52 percent are balanced (including households with no children aged 6–14, which represent 40 percent of the sample). The bottom panel of table A.3 shows the summary statistics for urban areas only, where fostering-out (-in) seems to be less (more) frequent than in rural areas. This is consistent with the idea that the quality of the network and the distance to urban areas matter for fostering.

In the empirical part, I estimate a probit model for the probability of fostering-in or -out a child on the number of biological sons and daughters as well as on the constructed variables for the gender imbalance (‘more sons 6–14’, and ‘more daughters 6–14’). Predictions are as follows. First, if fostering is motivated by the need for child labor, I expect that relative to balanced households, unbalanced households tend to foster more to achieve balance. Second, if there is preference for own biological sons, household fostering decisions should differ in response to an excess of male and female children.

6.2.2 Household fostering decisions, results

Table 10 shows the results for the probability of fostering-in a child. Column (1) shows that the higher the number of biological sons and daughters, the less likely a household is to receive a child (a Wald test accepts the null of equality of the coefficients on the number of boys and girls). Columns (2) and (3) show the results for the probability of fostering-in a girl or a boy. In columns (4) to (6) the sample is restricted to households with at least one child in the relevant age range. The results are similar to those obtained for the full sample but smaller in magnitude. These results are consistent with the child-labor hypothesis.

This evidence suggests that biological sons and daughters are both substitutes for foster children in household production. However, from this evidence it is not possible to infer household responses to gender imbalances, since the coefficients only indicate that fewer children are fostered in if there

⁴⁶Polygynous households in which the senior wife is older than 49 (thus, not interviewed) are also not considered.

⁴⁷The difference arises mainly from the fact that the sample does not include households in which there are no women eligible for interview and for whom the fertility history is unavailable. However, households with older women are the ones that host more children (i.e., fostering of grandchildren) and at the same time do not have children in the relevant age group (6–14). Among these households with missing information on the number of children, 11% host at least one child (compared to an average of 6% in the main sample). This explains most of the gap between the percentage of households receiving and sending children. Moreover, the definition of foster (in) children requires *both* parents to be alive and living elsewhere while, as above noted, there is no information on biological fathers of children living elsewhere.

Table 10: Household fostering-in decision: foster child, girl, or boy age 6–14

Households with:	full sample			subsamples			full sample			subsamples		
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)	(11)	(12)
dep. var:	foster child	foster girl	foster boy	foster child	foster girl	foster boy	foster child	foster girl	foster boy	foster child	foster girl	foster boy
# biol. sons	-0.015*** (0.002)	-0.011*** (0.002)	-0.004*** (0.001)	-0.007*** (0.002)	-0.005*** (0.002)	-0.001 (0.001)						
# biol. daughters	-0.020*** (0.002)	-0.014*** (0.002)	-0.006*** (0.001)	-0.011*** (0.002)	-0.007*** (0.002)	-0.004*** (0.001)						
more biol. sons												
more biol. daughters												
							-0.017*** (0.004)	-0.013*** (0.003)	-0.005* (0.003)	0.013** (0.005)	0.010** (0.005)	0.003 (0.003)
							-0.025*** (0.003)	-0.019*** (0.003)	-0.008*** (0.002)	0.005 (0.005)	0.004 (0.005)	-0.000 (0.003)
Observations	20578	20578	20578	12477	12477	12477	20578	20578	20578	12477	12477	12477
Pseudo R-squared	0.054	0.056	0.043	0.042	0.042	0.046	0.044	0.046	0.039	0.037	0.038	0.044

Notes. Probit estimates, marginal effects reported. Robust standard errors adjusted for clustering at the village level in parentheses. ***, **, and * indicate significance at 1%, 5% and 10% levels, respectively. The dependent variables are dummies for whether the household hosts at least one foster child, girl or boy, where a foster child is defined as being a resident child whose parents are alive and live elsewhere. Other controls include region fixed effects, urban dummy, wife (or woman head if female headed household) religion and ethnicity, years of education and age of the male head and his wife (or woman head), a wealth index, household land ownership, wife and husband working in agriculture, a dummy for polygynous household head and female headed household, # children ever born, # of daughters and sons alive age 0–5, # of household members older than 15 (male and female).

are more biological children of either sex (by keeping constant the number of children of the opposite sex). Columns (7) to (12) include indicators for the gender imbalance and reveal an asymmetric pattern. The negative and significant coefficients in columns (7) to (9) suggest that households with an imbalanced gender composition are less likely to foster-in children compared to balanced or childless households (i.e., the omitted category in columns 7, 8, and 9). Specifically, the excess of daughters is associated with a stronger reduction in the probability of receiving a child (column 7) and in particular a girl (columns 8 and 9), than the excess of sons.⁴⁸ To focus on the effect of actual gender imbalances on fostering-in decisions I restrict the sample to households with *at least* one child aged 6–14. In this case the omitted category (‘same number of boys and girls’) does not include *childless* households (i.e., the ones *more* involved in fostering-in). Column (11) shows that households in which there are more sons than daughters are 1 percentage point more likely to receive a girl (significant at 5% level and different from the coefficient on the excess of daughters by a Wald test). By contrast, households with an excess of daughters do not foster in more boys (column 12). This asymmetric response suggests that households with an excess of daughters do not foster in boys because of the desire for having their own biological sons. This motivation may lead them to save household resources (by avoiding hosting boys) and continue bearing children. Instead, households that with an excess of sons are significantly more likely to receive a girl, as girls are often needed in the house for performing domestic chores (Isiugo-Abanihe, 1985; Ainsworth 1996). This evidence is consistent with the idea that foster girls are substitutes for own biological daughters in fostering-in decisions, while boys are not substitutes for sons.

Table 11 reports the results for the fostering-out decision. Column (1) shows that the probability of sending away a child increases with the number of children, which is again consistent with the child-labor motivation for fostering. Moreover, having many daughters increases this probability *more* than having many sons (the Wald test rejects the null of equality of coefficients). This suggests that biological sons are fostered-out less than daughters, probably because parents tend to privilege having sons under their direct control more than daughters. Sons and daughters are not substitutes for each other in fostering-out, as shown in columns 2 and 3. The higher the number of biological daughters (sons), the more daughters (sons) are sent away: this is consistent with the idea of gender-specific roles within the household. Columns 4 to 6 focus on the subsample of households with at least one child, daughter or son, with similar results. Column (7) shows that unbalanced households send out children more than balanced or childless ones, even though a Wald test indicates that the coefficients of the variables for the excess of sons and daughters are not

⁴⁸A Wald test rejects the equality of the coefficients for the excess of sons and daughters separately in each of the regressions in columns 7 and 8, while accepts it in column 9.

Table 11: Household fostering-out decision: biological child, daughter or son age 6-14 living elsewhere

Households with:	full sample			subsamples			full sample			subsamples		
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)	(11)	(12)
dep. var:	child elsewhere	daughter elsewhere	son elsewhere	at least one child elsewhere	at least one dau elsewhere	at least one son elsewhere	child elsewhere	daughter elsewhere	son elsewhere	child elsewhere	daughter elsewhere	son elsewhere
# biol. sons	0.062*** (0.003)	0.003 (0.002)	0.060*** (0.002)	0.048*** (0.005)	-0.007 (0.005)	0.058*** (0.006)						
# biol. daughters	0.071*** (0.003)	0.068*** (0.003)	0.003* (0.002)	0.062*** (0.005)	0.078*** (0.006)	-0.007 (0.005)						
more biol. sons							0.150*** (0.009)	-0.021*** (0.005)	0.124*** (0.008)	-0.015 (0.012)	-0.043*** (0.015)	0.054*** (0.011)
more biol. daughters							0.162*** (0.009)	0.135*** (0.008)	-0.021*** (0.005)	-0.000 (0.012)	0.064*** (0.012)	-0.044*** (0.014)
Observations	20578	20578	20578	12477	9156	9246	20578	20578	20578	12477	9156	9246
Pseudo R-squared	0.218	0.257	0.233	0.083	0.072	0.063	0.193	0.213	0.201	0.068	0.061	0.058

Notes. Probit estimates, marginal effects reported. Robust standard errors adjusted for clustering at the village level in parentheses. Using survey weights. ***, **, and * indicate significance at 1%, 5% and 10% levels, respectively. The dependent variables are three dummies equal to one if at least one biological child, son or daughter of the household head lives elsewhere. Other controls include region fixed effects, urban dummy, wife (or woman head if female headed household) religion and ethnicity, years of education and age of the male head (or woman head), a wealth index, household land ownership, wife and husband working in agriculture, a dummy for polygynous household head and female headed household, # children ever born, # of daughters and sons alive age 0-5, # of household members older than 15 (male and female).

significantly different from each other. Columns (8)-(9) and (11)-(12) suggest that households tend to respond symmetrically to the imbalance of children of either gender in fostering-out decisions.

The asymmetries found for the household fostering decisions suggest that biological sons are preferred to non-biological boys, while this is not the case for female children. Alternatively, this is also partly consistent with the possibility that girls are in general more ‘productive’ or have a greater weight in the household production function. Consistent with this idea and considering that in rural areas boys may have a different (possibly greater) weight in the household production function if they work in the fields, the estimates obtained above should differ in urban and rural areas and between landed and landless households. Panel (a) and (b) of appendix table A.4 show that this is not the case, even though the asymmetries appear to be stronger in rural areas (in agreement with stronger son preference in more traditional rural areas). It is important to consider that the motivations for fostering are widely different in urban and rural areas and that these results deserve further examination.

7 Concluding remarks

This paper shows how fertility decisions are affected by the sex composition of children in a high-fertility setting where little is known about parental gender preferences and where sex ratios at birth are in the normal range. It is also among the first to explore how son preference interacts with complex social and family structures typical of a West African country, like polygyny and child fostering.

The findings indicate that women make fertility decisions that are motivated by the desire for giving birth to sons in Nigeria. Compared to women with a first-born son, women with a first-born daughter have and desire significantly more children, and use less contraceptives. Women with daughters among earlier-born children are also more likely to reduce the spacing between births, thereby increasing the risk of child and maternal mortality. Results suggest that son preference also has significant impacts on other aspects of women’s well-being. Women with first-born daughters are significantly more likely to end up in a polygynous union, to be divorced, and to be heads of the household. Moreover, the analysis of child fostering patterns reveals that daughters may be substitutes for foster girls, while the same does not hold for sons and foster boys, thus suggesting that the preference for *own* biological sons also interacts with household child fostering decisions.

This paper puts forward two channels that may contribute to explain excess mortality among adult women in Sub-Saharan Africa. The first is son-preferring fertility behavior which can have negative effects on women’s health by increasing their risk of maternal mortality and morbidity

through high fertility and short spacing. Although son-preferring fertility behavior is not found to be as strong as in other contexts (like India, where fertility is also lower), the risk of adverse consequences on women's health is significant due to strikingly high lifetime risk of maternal death in Nigeria. The paper provides evidence suggesting that adult mortality is in fact higher for women with first-born girls, especially among those of lower socioeconomic status. The second channel refers to the association between parental gender preferences and specific marital outcomes and living arrangements in which women are often found to be particularly vulnerable and disadvantaged.

The unavailability of accessible and affordable technologies for fetal sex detection may partly explain why sex ratios at birth are currently in the normal range in Sub-Saharan Africa. Nonetheless, the evidence of son preference uncovered in this study poses a potential issue of increasing sex ratios in the future.

Although these results cannot be generalized to all of Sub-Saharan Africa or to other contexts, they represent a starting point for further research studying the effects of gender preferences in a setting that has received little attention by the existing literature. This avenue of research would have straightforward policy implications. For instance, improvements in the quality and availability of health care services, especially for pregnant women, would help reduce the risk of maternal mortality. Other types of reforms include those of the inheritance systems giving women stronger inheritance rights and greater control over resources, thus increasing the 'value' of daughters in traditional patrilineal societies. While longer term reforms might take time to have a significant effect on women's lives, in the short-term social programs should be specifically aimed at improving the livelihoods of women with daughters.

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A Appendix

A.1 Alternative interpretations for declining sex ratio at birth

In this section I discuss two other potential explanations for patterns found in figure 1 and 2. These include biological factors and misreporting. In relation to the former, Mathews, Johnson, and Neil (2008) and Almond and Mazumder (2011) find that women exposed to adverse nutritional conditions are more likely to give birth to females. This is because of the biological ‘weakness’ of male fetuses (relative to female’s) that makes them more likely to die (Waldron, 1983). Therefore, it might be the case that the decreasing sex ratio reflects variations in nutritional conditions over time.⁴⁹ To check for the biological explanation, it is useful to understand whether the sex ratio is consistently high for births occurred during the period 1975–1994 by using data from surveys carried out in different periods. As noted in section 3, the decreasing trend in the sex ratios is common to all survey rounds, independently of the cohort of birth of the child, and not specific to a particular period (right panel of figure 1). This finding suggests that there might be other explanations for the systematically higher sex ratios for births that occurred further back in time from the year of the survey.

A second alternative explanation is that, being based on self-reported retrospective information, birth histories contained in the NDHS might suffer from misreporting and recall bias, especially for births of children who died in infancy (Smith, 1994; Byass et al., 2007). To check whether recall bias might be a gender-biased phenomenon (e.g., mothers more likely to under-report female than male births and deaths), I test if age at first birth predicts the sex at first birth. The idea is that if mothers do not recall their first-born daughters who died and report a later born son as being the first-born instead, they will look like they had a boy as first-born, but later. If this is the case, then age at first birth should be negatively associated with a female first-born. I estimate the probability that the first-born is female on age at first birth using a probit model (and controlling for region and birth year fixed effects, urban dummy, ethnicity, and religion). Table A.5 shows that age at first birth is positively (but not significantly) associated with a female first birth: this evidence is against the possibility of selective recall bias.

⁴⁹Nigeria experienced a severe civil war at the end of the 1960s (that mainly involved the South-East region populated by the Igbos) and several national-level oil shocks during the 1970s and the 1980s. If the biological explanation holds, these negative events would be consistent with lower sex ratio (more females born) for births occurred in the past, but this is not what is shown in the left panel of figure 1, suggesting that there may be other explanations. Moreover, in all regressions for fertility, spacing, and marital outcomes, I control for women’s birth year fixed effects and region as well as ethnic-specific time trends to control for time-specific events and regional or ethnic-specific trends that might be correlated with fertility and the probability of a female birth.

A.2 Figures

Figure A.1: Distribution of the length of the birth interval after the birth of each girl or boy (first four births to each woman)

Notes: Kolmogorov-Smirnov test for the equality of distribution calculated for range of interval 0-60 months.

A.3 Tables

Table A.1: Robustness: the effect of a first-born girl on the number of children ever born

dep. var:	# children ever born
first-born girl	0.065*** (0.024)
age	0.348 (0.354)
age sq.	-0.002 (0.005)
age at first birth	-0.189*** (0.004)
eduyrs	-0.025*** (0.004)
husband age	0.053*** (0.008)
husband age sq.	-0.001*** (0.000)
husband eduyrs	0.001 (0.003)
age at first marriage	-0.022*** (0.005)
first-born dead	0.448*** (0.038)
first-born girl x first-born dead	-0.054 (0.051)
Observations	17587

Notes. Negative binomial regression estimates, marginal effects reported. Robust standard errors adjusted for clustering at the household level in parentheses. Using survey weights. Includes six region dummies, birth year fixed effects, ethnicity and religion, urban dummy, and a wealth index. ***, **, and * indicate significance at 1%, 5% and 10% levels, respectively.

Table A.2: Length of the birth interval (# months), full sample

Women with:	3+ children	4+ children	5+ children
	(1)	(2)	(3)
girl*first-born girl	-0.112 (0.568)		
girl*G,G		-1.536* (0.816)	
girl*one boy one girl		-0.349 (0.653)	
girl*G,G,G			-3.320** (1.299)
girl*two girls one boy			-1.123 (0.992)
girl*two boys one girl			-0.428 (0.986)
girl	0.102 (0.370)	0.503 (0.543)	0.729 (0.890)
B.order dummies	Yes	Yes	Yes
Mother FEs	Yes	Yes	Yes
Observations	51081	45691	38584
R-squared	0.320	0.273	0.231

Notes. OLS estimates. Robust standard errors adjusted for clustering at the household level in parentheses. Using survey weights. ***, **, and * indicate significance at 1%, 5% and 10% levels, respectively. Include controls for dead child, the interactions between girl and dead child (as well as its interactions with the sex composition of earlier-born children in each subsamples), birth order (as well as its interactions with the sex composition of earlier-born children in each subsamples).

Table A.3: Fostering, summary statistics

Sample of households (urban and rural)					
	mean	sd	min	max	N
at least one child 6–14 living elsewhere	0.172	0.377	0	1	21501
at least one daughter 6–14 living elsewhere	0.106	0.308	0	1	21501
at least one son 6–14 living elsewhere	0.098	0.298	0	1	21501
at least one foster child 6–14 in the hh	0.061	0.239	0	1	21501
at least one foster girl 6–14 in the hh	0.041	0.198	0	1	21501
at least one foster boy 6–14 in the hh	0.025	0.156	0	1	21501
more daughters 6–14 (# daughters > #sons)	0.240	0.427	0	1	21501
more sons 6–14 (# daughters < #sons)	0.246	0.431	0	1	21501
same number (# daughters = #sons)	0.515	0.500	0	1	21501
no biol. children age 6–14	0.398	0.490	0	1	21501
# biological children 6–14	1.417	1.562	0	14	21501
# biological daughters 6–14	0.702	0.977	0	12	21501
# biological sons 6–14	0.715	0.987	0	10	21501
land (at hh level)	0.639	0.480	0	1	21482
woman working in agric	0.167	0.373	0	1	20874
husband working in agric	0.362	0.480	0	1	20874
Urban areas only					
at least one child 6–14 living elsewhere	0.133	0.340	0	1	6572
at least one daughter 6–14 living elsewhere	0.078	0.269	0	1	6572
at least one son 6–14 living elsewhere	0.077	0.266	0	1	6572
at least one foster child 6–14 in the hh	0.069	0.254	0	1	6572
at least one foster girl 6–14 in the hh	0.047	0.212	0	1	6572
at least one foster boy 6–14 in the hh	0.027	0.162	0	1	6572
# biological children 6–14	1.187	1.398	0	10	6572
# biological daughters 6–14	0.589	0.879	0	8	6572
# biological sons 6–14	0.598	0.888	0	6	6572
land (at hh level)	0.334	0.472	0	1	6559
woman working in agric	0.047	0.211	0	1	6446
husband working in agric	0.095	0.294	0	1	6446

Notes. 2008 NDHS. Using sample weights. Statistics are at the household level. Excludes households in which there is no woman eligible for the woman questionnaire (age 15–49) for which the number and sex of children is not available.

Table A.4: Household fostering decisions, urban and rural areas

	Fostering-in			Fostering-out		
	foster child	foster girl	foster boy	biol. child elsewhere	biol. girl elsewhere	biol. boy elsewhere
	(1)	(2)	(3)	(4)	(5)	(6)
Panel a. Urban and rural households						
more biol. sons	0.015** (0.006)	0.015*** (0.006)	0.001 (0.004)	-0.028** (0.014)	-0.051*** (0.016)	0.050*** (0.013)
more biol. daughters	0.007 (0.006)	0.006 (0.006)	0.000 (0.004)	0.000 (0.014)	0.062*** (0.013)	-0.034** (0.016)
more biol. sons*urban	-0.005 (0.011)	-0.011 (0.007)	0.009 (0.009)	0.045 (0.029)	0.030 (0.040)	0.012 (0.026)
more biol. daughters*urban	-0.006 (0.010)	-0.006 (0.008)	-0.002 (0.006)	-0.001 (0.027)	0.006 (0.026)	-0.041 (0.029)
urban	-0.002 (0.010)	0.004 (0.008)	-0.003 (0.006)	-0.006 (0.024)	-0.016 (0.025)	0.019 (0.024)
Observations	12477	12477	12477	12477	9156	9246
Pseudo R-squared	0.04	0.04	0.05	0.07	0.06	0.06
Panel b. Landed and landless households						
more biol. sons	0.016 (0.010)	0.010 (0.009)	0.006 (0.006)	0.003 (0.024)	-0.037 (0.032)	0.064*** (0.020)
more biol. daughters	-0.004 (0.010)	-0.002 (0.008)	-0.007 (0.005)	-0.004 (0.025)	0.043* (0.024)	-0.053* (0.027)
more biol. sons*land	-0.005 (0.011)	-0.000 (0.009)	-0.004 (0.006)	-0.025 (0.027)	-0.008 (0.037)	-0.015 (0.024)
more biol. daughters*land	0.014 (0.013)	0.009 (0.011)	0.011 (0.008)	0.006 (0.028)	0.031 (0.027)	0.012 (0.035)
land	0.003 (0.010)	0.000 (0.008)	-0.001 (0.006)	-0.007 (0.025)	-0.032 (0.025)	-0.007 (0.023)
Observations	12477	12477	12477	12477	9156	9246
Pseudo R-squared	0.04	0.04	0.05	0.07	0.06	0.06

Notes. Probit estimates, marginal effects reported. Robust standard errors adjusted for clustering at the village level in parentheses. Using survey weights. ***, **, and * indicate significance at 1%, 5% and 10% levels, respectively. Other controls include region fixed effects, urban dummy, wife (or woman head if female headed household) religion and ethnicity, years of education and age of the male head and his wife (or woman head), a wealth index, household land ownership, wife and husband working in agriculture, a dummy for polygynous household head and female headed household, # children ever born, # of daughters and sons alive age 0–5, # of household members older than 15 (male and female). Regressions in columns (1) to (4) are for the subsample of households with at least one child alive age 6–14; and the subsamples of hhs with at least one daughter and son age 6–14 in columns (5) and (6), respectively.

Table A.5: Age at first birth and sex of the first-born child

dep. var:	first-born girl
age at first birth	0.001 (0.001)
Observations	23455
Pseudo R-squared	0.002

Notes. Probit estimates, marginal effects reported. Robust standard errors adjusted for clustering at the household level in parentheses. Using survey weights. Includes six region dummies, urban dummy, birth year fixed effects, ethnicity and religion. ***, **, and * indicate significance at 1%, 5% and 10% levels, respectively.