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Gender Preferences in Africa: A Comparative Analysis of Fertility Choices

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Abstract

Exploiting Demographic and Health Surveys in Africa, we use a duration model of birth intervals to document variations in gender preferences across the continent. We find robust evidence of son preference in Northern Africa, taste for balance in South Africa and no revealed preference in Central Africa. Elsewhere, results are mixed. Then, son preference decreased over time in Sub-Saharan Africa, but increased in Northern Africa. Within country, we investigate the role of socioeconomic factors in shaping gender preferences. Traditional kinship structure is a good predictor, while religion is not. Modernization (wealth, female labor, contraceptive use) is associated with a strong taste for balance.

Keywords : Gender preferences, Fertility, Africa.

JEL Codes : I1, J13, J16, O55.

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In the early 90s, Sen (1990) coined the term « missing women » to draw attention on the excess mortality of women in Asia : he estimated that approximately an extra hundred million women would be there if men and women received similar care in health, medicine, and nutrition. Since then, a large literature focusing mainly on South Asia and East Asia has described the discrimination against girls, mentioning for instance sex-selective abortions (Sen 2001), differential child mortality (Rose 1999), or differential health status (Pande 2003). By contrast, Sub-Saharan Africa appears to do remarkably well. Sex ratios at birth are close to one, and survival rates as well as health outcomes are generally better for girls than for boys (Svedberg 1990, Anderson and Ray 2010). All this may explain why gender preferences for children in Sub-Saharan Africa are rarely studied. In this paper, we focus on fertility behavior as an alternative mechanism generating gender inequality, even when aggregate sex ratios are balanced.

In their seminal paper, Ben-Porath and Welch (1976) infer the existence of gender preferences from the correlation between the probability to stop having children and the gender composition of existing ones. The idea has given rise to formal models of differential *stopping* behavior in favor of sons, predicting that an average girl has more siblings than an average boy [1] (Jensen 2005). There might be important implications for gender inequality because girls would then face more competition for household resources. [2]

On the other hand, the analysis of differential *spacing* behavior is of special importance in the African context. Indeed, when couples have many children, gender preferences are

1. As a simple example, consider a population in which parents have only one child if the first-born is a boy, and have two children if she is a girl. 50% of couples would have one son, 25% one girl and one son, and 25% two girls. Sex ratio is perfectly balanced at the aggregate level, but at the household level, girls have always one sibling whereas boys have, on average, 1/3 sibling.

2. Differential stopping behavior has been tested and validated by an extensive literature focusing again on Asia : e.g. Clark (2000), Jensen (2005), and Basu and de Jong (2010) in India, Abrevaya (2009) in the Chinese and Asian Indian populations living in the US, and Hatlebakk (2012) in Nepal. There are also studies on OECD countries, but the focus is different : the taste for balance in the gender composition of children is generally used as an instrument to study the impact of the number of children on a given outcome, in particular on female labor supply (Angrist and Evans 1998).

more likely to lead to differences in birth intervals rather than in sibship size. Jensen (2005) and Basu and de Jong (2010) advocate looking at birth intervals to find evidence of gender preferences in a high-fertility context. Under son-preferring, differential spacing behavior imply that an average boy is breastfed longer than an average girl, [3] which may translate into inequality between boys and girls (Jayachandran and Kuziemko 2011). Another reason to consider birth intervals is to account for health risks related to spacing, and not only to the number of births : according to the medical literature on developing countries, short birth intervals are associated with adverse outcomes for mothers (Conde-Agudelo and Belizan 2000) and children (Conde-Agudelo, Rosas-Bermudez, and Kafury-Goeta 2006). The authors show that intervals lower than 24 months multiply the risk of infant death by 2.5, and intervals lower than 15 months multiply the risk of maternal death by 2. If gender preferences turn out to induce short birth spacing, they could be a significant cause of maternal and infant mortality in Africa.

Formally, we use a duration model of birth intervals : we test if the gender composition of current children influences the duration before the next birth. The main advantage of duration models is to deal properly with right-censored observations, i.e. families that are not yet complete by the time of the survey (Leung 1988, Leung 1991). We infer the existence of son (resp. daughter) preference when birth spacing is shorter for couples with fewer sons (resp. fewer daughters) ; and we deduce that preference for variety prevails when couples having a balanced mix of sons and daughters wait longer than couples having same-sex children. The conceptual framework underlying this strategy is a unitary model of the couple [4] choosing

3. Again, as a simple example, consider a population in which parents always have two children ; if the first-born is a girl, they try to have another child immediately, while they wait some time if the first-born is a boy. First-born girls, who represent half of the female population, are weaned prematurely while the entire male population is properly breastfed.

4. This is probably a strong assumption given the complexity of marital lives in Africa. In particular, we do not take into account that children might have different fathers, and that those fathers might also have children with other women. We will partly address this issue by comparing polygamous and monogamous women.

optimal spacing and stopping rules. People might have a taste for balance in the gender composition of children, or a girl/boy bias; then, costs and benefits may differ for sons and daughters. What is labeled as « gender preferences » is the outcome of a decision problem based on tastes and prices. One caveat of this strategy is that preferences are revealed if and only if couples have the means to control birth spacing and/or stopping. Using duration models of birth intervals, son preference has been extensively tested and validated in Asia, [5] but not in Africa. To our knowledge, there is no empirical study based on fertility behavior that documents systematically the variation in gender preferences in Africa. We contribute to fill in this gap using DHS data in 37 African countries.

The predominance of son preference in Northern Africa (Morocco, Tunisia and Egypt) is our most robust result. Then, South Africa is characterized by a strong taste for balance and Central Africa by the absence of revealed gender preferences. In the rest of the continent, results are rather mixed. We find weak evidence of son preference in Mali, Senegal and in the Great Lakes region; elsewhere, fertility behavior is consistent with either preference for variety or no preference. In any case, gender preferences do not seem to influence substantially fertility patterns in Sub-Saharan Africa. We also find that, overall, son preference has decreased over time in Sub-Saharan Africa, whereas it has increased in Northern Africa.
We further focus on Sub-Saharan Africa and investigate the role of socioeconomic factors in shaping gender preferences within country. Religion is not correlated with gender preferences : within a country, Muslims exhibit the same taste for balance as other religious groups. In contrast, traditional kinship structure is a strong predictor : son preference prevails in patrilineal ethnic groups only. Then, the preference for variety is stronger in wealthier households, while son preference is stronger when mothers do not work. Last, we find no evidence of gender preferences among women who do not use modern contraceptives; for them,

5. E.g. in China (Tu 1991), in Bangladesh (Rahman and DaVanzo. 1993), in the Chinese population of Malaysia (Pong 1994), in Vietnam (Haughton and Haughton 1995), in India (Arnold, Choe, and Roy. 1998), in South Korea (Larsen, Chung, and Gupta 1998), and in Taiwan (Tsay and Chu 2005).

we cannot tell whether they have no preferences or are not able to translate them into fertility choices.

All in all, we find that, if son preference changes considerably fertility patterns in Northern Africa, it does not seem to be the case in Sub-Saharan Africa. Beware that our results do *not* imply that there is no gender issues in this region. They suggest that women are not discriminated against from birth on, contrary to what may happen in other parts of the world. But it does not mean that such a discrimination would not appear later in life.

The outline of the paper is as follows. Section 1 provides background on theoretical motives for gender preferences and a review of empirical evidence in Africa. Section 2 presents the data and some descriptive statistics. Section 3 discusses the empirical strategy and the identification assumptions. The main results are reported in Section 4, and some robustness tests are described in Section 5. Section 6 concludes with directions for future research.

1 Gender preferences in Africa

1.1 Theoretical motives for gender preferences

The most important motive put forward by the literature on gender preferences is the traditional structure of family systems. In patrilineal [6] and patrilocal [7] family systems, men are the fixed points in the social order, so that investment in daughters is considered as investment in another family's daughters-in-law. In Asia, such a system has produced economic incentives to have sons. For instance, the money spent for a son's marriage remains in the family while the dowry paid for a daughter's marriage is a net expense. In the same vein, female labor force participation is only valued once the daughter is adult, hence benefiting the family-in-law. Last, sons act as old age insurance for their parents, because they are

6. Main assets are passed on through the male line whereas daughters are given movable goods.
7. Upon marriage, wives move to their husbands' abode.

the ones who remain in the family's house. They also act as widowhood insurance for their mother, because widows' claims on the late husband's resources enjoy a higher social legitimacy if they have sons (Agarwal 1994, Das Gupta, Zhenghua, Bohua, Zhenming, Chung, and Hwa-Ok 2003). Mothers, in particular, really need a son because their status improves substantially when their sons get married : they can exert their power over daughters-in-law. Ultimately, women play a dramatic role in the perpetuation and reinforcement of patriarchy. Demographers working on Africa have come to similar conclusions (Lesthaeghe 1989) : among the key factors shaping the reproductive regime in this region, they mention traditional inheritance patterns. In matrilineal societies, having daughters is necessary to perpetuate the lineage, whereas families need sons in patrilineal societies. But Africa is different from Asia along at least two dimensions. First, the system of brideprice prevails in almost all ethnic groups : the groom has to pay for the bride, contrary to what happens in a system of dowry. Second, the kinship structure is more flexible : adoptions and exceptions to allow daughters to inherit land in the absence of a son are not unusual in African patrilineal societies. Eventually, the imperative to have a biological son is weaker in Africa than in Asia.

Another motive specific to Africa is the depth of Islamic penetration. In Northern Africa, the influence of the Islamic law is strong. These societies are characterized by property concentration, endogamous marriages and women seclusion, which implies that women's security and status critically depend on their ability to have sons. In Sub-Saharan Africa, traditions and customs have generally advocated common land ownership, exogamous marriages, women labor participation and women's societies, which renders women less dependent on their sons.(Lesthaeghe 1989)

The last part of the literature focuses on the impact of modernization factors on gender preferences ; female education and labor participation, access to modern contraceptives, urbanization, economic growth and mass media are the most studied factors. The modernization hypothesis states that socio-economic development would equalize the value of daughters and sons to their parents, leading to preferences for variety. However, modernization also brings about birth control - promoting smaller family size and facilitating sex-selective reproductive behavior - which could intensify, at least in the short run, traditional gender preferences. So far, the debate is still open, since empirical studies have found mixed results, depending on the context, the indicator, and the empirical specification they look at.[8]

1.2 Empirical evidence so far

In her review of the empirical evidence on gender preferences, Fuse (2008) concludes that, although North Africa has not been subject to much research compared to East or South Asia, there is evidence of strong gender bias against girls. She further writes that « of all sub-regions in the world, it appears that the least is known about Sub-Saharan Africa ».

Cross-country analyses generally find evidence of son preference in Northern Africa, but not in the rest of the continent (see Arnold (1992) on declared preferences and fertility behavior and Chakravarty (2012) on breastfeeding duration). Sub-Saharan Africa is characterized by a female advantage in infant mortality (Anderson and Ray 2010), as well as in nutritional status and health outcomes (Svedberg 1990). It does not display any systematic gender differences in breastfeeding and health seeking behavior (Garenne 2003). Regarding declared preferences,[9] Fuse (2008) reports that most women in this region have no ideal gender composition, or would prefer to have the same number of sons and daughters. She points to some variation across countries, though : the proportion of women wanting more sons is much larger than that of women wanting more daughters in Burkina Faso, Chad,

8. For instance, modernization factors are associated with smaller son preference in studies on India (Bhat and Zavier 2003), Nepal (Barbar and Axinn 2004), China (Arnold and Liu 1986, Poston 2002), or Egypt (Vignoli 2006). But other studies on India (Das Gupta 1987, Basu 1999, Rajan, Sudha, and Mohanachandran 2000, Jayachandran 2014), South Korea (Edlund and Lee 2013), Egypt (Yount, Langsten, and Hill 2000) or Sub-Saharan Africa (Klasen 1996) have questioned this result.

9. Parents are supposed to have an ideal gender composition of children. Son (resp. daughter) preference is then defined as the ideal number of sons being strictly greater (resp. lower) than the ideal number of daughters.

Guinea, Ethiopia, Mali, Niger, Nigeria and Senegal, while the opposite is true in Gabon, Malawi, Namibia and Zambia. Deaton (1987) examined whether household expenditures on goods consumed only by adults declines more when an additional boy is born compared to an additional girl, and found no evidence of differential allocation of resources between boys and girls in Ivory Coast.

Still, some studies show that son preference may appear in Sub-Saharan Africa in case of income shocks. For instance, Flato and Kotsadam (2014) find that infant mortality increases more for girls than for boys during a drought ; they further explain that such a difference is due to discrimination, since the effect is larger in communities more likely to discriminate against daughters (strong declared son preference, preference for a small family size and low female employment). In the same vein, Friedman and Schady (2012) find that girls are more exposed than boys to mortality risk in case of aggregate economic shock.

Last, a specific study on Nigeria shows that women with first-born daughters are significantly more likely to end up in a polygynous union, to be divorced, and to be the head of the household ; they also have significantly more children (Milazzo 2014).

Only a few papers estimate duration models of birth intervals in an African context.[10] Gangadharan and Maitra (2003) find evidence of son preference in South Africa, but only among the Indian community. Anthropologists working on the Gabbra, a patrilineal and patrilocal society in Kenya, find that women with no son have shorter birth intervals than women with at least one son (Mace and Sear 1997). Last, Lambert and Rossi (2014) show that, in Senegal, women most at risk in case of widowhood substantially shorten birth spacing until they get a son. They relate son preference to women's needs for widowhood insurance.

Our paper contributes to the literature using duration models of birth intervals to test systematically for son preference in Africa. So far, evidence is quite limited in this region.

10. Some papers use duration models of birth intervals in Africa, but they are interested in the impact of socio-economic factors (e.g. mother's characteristics such as birth cohort, age at first marriage and at first birth, residence, education in Ghilagaber and Gyimah (2004)), not in son preference.

2 Data

2.1 Data

We use DHS surveys (Demographic and Health Surveys) that were collected from 1986 to 2012 in 37 African countries (surveys listed in Table 4 in the Appendix). DHS data contain stratified samples of mothers aged 15 to 49 who are asked about their reproductive history. In particular we know the date of birth and death of each of their born children, along with mothers and children's characteristics.

DHS data are provided with individual survey weights to ensure that the survey sample is representative of all mothers at the country level. Nonetheless, sample size of surveys is not proportional to population size. To obtain a representative sample of the 37 African countries studied, we reweighed the whole sample. [11]

The main advantage of these data is that we observe all births, for children either alive or dead at the time of the survey, and we know the year and month of birth of all children, which enables us to measure birth intervals in months. Also, surveys are similar across countries, with a large number of observations (cf. Table 4 in the Appendix), which allows a comparative analysis.

Nonetheless, the comparative analysis over space and time based on DHS data has two limits. On the one hand, surveys are not available in all African countries ; notably Algeria, Libya, Mauritania, Eritrea, Somalia, Angola and Botswana are missing. Still, our sample represents 92% of the whole African population in 2009. On the other hand, the surveys took place during a relatively long period of time, so that by pooling the surveys together, we are considering different periods in different countries. In Sub-Saharan Africa, the period of interest is quite homogeneous : in all countries, the majority of mothers are born in the

11. Using World Bank population statistics, we compute a sampling rate equal to the number of mothers in the survey implemented in country j and year i, divided by the total population of country j in year i. We also correct for the different number of surveys by country.

60s-70s. This is true also in Egypt, but not in other Northern countries : in Tunisia, Morocco and Sudan, most women are born before 1960. [12] We have to keep these caveats in mind when interpreting cross-country comparisons and time evolutions.

We exploit a second source of data to get information on family systems. We use Murdock's data on African ethnic groups (Murdock 1959) coded by Gray (1998) [13] to define which women belong to a matrilineal ethnic group. We opted for a conservative definition of matrilinearity, including only those ethnic groups listed by Gray that we found in the Dhs data. Patrilinearity is identified by default, and probably includes some matrilineal groups. [14] Such a measurement error in our classification of ethnic groups is likely to flatten the differences between matrilineal and patrilineal groups. So when comparing the two categories, we estimate a lower bound of the difference.

2.2 Descriptive statistics

Before estimating the duration model, we provide some descriptive statistics of the non-censored durations. The upper graph in Figure 1 represents the average birth spacing by gender composition of the first children. More precisely, we plot the average duration between births n and $(n+1)$ as a function of the proportion of boys among the first n children, for $n = 1$ to 7. [15] At each rank, birth spacing clearly displays an inverted U-shape : [16] it is lower

12. On the other hand, the fact that we do not observe recent cohorts in Morocco and Tunisia makes our sample more homogeneous : all countries are at a very early stage of the fertility transition.

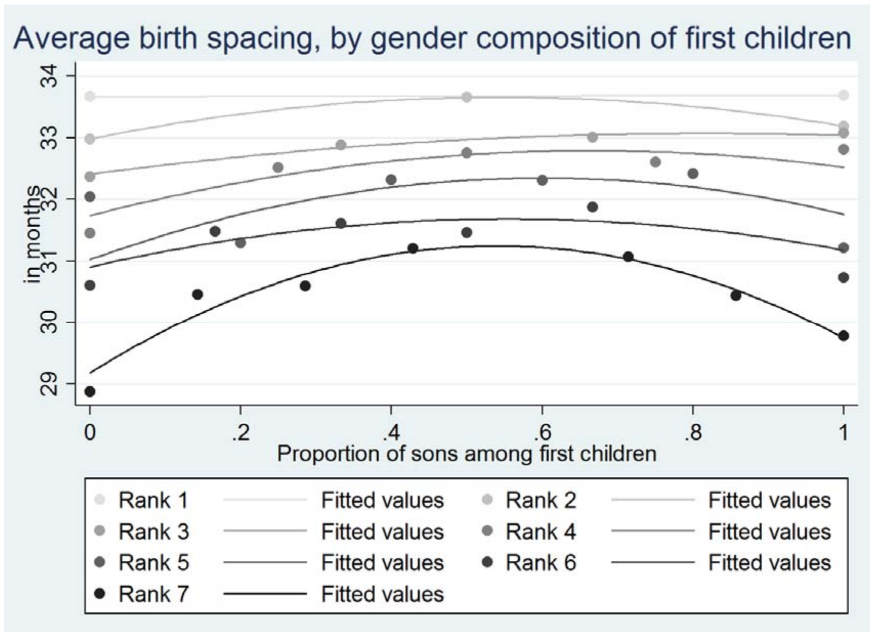
13. Gray (1998) provides a database that lists the characteristics of African ethnic groups reported by Murdock. Of particular interest for us is the variable $Descent : MajorType$ that indicates whether an ethnic group is patrilineal or matrilineal. Table 5 in Appendix gives details of our classification and explains how we matched Gray's and Dhs data.

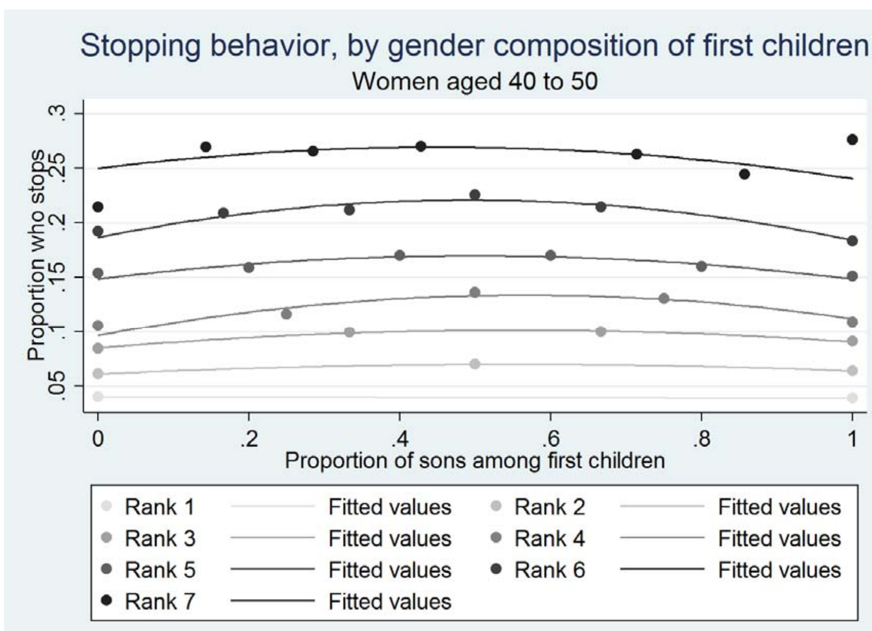
14. We know from Gray (1998) that matrilineal ethnic groups exist in some countries, but the ethnic group variable was missing in Dhs data (e.g. Nigeria, Sudan, Tanzania, Zimbabwe) or Dhs data types were too broad (e.g. Drc) so that we could not identify them.

15. At higher ranks, there are much fewer observations for each possible value of the proportion of boys, so that computing the average is not so meaningful.

16. When we regress the non-censored durations on the proportion of boys and the proportion squared, coefficients are significant and of expected sign, whatever rank we consider. We further plot the lines corres-

Figure 1: Descriptive statistics on non-censored observations





for couples with no son or no daughter, and higher when the sex ratio is balanced. The maximum is reached by couples having slightly more boys than girls.

We find the same pattern when we look at the stopping behavior (cf. lower graph in Figure 1). We restrict the sample to women over 40 years old to mitigate the issue of right-censoring, and we plot the proportion of women who stopped having children after the n^{th} birth as a function of the proportion of sons among the first n children. Here again, women are more likely to stop having children when they already have a balanced mix of boys and girls.

In Appendix, Table 6 reports some statistics on fertility stopping and spacing behaviors, by country. In our sample, women have on average 6.2 children, the average birth interval is 35 months, and one third of birth intervals are shorter than 24 months. But there is a lot of variation across the continent. Southern African countries stand out because of long intervals and relatively low numbers of children : less than one fourth of short birth intervals and approx. 4 children per woman. The opposite is true for the Sahel region : between 7 and 8 children per woman, and more than one third of short birth intervals. In Northern Africa, the proportion of short intervals is over 40%, although the number of children is not that high, between 5 and 6.

3 Empirical Strategy

3.1 A duration model of birth intervals

As explained in the introduction, we are mainly interested in differential spacing rules.[17] We use a duration model of birth intervals to infer the existence of gender preferences. Our variable of interest T is the duration between births n and $(n+1)$, measured in months,

ponding to the quadratic regressions on the graph, and they fit quite well.

17. As a robustness test, we also examine differential stopping rules, and as expected, we find much more scarce evidence of gender preferences (see Section 5.5).

where $n \geq 1$. Our coefficients of interest measure the impact of the gender composition of the first children on the subsequent birth interval. We estimate a Cox proportional hazard model (Cox 1972).

The main reason to prefer duration models to linear models is the issue of censoring : the former allow us to identify the distribution of a duration variable from potentially right-censored observations as long as the duration and the right-censoring variables are independent. This condition is very likely to be satisfied as the date of the survey is completely unrelated to the latest births.

An alternative strategy would be to estimate, on the one hand, the probability to have another child, and on the other hand, the duration before the next birth. In our strategy, we implicitly assume that the impact of the gender composition on both decisions is the same. The first reason for this choice is parsimony : we want to build a unique indicator of gender preferences, in order to compare it across countries, periods, socio-economic categories etc. Also, we would have to make some parametric assumptions to separate the stopping and spacing dimensions, while here, we are able to use a semi-parametric method of estimation. More importantly, in our context, it is not clear that fertility choices are a two-step decision process, in which people choose, first, if they want another child, and second, the timing of the birth. Indeed, if couples have more control over spacing out births than over stopping them, it may well be the case that they only decide to bring forward or to delay the next birth. The eventual number of births would then be mechanically determined by the successive decisions over timing together with the end of the couple's reproductive period.

3.2 Relating durations to the proportion of sons

We want to design a model that exploits the information on all ranks of birth, and not only intervals after a given rank.[18] To do so, we create a variable $Frac_n$ equal to the proportion

18. In Section 5.3, we investigate whether revealed gender preferences differ across ranks of birth.

of boys among the first n born children, and model the hazard function at each country level – the instantaneous probability to have another child at date t – as follows :

$$\lambda(t) = \lambda_0(t) \times \exp(\alpha_1.Frac_n + \alpha_2.Frac_n^2 + \theta.X_n)$$

Where $\lambda_0(t)$ is the baseline hazard function, common to all individuals, and X_n is a vector of mother's characteristics (birth cohort, age at birth n, age at birth n squared, religion, family system, union type, education, wealth, area of residence, employment status) ; [19] it also includes a dummy for each rank n, to control for potential differences between birth orders. In our specification, the unit of observation is not the mother, but the birth. We reweighed the observations to ensure that each woman counts once, irrespective of her number of children. [20] We also use robust standard errors clustered at the woman level to account for the correlation between the error terms related to the different birth intervals of the same woman.

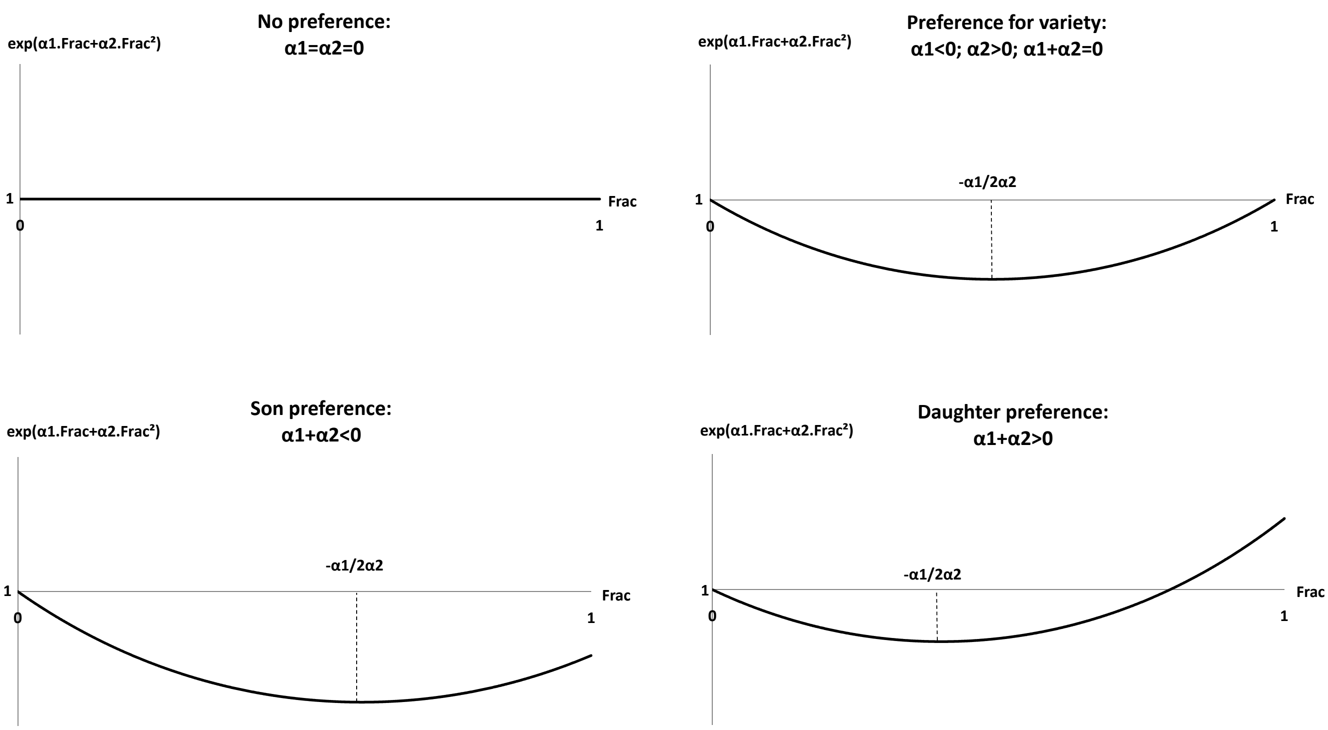
Under the proportional hazard assumption, $e^{\alpha_1+\alpha_2}$ measures the hazard ratio at any point in time between women having only sons vs. only daughters. If $\alpha_1 + \alpha_2 < 0$, it means that having only sons vs. only daughters decreases the hazard rate and hence increases the expected birth interval. In this case, we infer the existence of son preference. Conversely, if $\alpha_1 + \alpha_2 > 0$, we infer the existence of daughter preference.

We introduce the proportion squared to test for a taste for balance in the gender composition, as illustrated in Figure 2. We plot the multiplier on the baseline hazard as a function of $Frac$ for different values of α_1 and α_2. On the top left, we plot the trivial case in which $\alpha_1 = \alpha_2 = 0$, meaning that the gender composition of current children has no impact on

19. We introduce some controls to estimate more precisely the baseline hazard for different categories of mothers, thus reducing our standard errors. The magnitude of our estimates is unchanged if we remove the controls.

20. If we do not reweigh the observations, a woman with n births counts n times. So women having more children, meaning women with a taste for large families and older women, are over-represented.

FIGURE 2: Multiplier on the baseline hazard as a function of the proportion of sons



subsequent durations. Then, if $\alpha_1 < 0$ and $\alpha_2 > 0$, it implies that the hazard rate is lower for women having children of each gender. The lowest hazard rate is reached by women having a proportion of boys among their children that is exactly equal to $-\frac{\alpha_1}{2\alpha_2}$. Therefore, the longest duration is predicted to be observed (i) among couples having exactly the same number of boys and girls if $\alpha_1 + \alpha_2 = 0$ (graph on the top right) ; (ii) among couples having sons and daughters, but more sons than daughters, if $\alpha_1 + \alpha_2 < 0$ (graph on the bottom left) ; and (iii) among couples having sons and daughters, but more daughters than sons, if $\alpha_1 + \alpha_2 > 0$ (graph on the bottom right). Our statistic of interest is therefore $(\alpha_1 + \alpha_2)$.

We define the following classification :[21]

- No preference : α_1 and α_2 are not jointly significant.
- Preference for variety : $\alpha_1 < 0$, $\alpha_2 > 0$ and $\alpha_1 + \alpha_2 = 0$.
- Preference for boys : $\alpha_1 + \alpha_2 < 0$.
- Preference for girls : $\alpha_1 + \alpha_2 > 0$.

3.3 Identification assumptions

The main threat to identification is the prevalence of child mortality. In our sample, 15.7% of children died before turning 5 years old. In this context, when we analyse fertility choices, shall we consider the gender composition of the previous births or the gender composition of children alive at the time of the decision ? There is a trade-off between exogeneity and relevance. The composition that matters to parents is probably among children who survived ; but it is correlated to parents' choices regarding breastfeeding, nutrition, caring and health seeking behaviour. After all, the proportion of sons among survivors is also an outcome of gender preferences.

That is why we consider the gender composition of the first births in our empirical

21. We do not consider the case $\alpha_1 > 0$, $\alpha_2 < 0$ and $\alpha_1 + \alpha_2 = 0$ because we never observe it in our estimations.

strategy. The first key identification assumption is that there is no sex-selective abortion. We believe that it is likely to hold because sex ratio at birth in our sample is equal to 51.2%, which is the ratio observed in Western countries (Brian and Jaisson (2007), Ben-Porath and Welch (1976) in the US, Jacobsen, Moller, and Mouritsen (1999) in Denmark) and generally considered as the natural level. Also, abortions are extremely rare in Africa. Our strategy is close to an instrumental variable framework : we use the composition among births as an instrument for the composition among living children, and estimate the reduced form. [22]

The second identification assumption is that there is no sex-selective child mortality. [23] Otherwise, the coefficients in the reduced form capture both the reaction to the death of a child and the « true » impact of gender composition on the next birth. Consider, for instance, a population with no strategy in terms of fertility, and in which boys tend to die more than girls. Families with more sons at birth would be more likely to have lost one child. If parents intensify fertility after the death of a child, we would observe that families with more sons have shorter birth intervals, and deduce the existence of daughter preference. Note that such a conclusion would be driven by the female advantage in infant mortality, which is consistent with daughter preference. The error would be to infer the existence of gender preferences from fertility choices while the result of the test would be driven by sex-selective mortality. In our sample, the under-five mortality rate is larger for boys than for girls : 16.4% vs. 15.1%. [24] Consequently, if there was a mortality bias, we would tend to underestimate son preference and to overestimate daughter preference everywhere. So if we find evidence of son preference, it has to be driven by fertility choices. Sex-selective mortality alone could explain

22. Note that we cannot apply a 2SLS procedure because the outcome does not depend linearly on the instrumented variable.

23. Sex-selective adult mortality could also bias our estimates if parents form beliefs about the survival probability of their sons and daughters at adult age, and take fertility decisions according to these beliefs. However, qualitative evidence provided by demographers do not support the idea that people make such calculations about child loss (Randall and LeGrand 2003).

24. The female advantage in child mortality is observed in most countries in the world, with the notable exception of India and China. The world ratio of male to female under-five mortality rate is around 120 (United Nations 2011).

our results only when we conclude that daughter preference prevails.

The next question is whether the variation across countries arises mainly from the variation in sex-selective mortality. Consider two countries with the same fertility strategies, one with a female advantage in infant mortality, the other with no selective mortality. The mortality bias, and hence the underestimation of son preference, would be larger in the first country ; we would conclude that son preference is weaker than in the second country. Again, this is consistent with the fact that girls seem to be given more advantages in the first country. In robustness tests (Section 5.1), we try to get an order of magnitude of the mortality bias by focusing on parents who lost no child.

Another threat to our strategy is that the prevalence of maternal mortality could lead to sample selection. In particular, if mothers exhibiting specific gender preferences are more likely to die, older women would be selected regarding gender preferences, and our sample would not be representative of all mothers regarding gender preferences and fertility behavior. More precisely the kind of bias we have in mind is the following. In a son preference setting, mothers with more girls would have shorter birth intervals, which would increase their exposure to maternal mortality risk. Hence, when looking at older women, our sample would be biased by attrition and we will tend to underestimate son preference. Conversely, in a daughter preference setting, we would underestimate daughter preference. In a nutshell, we expect gender preferences to be underestimated if maternal mortality is correlated with gender preferences.[25] In our data, we do not have direct evidence of maternal mortality. However, we can get indirect evidence by looking at the sex ratio of the first born child. If maternal mortality selects women according to their gender preferences, this sex ratio should vary between younger and older women. In our sample, sex ratios of the first born child are around 0.51 in all countries. We do find that in some countries, this ratio increases (up to

25. Note that another mechanism could lead to the same bias : if mothers having preferences for sons are more likely to *forget* first-born girls who died in their first days of life than first-born boys, this sex-selective recall would also lead to an underestimation of gender preferences.

0.57 in Nigeria), or decreases (down to 0.45 in Sudan) for women older than 47 years old. These provide suggestive evidence that maternal mortality may be linked to gender preferences, and in different ways for different countries. We estimate an order of magnitude of the selection bias in Section 5.2.

4 Results

4.1 Comparative descriptive analysis

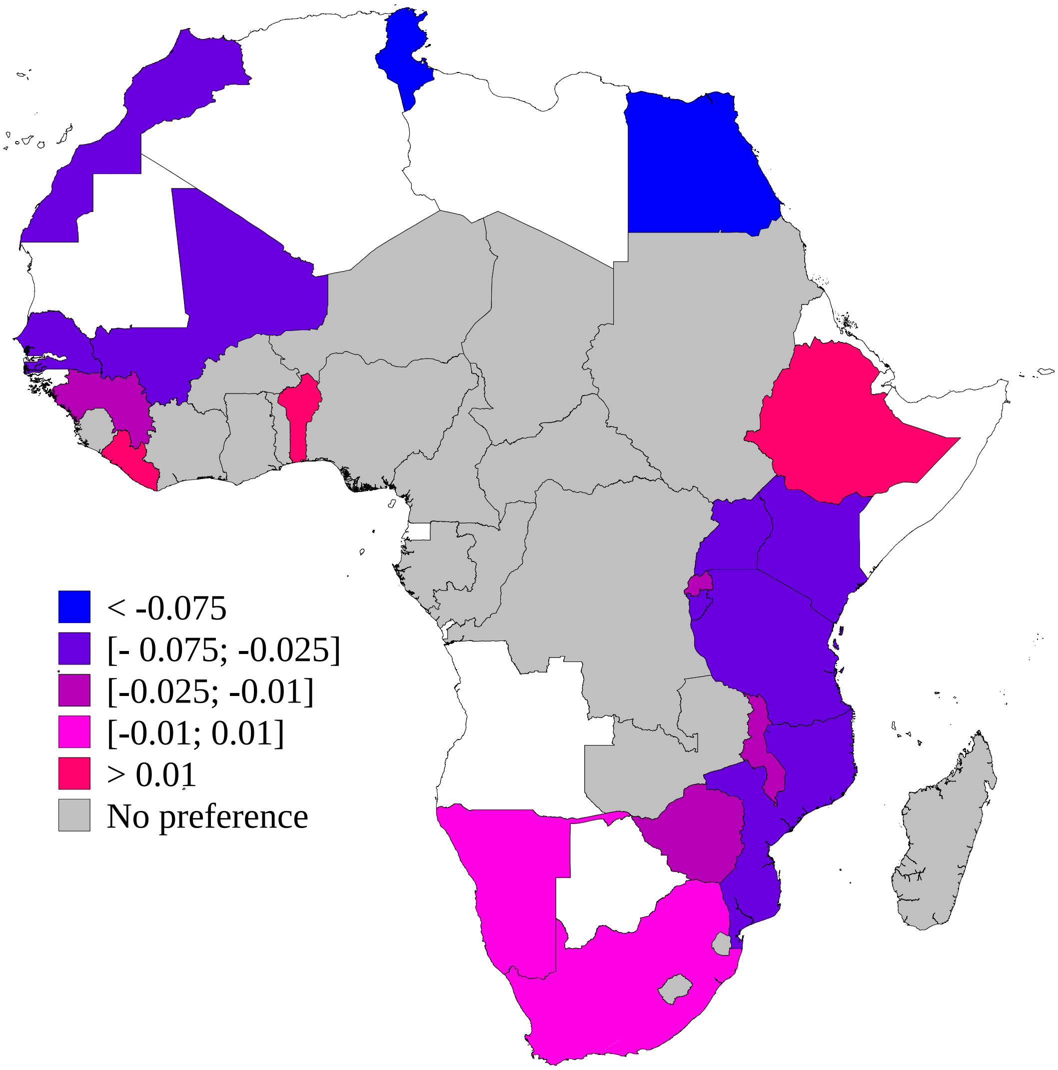
4.1.1 Heterogeneity over space

Figure 3 maps the magnitude of our indicator of gender preferences ($\alpha_1 + \alpha_2$) for countries in which α_1 and α_2 are jointly significant. Otherwise, countries are classified as « no preference ».[26] We find evidence of son preference in Northern Africa (Morocco, Tunisia, Egypt), Mali and Senegal, and also in the Great Lakes region (Burundi, Kenya, Uganda, Mozambique, Tanzania). Then, Southern Africa (Namibia and South Africa) is characterized by a preference for variety, and Central Africa (Cameroon, Chad, Congo, Congo DRC, Gabon, Central African Republic) by the absence of revealed gender preferences. In the rest of the continent, countries are divided into no gender preferences (Swaziland, Nigeria, Sierra Leone, Burkina Faso, Ghana, Togo, Niger, Sudan, Zambia, Lesotho and Madagascar, Cote d'Ivoire) and a taste for balance (Guinea, Liberia, Benin, Ethiopia, Rwanda, Malawi and Zimbabwe). No country displays daughter preference.

Such a simple sorting fails to give a sense of magnitude. When we predict the size of our coefficients α_1 and α_2 in a cross-country regression, the following patterns can be observed. The magnitude is stronger when the proportion of Muslims is higher, and when modern

26. The classification of countries can also be found in the Appendix in Table 7, in which we sort countries by our indicator of gender preferences ($\alpha_1 + \alpha_2$). We report the pvalue of the test $(\alpha_1 + \alpha_2) = 0$ and the pvalue of the test for joint significance of α_1 and α_2. The table is split in two between countries exhibiting gender preferences and countries with no preferences.

FIGURE 3: Gender preferences by country
Magnitude of our indicator $(\alpha_1 + \alpha_2)$



The map relates to our classification in the following way :

- No preference (grey) : α_1 and α_2 are not jointly significant
- Son preference (blue, dark purple) : $\alpha_1 + \alpha_2$ significantly negative. It is the case when $(\alpha_1 + \alpha_2) < -0.025$
- Preference for variety (purple, light purple, pink) : $\alpha_1 + \alpha_2$ not significantly different from zero
- Daughter preference (empty category) : $\alpha_1 + \alpha_2$ significantly positive

contraceptive methods are used more.[27] Then, we want to estimate by how much gender preferences impact fertility choices. We compute the predicted median birth spacing and the probability of short birth spacing ($\leq$ 24 months)[28] for (i) couples with no son ; (ii) couples with no daughter ; and (iii) couples having the optimal mix of sons and daughters.[29] Estimations are reported in Table 1 for countries in which we found evidence of gender preferences ; examining the magnitude is indeed meaningless when the proportion of sons has no significant impact on subsequent births.

Northern African countries stand out because they display the largest magnitudes : for instance, in Egypt, having no son is predicted to reduce the median birth spacing by 3 months as compared to having no daughter, and by 5 months as compared to the optimal mix of sons and daughters. Couples with no daughter, and couples with the optimal mix, are respectively 4.5 percentage points (a 13% decrease) and 7 percentage points (a 20% decrease) less likely to have short birth intervals than couples with no son. There, son preference has a strong impact on fertility patterns. Large magnitudes are also observed in South Africa. When the gender composition is perfectly balanced, couples are predicted to wait 7 months more than couples having only boys or only girls. The taste for balance therefore translates into sizeable differences between families. In the rest of Africa, gender preferences have a much weaker impact. Should they display preferences for boys or for variety, all countries exhibit very small differences in predicted birth spacing across our three categories of interest (0 or 1 month). In case of son preference, having only sons decreases by roughly 4% the probability of short birth spacing as compared to having only daughters.

27. Since contraceptive use is correlated with wealth and total fertility, we also find that the magnitude of preferences increases in wealth and decreases in the number of children per woman.

28. In the Cox model, one can derive an estimate of the survival function $\widehat{S}(t)$ (Box-Steffensmeier and Jones, 2004). The predicted median birth spacing is τ s.t. $\widehat{S}(\tau) = 0.5$ and the probability of short birth spacing is $\widehat{S}(24)$.

29. The optimal mix is the fraction of sons corresponding to the lowest hazard rate ; it is equal to $-\frac{\alpha_1}{2\alpha_2}$.

Table 1: Magnitude of gender preferences across countries

	Predicted median birth spacing (in months)			Probability of short intervals ($\leq$ 24 months)		
	No son	No daughter	Optimal mix	No son	No daughter	Optimal mix
Countries displaying son preference						
Egypt	29	32	34	35.4%	30.9%	28.5%
Tunisia	27	28	31	39.5%	36.5%	32.5%
Burundi	30	30	30	28.7%	27.1%	27.1%
Kenya	28	29	29	33.6%	32.1%	31.4%
Uganda	26	27	27	37.8%	36.2%	36.0%
Mali	27	27	28	35.9%	34.5%	33.8%
Mozambique	33	33	33	23.7%	22.7%	22.5%
Morocco	24	25	26	46.6%	45.0%	41.8%
Tanzania	33	33	33	22.5%	21.8%	21.4%
Senegal	30	31	31	27.7%	26.8%	26.8%
Countries displaying preferences for variety						
Zimbabwe	35	36	36	18.8%	18.4%	17.8%
Rwanda	28	28	29	34.0%	33.4%	32.1%
Guinea	33	33	34	22.9%	22.6%	20.7%
Malawi	33	33	34	22.1%	21.9%	21.3%
Namibia	35	35	36	23.1%	23.0%	21.9%
South Africa	58	58	65	13.5%	13.6%	12.2%
Benin	33	33	34	22.0%	22.3%	20.9%
Ethiopia	30	30	31	29.4%	29.7%	28.7%
Liberia	26	25	26	41.6%	42.4%	39.9%

We use the survival function and the parameters estimated in the Cox model. The optimal mix is the proportion of sons corresponding to the lowest predicted hazard rate ; it is equal to $-\frac{\alpha_1}{2\alpha_2}$.
The predicted median birth spacing and the probability of short birth spacing are computed for the median individual in each country.
We sort the countries according to the magnitude of our indicator of gender preferences ($\alpha_1 + \alpha_2$) (cf. Table 7 in Appendix). We do not provide estimations for countries displaying no gender preferences, because the question of the magnitude makes no sense in this case.

4.1.2 Heterogeneity over time

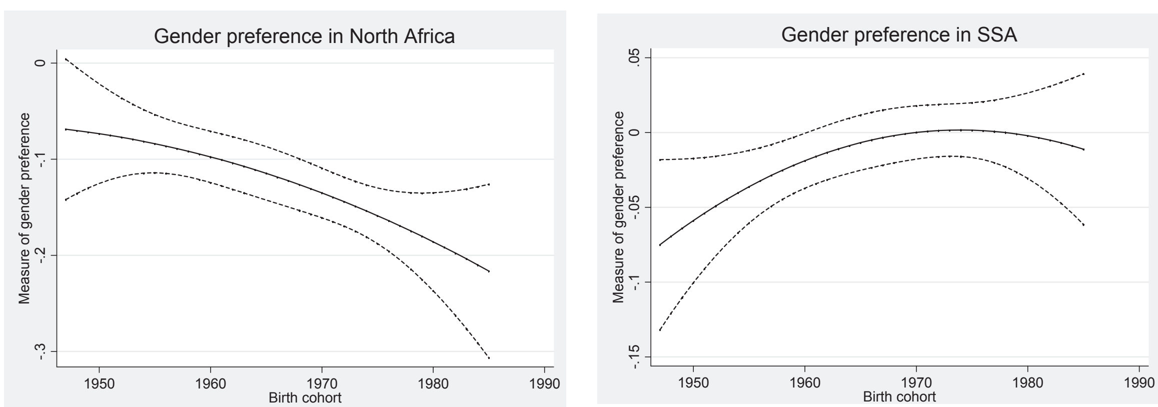
To study the evolution of gender preferences over time, we interact our variables of interest with the mother's birth cohort in the general model :

$$\begin{aligned}\lambda(t) &= \lambda_0(t) \times \exp(\lambda_1.cohort + \lambda_2.cohort^2 + \zeta_1.Frac_n + \zeta_2.Frac_n^2 + \phi_1.cohort.Frac_n + \\ &\quad \phi_2.cohort^2.Frac_n + \omega_1.cohort.Frac_n^2 + \omega_2.cohort^2.Frac_n^2 + \theta.X_n + \kappa.C)\end{aligned}$$

We pool all the surveys together, adding a vector of dummies for each country (C) to our main specification. From the estimates, we compute an indicator of cohort-by-cohort gender preference (our usual $\alpha_1 + \alpha_2$) :

$$\widehat{Pref} = \widehat{\zeta_1} + \widehat{\zeta_2} + \widehat{\phi_1}.cohort + \widehat{\phi_2}.cohort^2 + \widehat{\omega_1}.cohort + \widehat{\omega_2}.cohort^2$$

FIGURE 4: The evolution of $(\alpha_1 + \alpha_2)$ over time



Order 2 polynomial of mother's birth cohort with 5% confidence intervals. Cox estimation. Standard errors clustered at the mother level. Weights. Controls : country, rank of preceding birth, and mother's characteristics : age at preceding birth, religion, family system, union type, education, wealth, area of residence, employment status.

In Figure 4 we plot $\widehat{Pref}$ against the birth cohort, together with its 5% confidence intervals, in Northern Africa and in Sub-Saharan Africa. Our indicator is more and more negative over time in Northern Africa, meaning that son preference increases. When we break down this trend by country, we find that it is mainly driven by Egypt, especially in the most recent years. As explained in Section 2.1, we have few mothers born after 1970 in Morocco and Tunisia. Before that date, the trend observed in Morocco is similar to the Egyptian one, whereas it is rather flat in Tunisia – if anything, son preference tends to decrease.

In Sub-Saharan Africa, our indicator is increasing until 1970, and remains stable afterwards. Son preference prevails only among mothers born before 1960 ; from the 1960 cohort, $\widehat{Pref}$ is no longer significantly different from zero, and we can no longer infer the existence of son preference in Sub-Saharan Africa as a whole.

One way to explain why both trends differ is that most recent cohorts in Northern Africa have started to experience a fertility transition that may intensify traditional gender preferences.

4.2 Key drivers of gender preferences in Sub-Saharan Africa

Now that we have put forward heterogenous gender preferences over space and time in Sub-Saharan Africa, we want to investigate the role of socio-economic factors in shaping these preferences at the micro level. We examine whether women, in a given country and cohort, but belonging to different categories (e.g. educated vs. non-educated, rural vs. urban, matrilineal vs. patrilineal etc.) display different preferences.

4.2.1 Empirical strategy

We introduce interaction terms in our specification :

$$\lambda(t) = \lambda_0(t) \times \exp(\gamma_0.Alter + \gamma_1.Frac_n + \gamma_2.Frac_n^2 + \delta_1.Frac_n.Alter + \delta_2.Frac_n^2.Alter + \theta.X_n + \kappa.C)$$

Where *Alter* is a dummy equal to 0 if women belong to the reference category, and 1 if they belong to the alternative category. Again, we pool all the surveys together, and we add country and cohort dummies. Then, for each category, we compute our indicator of gender preference (using the notations of the general specification, it corresponds to $\alpha_1 + \alpha_2$) :

- Reference category : we test if $\gamma_1 + \gamma_2 = 0$
- Alternative category : we test if $\gamma_1 + \gamma_2 + \delta_1 + \delta_2 = 0$

We can conclude that gender preferences are different between the two categories if $\delta_1 + \delta_2$ is significantly different from 0.

4.2.2 Are religion and family systems shaping preferences ?

We start by considering two structural factors that appeared in the literature review as potential drivers of gender preferences : Islamic influence and traditional kinship structure. The first hypothesis we want to test is whether son preference is stronger among Muslims. The depth of Islamic penetration is indeed one way to explain the difference between Northern and Sub-Saharan African countries. The next question is whether, in a given country, Muslims and other religious groups have different preferences. In the first column of Table 2, we show that in Sub-Saharan Africa, Muslims exhibit the same taste for balance as Christians and animists. Coefficients on the interaction terms are not significant, and the indicator of gender preferences is the same in both categories. So religion may play a role at the macro level, by influencing family law, property rights, social norms etc., but in a given institutional setting, it does not seem to drive individual gender preferences.

The second hypothesis is that son preference would prevail in patrilineal ethnic groups, while daughter preference should be observed in matrilineal groups. As shown in Table 2, column 2, this hypothesis is validated : we find that our indicator of gender preferences is negative in patrilineal groups, which means son preference ; whereas it is positive in matrilineal groups, although not significant. Given the small proportion of the sample belonging to a matrilineal group (5.7%), we lack some power to take a definitive stance, but matrilineal groups seem to exhibit daughter preference. In any case, their preferences are significantly different from the ones in patrilineal

Table 2: Testing the impact of religion and family system in Sub-Saharan Africa

Reference category Alternative category	Christians and animists Muslims	Patrilineal Matrilineal	Monogamous Polygamous
$frac(\gamma_1)$	-0.142***	-0.167***	-0.203***
	(0.032)	(0.027)	(0.032)
$frac^2(\gamma_2)$	0.130***	0.155***	0.189***
	(0.030)	(0.026)	(0.031)
$frac \times Alter(\delta_1)$	-0.023	0.239***	0.223***
	(0.046)	(0.060)	(0.053)
$frac^2 \times Alter(\delta_2)$	0.028	-0.207***	-0.207***
	(0.042)	(0.056)	(0.050)
$(\gamma_1 + \gamma_2)$	-0.012	-0.011*	-0.015*
$(\gamma_1 + \gamma_2 + \delta_1 + \delta_2)$	-0.008	0.021	0.001
Pvalue test $(\delta_1 + \delta_2) = 0$	0.77	0.10	0.36
Observations	2590316	2694690	2246740

Dependent variable : duration between births n and $(n+1)$. $frac$: proportion of boys among the first n children. $Alter$: dummy for alternative category. Cox estimation, no hazard ratio. Standard errors clustered at the mother level. ***Significant at the 1 percent level. **Significant at the 5 percent level. *Significant at the 10 percent level. Weights. Controls : country, rank of preceding birth, and mother's characteristics : birth cohort, age at preceding birth, religion, family system, union type, education, wealth, area of residence, employment status.

groups.

When we further split the sample on the median wealth index, we find that our result on kinship structure is driven by relatively rich people. In the poorest half of the sample, couples in patrilineal and matrilineal groups exhibit the same preferences. One interpretation might be that inheritance rules impact gender preferences only when families have enough assets to bequest.

In the third column of Table 2, we test if gender preferences differ across union types : polygamous vs. monogamous. We find son preference in monogamous unions, but no evidence of gender preferences in polygamous unions. For them, the coefficients on $Frac$ and $Frac^2$ are very close to zero. Given the caveat that we mentioned in the introduction, this result is not surprising : our unitary model is well-suited to monogamous households, but it is less adequate to describe choices in more complex household structures.

4.2.3 How do preferences relate to modernization ?

Now, we turn to the modernization hypothesis, and examine individual indicators of development (see Table 3). More precisely, we compare the poorest half and the richest half of the sample (column 1) ; non-working and working women (column 2) ; women who do and do not use contraception (column 3) ; non-educated and educated women (column 4) ; rural and urban women (column 5) ; each time controlling for all other socioeconomic variables.

First, we find that wealth intensifies gender preferences, but has no impact on the preferred proportion of sons. In the richest half of the sample, coefficients on $Frac$ and $Frac^2$ are both significantly larger in absolute terms than in the poorest half ; but the indicator of preferences for boys vs. girls is the same and reveals preferences for variety. Wealth seems to strengthen the taste for balance, modifying the *magnitude*, but not the *nature* of gender preferences.

Our second result is that son preference is very strong for women who do not work, whereas working mothers exhibit preferences for variety. The difference between the two categories is significant. One explanation emphasizes insurance motives for the mother : non-working women are heavily dependent on their husband, and in case of widowhood, on their sons. The prevalence of son preference among non-working mothers also holds in matrilineal groups, suggesting that the

Table 3: Testing the impact of modernization factors in Sub-Saharan Africa

Reference category Alternative category	Poorest half Richest half	Not working Working	No contraception Contraception	No education Education	Rural Urban
$frac(\gamma_1)$	-0.101***	-0.129**	0.000	-0.137***	-0.173***
	(0.032)	(0.056)	(0.029)	(0.033)	(0.029)
$frac^2(\gamma_2)$	0.093***	0.092**	-0.010	0.129***	0.164***
	(0.031)	(0.047)	(0.027)	(0.031)	(0.028)
$frac \times Alter(\delta_1)$	-0.193***	-0.047	-0.952***	-0.042	0.061
	(0.048)	(0.077)	(0.052)	(0.043)	(0.048)
$frac^2 \times Alter(\delta_2)$	0.191***	0.090	0.945***	0.037	-0.065
	(0.045)	(0.059)	(0.048)	(0.040)	(0.045)
$(\gamma_1 + \gamma_2)$	-0.008	-0.037**	-0.010	-0.007	-0.009
$(\gamma_1 + \gamma_2 + \delta_1 + \delta_2)$	-0.010	0.005	-0.017	-0.012	-0.013
Pvalue test $(\delta_1 + \delta_2) = 0$	0.89	0.08	0.63	0.69	0.81
Observations	2421880	1907878	2694678	2693119	2694690

Dependent variable : duration between births n and $(n+1)$. $frac$: proportion of boys among the first n children. $Alter$: dummy for alternative category. Cox estimation, no hazard ratio. Standard errors clustered at the mother level. ***Significant at the 1 percent level. **Significant at the 5 percent level. *Significant at the 10 percent level. Weights. Controls : country, rank of preceding birth, and mother's characteristics : birth cohort, age at preceding birth, religion, family system, union type, education, wealth, area of residence, employment status.

insurance motive might prevail over the lineage motive.

Then, the correlation with contraceptive use is in line with expectations. Here, the variable *Contraception* is equal to one if the woman uses a modern method of contraception at the time of the survey. It does not imply that she was using it during her whole birth history, but it is a good proxy for how much control she has over fertility choices. For women who do not use modern contraceptives, the coefficients on $Frac$ and $Frac^2$ are close to zero : the gender composition of the first children does not influence subsequent birth intervals. Either they have no gender preferences, or they lack control over fertility to translate their preferences into actions.

Last, we find no correlation between gender preferences and either education or area of residence.

5 Robustness Tests

5.1 Testing the child mortality bias

As mentioned in Section 3.3, we test if sex-selective child mortality introduces a bias in our estimates. The idea is to isolate couples who lost at least one child among the first n births, and to focus on couples who lost no child. If we find evidence of gender preferences for the latter, they cannot be driven by differential mortality, they have to be driven by differential fertility rules. In Table 8 in Appendix, we interact $Frac$ and $Frac^2$ with a dummy equal to one if at least one of the first n children died. As expected, we find that the sex-selective mortality leads us to underestimate the extent of son preference in Africa : our indicator $(\alpha_1 + \alpha_2)$ is more negative among women who did not lose any children (column 2) than in the baseline (column 1). However, the difference is small, around 10%. A further result is that couples who lost at least one child shorten subsequent birth spacing, and shortening is the strongest when siblings were of both gender, and then stronger when siblings were all boys than when they were all girls. So there is evidence of son preference among those who lost a child as well.

Then, we check if our baseline classification of countries into the « son preference group » and the « preference for variety group » remains valid.[30] In Table 9 in Appendix, we find, again, that focusing on couples who lost no child slightly shifts our results towards more son preference. However, $(\alpha_1 + \alpha_2)$ remains not significantly different from 0 in the group classified as « preferences for variety », implying that our findings are robust.

To better understand the magnitude of the bias, we compute the predicted median birth spacing and the probability of stopping for couples who did not lose a child. In absolute values, birth spacing increases by one month, and the probability of stopping by roughly one percentage point, compared to the magnitudes discussed in section 4.1. But in relative terms, the differences across categories remain very stable.

The last test is to look how our sorting of countries is affected by mortality. Going on with same specification, we compute $(\alpha_1 + \alpha_2)$ among people who lost no child in each country, and compare it to our baseline $(\alpha_1 + \alpha_2)$. The correlation between both indicators is 0.97. Then, we sort the countries according to each indicator, and the rank correlation is equal to 0.95. In the end, the variation in sex-selective mortality across countries does not seem to drive the variation we observe.[31]

5.2 Testing the sample selection of mothers

We further deal with a potential selection bias. Our strategy is to run our model on mothers aged less than 40 who are less likely to have died or forgotten their first-born children.[32] For this sample of younger mothers, we find that the gender of the first born is exogenous to socioeconomic characteristics, which supports the assumption that the sample is not selected. We find that the correlation between the measure computed on younger mothers and our baseline measure is 0.94. Furthermore, the rank correlation between both classifications is 0.88. These strong correlations are reassuring regarding the fact that our results are not driven by differences between countries in maternal mortality or recall bias. Last, we checked, as in Table 9, that our classification of countries

30. See section 4.1 for the listing of countries.
31. Another piece of evidence is given by the low rank correlation (0.16) between our sorting and the sorting of countries on the child mortality ratio between girls and boys.
32. Results not shown, but available on request.

into « preference for son » and « preference for variety » remains robust once we restrict our sample to younger mothers.

One can argue that if maternal mortality is very high *and* affect women before age 40, our strategy is not enough to account for the selection bias. In particular, one can fear that when we find no gender preferences in some countries, it might be due to the fact that women having stronger son preference massively died before 40 years old. Let us consider simple back of the envelope calculations to get an order of magnitude of such a downward bias. In countries classified as « no preference », let us assume that there are in fact two groups of women. The first one, accounting for $M\%$ of the population, has very strong son preference ; we attribute to them the largest magnitude found in our sample ($\alpha_1+\alpha_2 = -0.17$ in Egypt). The second group has no gender preference ; for them, $\alpha_1 = \alpha_2 = 0$. The scenario that would lead to the most extreme selection bias is that all women having a preference for boys die, while all women having no preference survive. In this case, M represents the mortality rate. Following our baseline strategy, we would only observe surviving women and compute an indicator of gender preferences equal to 0, whereas the true indicator for the whole population would be around $-0.17 \times M$.[33] These countries could therefore reach the lowest magnitude of son preference observed in our sample ($\alpha_1 + \alpha_2 = -0.04$ in Senegal) if they had a maternal mortality rate at least equal to $M = \frac{0.04}{0.17} = 23.5\%$. It amounts to a lifetime risk of 1 out of 4, a magnitude never reached in Sub-Saharan Africa.[34] To conclude, maternal mortality would need to reach unlikely high levels in order to drive our « no preference » results.

5.3 Investigating heterogenous effects across ranks of birth

In our model, we decided to pool all ranks of birth together in order to have more power and to build a single indicator of gender preferences. Now, we investigate if gender preferences have an

33. Under the technical assumption that the weighted average of our indicator in both groups is a good proxy for the global indicator.

34. By comparison, it is almost ten times higher than the average risk in Sub-Saharan Africa in 2013 (1 out of 38), and four times larger than the largest risk (1 out of 15 in Chad) (WHO, UNICEF, UNFPA and The World Bank 2014).

heterogenous effect across ranks of birth. We estimate a model similar to Section 4.1.2 : we interact our variables of interest with the child's rank of birth, and we compute an indicator of rank-by-rank gender preference. In Figure 5, in Appendix, we plot this indicator together with its 5% confidence intervals in Northern Africa and in Sub-Saharan Africa. Both graphs display a U-shape, meaning that son preference is the strongest between ranks three and six. At lower and higher ranks, it is much weaker in Northern Africa, and tends to disappear in Sub-Saharan Africa.

There are two non-mutually exclusive ways to explain this pattern. First, the intensity of gender preferences, for the same couple, may vary across birth orders. It could be low at lowest ranks, because parents still have time for other tries in the future. Then the intensity could increase as time passes by, and parents start worrying about the eventual gender composition of their children. Last, the trend could revert at highest ranks if parents can already count on their eldest children. The second explanation is that gender preferences may be heterogenous across couples depending on the family size. Such a U-shape is consistent with (i) preference for variety in small size families, (ii) son preference in middle size families, and (iii) no preference in large size families. [35]

These results raise some concern about the implicit assumption in our model that birth orders have a multiplicative effect on the baseline hazard. To remove any doubt, we estimated our model rank by rank, for each country. We retrieved coefficients α_1 and α_2 for each rank, and we computed a weighted average, taking into account the number of observations at each rank. We were thus able to build a new indicator of gender preferences for each country, and we checked that it was strongly correlated to our baseline indicator (0.86). The restriction we made on birth order effects in our main specification does not change qualitatively our results.

35. When we interact $Frac$ and $Frac^2$ with family size instead of birth order (using the subsample of women over 40 years old), we also find a U-shape, which provides support for this explanation.

5.4 In which countries does the gender composition of the first two children predict subsequent fertility choices ?

One extension of our analysis is to consider the gender composition of the first children as an instrument for future fertility choices in order to estimate the impact of fertility on another outcome. In which African countries are gender preferences strong enough for the first stage to hold ? We examine the situation after the second birth. Since $Frac_n$ is not at all continuous when $n = 2$, we estimate an alternative proportional hazard model :

$$\lambda(t) = \lambda_0(t) \times \exp(\beta_{girls}.Girls + \beta_{boys}.Boys + \theta.X_2)$$

Where $Girls$ is a dummy equal to 1 if the first two children are girls, $Boys$ is a dummy equal to 1 if the first two children are boys. We do not reweigh the observations here, because the unit of observation is the mother. Robust standard errors are clustered at the finest geographical level defined in DHS (DHS cluster). Women having exactly 1 boy and 1 girl are the reference category. If $\beta_{girls} > 0$ and $\beta_{boys} > 0$, there is evidence of a taste for balance, because the lowest hazard rate – hence the longest expected interval before the third birth – is reached by women having children of both sex. Then, from the relative values of β_{boys} and β_{girls}, we can infer the existence of son or daughter preference.

Not surprisingly, the gender composition of the first two children is a strong predictor of next birth spacing in Northern Africa : Morocco, Tunisia and Egypt systematically display a strong son preference. Couples in Mali, Nigeria and Zimbabwe also wait significantly less before the third birth when they have at first two daughters vs. two sons. Then, the instrument would work in Benin, Guinea, Rwanda, Malawi and South Africa : they exhibit preferences for variety. Interestingly, some countries display daughter preference after the second birth : Ethiopia, Namibia, and Cote d'Ivoire. But in the vast majority of countries, the gender composition of the first two children does not influence the duration before third birth.

How to explain the discrepancy between the classification mentioned above and the one illustrated in Figure 3 ? When we use only intervals after the second birth, we lose some power to detect small magnitudes, as compared to the specification exploiting all parities. Mechanically, there are more countries in which we find no evidence of gender preferences at rank 2. Furthermore, as shown in section 5.3, the impact of gender preferences is weaker at rank 2 than at higher ranks. However, two specific cases are worth mentioning : in Nigeria and Cote d'Ivoire, gender preferences (respectively, son preference and daughter preference) are significant at rank 2 but disappear at higher ranks.

5.5 Evidence from differential stopping rule

Can we infer the same type of gender preferences by examining only stopping rules, and not spacing rules ? To avoid the issue of censoring, we restrict our sample to women over 40 years old, and we consider a linear probability model predicting the probability to have another child as a function of the proportion of sons and the proportion squared ; we include the same controls as in the duration model : a dummy for each rank, and mother's characteristics. We sort the countries according to the same criteria we used in the duration model. [36] We find evidence of son preference in Egypt, Mozambique and Gabon, daughter preference in Sudan and preference for variety in Congo DRC, Guinea, Kenya, Morocco, Namibia and Ethiopia. But in the vast majority of countries, we cannot detect any differential stopping rule depending on the gender composition of the first children.

5.6 At least one son or more sons ?

One may wonder if son preference means that parents want many sons, or that they want at least one son. In the literature, this interpretation has been called the « funeral pyre » hypothesis, in reference to the Hindu tradition that requires a son for lighting the funeral pyre of the parents (Arnold, Choe, and Roy. 1998). Similarly, preferences for variety may refer to a preference for at least one child of each gender, or to a taste for a perfect balance in the gender composition of

36. Results not shown, but available upon request.

children. In a low fertility setting, both interpretations generally coincide. But in our context, it is worth examining if there are gender preferences above and beyond « having at least one son and/or one daughter ».

In Table 10 in Appendix, we test if women having at least one child of each gender keep displaying gender preferences. The answer is positive : the quadratic relationship between the proportion of sons and the hazard rate still holds. Interestingly, coefficients are even larger in absolute values, meaning that the magnitude of gender preferences is stronger in the sub-population having at least one son and one daughter. A tentative explanation is that some people would believe that the probability to give birth to a boy vs. a girl might vary across couples, and they may try to infer their own probability from past outcomes. Couples having only sons or only daughters in the past may therefore believe that they would never have a child of the other sex, and stop accelerating births for another try.

In the end, in such a high fertility setting, son preference means that parents want more sons than daughters; and preferences for variety mean that they want the same number of boys and girls.

6 Conclusion

All in all, we find robust evidence that son preference influences fertility patterns in Northern Africa : people tend to shorten birth spacing and to have additional children as long as they have not had enough sons. This has strong implications for gender inequality : an average girl would be weaned sooner, and would face more competition from her siblings, than an average boy. Moreover, women, as mothers, would put their own lives in jeopardy to ensure that enough sons are born.

We cannot draw the same conclusion for Sub-Saharan Africa : there is some evidence of son or daughter preference in given countries, but our results are not always robust, and in any case, of small magnitude. If anything, fertility behavior is rather consistent with a taste for balance in the gender composition of children, especially in South Africa. So that globally, the impact of gender preferences on fertility patterns does not seem substantial enough to induce gender inequality.

But there are plenty of mechanisms by which gender preferences prevailing in a society may translate into inequality, beginning with family law and property rights. On many dimensions, it might be argued that women have a subordinate status in Sub-Saharan Africa. Anderson and Ray (2010) show that, in this region, there are « missing women », too. But contrary to India and China, they are in majority of adult age ; HIV/AIDS and maternal deaths are the two main sources of female excess mortality. In the same vein, our results suggest that there is no strong discrimination against female foetus or little girls. The discrimination against women would appear later in life : at puberty ? at marriage ? at motherhood ? at widowhood ? Understanding when and why remains an open question.

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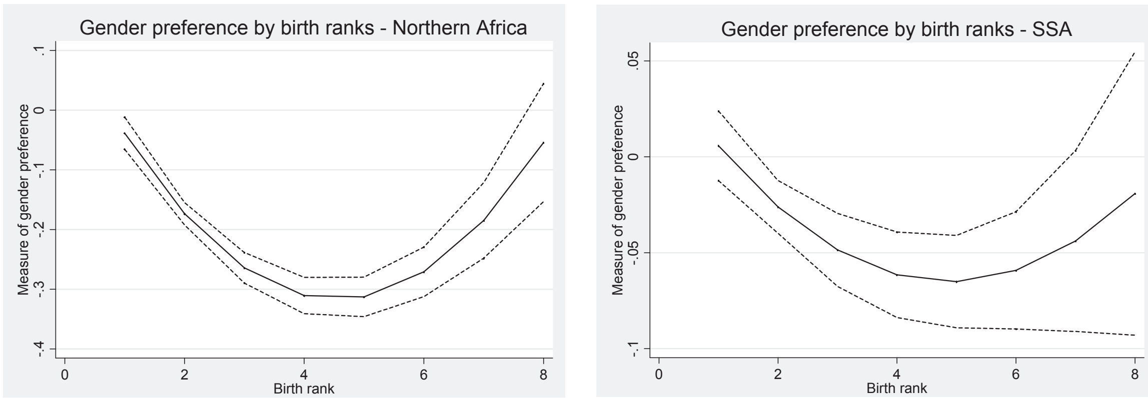
Appendix : For online publication

Figure 5: $(\alpha_1 + \alpha_2)$ by birth ranks



Order 2 polynomial of child's birth rank with 5% confidence intervals. Cox estimation. Standard errors clustered at the mother level. Weights. Controls : country and mother's characteristics : birth cohort, age at preceding birth, religion, family system, union type, education, wealth, area of residence, employment status.

Table 4: Survey waves and years

Country	Survey years	Nb births	Nb women
Benin	1996, 2001, 2006, 2011	143,141	35,210
Burkina Faso	1992, 1998, 2003, 2010	140,428	32,396
Burundi	1987, 2010	36,406	8,731
Congo Dem. Rep.	2007	29,548	7,148
Cameroon	1991, 1998, 2004, 2011	98,566	25,266
Central African Rep.	1994	16,936	4,388
Chad	1996, 2004	47,187	10,508
Rep. of the Congo	2005, 2011	48,635	13,339
Cote d'Ivoire	1994, 1998, 2011	60,656	15,654
Egypt	1988, 1992, 1995, 2000, 2005, 2008	294,830	76,897
Ethiopia	2000, 2005	129,113	30,263
Gabon	2000, 2012	39,987	10,882
Ghana	1988, 1993, 1998, 2003, 2008	67,676	17,748
Guinea	1999, 2005, 2012	77,741	18,622
Kenya	1989, 1993, 1998, 2003, 2008	117,031	28,606
Lesotho	2004, 2009	29,137	10,023
Liberia	1986, 2007	39,387	9,932
Madagascar	1992, 1997, 2003, 2008	109,847	28,417
Malawi	1992, 2000, 2004, 2010	164,935	41,394
Mali	1987, 1995, 2001, 2006	150,720	32,570
Morocco	1987, 1992, 2003	80,669	18,970
Mozambique	1997, 2003, 2011	101,179	27,154
Namibia	1992, 2000, 2006	47,840	15,126
Niger	1992, 1998, 2006, 2012	131,290	27,403
Nigeria	1990, 1999, 2003, 2008	179,246	40,960
Rwanda	1992, 2005, 2010	82,151	19,838
Senegal	1986, 1992, 1997, 2005, 2011	144,101	33,956
Sierra Leone	2008	21,136	5,876
South Africa	1998	22,934	8,223
Sudan	1989	25,805	5,277
Swaziland	2006	11,410	3,488
Tanzania	1991, 1996, 1999, 2004, 2010	126,319	30,830
Togo	1988, 1998	37,051	8,825
Tunisia	1988	16,463	3,856
Uganda	1988, 1995, 2000, 2006, 2011	120,935	27,339
Zambia	1992, 1996, 2001, 2007	92,092	22,423
Zimbabwe	1988, 1994, 1999, 2005, 2010	82,134	24,606

Table 5: Matrilineal ethnic groups : matching the Ethnographic Atlas and DHS data

Ethnic group in Gray (1998)	Country	Year of DHS survey	Code in DHS survey	Comment
Dorosie / Voltaic people	Burkina Faso			Not listed in DHS
Lobi / Voltaic people	Burkina Faso	1992, 1998, 2003, 2010	6	
Tuareg	Burkina Faso	1992, 1998, 2003, 2010	9	
Udalan / Plateu Nigerians - Chadic	Burkina Faso			Not listed in DHS
Fur / Darfur	Chad			Not listed in DHS
Sundi / Central Bantu, basundi	Congo	2005	9	
Teke / Northwestern Bantu, bateke	Congo	2005	22, 23	
Teke / Northwestern Bantu, bateke	Congo	2011	5	
Yombe / Central Bantu, bayombe	Congo	2005	12	
Bembas (Luapula, Lamba, Buye, Kaonde, Lala)	DRC			Definition of ethnic groups too broad
Kasai (Bunda, Yanzi, Dzing, Sakata, Kuba, Lele)	DRC		7	Definition of ethnic groups too broad
Kongos (Sundi, Yombe)	DRC			Definition of ethnic groups too broad
Kwango (Pende, Suku, Yaka)	DRC			Definition of ethnic groups too broad
Lunda (Luvale, Ndembu, Chokwe)	DRC		8	Definition of ethnic groups too broad
Teke / Northwestern Bantu	DRC			Definition of ethnic groups too broad
Wodaabe, Bororos	DRC			Definition of ethnic groups too broad
Mpongwe / Northwestern Bantu	Gabon			Not listed in DHS
Shogo / Northwestern Bantu	Gabon	2000, 2012	6	
Teke / Northwestern Bantu	Gabon	2000, 2012	3	
Twi Akan (Akyem, Anyi, Brong)	Ghana	1988	1,2,3	
Twi Akan (Akyem, Anyi, Brong)	Ghana	1993	1,2,3,4	
Twi Akan (Akyem, Anyi, Brong)	Ghana	2003, 2008	1	
Twi Lagoon (Assini)	Ghana	1988	1,2,3	
Voltaic people (Lobi, Kulango)	Ghana			Not listed in DHS
Tenda	Guinea			Not listed in DHS
Twi Akan (Baule, Anyi, Brong)	Ivory Coast	1998	1	
Twi Akan (Baule, Anyi, Brong)	Ivory Coast	2011	101-106, 108-111, 114	
Twi Lagoon (Avikam, Assini)	Ivory Coast			Not listed in DHS
Voltaic people (Lobi, Kulango)	Ivory Coast	1998	5	Ethnic group variable is missing in 1994 survey
Voltaic people (Lobi, Kulango)	Ivory Coast	2011	140, 146	Ethnic group variable is missing in 1994 survey
Chewas	Malawi	2000, 2004, 2010	1	
Nyanja / Maravi	Malawi	2010	12	
Nyasa / Maravi	Malawi			Not listed in DHS
Yao	Malawi	2000, 2004, 2010	5	
Antessar / Tuareg	Mali			Not listed in DHS
Udalan / Plateu Nigerians - Chadic	Mali			Not listed in DHS

Table continued

Ethnic group in Gray (1998)	Country	Year of DHS survey	Code in DHS survey	Comment
Chewas, chichewas	Mozambique	2003	10	
Chewas, chichewas	Mozambique	1997	14	
Chewas, chichewas	Mozambique	2011	11	
Kunda / Maravi	Mozambique			Not listed in DHS
Makonde / Yao	Mozambique	2011	15	
Makonde / Yao	Mozambique	1997	50	
Nyanja / Maravi	Mozambique	2003	26	
Nyanja / Maravi	Mozambique	1997	21	
Nyasa / Maravi	Mozambique			Not listed in DHS
Sena / Maravi, chisena	Mozambique	2011	4	
Sena / Maravi, chisena	Mozambique	2003	5	
Sena / Maravi, chisena	Mozambique	1997	25	
Yao	Mozambique	2011	12	
Ambo / Southwestern Bantu	Namibia	1992	3	
Ambo / Southwestern Bantu	Namibia	2000	7	
Ambo / Southwestern Bantu	Namibia	2006	7	
Tuareg (Azjer, Ahaggaren, Asben)	Niger	1992, 1998	8	Ethnic group variable is missing in 2012 survey
Tuareg (Azjer, Ahaggaren, Asben)	Niger	2006	7	Ethnic group variable is missing in 2012 survey
Udalan / Plateu Nigerians - Chadic	Niger			Not listed in DHS
Daka / Eastern Nigritic	Nigeria			Ethnic group variable is always missing
Kurama, Gure /Bantoid People	Nigeria			Ethnic group variable is always missing
Longuda / Eastern Nigritic	Nigeria			Ethnic group variable is always missing
Ndoro /Bantoid People	Nigeria			Ethnic group variable is always missing
Tenda	Senegal			Not listed in DHS
Sherbro / Ku and Peripheral Mande	Sierra Leone	2008	16	
Fur / Darfur	Sudan			Ethnic group variable is always missing
Nuba / Nubians (Midobi, Tumtum)	Sudan			Ethnic group variable is always missing
Makonde / Yao	Tanzania			Ethnic group variable is always missing
Nyasa / Maravi	Tanzania			Ethnic group variable is always missing
Ranji / Rift	Tanzania			Ethnic group variable is always missing
Sagara / Rufiji	Tanzania			Ethnic group variable is always missing
Zigula (Luguru, Nguru, Kwere)	Tanzania			Ethnic group variable is always missing
Bemba (Lamba, Lala, Kaonde, Luapula)	Zambia	1996, 2001, 2007	1, 2, 3, 26, 27	
Bemba (Lamba, Lala, Kaonde, Luapula)	Zambia	1992	1	
Lunda (Luvale, Ndembu, Luchazi)	Zambia			Not listed in DHS
Maravi (Chewa, Kunda, Nyanja)	Zambia	1996, 2001, 2007	48, 51, 52, 53	
Maravi (Chewa, Kunda, Nyanja)	Zambia	1992	5	
Tonga / Middle Zambesi Bantu	Zambia	1992	2	
Tonga / Middle Zambesi Bantu	Zambia	1996, 2001, 2007	19	
Tonga / Middle Zambesi Bantu	Zimbabwe			Ethnic group variable is always missing

Table 6: Number of children and birth spacing, by country

Country	**Number of children**	**Average birth intervals (in months)**	**Proportion of short intervals (≤ 24 months)**
Benin	6.3	35.5	26.7
Burkina Faso	7.0	34.8	25.4
Burundi	6.9	32.5	32.0
CDR	6.5	34.1	32.3
Cameroon	6.3	34.1	32.9
Centrafrique	6.2	32.9	35.2
Chad	7.2	31.2	35.2
Congo-Brazza	5.2	43.7	21.5
Cote d'ivoire	6.4	37.4	27.6
Egypt	5.3	33.3	40.3
Ethiopia	6.1	34.6	31.1
Gabon	5.5	41.1	29.9
Ghana	5.8	40.0	23.0
Guinee	6.4	37.0	23.5
Kenya	6.5	34.4	34.8
Lesotho	4.5	45.1	16.7
Liberia	6.2	37.2	32.2
Madagascar	6.2	33.8	38.3
Malawi	6.8	34.7	28.7
Mali	7.5	31.9	36.8
Maroc	6.2	33.7	40.1
Mozambique	6.0	36.3	28.2
Namibia	5.1	42.1	25.6
Niger	7.9	30.9	36.7
Nigeria	6.7	33.1	35.3
Rwanda	6.7	33.0	33.2
Senegal	6.7	34.3	30.6
Sierra Leone	5.6	38.6	28.1
South Africa	3.9	48.3	21.8
Sudan	7.6	29.1	45.5
Swaziland	5.5	41.7	26.2
Tanzania	6.6	36.1	26.8
Togo	6.7	35.9	24.8
Tunisia	6.3	29.4	46.9
Uganda	7.4	31.4	37.9
Zambia	7.1	34.4	28.2
Zimbabwe	5.6	40.7	21.3
Total	6.2	34.7	33.0

Weights. The number of children is computed on the sample of women over 40 years old

Table 7: Which African countries exhibit gender preferences ?

Country	$\alpha_1 + \alpha_2$	pvalue of the sum	pvalue joint significance
α_1 and α_2 are jointly significant			
Egypt	-0.168	0.000	0.000
Tunisia	-0.099	0.013	0.000
Burundi	-0.064	0.022	0.042
Kenya	-0.057	0.001	0.000
Uganda	-0.055	0.001	0.001
Mali	-0.050	0.005	0.000
Mozambique	-0.049	0.032	0.040
Morocco	-0.048	0.017	0.000
Tanzania	-0.038	0.047	0.015
Senegal	-0.038	0.018	0.045
Zimbabwe	-0.025	0.204	0.021
Rwanda	-0.022	0.228	0.000
Guinea	-0.017	0.369	0.000
Malawi	-0.014	0.389	0.032
Namibia	-0.002	0.934	0.097
South Africa	0.004	0.915	0.017
Benin	0.011	0.465	0.000
Ethiopia	0.012	0.516	0.100
Liberia	0.026	0.385	0.062
α_1 and α_2 are not jointly significant			
Swaziland	-0.051	0.288	0.366
Gabon	-0.019	0.576	0.754
Zambia	-0.017	0.338	0.547
Cameroon	-0.015	0.422	0.157
Nigeria	-0.012	0.500	0.243
Ghana	-0.009	0.652	0.673
Madagascar	-0.005	0.793	0.157
Burkina Faso	-0.002	0.923	0.798
Niger	0.002	0.887	0.569
Congo DRC	0.005	0.894	0.259
Congo	0.010	0.738	0.278
Sierra Leone	0.012	0.745	0.505
Sudan	0.017	0.638	0.713
Lesotho	0.017	0.588	0.682
Togo	0.035	0.226	0.221
Chad	0.035	0.218	0.125
Cote d'Ivoire	0.044	0.112	0.265
Central African Republic	0.055	0.161	0.303

Cox estimation. Weights. Controls : country, rank of preceding birth, and mother's characteristics : birth cohort, age at preceding birth, religion, family system, union type, education, wealth, area of residence, employment status.

Table 8: Testing the mortality bias

	Baseline	Test
$frac(\alpha_1)$	-0.234***	-0.321***
	(0.023)	(0.025)
$frac^2(\alpha_2)$	0.209***	0.293***
	(0.022)	(0.024)
dead child		0.009
		(0.016)
$frac \times$ dead child		0.257***
		(0.059)
$frac^2 \times$ dead child		-0.236***
		(0.054)
$(\alpha_1 + \alpha_2)$	-0.025***	-0.028***
Observations	3105217	3105217

Dependent variable : duration between births n and $(n+1)$. $frac$: proportion of boys among the first n children. dead child : dummy for at least one dead child among the first n children. Cox estimation, no hazard ratio. Standard errors clustered at the mother level. ***Significant at the 1 percent level. **Significant at the 5 percent level. *Significant at the 10 percent level. Weights. Controls : country, rank of preceding birth, and mother's characteristics : birth cohort, age at preceding birth, religion, family system, union type, education, wealth, area of residence, employment status.

Table 9: Testing the mortality bias – by groups of countries

	Preference for variety		Son preference	
	Baseline	Test	Baseline	Test
$frac(\alpha_1)$	-0.200***	-0.166***	-0.418***	-0.591***
	(0.050)	(0.055)	(0.023)	(0.025)
$frac^2(\alpha_2)$	0.203***	0.163***	0.340***	0.512***
	(0.048)	(0.053)	(0.022)	(0.024)
dead child		0.025		-0.035**
		(0.034)		(0.016)
$frac$ × dead child		-0.174		0.704***
		(0.126)		(0.062)
$frac^2$ × dead child		0.210*		-0.693***
		(0.111)		(0.058)
$(\alpha_1 + \alpha_2)$	0.003	-0.003	-0.078***	-0.079 ***
Observations	773577	773577	1167669	1167669

Dependent variable : duration between births n and $(n+1)$. $frac$: proportion of boys among the first n children. dead child : dummy for at least one dead child among the first n children. Cox estimation, no hazard ratio. Standard errors clustered at the mother level. ***Significant at the 1 percent level. **Significant at the 5 percent level. *Significant at the 10 percent level. Weights. Controls : country, rank of preceding birth, and mother's characteristics : birth cohort, age at preceding birth, religion, family system, union type, education, wealth, area of residence, employment status. Countries in group « Preference for variety » : Zimbabwe, Rwanda, Guinea, Malawi, Namibia, South Africa, Benin, Ethiopia and Liberia. Countries in group « Son preference » : Egypt, Tunisia, Burundi, Kenya, Uganda, Mali, Mozambique, Morocco, Tanzania and Senegal.

TABLE 10: Are there gender preferences beyond « at least one son and/or one daughter » ?

	Baseline	Isolating same-sex siblings
$frac(\alpha_1)$	-0.234***	
	(0.023)	
$frac^2(\alpha_2)$	0.209***	
	(0.022)	
$fraction(\lambda_1)$		-0.535***
		(0.076)
$fraction^2(\lambda_2)$		0.437***
		(0.075)
girls		-0.096***
		(0.020)
boys		-0.115***
		(0.020)
$(\alpha_1 + \alpha_2)$	-0.025***	
$(\lambda_1 + \lambda_2)$		-0.098***
Observations	3105217	3105217

Dependent variable : duration between births n and $(n + 1)$. $frac$: proportion of boys among the first n children. $girls$: first n children are all girls. $boys$: first n children are all boys. $fraction$ is equal to $frac$ iif $frac < 1$. $fraction$ and $fraction^2$ capture the impact of the gender composition among women having at least one child of each gender. Cox estimation, no hazard ratio. Standard errors clustered at the mother level. ***Significant at the 1 percent level. **Significant at the 5 percent level. *Significant at the 10 percent level. Weights. Controls : country, rank of preceding birth, and mother's characteristics : birth cohort, age at preceding birth, religion, family system, union type, education, wealth, area of residence, employment status.