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Abstract

Exploiting original data from a Senegalese household survey, we provide evidence that fertility choices are partly driven by women's needs for widowhood insurance. We use a duration model of birth intervals to show that women most at risk in case of widowhood intensify their fertility, shortening birth spacing, until they get a son. Insurance through sons might entail substantial health costs since short birth spacing raises maternal and infant mortality rates.

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In many developing countries, the lack of insurance markets and social safety nets is mitigated by family. A large literature describes the use of family as insurance, examining for instance family income diversification strategy through migration (Rosenzweig, 1988), or children acting as old-age insurance for their parents (Nugent, 1985 ; Hoddinott, 1992). These mechanisms have been thoroughly described, however there is very scarce evidence on the costs to the beneficiaries of those family-based insurance systems. The exchange is likely to be non-monetary and not strictly *quid pro quo* – facts that contribute to the costs remaining hidden. Exposing the existence and extent of these costs could temper the general optimism towards informal insurance mechanisms.

In this paper, we analyze such a family based insurance mechanism – one whereby women in Senegal rely on sons as insurance in case of widowhood. As detailed below, we exploit the variation in widows' vulnerability that arises from the rivalry for inheritance between the husband's children. When the husband already has children from ex-wives, whether divorced or deceased, the current wife's best way to secure future access to his resources – in particular his house – is to have a son.[1] Hence, we investigate whether fertility choices are partly driven by women's insurance strategies. Our hypothesis is that the existence of children of ex-wives (henceforth we call these children "rivals") should exacerbate son preference and intensify fertility. This hypothesis is supported by the raw data on birth intervals. Building on demographers' methodology, we use semi- and fully parametric estimations of a duration model of birth intervals to provide evidence that women with rivals and no son do indeed tend to have another child more quickly. We find that when there are rivals, having only girls versus having at least one son multiplies by 1.5 to 2 the probability of short birth spacing (intervals shorter than 24 months). In the absence of rivals, the gender composition of first births does not affect subsequent birth intervals. We show that the effect is clearly evidenced for women who co-reside with their husband, in particular for the poorest half of the population, justifying the prior that the house is the main stake in the bequest process. Also, fertility choices are more strongly affected after the third child.

We focus on the case of Senegal, but such a mechanism may theoretically be at play in other countries as soon as (i) widows run the risk of being dispossessed, and (ii) having sons could mitigate this risk. According to a UN report on widows worldwide, the first statement

1. Rivalry may also arise in polygamous unions between children of co-wives, but we choose not to consider them for two reasons. First, our identification strategy fails to deal with simultaneity issues : in polygamous unions, women might adjust the number of their children to react to a co-wife's births, so that the rivals of a given wife cannot be considered as exogenous to her own fertility choices. Second, it can be argued that co-wives have an alternative insurance strategy based on cooperation, since they may agree to continue to live together in the late husband's house (see qualitative evidence in Lambert and van de Walle, 2012).

holds throughout the developing world (United Nations, 2001). The identity of rivals for inheritance varies from one country to another : they might be husband's brothers, parents, nephews, or as in Senegal, children from previous unions. The second statement is more difficult to assess, but several studies report that having sons attenuates the risk of being dispossessed in various parts of the world ; e.g. in rural South Asia (Cain, 1986, Agarwal, 1994), in Indonesia (Carranza, 2012), in Nigeria (United Nations, 2001), in Uganda (Akiiki Asiimwe, 2009), in Zambia and Zimbabwe (Kevane, 2004).

In Senegal, women are highly exposed to the risk of losing their husbands, in particular because of the age difference between spouses : 10 years on average, and often much more. The question of the means available to women to cope with the impact of widowhood on living standards is a very urgent one. Formal mechanisms are weak : finance and insurance markets are very imperfect, social safety nets are missing, and opportunities for self-support are limited due to labor market restrictions and constraints stemming from social norms. Hence, widows are dependent on their access to their late husband's resources.

Furthermore, as 95 percent of the population is Muslim, inheritance practices are ruled by Islamic and customary patrilineal laws. In practice, wives are excluded from a bequest after the death of the husband ; the inheritance is to be shared among the husband's children, sons inheriting more, and more frequently, than daughters (Lambert, Ravallion and van de Walle, 2014). Children, especially sons, therefore turn out to be women's best claim to their late husbands' resources in case of widowhood. Any children the husband had with other women may compete for the man's inheritance. Therefore, the number and gender composition of the children of the current wife, relative to other children the husband might have, determine the share of inheritance she might control *de facto* ; it influences, in turn, her probability to carry on residing in the late husband's house. Women might therefore have incentives to be pregnant often in order to maximize the number of sons they have.

Demonstrating such a mechanism may matter if we are to better understand the costs of informal family insurance. Indeed, these costs are rarely studied ; until recently, it was assumed that taking part in family insurance networks has no cost other than reciprocity. Recent papers have focused on the monetary costs for relatively high-income members, and on the distortions generated by the "kin tax" in African economies (Jakiela and Ozier, 2015). The monetary costs are often considered negligible for low-income members. Non-monetary costs have been totally overlooked, and yet they might be important from a welfare perspective. The cost of widowhood insurance through sons is potentially huge : for both mother and child, health risks related to frequent pregnancies are well-documented in the medical literature on developing countries (e.g. Conde-Agudelo et al., 2006 ; Conde-Agudelo and Belizan, 2000 ; DaVanzo et al., 2004). Increasing the interval between births to 24 months is a goal for many family planning programs. [2] In the economics literature, Jayachandran and Kuziemko (2011) argue that Indian mothers breastfeed girls less than boys because they want to try again for a son quickly after the birth of a daughter. They show that son preference translates into a female disadvantage in child survival rates. In Nigeria and in India, Milazzo (2012 and 2013) finds that son preference induces short birth spacing, and might be a significant cause of female excess adult mortality.

In Senegal, maternal and infant mortality rates are a matter of concern : out of each 1,000 births, four mothers died as well as 47 children before turning a year old – 29 of those in the first month, over the period 2007-2011. Further, children born less than two years after a sibling are 2.4 times more likely to die than the others, whatever mortality rates we consider : neonatal, post-neonatal, infant, juvenile, infant-juvenile (ANSD and ICF International, 2012). [3] Despite important supply-side interventions, mortality rates remain high, in part because frequent pregnancies persist at a large scale. [4] Women's needs for widowhood insurance might help explaining why this practice is so difficult to deter.

Exploiting original data from a nationally representative household survey, we develop a strategy that is close to a difference-in-difference framework. We analyze how the gender composition of the first children influences subsequent choices of women with and without rivals. While the two groups might have other differences, what matters to our identification strategy is that (i) the gender of the children they give birth to is exogenous, and (ii) in the absence of rivals, both groups would have reacted similarly to this exogenous variable : having one son vs. one daughter. Using such a strategy to identify the existence of a widowhood insurance motive for fertility is very similar in spirit to Nugent's (1985) suggestion : in order to evidence old-age security motive, one has to compare similar populations with different intensity of the motive. Here, among women with rivals, we compare those having a son and those having only daughters ; they are similar, but the strength of their motivation is different. Then, we factor in women without rivals to cancel out any son preference that

2. Some programs such as the Optimal Birth Spacing Initiative, created under the auspices of USAID, even recommend an optimal birth spacing of three to five years.

3. Ronsmans (1996) shows that in rural Senegal, short birth intervals also increase the odds of dying for the preceding child : if the mother delivers another child within two years, the risk of mortality for the index child is four times higher in the second year of life than if the next birth takes place more than two years after his birth. However, this correlation is not necessarily causal : short subsequent birth intervals may be either a cause (abrupt weaning) or a consequence (willingness to "replace" a dead child) of child mortality.

4. Budget of the Health Ministry was multiplied by 3.5 between 2000 and 2010 and it now reaches 10,4 percent of national budget (ANSD and ICF International, 2012).

prevails in the population but is not related to widowhood insurance.

All in all, we find that son preference only appears in women with rivals, and especially when these women have no independent housing. Women's needs for widowhood insurance seem therefore to play a role in fertility choices, leading to more frequent pregnancies for the most exposed women : those with rivals and no son. This suggests that the lack of formal widowhood insurance hampers birth control efforts and, even more worryingly, imposes a potentially heavy health cost on both women and children, whether widowhood occurs or not. Such non-monetary costs underline the fact that, even if an informal insurance system plays a role in granting access to resources in case of widowhood, it is a poor substitute for a formal insurance scheme. The benefits of implementing some public widowhood insurance scheme, or changing inheritance laws, would thus be far reaching.

The outline of the paper is as follows. Section 1 provides background on widowhood and inheritance practices in Senegal. Section 2 presents the data and some descriptive statistics. Section 3 introduces a simple theoretical model to clarify hypotheses and mechanisms at play. Section 4 discusses the empirical strategy, the identification assumptions and the estimation methods. Section 5 reports the main results, and Section 6 concludes with policy implications.

1 Widowhood and Bequest in Senegal

1.1 An ex-ante risk-coping strategy based on children

An idea has emerged from several disciplines, that women's vulnerability in Sub-Saharan Africa is conducive to large family size. [5] Family law renders women dependent on spouses, or on adult children in case of death or abandonment by the husband. Women therefore have an interest in early marriage and high fertility. Anthropologists such as Bledsoe (1990) also endorse the idea that in all societies, people make efforts to construe marital status and parenthood to their advantages. She claims that in Africa, as long as children remain women's best claim to male resources, women will continue to want many children.

Demographers working on Sub-Saharan Africa have come to the same conclusion : among the key factors shaping the reproductive regime in this region, they mention inheritance patterns and women's subordinate status. In patrilineal societies, in which wives "belong" to the male kin, women's security and status critically depend on their ability to have sons (Lesthaeghe, 1989).

5. In Senegal, fertility is still high : five is both the average and median number of children among women older than 45 years (cf. Figure A.1 in Appendix A).

All this requires that fertility is an active choice of women. In Senegal, modern contraceptive methods are widely known, but not much used : over 92 percent of married women can name at least one modern method ; but only 12 percent – mostly urban, educated and wealthy women – report using one of them. Still, only 7.5 percent of married women report unmet needs for birth control to limit the number of children. Unmet needs for birth control to increase intervals between births are higher, though (22 percent). Another way to assess whether women do indeed have some control over their fertility is to look at the percentage of unwanted or unplanned pregnancies. Only four percent of pregnancies that occurred during the period 2005-2010 are reported by the mothers as unwanted and 20 percent as unplanned (ANSD and ICF International, 2012).

A majority of Senegalese women therefore claim to be choosing the number of children they have and, to a lesser extent, choosing the timing of births. One explanation is that birth spacing is partly determined by breastfeeding amenorrhea and abstinence, and the period of breastfeeding is in the domain of the mother. So women may have some control over the spacing of births, even in the absence of modern contraceptive methods, as argued by some anthropologists Mace and Sear (1997). In Senegal, the median postpartum non-susceptible period (related to amenorrhea and/or abstinence) is 14 months, and the average 15 months (ANSD and ICF International, 2012).

1.2 The importance of having a son

In Senegal, women have little access to inheritance as daughters, and even less as wives. Indeed, as described in Lambert, Ravallion and van de Walle (2014), the Islamic inheritance law is the dominant rule in the country, despite the existence of an alternative possibility ; in particular, wills and testaments are rarely made, and even less often implemented. *De jure*, the Islamic rule entitles wives to one eighth of the total bequest, to be shared among all wives in case of polygamous union. Further, this inheritance system splits the bequest between all the children recognized by the man,[6] limiting the inheritance of daughters to half of that of sons (RADI, 2010). In addition, patrilineal tradition, reinforced by Islam, favors sons for inheriting assets (house and land), as assets are transmitted within the paternal lineage. Since daughters typically move to their husband's abode, they are supposedly compensated by their brothers, with money or other forms of wealth, for what would have otherwise been their share of inheritance. But in practice, the compensation does not always take place. In

6. A man's extended family (uncles, cousins, nephews, etc.) is entitled to a share of the bequest only if he died childless.

fact, in households where there is very little wealth apart from the assets that have to remain in the lineage, sons often do not have much to give to their sisters in terms of compensation. [7] In our data, among people who lost their father, we find that daughters are significantly less likely to report that the father left a bequest, and among those who do, daughters are 2.5 times more likely to have not received anything (10 percent vs. 4 percent) ; they are also less likely to inherit a share of land or a share of the house, but not significantly more likely to inherit money, which contradicts the compensation story.

Another reason why daughters might be deprived of their rights is the non-implementation of the inheritance process. Typically, brothers and their own families may stay together in the father's house after his death ; as long as they are willing to live together, they do not feel the need of an actual division of property to take place and may consider that the father left no bequest. Among people who reported no bequest, we find that sons are significantly more likely to live in the household of their late father. If we consider that in such a case the house is transmitted de facto, we can estimate the proportion of people who inherited a share of the house : 60 percent of sons vs. 49 percent of daughters. This might not seem such a large imbalance. Nevertheless, when one looks at women at the time of their own death rather than at the time of their parents' death, it appears that women are vastly less likely than men to bequeath a house : less than a quarter of adults whose mother died have inherited anything from her, and only 22 percent of those who did inherited part of a house. By contrast, 72 percent of adults whose father died have inherited something, and 71 percent of them inherited a share of the house. This significant gap between men and women in the possibility to transfer property rights at the time of their own death suggests that, even if a large fraction of women claim to have inherited a house and hence appear to feel entitled to part of the house property, this might be only a temporary situation. In fact, by the time they die, they seem to have lost any access to it.

In total, men seem to be more successful than women at claiming their inheritance rights and in holding on to their property. As a result, the share of the bequest that will accrue to a woman's children will depend not only on the number but also on the gender composition of her offspring, relative to rivals. [8] Because of patrilinearity, having at least one son might

7. For wives, the situation is fairly similar to that of daughters. If the house and land go to sons that the late husband had from previous marriage, these sons might be poor enough to not be able to compensate her or to legitimate their lack of will to give her any compensation.

8. A similar situation is observed in Uganda. The International Fund for Agriculture Development states that "under customary law, [...] a widow is allowed only 25 percent of the estate of the deceased husband. All children, even if they are illegitimate, are entitled to 75 percent. In theory this applies to girls as well as to boys. However, reports indicate that, in practice, female children may often not inherit"

be a necessary condition to secure some access to the house when there are rivals. How many sons does a woman need to have to ensure that she is not dispossessed of the house by rivals is an empirical question. In fact, at the death of the husband, a son who already resides in the house might well be able to appropriate the house, even if his rivals represent more inheritance shares. [9] We will show that this is what our results suggest.

Further, having a son might not only raise the likelihood of remaining in the house, but could also provide an alternative in case of eviction. Indeed, widows are more likely to be hosted by a son than a daughter : among widows hosted by their children, 92 percent live with one son vs. 34 percent with one daughter. If we further consider widows living with exactly one child, 68 percent live with one son vs. 32 percent with one daughter, and both situations are very different : widows are hosted by their sons in two thirds of the cases, while they host their daughters in two thirds of the cases (being head of household, or remarried and co-residing with the new husband).

1.3 The vulnerability of widows

So far, the economic literature on vulnerability has paid little attention to the situation of widows in Sub-Saharan Africa. Yet, a few papers highlight that women are extremely vulnerable to the loss of their husbands : households headed by widows are significantly poorer than male-headed and other female-headed households (Appleton, 1996 in Uganda ; Horrell and Krishnan, 2007 in Zimbabwe). In Mali, van de Walle (2013) finds that widowhood has a strong, lasting impact on women's living standards and welfare indicators : the detrimental effects persist over time even if women remarry, and the disadvantage is passed on to children. The vulnerability of widows stems from women's subordinate status regarding legal protection, individual endowments, social norms and access to and control over resources. Historians have also documented that women who lack male support, like widows, are more vulnerable in case of famine (Vaughan, 1987 in Malawi).

Vulnerability is exacerbated when women have rivals for their husbands' resources. The question of wives' rivalry has only been studied in the context of polygamous households. Researchers working on Mali have documented that co-wife rivalry adversely affects child survival, especially that of sons since they remain to compete for land (Strassmann, 1997). In this paper we are interested in the competition for inheritance with children of ex-wives, rather than with co-wives. The existence of children of at least one ex-wife is a common fea-

(http ://www.ifad.org/gender/learning/challenges/widows/55.htm).

9. Legal dispute would be too expensive for most households, and is likely to take many years.

ture in the Senegalese context where widowhood, divorce and remarriage are very frequent for both men and women. Antoine and Dial (2003) point to the lack of reliable data on divorce in Africa : cross-sectional studies often underestimate the frequency of divorce because remarriage happens quickly. Using a biographical approach, they find that one union in three ends with a divorce in Dakar ; 25 percent of divorces occur after less than seven years of marriage, and 50 percent of divorced women are remarried five years later. To our knowledge, there are no studies about the impact of past marriage of the husband on current wife's well-being.

For current wives, the main stake in case of widowhood is probably the husband's house. First, notably for poor households, the house is likely to be the main asset. In such a case, as indicated above, there might be very little else that would permit monetary compensation to the widow and that could be as worthy to her as the use right of the house. Second, according to qualitative evidence, women fear being thrown out of the house by the children of an ex-wife and having nowhere to go (Lambert and van de Walle, 2012). For instance, a childless widow reported how she was confronted by the children of her husband after his death : they offered her a room to rent in their father's house at a steep price that she could not pay, and she was forced to leave. Remaining in the house of the late husband is crucial for two reasons : avoiding the obligation of remarrying, and in case of remarriage, keeping a relatively high degree of autonomy in this new marriage. Interviews and data support this claim as illustrated by Table 1, in which we provide some statistics on widows, broken down by the place of residence after the husband's death. Widows in our data were on average 46 years old when they lost their husband and children were born from this union in the vast majority of cases (88 percent). Two out of three widows were able to remain in the same household after the death of their husband, while the other third joined another household. The proportion of widows who remained in the husband's house depends on children's gender : it is 71 percent for women having at least one son with the deceased husband and 50 percent for women having only daughters. This is circumstantial evidence that having a son matters. Among those who could remain in the same household, 91 percent still reside there at the time of the survey : they became head of household, or stepped aside for one of their children. Only 19 percent got remarried, and among those who did, only one third live with their new husband. The story is completely different for widows hosted by another household at the death of their husband. They generally joined up with their parents, a child or a sibling. The host family provided only temporary accommodation : in 76 percent of the cases, widows had moved to another place at the time of the survey, after a stay shorter than two years for half of them. Consistently, the fraction of remarried women living with the new husband is much higher in this population : 49 percent remarry, and among those, two thirds live with their new husband.

Remarriage is an alternative widowhood insurance mechanism, which seems to be only a second-best for women. Qualitative evidence show that extremely vulnerable widows wish to remarry hoping to receive some help. But as soon as they can be self-reliant or rely on their own family, in particular on their adult sons, to host and to support them, widows clearly state that they do not want to remarry (Lambert and van de Walle, 2012). One reason is that they would be very likely to end up as a second or third wife in a polygamous union : in our data, three quarters of remarried widows are married to a polygamous husband. For those who could not avoid remarriage, being able to keep their own dwelling is associated with a gain in autonomy for the rest of their life. In Senegal, non-coresidence is not unusual : nearly a quarter of married women do not co-reside permanently with their husband. It is in particular the case for women who have their own house, which may be inherited from a previous husband now deceased. When spouses do not co-reside, the husband has another dwelling and comes for regular visits. For married women, it seems to be a favorable situation. Women who do not co-reside benefit from greater autonomy and enjoy relatively higher per-capita consumption levels (De Vreyer and Lambert, 2013). Also, in case of polygamous unions, non-coresidence mitigates conflicts between co-wives.

It is important to note that the argument that women need a son to guarantee access to the husband resources is less relevant for women without rivals, even if the probability of termination of this union is clearly positive. In Senegal, there is no legal provision for alimony and maintenance allowance for children. Hence, the divorcee does not obtain anything from the ex-husband, whatever the number of children.[10] In addition, at the death of the ex-husband, not only is the ex-wife not entitled to any share of the inheritance, but even if her children obtain the house, she would not return to live in the abode of a man she divorced, or who repudiated her. First, she is likely to be remarried and would keep the living arrangements she has with her current husband. And second, in such a patrilineal society, it would not be conceivable that she returns as household head in the house of a former marriage that belongs to another lineage. Hence, a wife who has no rivals has no incentive to adjust her fertility : whatever the gender composition of her children, either she

10. In our data, men having children from previous unions are not more likely to transfer money outside of the household than those who do not (39.5 percent vs. 42.4 percent, p-value=0.55, 705 observations). And among those who do transfer, men with previous children do not transfer a larger amount (112K FCFA vs. 137K FCFA, p-value=0.48, 295 observations).

remains married to her husband until his death, which grants her some access to the bequest, or she divorces and gives up on these resources. [11]

2 Data

2.1 The survey "Poverty and Family Structure"

The data used here come from an original survey entitled "Poverty and Family Structure" (PSF) conducted in Senegal in 2006-2007 (description in DeVreyer et al., 2008). [12] It is a nationally representative survey conducted on 1,800 households spread over 150 primary sampling units drawn randomly among the census districts. About 1,750 records can be exploited.

In addition to the usual information on individual characteristics, the survey obtained a comprehensive description of the household structure. Of particular relevance for our purposes is the fact that the survey registers for each child how many half-siblings he has, counting separately siblings from the same father only, siblings from the same mother only and siblings from both the same parents. Therefore, we are able to identify the children from previous unions on the mother's side, and the rivals on the father's side : since we consider monogamous unions only, all siblings having the same father but a different mother than the current wife, are children of an ex-wife. There is also information on all children below age 25 of household members, even if they do not live with their parents. Children who died are also reported but there is no information on the timing of births. As a result, a woman's complete birth history for surviving children is available only if all her children are under 25 years old. Among the children registered, only 59 percent live with both their parents ; the figure is similar to the national average (ANSD and ICF International, 2012). When parents do not live together, children live with their mothers in two thirds of the cases. That is why standard household surveys would largely underestimate the number of rivals, or even ignore their existence. Moreover, detailed information is collected on all current spouses, whether they co-reside or not, and on past marital history of all members. This is relevant since having previous unions, or children from previous unions, might play a role in determining

11. Women with and without rivals end up in the same situation if they eventually get divorced. What changes is the situation if they remain married : it is only in presence of rivals that having a son matters.

12. Momar Sylla and Matar Gueye of the Agence Nationale de la Statistique et de la Démographie of Senegal (ANSD), and Philippe De Vreyer (University of Paris-Dauphine and IRD-DIAL), Sylvie Lambert (Paris School of Economics-INRA) and Abla Safir (now with the World Bank) designed the survey. The data collection was conducted by the ANSD.

women's current fertility. Lastly, we use information on education, occupation and income, as additional controls.

2.2 Descriptive statistics on women

The sample consists of 936 women under 39 years old, engaged in monogamous unions, with at least one child from the current union. We had to restrict the sample of interest to relatively young women to ensure that their complete birth history is known, as mentioned above. In our data, the lower bound for age at first birth is 13, so children of women below 39 years old are all younger than 25. Since birth cohort is orthogonal to other variables of interest, this selection will not induce any estimation bias. Further, we focus on a given woman's children with her current husband, because they are the ones at stake during the bequest process.

Table 2 gives an overview of important characteristics of these women. For some of them, observations are missing for one or more variables. We have all the information for only 761 women, on whom we perform the semi- and fully parametric estimations. [13] The table provides the descriptive statistics for both the full sample and the subsample with complete information. They do not differ in any notable way. Non-parametric estimations are carried out on the full sample of 936, and we comment below on the descriptive statistics on this sample. Women have on average 2.7 living children with their current husband ; the number ranges from one to nine (Figure A.1 in Appendix A displays the distribution). Roughly one woman in five has lost at least one child, and four percent have children from previous unions. Of course, since they are in the middle of their reproductive lives, we do not observe their total fertility. The average age in the sample is 28, while the median ages at births 2 and 3 are respectively 25 and 28 years old. [14] We will therefore focus on children of rank 1, 2 and 3 ; only a small fraction in our sample has already given birth to more children. Fertility varies significantly across places of residence, education levels and employment status ; in our sample, 48 percent of women live in rural areas, 44 percent received no education at all, and 33 percent work. Regarding their marital lives, the vast majority of these women (87 percent) are in their first union ; the remainder have only one broken union. Women marry for the first time at age 19 on average, and they marry a much older husband : the age difference is 10 years on average, and up to 23 years in the top decile of the distribution of age differences. The threat of becoming a young widow is therefore probably a widespread

13. Table A.1 in Appendix A provides the detail of the number of observations with missing values.
14. The median ages at births 2 and 3 are derived from the sample of women above 45 years old.

matter of concern. Only seven percent of husbands work in the public sector, so the public system of widowhood pensions covers a very small part of the population. The average annual income is 1.6 million FCFA (approx. 2,500 euros), but it is driven by the right tail of the distribution as 50 percent of the husbands earn less than 500 K FCFA annually (approx. 750 euros). Turning to statistics about rivals, marrying a man who already has children with ex-wives is not unusual : 17 percent of women have at least one rival, and 12 percent at least one male rival. The number of rivals in inheritance shares (one share = one boy = two girls) is on average 1.7, but the variance is large with numbers ranging from 0.5 to 7.5 shares. 10 percent have strictly more rivals than own children at the time of the survey.

Women with rivals are described in Panel A of Table 3. They have themselves more complex marital lives : they are older and more likely to have broken unions, in which children were born. The age difference with their husband is significantly higher, by four years. They are also more likely to live with their current husband, and to work.[15] A first hint that the presence of rivals may have an impact on fertility is that women with rivals have on average more children with their current husband, and in quicker succession. As predicted by the medical literature, shorter birth intervals are correlated to a higher mortality rate for those children.

A subsample of particular interest will be that of co-residing wives.[16] As explained in the preceding section, rivalry for inheritance seems to be mostly about the husband's house ; hence, the threat of being thrown out of the house by rivals only jeopardizes women who actually live in their husband's house. It is therefore natural to look at the potentially most exposed women to find evidence of insurance strategies. As shown in Panel B of Table 3, only three married women in four live with their husband. Non-coresiding wives are very similar to other women in terms of socioeconomic characteristics (age, area of residence, education, occupation), but they tend to have fewer children, dead or alive, with their current husband.

Note that we cannot exhibit any descriptive statistics on widows with rivals and no son, since, consistently with our claim, there are very few of them (eight observations) in the

15. The same table replicated for the subsample of co-residing wives displays very similar patterns.

16. One caveat is that co-residence status is observed at the time of the survey, and not when fertility choices were made. Consequently, some women are considered as co-residing, though they were housed independently when their children were born. If they anticipated that they would eventually join their husband's house, they should indeed be included in the category of most exposed women. Conversely, some women are considered as non-coresiding, while they used to live with their husband. In our sample, out of 231 non-coresiding women, 46 actually remained with their family-in-law after the husband, generally for professional reasons, left the household. Those women clearly face the same risk of eviction as co-residing ones. We find that, indeed, they behave similarly : when we include them in the subsample of co-residing wives, the magnitude and significance level of our coefficients remain completely unchanged.

sample. If we believe that the widowhood insurance mechanism is at play, such a situation would indeed be rare.

3 A model of fertility choices with inheritance concerns

3.1 General framework

Following the narrative of the previous sections, we model the choice of the wife regarding the time interval between the last and the next birth. To ensure tractability, we limit ourselves to a static model at parity n and represent the following trade-off : a long interval is good for health since shorter birth spacing implies riskier pregnancy, but it reduces the likelihood of a birth if one of the spouse dies or becomes infertile in the meantime. [17]

Formally, the wife maximizes her expected utility over t :

$$v(t) = p(t).(u(n+1) - c(t)) + (1 - p(t)).u(n)$$

Where $p(t)$ is the probability that the couple can conceive a child at date t, and it is decreasing with t ; $c(t)$ is the health cost of a birth at date t, and it is decreasing with t ; $u(n)$ is the utility to the mother of having n children, and it is increasing and concave in n.

We consider the following functional forms to get a simple closed form for t^* : $p(t) = e^{-\lambda t}$ and $c(t) = a - bt$, where a is very large and b small enough to ensure that $c(t) > 0\ \forall t$ and that $t^* > 0$. The first order condition of this maximisation program gives :

$$t^* = \frac{1}{\lambda} + \frac{a - \Delta u_n}{b} \tag{1}$$

Where $\Delta u_n = u(n+1) - u(n)$. Equation 1 has a fairly straightforward interpretation. Indeed, it states that the optimal birth spacing is shorter when the marginal utility of an additional child is higher. The intuition is that, since the cost of a birth is decreasing over time, there is a moment when this cost is offset by the benefits of having an additional child. If the benefits are high, that moment comes sooner.

In this general framework, the gender of children and the presence of rivals play no part. They are introduced sequentially in the following sections.

17. In such a static framework, the trade-off is between health and the likelihood of *one* birth. This is a simplification of the actual trade-off, which is between health and the likelihood of *several* births. Indeed, shorter birth spacing increases the likelihood of a birth during this period, but also the possibility to try again in the next period, and in the next one etc.

3.2 Son preference

Let $u(n)$ now depend on the gender composition of children. We denote s the number of sons and, following Jayachandran and Kuziemko (2011), we model son preference as follows :

$$u(n, s) = q(n) + g(s)$$

Where $q(n)$ is increasing and concave and $g(s)$ is positive, increasing and concave. $g(s)$ captures the additional gain from having sons, holding the number of children constant. With this hypothesis, sons are more valuable than daughters and all the more valuable as there are few of them.

Since the gender of the potential next birth is unknown and a son has a one chance over two to be born, the optimal birth spacing is the same as in Equation 1, except that :

$$\Delta u_n = q(n+1) - q(n) + \frac{g(s+1) - g(s)}{2}$$

If we compare the optimal choice in the absence of a son and the optimal choice with at least one son, holding n constant, we get : [18]

$$t_0^* - t_{s>0}^* = -\frac{1}{2b}[(g(1) - g(0)) - \mathbb{E}(g(s+1) - g(s)|s > 0)] \tag{2}$$

We find that $t_0^* - t_{s>0}^* < 0$ iff $g(.) > 0$, strictly increasing and strictly concave. Hence, in the presence of son preference as modelled here, the next birth will take place sooner in the absence of a boy among the first n children.

3.3 Risk of dispossession

In order to capture the risk of losing access to the husband's house after his death and the way it is related to the presence of rivals and to the gender composition of one's offsprings, we introduce a new term in $u(n, s)$. We write :

$$u(n, s) = q(n) + g(s) - \pi L$$

Where L is the amount of the loss, and π is the probability of the loss. π depends on two factors. First, the presence ($R = 1$) or absence ($R = 0$) of rivals : we assume that the risk

18. We need the expectation operator because we pool different values of s in "at least one son".

of being dispossessed exists iff there are rivals i.e. $\pi(R=1)>0$ and $\pi(R=0)=0$. Second, the number of children, weighted by gender : $M = s+\delta.(n-s)$. We assume that $\pi(M)$ is decreasing and convex in M and that $\delta < 1$. In other words, having children decreases the risk of dispossession ; sons and first-born children contribute more than daughters and last-born children to this risk reduction.

In presence of rivals, we compare the choice without a son ($M_0 = \delta n$) and the choice with at least a son ($M_s = s+\delta.(n-s)$), holding n constant. We get :

$$\tilde{t}_0^* - \tilde{t}_{s>0}^* = -\frac{1}{2b}[(g(1)-g(0)) - \mathbb{E}(g(s+1)-g(s)|s>0)] + \frac{L}{b}(\Delta\pi(M_0) - \mathbb{E}(\Delta\pi(M_s)|s>0)) \quad (3)$$

Where $\Delta\pi(M) = \frac{\pi(M+1)+\pi(M+\delta)}{2} - \pi(M)$ is the expected variation in the risk of losing the house triggered by the additional birth. Note that $\Delta\pi(M) \leq 0$. In the absence of rivals, $\pi = 0$. Hence, the difference in the optimal choice with and without a son is still given by Equation 2.

The difference in difference term obtained by taking the difference between Equations 3 and 2 captures a widowhood insurance motive if and only if the function $g(s)$ is the same for women with and without rivals. This is what we call the "common intrinsic son preference assumption".[19] Under this assumption, the difference in difference term (henceforth the DiD term) writes :

$$DiD = (\tilde{t}_0^* - \tilde{t}_{s>0}^*) - (t_0^* - t_{s>0}^*) = \frac{L}{b}(\Delta\pi(M_0) - \mathbb{E}(\Delta\pi(M_s)|s>0)) \quad (4)$$

The term is negative because the reduction in risk accruing from an extra birth is lower in absolute value when the woman already has sons (she starts with M_s) than when she does not (starts with M_0). Indeed, $\pi(M)$ is convex and $M_0 < M_s$.

We do not need to make more specific assumptions on $\pi(M)$, and its exact shape is an empirical question. One way to know more about $\pi(M)$ and δ is to look at DiD for different values of s, holding n constant. We will do this in Section 5.3. If the risk of losing the house hardly decreases when passing from one to two or three sons (i.e. $\Delta\pi(M_1) \approx \Delta\pi(M_2) \approx \Delta\pi(M_3)$ for each n), it will mean that one son is enough to feel insured. In such a situation, it is possible to test the common intrinsic son preference assumption by comparing women with exactly one son and women with two sons and more, when $n \geq 2$. Let us denote $g(.)$ and $\tilde{g}(.)$ the son preference functions for women without and with rivals, respectively. Since

19. Note that we do not need any assumption on $q(n)$. Both groups of women may value children differently.

the $\Delta\pi$ terms cancel out, the double difference term writes :

$$(\tilde{t}^*_{s=1} - \tilde{t}^*_{s>1}) - (t^*_{s=1} - t^*_{s>1}) = -\frac{1}{2b} \quad \begin{aligned}&[(\tilde{g}(2) - \tilde{g}(1)) - \mathbb{E}(\tilde{g}(s+1) - \tilde{g}(s)|s > 1) \\ &-(g(2) - g(1)) + \mathbb{E}(g(s+1) - g(s)|s > 1)]\end{aligned}$$

If the common intrinsic son preference assumption holds, then $g(.) = \tilde{g}(.)$, hence this double difference should be nil. This is the rationale of the Placebo test presented in Section 5.3.

3.4 Summary of assumptions and predictions

The model clarifies what is called son preference here. It is simply that there is a positive additional utility gain to having a son, for a given number of children ($g(.) > 0$), and this gain is strictly increasing and strictly concave. Moreover, the key assumption for the interpretation of the double difference as the causal impact of rivals is that $g(.)$ is the same for women with and without rivals. Last, the only assumptions about the bequest process is that the probability of being dispossessed of the house is positive in the presence of rivals and nil otherwise, and this risk decreases with the number of children, faster with low parities than with high parities, and faster with the number of sons than with the number of daughters.

The model predicts that a woman without sons will accelerate more her fertility relative to a woman with at least one son in the presence of rivals than in the absence of such threat.

One limit of this simple framework is to overlook dynamics. In particular, we are unable to predict that the double difference should be larger for higher parities. In the empirical analysis, we observe that the effect increases in n. We explain this result by women's anticipations on how many additional tries for a son they are likely to have in the future. Our model does not allow us to discuss this type of dynamics.

4 Empirical Strategy

4.1 A duration model of birth intervals

Consistent with the above model, the idea underlying our empirical strategy is to compare fertility choices and son preference of wives with and without rivals. Building on the literature on son preference, we use a duration model of birth intervals (Leung, 1988). The existence of son preference is confirmed when the length of the birth interval before the next child is longer for couples with more sons.

The advantage of duration models is their ability to deal properly with right-censored observations, i.e. families that are not yet complete by the time of the survey.[20] Indeed, they 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 here, as the date of the survey is completely unrelated to the latest births. Another rationale for considering birth intervals is to account for health risks related to the spacing, and not only to the number, of births.

Using duration models, son preference has been tested and validated by an extensive literature focusing mainly on Asia.[21] Only a few papers examine this question in an African context :[22] Gangadharan and Maitra (2003) find evidence of son preference in South Africa, but only among the Indian community. To our knowledge, the only conclusive paper on African data using duration models was written by anthropologists working on Gabbra pastoralists in the North of Kenya ; they find that women with no son have significantly shorter subsequent birth intervals than women with at least one son (Mace and Sear, 1997). They relate son preference to the patrilineal and patrilocal family system prevailing among the Gabbra. Hence, evidence on son preference is quite limited in Sub-Saharan Africa.

4.2 The model with proportional hazard

To test whether son preference is related to the presence of rivals, we use a model with proportional hazard.[23] Our variable of interest T is the duration between successive births, measured in months. Our coefficients of interest measure the impact of the gender composition of the first children on the subsequent birth interval, among wives with and without

20. We have only limited information on the children of women with complete fertility : among women over 45 years old who declare having at least one child, not a single child is found in our dataset in almost half of the cases ; all the children are found in only seven percent of the cases. On average, a woman declares five children and only 1.3 is registered. As a result, we do not know the rank and gender of each child, so we cannot infer the gender composition of the firstborns.

21. For instance, in China (Tu, 1991), Bangladesh (Rahman and Da Vanzo, 1993), the Chinese population of Malaysia (Pong, 1994), Vietnam (Haughton and Haughton, 1995), India (Arnold et al, 1998), South Korea (Larsen et al., 1998), and Taiwan (Tsay and Cyrus Chu, 2005).

22. Some papers use duration models of birth intervals in Africa, but they are interested in the impact of socioeconomic factors (e.g. Ghilagaber and Gyimah, 2004), not in son preference. Other papers deal with son preference in Africa, but do not estimate duration models ; for instance, Milazzo (2012) estimates a linear model for the probability of short birth interval to convincingly underline the role of son preference in the fertility choices of Nigerian women.

23. We test the proportional hazard assumption following the procedure developed by Grambsch and Therneau (1994) and based on the Schoenfeld partial residuals (Schoenfeld, 1982), and we fail to reject it.

rivals. We consider intervals between births n and $(n+1)$ for $n = \{1, 2, 3\}$.[24] Our main specification is a pooled regression of all intervals ; as a robustness check, we run separate regressions for each of the three parities to determine the one at which hazard rates diverge (see Section 5.3).

The key element of duration models is the hazard function : the instantaneous probability to have another child at date t. In the specification with proportional hazard, the hazard function is modeled as follows :

$$\lambda(t) = \lambda_0(t) \times \exp(X_n'\beta)$$

Where $\lambda_0(t)$ is the baseline hazard function, common to all individuals, and X_n is a vector of individual covariates at birth n susceptible to influence the hazard function. Then the survival function – the probability not to have another child at least until t – is as follows :

$$S(t) = \Pr(T > t) = \exp(-\int_0^t \lambda_0(u)du \times \exp(X_n'\beta))$$

In our case, X_n includes *Girls*, a dummy equal to one if the first n children are all girls ; *Rivals*, a dummy equal to one if there is at least one rival ;[25] *Girls* × *Rivals*, the interaction term between *Girls* and *Rivals* ; the rank of preceding birth to control for potential differences between parities ; additional characteristics of the woman related to fertility : age at birth n, area of residence, level of education, age at first marriage, whether in her first marriage, living or not with her husband, having at least one dead child,[26] having at least one child from previous unions.

The fact that we do not observe the birthdate of deceased children might lead us to overestimate the true duration between all successive births, and hence to underestimate son preference. If women do indeed shorten birth spacing until they get a son, we would wrongly assign a large birth interval to those who lost a child as a consequence. Such a bias would be larger for women with a stronger son preference, because the impact on child mortality would be higher. Our coefficients can therefore be viewed as a lower bound for son preference, and as a lower bound for the difference in son preference between women with and without rivals.

24. As a robustness check, we will include all parities instead of limiting up to parity 3. See Section 5.3.
25. As robustness checks, we consider alternative definitions of rivals, such as the number of rivals or a dummy for at least one male rival.
26. This is the best we can do to mitigate the fact that we do not observe all successive births ; results are qualitatively unchanged if the dummy is removed.

Also, one might worry about selective mortality : those women who intensify their fertility when they have only daughters are more at risk of maternal death, and therefore more likely to be under-represented in the sample. But this would again lead us to underestimate son preference. [27]

4.3 Identification assumptions

Identification relies on the assumption that gender of first born children is exogenous to all characteristics, and in particular to whether one has rivals or not. In fact, in a simple double difference framework, the estimation amounts to computing the empirical counterpart of the DiD term of our theoretical framework :

$$\begin{aligned}\widehat{DiD} =&(\mathbb{E}[t|Girls=1,Rivals=1]-\mathbb{E}[t|Girls=0,Rivals=1]) \\ &-(\mathbb{E}[t|Girls=1,Rivals=0]-\mathbb{E}[t|Girls=0,Rivals=0])\end{aligned} \tag{5}$$

It is crucial that $Girls$ is exogenous, and in particular orthogonal to $Rivals$, for each part of the $\widehat{DiD}$ term to give the estimate of the causal impact of the treatment (having a first born female rather than male) on birth interval in each of the subgroups, women with rivals and women without rivals. The double difference provides an estimate of the heterogeneity of this impact across women with and without rivals.

This assumption is likely to hold insofar as there is no evidence of sex-selective abortion or infant mortality in Senegal (ANSD and ICF International, 2012). In our sample, sex ratios among children are perfectly balanced (cf. Figure A.2 in Appendix A). Moreover, we test that women having n daughters are indeed similar to women having at least one son for $n = \{1, 2, 3\}$: balancing tests presented in Appendix A, Table A.2, provide some level of reassurance that the $Girls$ dummy is exogenous. [28]

Under this assumption, the second row of Table 4 gives the estimation of the first term in brackets in Equation (5). We find that indeed, in presence of rivals, the absence of a son reduces birth interval by about six months. The first row computes the second term and gives the same difference for women without rivals. It can be seen that in this case, no significant difference in birth interval is driven by the gender of the first born children.

27. Among children younger than 15 years old (6,150 observations), 2.4 percent of boys and 2.5 percent of girls have lost their mothers. Nevertheless, the difference between boys and girls is not significant (p-value=0.82).

28. We conduct univariate t-tests to compare covariates' means across both subsamples, and we find no statistically significant difference at 5 percent when $n = \{1, 2\}$ and only one (in 17 covariates) when $n = 3$.

Hence, the empirical difference-in-difference clearly shows that women with rivals and no son display significantly shorter birth intervals (by eight months) than all other women (see Table 4).[29] In the same way, the proportion of women giving birth with intervals shorter than 24 months is significantly greater (by 22 percentage points) among those women. The estimations provided below only confirm and specify these raw results.

Turning to the proportional hazard model, e^{β_k} is the hazard ratio at any point in time between two individuals that only differ by one unit of X_k. In our case, β_1, β_2 and β_3 are the coefficients on $Girls$, $Rivals$ and $Girls \times Rivals$, respectively. e^{β_1} measures the hazard ratio between women having only daughters and women having at least one son, among those without rivals. As discussed above, β_1 identifies a causal impact if the gender composition of the first children is truly exogenous. If we find that $e^{\beta_1} > 1$ (in other words, having only daughters vs. at least one son increases the hazard rate and hence decreases the expected birth interval) then we can infer the existence of son preference among women without rivals.

$e^{\beta_1+\beta_3}$ measures the same ratio among women with rivals. Therefore, e^{β_3} is the odd-ratio that captures the difference in son preference between women with and without rivals. It corresponds to the DiD term in our theoretical framework. If our hypothesis is true, then we expect $e^{\beta_3} > 1$; in other words, having only daughters vs. at least one son should increase *more* the hazard rate and decrease *more* the expected birth interval when there are rivals. β_3 can be interpreted as the effect of an insurance strategy through sons as long as both groups of women have the same intrinsic son preference, as spelled out in the model.

Of course, as described in Section 2.2, wives with and without rivals are different ; even after controlling for observable differences, some unobservable characteristics may drive both fertility choices and the decision to marry into a union with rivals.[30] However, as clarified in the model, what matters for our interpretation to hold, is that in the absence of rivals, both groups would have reacted similarly to an exogenous shock : having one son vs. one daughter. Under this assumption, if wives with rivals eventually display a stronger son preference than wives without rivals, the difference would be caused by the presence of rivals. In the results presented below, we control for all observational differences between these two groups of women, so that the necessary assumption is that of common son preference conditional on the whole set of observables. In Section 5, we run some placebo tests to ensure that (i) the common intrinsic son preference is plausible, and that (ii) an insurance-based interpretation explains the results better than potential alternative mechanisms. Other differences between

29. Statistics are computed on non-censored durations, using the sample of co-residing wives.

30. It is worth underlining that, given the general complex patterns of marital lives in Senegal, marrying a divorced man or a widower is likely to happen to most women at some point in their lives.

women with and without rivals are wiped out by the double difference approach.

As a first placebo test, we compute the double difference on the raw data, replacing the variable *Rivals* by the dimensions on which women with rivals differ from other women : *Age*, woman's age at the time of the survey ; *Work*, a dummy indicating whether the woman works ; *BrokenUnion*, a dummy indicating whether the woman had previous unions ; *Prev.Child*, an indicator for having children from previous unions ; *Nb.cur.Child*, the number of children in the current union ; and *Dead.Child*, an indicator for having one dead child from the current union. The idea is to see whether there is any heterogeneity in son preference along dimensions that are correlated with the existence of rivals. In all but one of the tests, the coefficient on the interaction term is not significant. It is significant when we consider the dead child dummy, but the sign is in the opposite direction. Since *Dead.Child* is positively correlated with *Rivals*, our result cannot be driven by the higher probability to have lost a child (cf. Table E.1 in Appendix E).

One limitation of our empirical strategy is that we are dealing with rather small samples. Table A.1 in Appendix A gives details about the number of observations for each estimation, as well as the exact number of observations in each cell of interest.

4.4 Descriptive statistics on birth intervals

Table 5 presents some descriptive statistics of our duration variables. When we pool all durations together, durations after birth 1 account for almost half of the observations, durations after birth 2 for one third, and the remaining durations are observed after birth 3. A bit more than one third of the durations are right-censored, a proportion that logically increases in parity. Among non-censored observations, the distribution is quite similar for the three parities, with a mean around 36 months, and a median around 30 months. More than a third (37 percent) of the births would be considered as risky with a birth spacing lower than 24 months. Note that duration between births 1 and 2 displays a larger variance due to "extreme" values.[31]

Before estimating the proportional hazard model, the first step to get a picture of the survival function is to compute the Kaplan-Meier estimator of $S(t)$ (Kaplan and Meier, 1958). The two left hand side graphs of Figure 1 show it plotted when durations are pooled together and when they are separated by parity. Both graphs display a common shape with

31. Approximately two percent of birth intervals are larger than eight years (mostly observed after birth 1). In half of the cases, women have in fact lost one or two children in the meantime ; the remaining cases are mostly urban, educated, working women living in relatively rich households.

three important features :

1. $\widehat{S}_{KM}(t) = 1$ if $t < 8$: the shortest birth interval is equal to eight months.
2. The exit rate is increasing until approximately 50 months and then decreasing.
3. $\widehat{S}_{KM}(t) > 0$ even if t is large i.e. the distribution of T is defective ; in other words, some women will never have an additional child.

Also, note that we observe "jumps" in the survival function at 12, 24, 36 etc. months. They are due to measurement approximations : in roughly one third of the non-censored cases, we do not observe the exact dates of birth for the children n and $(n + 1)$, hence we approximate the birth interval by the age difference multiplied by 12, to get the duration in months. The same approximation is implemented on censored observations in 20 percent of the cases, when the exact birthdate of the latest child is missing.

As a consequence, when we observe t_n, the information we derive is more or less precise, depending on the type of observation : [32]

- Not censored, precisely measured ; we infer that $T_n \in [t_n - 1; t_n + 1]$.
- Not censored, imprecisely measured ; we infer that $T_n \in [t_n - 12; t_n + 12]$.
- Censored, precisely measured ; we infer that $T_n > t_n - 1$.
- Censored, imprecisely measured ; we infer that $T_n > t_n$.

This will be dealt with when we estimate a fully parametric version of the model.

Using the Kaplan-Meier estimate of the survival function and its variance, it is possible to test for the equality of distributions between two sub-populations (log-rank tests). [33] As shown by the right hand side graphs of Figure 1, the survival function of women having only daughters and rivals is below that of the others, which means shorter birth intervals, in two cases : first, in the pooled specification, when we restrict the sample to co-residing wives, we reject the null (at 3 percent) that women with only daughters and rivals have the same survival function as the rest of the population. Second, when $n = 3$, although we cannot reject the null at any conventional level (the p-value is 0.11).

The non-parametric estimation provides a first hint that women with the highest need for widowhood insurance also make different fertility choices.

32. See Appendix B for further details about the approximation and the number of observations of each type.

33. We compare women with only daughters vs. those with at least one son ; women with rivals vs. those without any ; and finally women with both only daughters and rivals vs. others.

4.5 Estimation methods

4.5.1 Semi-parametric estimation

We start by considering a semi-parametric estimation of the Cox proportional hazard model (Cox, 1972). Using a partial likelihood, the method of estimation makes it possible to estimate the vector of coefficients β without imposing a functional form on $\lambda_0(t)$. The idea is to focus on the ordering of events rather than on the exact duration for each observation. We use robust standard errors clustered at the woman level to account for the correlation between the error terms related to the different birth intervals for the same woman.[34]

One can derive an estimate of the baseline survival function, $\widehat{S}_0(t^*_{(j)})$, constant by pieces on the interval $[t^*_{(j)}; t^*_{(j+1)}[$, where $t^*_{(1)} < ... < t^*_{(k)}$ are the k ordered distinct values among non-censored observations (Box-Steffensmeier and Jones, 2004). We use it to compute the expected duration conditional on X for the women who eventually have another child :[35]

$$\mathbb{E}(T \mid X, T \leq t^*_{(k)}) = \Big(\sum_{j=1}^{k-1} \widehat{S}_0(t^*_{(j)})^{\exp(X'\hat{\beta})} \times (t^*_{(j+1)} - t^*_{(j)}) \Big) - t^*_{(k)} \widehat{S}_0(t^*_{(k)})^{\exp(X'\hat{\beta})}$$

Another quantity of interest is $\big(1 - \widehat{S}(24 \mid X)\big)$: the probability of a short birth interval for women with characteristics X. We define "short" as less than or equal to 24 months because it is the minimum length advocated by family planning programs and used in health statistics (e.g. the Demographic and Health Surveys (DHS) Program) below which a birth is considered risky.

Nonetheless, the semi-parametric estimation has three main limitations. First, it is not well-designed to model individual-level unobserved heterogeneity, which would be useful here to capture the unobserved determinants of fertility across women (e.g. fecundity level). Second, the specific measurement approximations described above cannot be accounted for with the standard Cox estimation. Third, the estimation does not explicitly take into consideration the defective distribution of our durations.

34. We use the Breslow method to handle ties among non-censored durations since the exact marginal-likelihood method (more accurate) is not available when standard errors are clustered. In the regressions separated by parity, we use the exact marginal-likelihood method.

35. In the Cox model, some women will never have another child, so the global expected duration would be equal to infinity. That is why our quantity of interest is in fact the expected duration given that $T \leq t^*_{(k)}$.

4.5.2 Fully parametric estimation

To account for (i) unobserved heterogeneity, (ii) measurement approximations, and (iii) a defective distribution, we estimate a fully parametric multispell model. The details of the specification are given in Appendix C.

Unobserved heterogeneity (ν) is introduced in the hazard function. The baseline hazard is defined so as to reproduce two characteristics of the durations observed in our sample : (i) there is no exit before eight months, and (ii) the inverted U shape of the exit rate arises from the combination of an increasing baseline hazard together with unobserved heterogeneity.[36]

Further, the contribution to the likelihood of each observation is specified in a way that takes into account measurement approximations ; it uses a confidence interval that depends on how precisely the duration is measured, instead of the density and survival function.

Finally, we model explicitly the probability to stop having children after the n^{th} birth. We are thus able to disentangle the effect of covariates, in particular $Girls$, $Rivals$ and $Girls \times Rivals$, on the number of births and on the spacing of births.

As in the Cox estimation, we are able to compute the probability of having another child, the expected birth interval and the probability of short birth spacing, for different categories of women.

5 Results

5.1 Semi-parametric estimation : Results

The results of the Cox estimation are presented in Table 6. When we consider the whole sample of monogamous women under 39 years old having at least one child from the current union, there is no ratio significantly different from one ; we notice that in magnitude, e^{β_1} and e^{β_2} are indeed very close to one, whereas e^{β_3} is higher, but imprecisely estimated.

In accordance with the hypothesis that the crux of the matter lies with the access to the house, we consider the subsample of co-residing wives. For them, we have information about the husband's characteristics (income, sector of activity, age difference) that we include in the regression as additional controls ; we also add a dummy indicating whether the woman works. e^{β_1} remains insignificant, so we find no detectable evidence of son preference among

36. Women with a high hazard rate exit at the beginning, and after some time only women with a low hazard rate remain. We do not investigate further whether the inverted U shape results from the exit process *per se*, or from the variation in the composition of the population, because in this study we are not mainly interested in duration dependence.

women without rivals, at least up to their third child. e^{β_3} is higher in this subsample, and now significantly different from one (at the five percent level). Hence, among women co-residing with their husband and with at least one rival, the hazard rate is multiplied by $e^{\beta_1+\beta_3} = 1.6$ for those having only daughters as compared to those having at least one son.

We further investigate whether the presence of rivals has a heterogeneous impact depending on the socioeconomic status of the household ; to do so, we split the sample of co-residing wives on the median income.[37] In the poorest half of this population, e^{β_3} is very high and significant (at one percent) : having only daughters multiplies the hazard rate by 2.4 for women with rivals, while it has no significant impact for women without rivals. In the richest half of the population, e^{β_3} is not significantly different from one.[38] At first sight, such a result might seem at odds with standard economic predictions that the larger the bequest, the stronger the incentives to take hold of it. However, what matters in our context is probably not the absolute value of the bequest as much as the relative value of the house in the bequest. As discussed in Section 1, wives in poor households cannot be compensated with money or movable goods for what would be their share of the house. Moreover, women in those households have hardly any alternative in case of eviction because they cannot rely on personal resources or on their extended family to support them.

Another group of women particularly exposed to the risk of widowhood would be those married to an old husband. However, we would need more women married to an old enough husband to be able to estimate the effect of husband's age.[39]

Also, note that e^{β_2} is never different from 1. Hence, after controlling for observable differences between women with and without rivals, both groups make similar fertility choices once they have a son. Regarding the controls, hazard ratios are in line with expectations. The hazard rate decreases with parity : durations are longer between the latest children. It is higher for women in their first union and who married younger. Also, it is higher for

37. Results are qualitatively unchanged if we take education instead of income as an indicator of socioeconomic status : e^{β_3} is much higher for non-educated women than for educated ones. The same is true if we split the sample according to the rural/urban divide. In this case, we find that e^{β_3} is very large for rural women, while none of the coefficients is significantly different from one in the urban sub-sample.

38. If we interact a *Poor* dummy with our variables of interest instead of splitting the sample, the difference between rich and poor is confirmed : the coefficient on the triple interaction $Girls \times Rivals \times Poor$ is equal to two and significant at 10 percent.

39. The distributions of husband's age at birth n are : $Q_1 = 27, Q_2 = 31, Q_3 = 36$, among women without rivals ; and $Q_1 = 30, Q_2 = 36, Q_3 = 42$, among women with rivals. They are vastly overlapping. The impact of husband's age on women's choices is likely to be non-linear, but the relevant thresholds are difficult to figure out. Indeed, the mortality risk is increasing with age, but maybe not in a way that is very perceptible by wives : 0.41 percent for men aged 35 to 39, 0.55 percent for those aged 40 to 44 and 0.58 percent for those aged 45 to 49 (ANSD and ICF International, 2012).

co-residing wives and for rural mothers. When we restrict the sample to co-residing wives, we find that women married to civil servants, and therefore entitled to a public system of widowhood pensions, have lower hazard rates. Age difference and employment status of the woman are never significantly correlated with the hazard rate.

To better understand the consequences of these results in terms of health, we estimate (i) the expected birth spacing and (ii) the probability of short birth spacing for the two populations in which we detected an effect : co-residing wives and the poorest half of co-residing wives (cf. Table 7). We consider the median individual as the reference individual, by setting the values of the controls to the median of the sample. [40] Among co-residing wives, we find that, holding everything else constant, the expected birth spacing is quite similar for all women when there is no rival. But in presence of rivals, having only daughters vs. at least one son is predicted to reduce birth spacing by 10 months and to multiply the probability of short birth spacing by 1.5. These same figures are 11 and 2, respectively, when we consider the poorest half of the population. These estimates are very comparable to what is observed in the raw data. [41] The probability of very short birth spacing (below 15 months) is multiplied by 1.5 for co-residing wives and by 2.3 for women in relatively poor households. [42]

5.2 Fully parametric estimation : Results

The estimates of the fully parametric specification are reported in Table 8 ; they confirm, strengthen and specify the results of the Cox model. The first clarification is that, on the one hand, our variables of interest $Girls$, $Rivals$ and $Girls \times Rivals$ have no significant impact on the probability to stop having children, whatever sample we consider. [43] But on the other hand, the effect of $Girls \times Rivals$ on durations is much higher and significant than estimated in the Cox model ; e^{β_3} turns out to be significant even for the whole sample. The ranking across subsamples remains the same : e^{β_3} is still larger for co-residing wives, in particular for the poorest half. Otherwise, e^{β_1} and e^{β_2} are similar in magnitude to the Cox estimates, and in the same way, never significantly different from 1.

Hence, estimations based on the explicit modeling of a defective distribution suggest that

40. We check that results are qualitatively unchanged if we consider the average individual (not shown).

41. Excluding extreme values of birth intervals (above 96 months) does not change the predicted probability of short birth spacing, but mechanically reduces the expected birth spacing and flattens out differences between categories : the gap between women having rivals and only daughters vs. at least one son amounts to six and eight months, for co-residing wives and for the poorest half of co-residing wives, respectively.

42. Same computation as for the probability of birth spacing below 24 months (not shown).

43. The probability of having another child is estimated to be around 98 percent for the median individual, irrespective of the gender composition of the first children and/or the presence of rivals.

women's needs for a son influence the spacing, but not the number of births – at least up to the third child.[44] Our results are in line with the literature on son preference : Jensen (2005) and Basu and De Jong (2010) point out that son preference does not necessarily lead to differences in sibship size when fertility is high ; in this case, they advocate looking at birth intervals to find evidence of son-preferring behavior.

Further clarification is also obtained on the controls. Living in a rural area, being married for the first time, and co-residing with the husband, all are negatively correlated with stopping births. Other controls relate to birth spacing : durations are positively correlated with the age at first marriage, the level of education, the rank of the child and having children from previous unions ; and they are negatively correlated with the age at the previous birth. Co-residence is the sole variable connected with both the number of children and their spacing. When we introduce controls related to the husband, wives of civil servants display a lower probability to stop, but longer birth intervals. Age difference and women's employment remain insignificant.

Our quantities of interest capturing the magnitude of the effect display larger differences across groups than those derived from the Cox estimation :[45] among co-residing wives (resp. among the poorest half), having only daughters vs. at least one son multiplies the probability of short birth spacing by 2 (resp. 3.5) and reduces the expected by birth spacing by 11 months (resp. 17 months) when there are rivals. But just as in the Cox model, birth spacing is predicted to be quite similar for all women when there is no rival (cf. Table 9).

Regarding the baseline survival function, coefficients are always precisely estimated. The baseline survival function of co-residing wives is below the one of the whole sample, reflecting their increased fertility (cf. upper graph in Figure D.1, Appendix D). Last, in all samples, estimates of the variance of the unobserved heterogeneity are significant at the one percent level and of the same order of magnitude.[46]

44. We cannot formally test what happens at higher parities because the number of women having only girls and rivals is too low. Still, we can provide some circumstantial evidence that the presence of rivals increases the likelihood of having an adult son. Indeed, for a subsample of widows, we have information on the presence or absence of rivals in their latest union and on the presence or absence of one surviving son with the deceased husband (299 observations). It turns out that among widows with rivals, only 8 out of 167 (4.8 percent) have no surviving son whereas among widows without rivals, 15 out of 132 (11.4 percent) have no surviving son. The difference between the two groups is significant at 10 percent.

45. Here again, we compute these quantities for the median individual (setting $\nu = 0$).

46. We find $\sigma = 0.68$ for the whole sample and 0.75 for co-residing wives.

5.3 Robustness and placebo tests

We further perform robustness and placebo tests reported in Appendix E.

First, we disentangle the results by parity. We start by focusing on durations after birth 1. [47] As shown in Table E.2, results of the pooled Cox regression still hold when we consider only durations after birth 1 : e^{β_3} remains of the same order of magnitude and significance level. So the mechanism seems to be already at play after birth 1 and the impact of insurance strategies on fertility is detectable for the most exposed women.

Due to small sample size, we cannot investigate the issue for parities higher than one for the sample of co-residing wives. We can nevertheless do it for the whole sample. [48] Remember that, in that sample, results of the pooled Cox estimation exhibit no visible sign of son preference whether women have rivals or not. Coefficients of the estimation by parity are reported in Table E.3 for the semi-parametric estimation, and in Table E.4 for the fully parametric estimation (the baseline survival functions corresponding to each parity are plotted in the lower graph in Figure D.1, Appendix D). Both methods give the same result : the only ratio significantly different from one is e^{β_3} when $n = 3$. So women with rivals and no son tend to shorten birth spacing particularly after the birth of their third daughter. In terms of magnitude, the semi-parametric estimation predicts that having three daughters vs. at least one son reduces birth spacing by 13 months and doubles the probability of short birth spacing among women with rivals, whereas it has almost no impact among women without rivals. The fully parametric estimation is even more alarmist by predicting a reduction by 21 months and a multiplier of 4. Hence, after birth 3, the pressure seems to be so much stronger that the effect becomes visible in the whole population.

The average ideal family size among married Senegalese women is five children (ANSD and ICF International, 2012). Hence, the vast majority of women probably expect to have other "tries" for a son, when at first they have one or two daughters. The pressure to have a son would start increasing after the third "missed try," once some women have reached their

47. We did the same exercise on durations after birth 2 : estimates are line with the pooled results, even though not significant, but we should be careful with the interpretation due to the small sample size when $n = 2$.

48. Note that we reject the proportional hazard assumption when $n = 1$, both in the global test and in the tests for education and dead children variables. When we split the sample according to the median duration between births 1 and 2, we find that, indeed, the effect of those two variables are not constant in time. The lack of education has a positive impact on hazard rates only for shorter durations. And having one dead child has a negative impact on hazard rates only for larger durations : this seems consistent as the death of a child has a mechanical positive impact on the duration between births 1 and 2 only if it occurred between them – and this cannot be the case if the duration is short. But reassuringly, β_1, β_2 and β_3 are not different for shorter and longer durations.

ideal family size. We predict that the effect should keep growing at parities 4 and 5, but unfortunately, our sample size does not allow us to test this prediction.[49]

Second, we deal with our main concern and try to test for common intrinsic son preference of women with and without rivals. As already underlined, this is crucial to be able to interpret our results as the effect of an insurance strategy and not as resulting from systematic difference in preferences. We proceed in two steps, following the approach developed at the end of Section 3.3. In the first step, we explore how the risk of being dispossessed varies with the number of sons, which comes down to studying the shape of the $\pi(M)$ function. In our main specification, we break down the interaction term by number of sons s. Table E.5 shows that the impact of having at least one son does not depend on the number of sons.[50] In the theoretical framework, it means that $\Delta\pi(M_s)$ does not differ much across s while $\Delta\pi(M_0)$ is large in absolute terms. This pattern is consistent with $\delta \approx 0$ (i.e. daughters are of very little help when it comes to keeping the house) and $\pi(M)$ very convex. In other words, it suggests that having one son is enough to get rid of the bulk of the risk.

In the second step, we estimate the double difference term for women who already have a son, comparing those with exactly one son and those with more than one. If the assumption holds, it should be nil. The intuitive reasoning is the following : if one son is already enough to feel insured, women with rivals should display the same son preference as women without rivals from the moment they give birth to their first son. Restricting the sample to women of parity 2 or 3 with at least one son, we find indeed that the impact of having exactly one son vs. having more than one son is the same in the two groups (Table E.6).

Third, we investigate an alternative interpretation for higher son preference in the presence of rivals, which emphasizes the role of the husband. One story is that husbands would be more likely to divorce, or repudiate, a first wife if she had not given birth to a son ; the next wife would then feel the pressure to have a son quickly to avoid divorce or repudiation. Such a mechanism is very improbable in our case since there are slightly more boys than girls among rivals, and also slightly more ex-wives with at least one son than ex-wives with at least one daughter.

49. Jayachandran and Kuziemko (2011) have also put forward the hypothesis that gender is most predictive of subsequent fertility near ideal family size in India. They build a formal model of fertility choices under son preference, in which breastfeeding acts as a contraceptive. They predict that the probability of breastfeeding should be lowest for daughters without older brothers ; and the gender gap is expected to rise once ideal family size is reached.

50. The interaction term is significant when $s = 1$ and not when $s = 2$ or $s = 3$, because there are fewer observations with two or three sons. But in terms of magnitude, the coefficients on the three interactions terms are not significantly different from each other.

Another story assumes that men with a strong son preference would be more likely to take another wife in order to maximize the number of sons they have, and the arrival of a new wife could prompt the first one into divorcing ; $e^{\beta_3} > 1$ would simply capture the son preference of the husband, not of the wife. If this were true, we should observe evidence of son preference among the rivals, since the husband would have put pressure on his first wife as well. We have no systematic, detailed information on the timing of births of the rivals so we cannot use a duration model of birth intervals. Instead, we test a standard prediction of son-preferring, differential stopping fertility models : under son preference, an average girl is predicted to have more siblings than an average boy (Jensen, 2005, Basu and De Jong, 2010). It is not the case for the rivals in our sample : boys and girls have on average 2.5 siblings. Hence, husband's son preference does not seem to drive the results. [51]

Moreover, even if the son preference of husbands is the same in the two groups, one might still fear that women with rivals are more influenced by husbands' preference than women without rivals, because they have a lower bargaining power. It would be the case if marrying into a family where rivals exist were a sign of vulnerability. One way to rule out this story is to compare women married to the household head and women married to another member of the household. If it were about bargaining power, the effect should be driven by the most vulnerable women, those married to non household heads (living with their in-laws in particular). On the contrary, if it were about rivalry for inheritance, the effect should be driven by women married to the household head, because house ownership lies in the household head's hands. Results are reported in Table E.7. The effect is entirely driven by the wives of the household heads while it is hardly present for the other wives, which is supportive of the rivalry for inheritance story.

Fourth, we consider alternative definitions of rivals to capture potential effects of the gender composition and the number of rivals. So far, our results indicate that the presence of one son plays a crucial part in the inheritance process. If we turn the reasoning around, rivals should represent a threat (i) only when there are male rivals and (ii) as soon as there is one male rival.

We begin with defining rivals as a dummy equal to one if there is at least one male rival, thus excluding women having only female rivals from the category. As shown in the first

51. Further reassurance in this respect can be obtained by looking at the DHS data for Senegal, where both members of the couple are asked to declare their fertility preferences. Using this data on co-residing couples, we find no difference in the son preference declared by fathers in the presence and in the absence of rivals. In addition, husbands' preferences seem to be slightly more correlated with actual fertility outcome than wives' preferences, but not differentially so according to the presence or absence of rivals.

column of Table E.8, the results observed in our sample of interest – co-residing wives – still hold. Also if we compare the effect of having at least one male rival vs. at least one female rival, e^{β_3} is smaller and no longer significant in the female case. When we introduce both, the coefficient on $Girls \times Rivals\_male$ is 35 percent higher than the coefficient on $Girls \times Rivals\_female$, even though neither is significant (Table E.8).[52] This is a first hint that the threat to inheritance is constituted by male rivals, as is confirmed below.

We then test whether the effect is linear in the number of rivals (Table E.9). In order to do so, we create a variable equal to the number of rivals in inheritance shares. We find that among co-residing wives, each additional inheritance share for the rivals increases the hazard ratio between women having only daughters and women having at least one son by 23 percent (column one). However, the impact of the number of rivals is non-linear : most of the effect takes place between zero and one rival ; each additional rival raises the ratio only slightly (column two). Among women with exactly one rival, having only daughters doubles the hazard rate (column three). In accordance with our claim, the effect is entirely driven by women having one male rival ; it is zero for women having one female rival (column four).

Fifth, we check that our estimates do not capture preferences for diversity rather than son preference, by replacing the *Girls* dummy by a *Boys* one : we find that having only sons vs. at least one daughter has no significant impact on the hazard rate, either for women with or without rivals ; if anything, the ratio is always lower than one, which refutes the hypothesis of preferences for diversity (Table E.10).

Our last concern is to check that hazard ratios in our main subsample of interest – co-residing wives – are robust to alternative specifications : (i) including all parities instead of limiting up to parity 3 ; and (ii) adding controls for ethnic groups (Table E.11). If anything, when all parities are included, e^{β_1} increases slightly and borders significance (p-value=0.16), suggesting that son preference may appear at higher parities even in the absence of rivals. Nevertheless, son preference remains significantly stronger for women with rivals. Coefficients are also unchanged when we control for previous durations in the regressions by parity (not shown).

52. Both coefficients are not significantly different from each other, which can be explained by the large standard errors and by the fact that having at least one female rival does not prevent from having also one male rival.

6 Conclusion

All in all, we find that wives with rivals for inheritance need to have one son, while there is no detectable sign of son preference for wives without rivals, at least up to their third child. These findings support the idea that sons play a key role in insuring women in case of widowhood. However, the impact of insurance strategies on fertility only materializes for the most exposed women : those with no independent housing, with rivals and without a son among the first children. The cost of insurance through sons in terms of health is therefore borne by a small fraction of women : in cross-section, they represent only seven percent of our sample.[53] But, from a longitudinal perspective, the fraction of women concerned at some point in their life should be much higher given the high rates of divorce and remarriage in Senegal. For instance, among the older women in our sample (35-39 years old), 27 percent have rivals. If we further consider all women in PSF data who have lost their husband, excluding polygamous unions, it appears that, eventually, one third face rivals.

For these women and their children, the cost is huge. We showed that the probability of short birth spacing was multiplied by 1.5 to 2 ; and we know from the medical literature that short birth spacing multiplies the risk of infant death by 2.4 (ANSD and ICF International, 2012) and doubles the risk of maternal death (Conde-Agudelo and Belizan, 2000).

In Senegal, cutting down maternal and child mortality as well as implementing birth control are two key objectives of the national health plan. Our results suggest that reducing women's reliance on children to eventually disassociate fertility choices from insurance considerations could help achieve these objectives. The general consensus is that improving women's status is a precondition for the fertility transition. In a chapter entitled "Women's agency and social change," Amartya Sen argues that women's empowerment will bring about lower fertility rates and lower infant mortality (Sen, 1999). However, Sen maintains that the main drivers of fertility transition are (i) a change in women's aspirations and bargaining power, and (ii) access to birth control methods. We argue that both conditions are necessary, but not sufficient : as long as women have no alternative insurance strategies, they will continue to have many children to mitigate downside risks, starting with widowhood.

One might think of three, non-mutually exclusive, ways ahead. First, reforming inheritance rules to grant wives, irrespective of the number of children, a significant percentage of the husband's wealth, or at least the usufruct of the house for some years. Second, expanding social benefits to protect a larger share of the population : in the Senegalese case, it would be

53. 75 percent of women currently live with their husband and among them, 18 percent have rivals and 50 percent have a female firstborn.

broadening to all widows the current system of pensions paid to the widows of civil servants. Third – but in the longer run, given the difficulties encountered by insurance companies in developing countries – promoting formal insurance and financial markets.

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References

Agarwal, Bina. 1994. *A Field of One's Own : Gender and Land Rights in South Asia.* Cambridge : Cambridge University Press.

Agence Nationale de la Statistique et de la Démographie (ANSD) and ICF International. 2012. "Demographic and Health Survey – Multiple Indicator Cluster Survey (DHS-MICS) 2010–2011." Calverton, Maryland : ANSD et ICF International.

Akiiki Asiimwe, Florence. 2009. "Statutory Law, Patriarchy and Inheritance :Home ownership among Widows in Uganda." *African Sociological Review* 13(1) 124–142.

Antoine, Philippe, and Fatou Binetou Dial. 2003. "Mariage, divorce et remariage à Dakar et Lomé." Actes du colloque de l'AUF : Familles du Nord, familles du Sud, Marseille 23–26 juin 2003.

Appleton, Simon. 1996. "Women-Headed Households and Household Welfare : An Empirical Deconstruction for Uganda." *World Development* 24(12) 1811–1827.

Arnold, Fred, Minja K. Choe, and T.K. Roy. 1998. "Son Preference, the Family-Building Process

and Child Mortality in India.” *Population Studies* 52(3) 301–315.
Basu, Deepankar, and Robert de Jong. 2010. ”Son targeting fertility behavior : Some consequences and determinants.” *Demography* 472 : 521–536.
Bledsoe, Caroline. 1990. ”Transformations in Sub-Saharan African Marriage and Fertility.” *Annals*, AAPSS 510 : 115–125.
Box-Steffensmeier, Janet, and Bradford Jones. 2004. ”Timing and Political Change : Event History Modeling in Political Science.” Ann Arbor : University of Michigan Press.
Cain, Mead. 1986. ”The Consequences of Reproductive Failure : Dependence, Mobility, and Mortality among the Elderly of Rural South Asia.” *Population Studies : A Journal of Demography* 40(3) 375–388.
Carranza, Eliana. 2012. ”Islamic inheritance law, son preference and fertility behavior of Muslim couples in Indonesia. ” World Bank Policy Research WP 5972, World Bank, Washington DC.
Conde-Agudelo, Agustin, and Jose M. Belizan. 2000. ”Maternal Morbidity and Mortality Associated with Interpregnancy Interval : Cross Sectional Study.” *British Medical Journal* 321 1255–1259.
Conde-Agudelo, Agustin, Anyeli Rosas-Bermúdez, and Ana Cecilia Kafury-Goeta. 2006. ”Birth Spacing and Risk of Adverse Perinatal Outcomes : A Meta-analysis.” *The Journal of the American Medical Association*, 295(15) 1809–1823.
Cox, David Roxbee. 1972. ”Regression Models and Life-Tables (with discussion).” *Journal of the Royal Statistical Society Series B* 34 : 187–220.
DaVanzo, Julie, Abdur Razzaque, Mizanur Rahman, and Lauren Hale. 2004. ”The Effects of Birth Spacing on Infant and Child Mortality, Pregnancy Outcomes, and Maternal Morbidity and Mortality in Matlab, Bangladesh.” Rand WP, WR-198. Santa Monica, California : Rand.
De Vreyer, Philippe, and Sylvie Lambert. 2013. ”Household Risk Management in Senegal.” Background paper for the 2014 World Development Report, *Managing Risk for Development*, World Bank.
De Vreyer, Phillippe, Sylvie Lambert, Abla Safir, and Momar B. Sylla. 2008. ”Pauvreté et structure familiale, pourquoi une nouvelle enquête ?” *Stateco*, 102, 261–275.
Gangadharan, Lata, and Pushkar Maitra. 2003. ”Testing for Son Preference in South Africa.” *Journal of African Economies* 12(3) 371–416.
Ghilagaber, Gebrenegus, and Stephen Obeng Gyimah. 2004. ”A Family of Flexible Parametric Duration Functions and their Applications to Modelling Child-Spacing in Sub-Saharan Africa.” *PSC Discussion Papers Series* 18(1).
Grambsch, Patricia M., and Terry M. Therneau. 1994. ”Proportional hazards tests and diagnostics based on weighted residuals.” *Biometrika* 81(3) 515–526.
Haughton, Dominique, and Jonathan Haughton. 1995. ”Son Preference in Vietnam.” *Studies in Family Planning* 26(6) 325–337.
Hoddinott, John. 1992. ”Rotten Kids or Manipulative Parents : Are Children Old Age Security in

Western Kenya ?" *Economic Development and Cultural Change* 40(3) University of Chicago Press, 545–65.
Horrell, Sara, and Pramila Krishnan. 2007. "Poverty and Productivity in Female-Headed Households in Zimbabwe." *Journal of Development Studies* 43(8) 1351–1380.
Jakiela, Pamela, and Owen Ozier. 2015. "Does Africa Need a Rotten Kin Theorem ? Experimental Evidence from Village Economies." *Review of Economic Studies*, forthcoming.
Jayachandran, Seema, and Ilyana Kuziemko. 2011. "Why Do Mothers Breastfeed Girls Less Than Boys ? Evidence and Implications for Child Health in India." *The Quarterly Journal of Economics* 126 : 1485–1538.
Jensen, Robert. 2005. "Equal treatment, unequal outcomes ? Generating sex inequality through fertility behavior." Unpublished document. JFK School of Government, Harvard University.
Kaplan, E.L., and Paul Meier. 1958. "Nonparametric estimation from incomplete observations." *Journal of the American Statistical Association* 53 : 457–481.
Kevane, Michael. 2004. "Women and devlopment in Africa : how gender works." Lynne Rienner Publishers, Inc.
Lambert, Sylvie, and Dominique van de Walle. 2012. "Transcription des entretiens menés avec des veuves et des divorcées au Sénégal en Novembre 2011." Unpublished document, World Bank.
Lambert, Sylvie, Martin Ravallion, and Dominique van de Walle. 2014. "Intergenerational Mobility and Interpersonal Inequality in an African Economy." *Journal of Development Economics* 110 : 327–344
Larsen, Ulla, Woojin Chung, and Monica Das Gupta. 1998. "Fertility and son preference in Korea." *Population Studies* 52 : 317–325.
Lesthaeghe, Ron J. (editor). 1989. "Reproduction and Social Organization in Sub-Saharan Africa." Berkeley : University of California Press.
Leung, Siu Fai. 1988. "On tests for sex preferences." *Journal of Population Economics* 1 : 95–114.
Mace, Ruth, and Rebecca Sear. 1997. "Birth interval and the sex of children in a traditional African population : An evolutionary analysis." *Journal of Biosocial Science* 29(4) 499–507.
Milazzo, Annamaria. 2014a. "Son Preference, Fertility and Family Structure : Evidence from Reproductive Behavior among Nigerian Women." World Bank Policy Research WP 6869, World Bank, Washington DC.
Milazzo, Annamaria. 2014b. Why are Adult Women Missing ? Son Preference and Maternal Survival in India", World Bank Policy Research WP 6802, World Bank, Washington DC.
Nugent, Jeffrey B. 1985. "The Old-Age Security Motive for Fertility." *Population and Development Review* 11(1) 75–97.
Pong, Suet-ling. 1994. "Sex preference and fertility in Peninsular Malaysia." *Studies in Family Planning* 25(3) 137–48.
RADI. 2010. "Livret d'information juridique : les successions en droit sénégalais", Réseau Africain

pour le Développement Intégré, Ministère de la justice.

Rahman, Mizanur, and Julie DaVanzo. 1993. "Gender Preference and Birth Spacing in Matlab, Bangladesh." *Demography* 30(3) 315-332.

Rosenzweig, Mark R. 1988. "Risk, Private Information, and the Family." *The American Economic Review*, 782. Papers and Proceedings of the One-Hundredth Annual Meeting of the American Economic Association 245–250.

Ronsmans, C. 1996. "Birth spacing and child survival in rural Senegal." *International Journal of Epidemiology* 25(5) 989–97.

Schoenfeld, David. 1980. "Chi-squared goodness-of-fit tests for the proportional hazards regression model." *Biometrika* 67(1) 145–153.

Sen, Amartya. 1999. *Development as Freedom.*, Oxford : Oxford University Press.

Strassmann, Beverly. 1997. "Polygyny as a Risk Factor to Child Mortality among the Dogon." *Current Anthropology* (38) 688–695.

Tsay, Wen-Jen, and C.Y. Cyrus Chu. 2005. "The pattern of birth spacing during Taiwan's demographic transition." *Journal of Population Economics* 18 : 323–336.

Tu, Ping. 1991. "Birth spacing patterns and correlates in Shaanxi, China." *Studies in Family Planning* 22(4) 255–63.

United Nations. 2011. "Widowhood : invisible women, secluded or excluded." United Nations publication.

Vaughan, Megan. 1987. *The Story of an African Famine : Gender and Famine in Twentieth-Century Malawi.*, Cambridge : Cambridge University Press.

van de Walle, Dominique. 2013. "Lasting Welfare Effects of Widowhood in Mali." *World Development* 51 : 1–19.

Figure 1: Non-parametric estimation of survival functions (pooled and separated by parity and by sub-population)

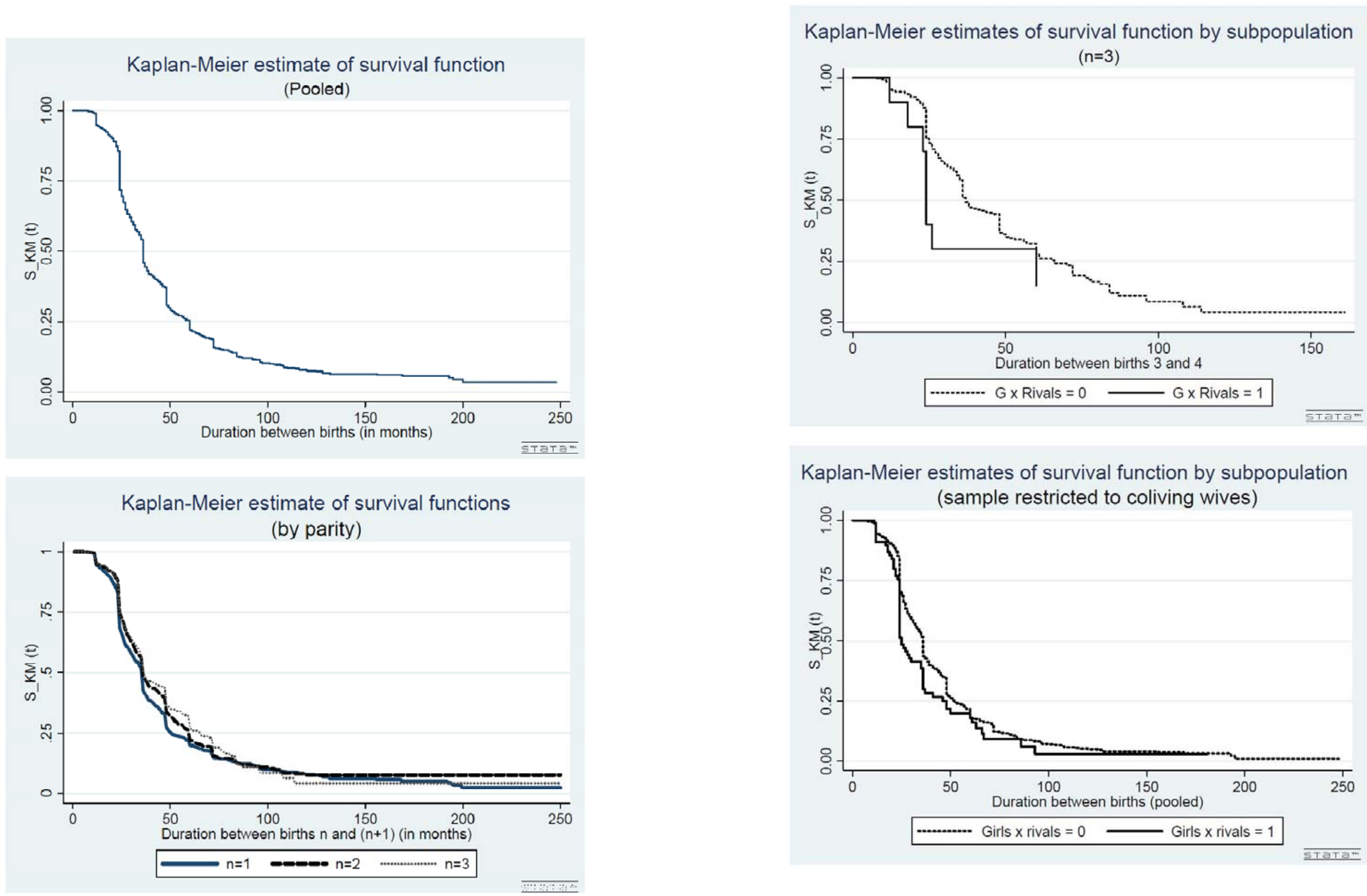


Table 1: Widows' characteristics, by place of residence after the husband's death

		Remained in the same household	Hosted by another household
Observations		345	172
Average age at husband's death		49.4	37.7
Had at least one son with the deceased husband†		89%	77%
Still in the household (at the date of the survey)		91%	24%
If not : median length of stay (in years)		7	2
Household head	The widow herself	45%	19%
	A child	29%	11%
	A parent	4%	26%
	A sibling	4%	12%
	Other	18%	32%
Remarried		19%	49%
If remarried, co-residing (with the new husband)		35%	68%

Sample : women whose latest broken union ended with the husband's death (statistics computed for the latest period of widowhood) ; only 20 percent of them have experienced more than one broken union.

† The proportion of women having at least one son with the deceased husband is computed on the sub-sample of widows leaving with at least one child and having exactly one broken union ; otherwise, we have no information on the gender composition of children born during the previous marriage (346 observations). Using these numbers, one infers that 71% of women having one son remained in the husband's house, whereas 50% of women having only daughters did.

Table 2: Descriptive statistics : Sample of women under 39 years old, engaged in a monogamous union, having at least one child from current union

	All women	**Women with no missing observations**
Number of observations	936	761
Number of children from current union	2.7 (1.63)	2.7 (1.64)
At least one child from previous unions	4.1%	3.9%
At least one dead child (from current union)	19.8%	19.8%
Age	27.9 (6.02)	28.0 (6.05)
Rural	48.4%	47.6%
No education	43.8%	43.9%
Work	32.8%	32.8%
First union	86.6%	88.3%
If not : nb broken unions	1.1 (0.38)	1.1 (0.35)
Age at first marriage	18.9 (4.30)	19.0 (4.35)
Live with their husband	75.3%	77.9%
If co-residing : Age difference	10.2 (6.54)	10.2 (6.33)
If co-residing : Public sector	7.3%	7.8%
If co-residing : Annual income K FCFA (median)	500	500
At least 1 rival	17.1%	16.8%
At least 1 male rival	11.9%	11.3%
Average nb of rivals (inheritance shares)	1.7 (min 0.5 ; max 7.5 ; s.d. 1.38)	1.6 (min 0.5 ; max 7.5 ; s.d. 1.38)
Strictly more rivals than children (inheritance shares)	9.7%	9.1%

Standard deviations are in parentheses. The second column presents statistics on women for whom there is no missing observation ; we perform the semi- and fully parametric estimations on this subsample.

TABLE 3: Descriptive statistics : Comparisons

	A. Wives with and without rivals			**B. Co-residing and non co-residing wives**		
	At least one rival	**No rival**	**p-value**	**Co-residing**	**Non co-residing**	**p-value**
Number of observations	159	770		705	231	
Number of children from current union	3.1	2.6	0.01***	2.8	2.1	0.01***
At least one child from previous unions	10.7%	2.6%	0.01***	3.4%	6.0%	0.12
At least one dead child (from current union)	26.7%	18.5%	0.04**	21.2%	15.2%	0.04**
Interval between births 1 and 2 (in months)	33.3	36.8	0.10*	35.3	40.6	0.05**
Age	30.2	27.5	0.01***	27.9	28.0	0.93
Rural	44.7%	49.5%	0.27	48.9%	46.8%	0.56
No education	40.6%	44.6%	0.37	44.0%	43.0%	0.79
Work	40.5%	30.8%	0.02**	32.8%	32.9%	0.98
First union	78.5%	88.8%	0.01***	92.9%	66.5%	0.01***
If not : nb broken unions	1.2	1.05	0.14	1.0	1.1	0.08*
Age at first marriage	19.2	18.8	0.29	18.7	19.4	0.04**
Live with their husband	81.2%	74.8%	0.07*			
If co-residing : Age difference	13.7	9.5	0.01***			
If co-residing : Public sector	8.7%	7.0%	0.52			
If co-residing : Annual income K FCFA (median)	500	500	0.83			
At least 1 rival				18.3%	13.4%	0.07*
At least 1 male rival				12.9%	8.9%	0.08*
Average nb rivals (inheritance shares)				1.7	1.7	0.81

We conduct t-tests comparing the means of each covariate, and a nonparametric equality-of-medians test for the income : a low p-value indicates that both subsamples are statistically different with respect to the corresponding covariate.

*** Significant at the 1 percent level. **Significant at the 5 percent level. * Significant at the 10 percent level.

The total number of observations in panel A is different from 936 because there are 7 missing values for the presence of rivals.

TABLE 4: Birth intervals and proportion of short birth spacing, by gender of first born children and presence of rivals

Birth intervals in months	$Girls = 0$	$Girls = 1$	Difference
$Rivals = 0$	34.9	36.5	1.6
	(0.89)	(1.14)	(1.44)
nb obs.	519	314	
$Rivals = 1$	37.1	31	-6.1**
	(1.78)	(2.54)	(3.1)
nb obs.	128	63	
Difference	2.2	-5.5**	**-7.7****
	(1.99)	(2.78)	(3.42)
Proportion of birth interval shorter than 24 months	$Girls = 0$	$Girls = 1$	Difference
$Rivals = 0$	0.376	0.369	-0.006
	(0.02)	(0.03)	(0.03)
nb obs.	519	314	
$Rivals = 1$	0.344	0.556	0.212***
	(0.04)	(0.06)	(0.74)
nb obs.	128	63	
Difference	-0.032	0.187***	**0.218*****
	(0.05)	(0.07)	(0.08)

Standard errors are in parentheses. Significance levels :*** Significant at the 1 percent level. **Significant at the 5 percent level. * Significant at the 10 percent level.

Girls : first children are only girls ; *Rivals* : there is at least one rival. Sample of co-residing wives. Non-censored durations.

TABLE 5: Statistics on non-censored durations

	Pooled	Parity #1	Parity #2	Parity #3
Total number obs	*1939*	*904*	*644*	*391*
Number non-censored obs	1250	634	395	221
Number of months				
Mean	36.2	36.4	35.9	36.2
Std dev.	21.3	23.3	18.5	19.8
Min	8	8	8	8
Q1	24	24	24	24
Q2	30	30	31	30
Q3	45	44	47	48
Max	200	200	120	114
Number obs $\geq$ 96 months	29	20	5	4

Statistics computed on non-censored durations, expressed in months.

Approx. 2% of birth intervals are larger than 8 years : in half of the cases, women have in fact lost 1 or 2 children in the meantime. The remaining cases are mostly urban, educated, working women living in relatively rich households. Results are qualitatively unchanged if we exclude extreme values (above 96 months) from the sample.

Table 6: Cox estimation

Hazard ratios $= e^{\beta_k}$	Whole sample (1)	Co-residing wives (2)	Poorest half (3)	Richest half (4)
Girls	1.00	1.07	1.08	1.04
	(.069)	(.088)	(.118)	(.137)
Rivals	1.03	0.94	0.78	1.01
	(.108)	(.112)	(.133)	(.164)
Girls $\times$ *Rivals*	1.23	1.54**	2.26***	1.21
	(.217)	(.334)	(.612)	(.395)
Controls	Yes	Yes	Yes	Yes
Log-likelihood	-6569	-4607	-2111	-1956
Observations	1589	1117	569	548
Clusters	761	515	260	255
Obs in cell : *Girls* $\times$ *Rivals* $= 1$	93	64	36	28

Robust standard errors of e^{β_k} are in parentheses (clustered at the woman level) - Significance levels (for $e^{\beta} \neq 1$) :*** Significant at the 1 percent level. **Significant at the 5 percent level. * Significant at the 10 percent level. Breslow method to handle ties among non-censored durations.

Column (2) : sample restricted to wives living with their husband ; this sample is then split on the median income into the poorest half in column (3) and the richest half in column (4).

Girls : first children are only girls ; *Rivals* : there is at least one rival. Controls include characteristics of the woman : age at preceding birth, place of residence, level of education, age at first marriage, being or not in her first marriage, living or not with the husband, having at least one dead child, having at least one child from previous unions, rank of preceding birth. In columns (2) to (4), we include additional controls related to the husband : sector of activity of the husband, income of the husband, age difference with the husband as well as the occupation of the wife.

Table 7: Quantities of interest derived from Cox estimation

Co-residing wives	*Girls* $= 0$ *Rivals* $= 0$	*Girls* $= 1$ *Rivals* $= 0$	*Girls* $= 0$ *Rivals* $= 1$	*Girls* $= 1$ *Rivals* $= 1$
Expected birth spacing in months	38.2	36.7	39.5	29.8
Probability of birth spacing ≤ 24 months	33.0%	35.9%	32.4%	47.6%

Poorest half (of co-residing wives)	*Girls* $= 0$ *Rivals* $= 0$	*Girls* $= 1$ *Rivals* $= 0$	*Girls* $= 0$ *Rivals* $= 1$	*Girls* $= 1$ *Rivals* $= 1$
Expected birth spacing in months	36.0	35.1	38.5	27.5
Probability of birth spacing ≤ 24 months	33.8%	36.0%	27.5%	54.3%

Expected birth spacing $= \mathbb{E}(T \mid X, T \leq t^*_{(k)})$. Probability of birth spacing ≤ 24 months $= 1 - \widehat{S}(24 \mid X)$

We use the baseline survival functions estimated in the Cox model. We set the values of the controls to the median of the sample. We check that results are very similar if we set the values of the controls to the sample mean (not shown).

TABLE 8: Fully parametric estimation

Hazard ratios $= e^{\beta_k}$	Whole sample (1)	Co-residing wives (2)	Poorest half (3)	Richest half (4)
Girls	1.01	1.03	1.19	0.97
	(.106)	(.124)	(.195)	(.160)
Rivals	0.97	1.02	0.69	1.08
	(.147)	(.179)	(.172)	(.288)
Girls $\times$ *Rivals*	1.60**	2.29***	3.87***	1.79
	(.387)	(.653)	(1.440)	(.863)
Controls	Yes	Yes	Yes	Yes
Log-likelihood	-2950	-2134	-908	-1203
Observations	761	515	260	255

Standard errors of e^{β_k} are in parentheses. Significance levels (for $e^{\beta} \neq 1$) : *** Significant at the 1 percent level. **Significant at the 5 percent level. * Significant at the 10 percent level.

Column (2) : sample restricted to wives living with their husband; this sample is then split on the median income into the poorest half in column (3) and the richest half in column (4).

Girls : first children are only girls; *Rivals* : there is at least one rival. Controls include characteristics of the woman : age at preceding birth, place of residence, level of education, age at first marriage, being or not in her first marriage, living or not with the husband, having at least one dead child, having at least one child from previous unions, rank of preceding birth. In columns (2) to (4), we include additional controls related to the husband : sector of activity of the husband, income of the husband, age difference with the husband as well as the occupation of the wife.

TABLE 9: Quantities of interest derived from fully parametric estimation

Co-residing wives	$Girls = 0$, $Rivals = 0$	$Girls = 1$, $Rivals = 0$	$Girls = 0$, $Rivals = 1$	$Girls = 1$, $Rivals = 1$
Expected birth spacing in months	38.1	37.6	37.8	27.1
Probability of birth spacing $\leq$ 24 months	20.5%	21.1%	20.8%	42.7%

Poorest half (of co-residing wives)	$Girls = 0$, $Rivals = 0$	$Girls = 1$, $Rivals = 0$	$Girls = 0$, $Rivals = 1$	$Girls = 1$, $Rivals = 1$
Expected birth spacing in months	35.5	33.3	40.8	23.6
Probability of birth spacing $\leq$ 24 months	22.8%	26.4%	16.4%	56.1%

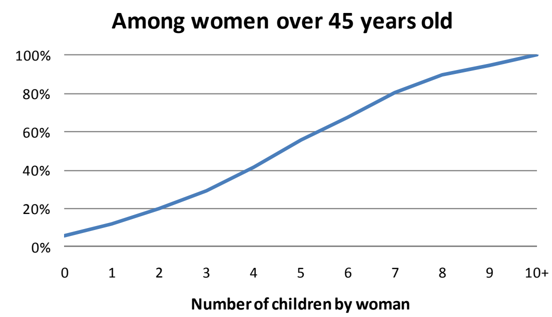
Expected birth spacing $= \mathbb{E}(T \mid X, \nu = 0, \hat{\theta})$. Probability of birth spacing ≤ 24 months $= 1 - S(24 \mid X, \nu = 0, \hat{\theta})$

We use the survival function and the parameters estimated in the fully parametric specification. We set the values of the controls to the median of the sample. We check that results are very similar if we set the values of the controls to the sample mean (not shown).

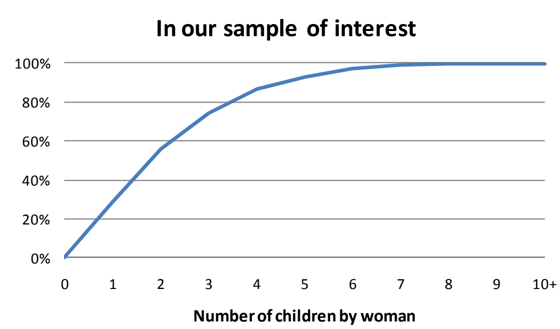
Appendix : For online publication

Appendix A : Additional descriptive statistics

Figure A.1: Cumulative distribution of the number of children by woman

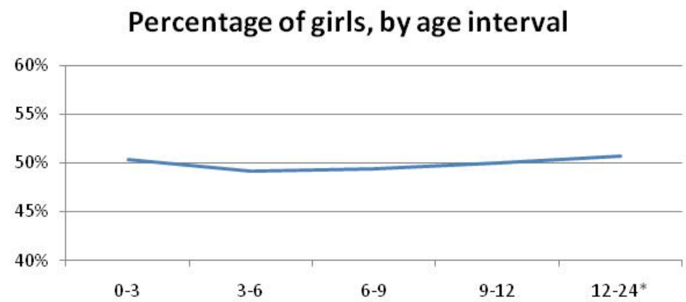


Sample: Women over 45 years old in the PSF survey (1060 observations)
Distribution in line with national statistics reported in DHS-MICS, 2010-11.



Sample: Women below 39 years old, engaged in a monogamous union, having at least one child from current union (936 observations).

Figure A.2: Sex ratio by age, among children from current union



Sample: All children from current union of the 936 women in our baseline sample.
Percentages are unchanged if we include the 72 children from previous unions.

* The last age interval is larger to reach a sufficient number of observations (see below).

Age interval	0-3 years	3-6 years	6-9 years	9-12 years	12-24 years	Total
Observations	739	632	441	276	363	2451

Table A.1: Number of observations : Breakdown by missing variables and cells of interest

		Pooled	Parity #1	Parity #2	Parity #3
Initial number obs		**2016**	**936**	**666**	**414**
Twins		27	9	7	11
Missing duration		50	23	15	12
Number observations	**(non-parametric estimation)**	**1939**	**904**	**644**	**391**
At least 1 missing obs		350	167	116	67
	sex children	2	2	0	0
	rivals	20	10	6	4
Missing	education	68	29	23	16
	age first marriage	99	54	30	15
	first union	125	56	45	24
	dead children	112	50	41	21
Number observations	**(parametric estimation)**	**1589**	**737**	**528**	**324**
	$Girls = 1$	548	365	140	43
Among which :	$Rivals = 1$	270	112	91	67
	$Girls \times Rivals$=1	93	57	26	10

The ”pooled” column brings together all birth intervals, which are then split by parity ; the 1589 observations correspond to 761 different women. But the duration between children 1 and 2 is missing for a few of them, that is why we have only 737 observations when $n = 1$. Sample : Women below 39, in a monogamous union, having at least one child from current union.

TABLE A.2: Balancing tests : Having only daughters vs. having at least one son

	Parity #1			Parity #2			Parity #3		
	G=0	**G=1**	**p-value**	**GG=0**	**GG=1**	**p-value**	**GGG=0**	**GGG=1**	**p-value**
Nb observations	372	365		388	140		281	43	
Nb children (in current union)	2.6	2.7	0.39	3.3	3.5	0.17	4.1	4.4	0.13
Children from previous unions ≥ 1	4.0%	3.8%	0.89	2.6%	5%	0.23	2.8%	2.3%	0.84
Dead children (in current union) ≥ 1	18.8%	20.8%	0.50	25%	18.6%	0.11	26%	34.9%	0.25
Age	27.9	28.1	0.59	29.3	29.6	0.60	31.1	30.8	0.71
Rural	49.7%	44.3%	0.15	50.3%	50.7%	0.93	52.7%	58.1%	0.50
No education	45.7%	42.7%	0.42	45.6%	43.6%	0.68	46.3%	55.8%	0.25
First union	90.1%	86.6%	0.14	91.5%	85.7%	0.08*	92.2%	97.7%	0.05**
If not : nb broken unions	1.1	1.05	0.62	1.1	1.05	0.87	1.1	1.0	0.08*
Age at first marriage	19.0	19.0	0.89	18.7	18.3	0.28	18.4	18.2	0.76
At least 1 rival	15.3%	16.7%	0.61	17.5%	18.6%	0.78	21%	25.5%	0.52
At least 1 male rival	9.1%	12.0%	0.20	11.1%	15%	0.25	12.8%	18.6%	0.36
Average nb rivals (in inheritance shares)	1.4	1.8	0.13	1.5	2.1	0.09*	1.6	2.0	0.41
Currently working	32.8%	32.6%	0.96	32.9%	32.1%	0.86	33.3%	27.9%	0.47
Co-reside with the husband	76.3%	79.4%	0.31	82.0%	85.0%	0.40	83.6%	88.4%	0,38
If co-residing : Age difference	10.0	10.4	0.36	10.1	10.8	0.28	10.2	12.2	0.07*
If co-residing : Husband in public sector	6.6%	9.5%	0.21	6.9%	8.6%	0.57	4.8%	8.1%	0.49
If co-residing : Annual income K FCFA (median)	480	600	0.07*	500	600	0.50	500	310	0.43

We conduct t-tests comparing the means of 16 covariates, and a nonparametric equality-of-medians test for the income : a low p-value indicates that both subsamples are statistically different with respect to the corresponding covariate. *** Significant at the 1 percent level. **Significant at the 5 percent level. * Significant at the 10 percent level.

A F-test rejects the joint significance across variables : the p-values are 0.56 (Parity #1), 0.33 (Parity #2) and 0.82 (Parity #3).

G, GG, GGG : first n children are girls.

Appendix B : Measures of birth intervals

We have four types of observations. Table B.1 below gives further detail about the number of observations of each type :

1. Not censored, precisely measured ; we observe :

$$t_n = (\text{birth year of child } (n+1) - \text{birth year of child } n) \times 12 + \text{birth month of child } (n+1) - \text{birth month of child } n$$

   So depending on the birth days of both children, we infer that $T_n \in [t_n - 1; t_n + 1]$.

2. Not censored, imprecisely measured ; we observe :

$$t_n = (\text{age of child } n - \text{age of child } (n+1)) \times 12$$

   So depending on the exact birth dates of both children, we infer that $T_n \in [t_n - 12; t_n + 12]$ [54].

3. Censored, precisely measured ; we observe :

$$t_n = (\text{year of survey - birth year of child } n) \times 12 + \text{month of survey - birth month of child } n$$

   We infer that $T_n > t_n - 1$.

4. Censored, imprecisely measured ; we observe :

$$t_n = (\text{age of child } n) \times 12$$

   We infer that $T_n > t_n$.

54. E.g. if one child is 4 years old and the next one is 2 years old, we measure $t = 24$ and the two extreme cases are given by (i) the eldest will turn 5 the day following the survey while the youngest turned 2 the day before : the real t is 36 ; and (ii) the youngest will turn 3 the day following the survey while the eldest turned 4 the day before : the real t is 12.

Table B.1: Number and frequence of observations by censoring and measurement status

Pooled	Precisely measured	Imprecisely measured	Total
Non-censored	621 (39%)	413 (26%)	1034 (65%)
Censored	441 (28%)	114 (7%)	555 (35%)
Total	1062 (66%)	527 (34%)	1589 (100%)

n=1	Precisely measured	Imprecisely measured	Total
Non-censored	314 (43%)	203 (28%)	517 (70%)
Censored	176 (24%)	44 (6%)	220 (30%)
Total	490 (66%)	247 (34%)	737 (100%)

n=2	Precisely measured	Imprecisely measured	Total
Non-censored	198 (37%)	132 (25%)	330 (63%)
Censored	156 (30%)	42 (8%)	198 (38%)
Total	354 (67%)	174 (33%)	528 (100%)

n=3	Precisely measured	Imprecisely measured	Total
Non-censored	109(34%)	78 (24%)	187 (57%)
Censored	109(34%)	28 (9%)	137 (43%)
Total	218 (67%)	106 (33%)	324 (100%)

A duration is considered as imprecisely measured when it is derived from the age of the child instead of the exact birthdate.

Appendix C : Fully parametric duration model

We start by defining a dummy R_n for long-term survivors at parity n, i.e. women who stop having children after the n^{th} birth. The probability of stopping is likely to depend on the characteristics of women after birth n summarized in the vector X_n ; we estimate the probability with a logit :

$$Pr(R_n = 1 \mid X_n) = p(X_n) = \frac{\exp(X_n'\alpha)}{1 + \exp(X_n'\alpha)}$$

For the women who have another child, the hazard function is now given by :

$$\lambda(t \mid X_n, \nu) = \lambda_0(t) \times \exp(X_n'\beta + \nu)$$

Where :

- $\nu \sim \mathcal{N}(0, \sigma)$: ν has a normal distribution such that $\mathbb{E}(\nu) = 0$ and $Var(\nu) = \sigma^2$.
- The baseline hazard function has a Weibull distribution with parameters λ and a.

$$\lambda_0(t) = \begin{cases} \lambda a(\lambda(t-8))^{a-1} & \text{if } t \geq 8 \\ 0 & \text{if } t < 8 \end{cases}$$

This hazard function replicates two characteristics of the durations observed in our sample : (i) there is no exit before 8 months, and (ii) the inverted U-shape of the exit rate arises from the combination of an increasing baseline hazard together with unobserved heterogeneity.

The survival function conditional on R_n, X_n and ν allows us to retrieve the aggregate survival function conditional on X and ν.

$$S_1(t \mid R_n = 1, X_n, \nu) = 1$$

$$S_1(t \mid R_n = 0, X_n, \nu) = S(t \mid X_n, \nu) = \begin{cases} \exp\Big(-(\lambda(t-8))^a \exp(X_n'\beta + \nu)\Big) & \text{if } t \geq 8 \\ 1 & \text{if } t < 8 \end{cases}$$

We derive the aggregate survival function conditional on X and ν :

$$S_1(t \mid X_n, \nu) = p(X_n) + (1 - p(X_n)) \times S(t \mid X_n, \nu)$$

Therefore, $S_1(t \mid X_n) \to p(X_n) > 0$ when $t \to \infty$. Using this specification, we are able to disentangle the effect of covariates, in particular *Girls*, *Rivals* and *Girls* $\times$ *Rivals*, on the number of births and on the spacing of births. The vector α captures the impact on the probability to stop having children, whereas the vector β captures the impact on durations.

We estimate a multispell model in which the individual contribution to the likelihood $L_i(\theta \mid t_i, X_i, \nu_i)$ is given by :[55]

1. $p(X_{1,i}) + (1 - p(X_{1,i})) \times s(t_{1,i} \mid X_{1,i}, \nu_i)$ if woman i has exactly 1 child.
2. $\Big((1 - p(X_{1,i})) \times f(t_{1,i} \mid X_{1,i}, \nu_i)\Big) \times \Big(p(X_{2,i}) + (1 - p(X_{2,i})) \times s(t_{2,i} \mid X_{2,i}, \nu_i)\Big)$ if woman i has exactly 2 children.
3. $\Big((1 - p(X_{1,i})) \times f(t_{1,i} \mid X_{1,i}, \nu_i)\Big) \times \Big((1 - p(X_{2,i})) \times f(t_{2,i} \mid X_{2,i}, \nu_i)\Big) \times \Big(p(X_{3,i}) + (1 - p(X_{3,i})) \times s(t_{3,i} \mid X_{3,i}, \nu_i)\Big)$ if woman i has exactly 3 children.
4. $\Big((1 - p(X_{1,i})) \times f(t_{1,i} \mid X_{1,i}, \nu_i)\Big) \times \Big((1 - p(X_{2,i})) \times f(t_{2,i} \mid X_{2,i}, \nu_i)\Big) \times \Big((1 - p(X_{3,i})) \times f(t_{3,i} \mid X_{3,i}, \nu_i)\Big)$ if woman i has 4 children and more.

Where $t_{n,i}$ is the duration (potentially right-censored) between births n and $(n+1)$ for woman i ; $f(.)$ and $s(.)$ are defined to explicitly allow for measurement approximations :

$$f(t_{n,i}) = (S(t_{n,i} - 1) - S(t_{n,i} + 1))^{P_{n,i}} \times (S(t_{n,i} - 12) - S(t_{n,i} + 12))^{(1 - P_{n,i})}$$

$$s(t_{n,i}) = S(t_{n,i} - 1)^{P_{n,i}} \times S(t_{n,i})^{(1 - P_{n,i})}$$

Where $P_{n,i}$ is a dummy equal to 1 if the duration n is precisely measured for woman i. Hence, the contribution to the likelihood uses a confidence interval that depends on how precisely the duration is measured, instead of the density and survival functions.

Now, $\theta \equiv \{\alpha, \beta, \lambda, a, \sigma\}$ is the vector of parameters to be estimated. In our main specification, we constrain the parameters to be the same for all parities and compare the estimates to those found in the pooled Cox regression. Then, in robustness checks by parity, we allow the parameters[56] to vary depending on n, and compare the estimates to those found in the three separate Cox regressions. However, in the last specification, we constrain the model in order to keep enough degrees of freedom. In the logit, we keep only the strongest predictors of the probability to have another child : place of residence, being or not in her first marriage, living or not with the husband. In the survival function, we keep our variables of interest and the following controls : age at birth

55. The level of analysis is no longer the births as in the Cox model, but the women. Here, the number of observations is exactly equal to the number of clusters in the Cox model (=761).

56. Except σ, the standard deviation of the unobserved heterogeneity. This is precisely the strength of the multispell model : random draws from the normal distribution are attributed once and for all to each individual, explaining simultaneously all durations.

n, level of education, age at first marriage, living or not with the husband, having at least one dead child, having at least one child from previous unions.

To compute the standard errors of e^{β_k}, we use the following approximation :

$$se(e^{\beta_k}) = e^{\hat{\beta}_k} \times se(\beta_k)$$

Since there is no closed form expression for the expected likelihood, we use the technique of simulated maximum likelihood. We estimate $\mathbb{E}(L_i(\theta \mid t_i, X_i))$ with :

$$\widetilde{L}_i(\theta \mid t_i, X_i) = \frac{1}{H} \sum_{h=1}^{H} L_i(\theta \mid t_i, X_i, \nu_{i,h})$$

Where $H = 60$ is the number of random draws from a normal distribution for each individual.

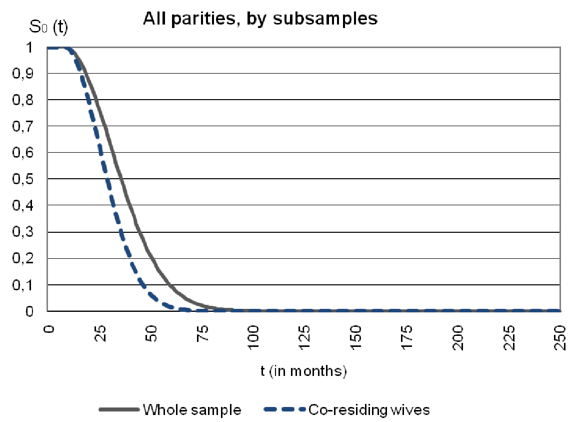
Last, we are able to compare across different categories of women, the probability of having another child, $(1 - p(X))$, and the expected duration conditional on X :

$$\mathbb{E}(T \mid X, \nu, \hat{\theta}) = \int_0^{\infty} S(u \mid X, \nu, \hat{\theta}) du$$

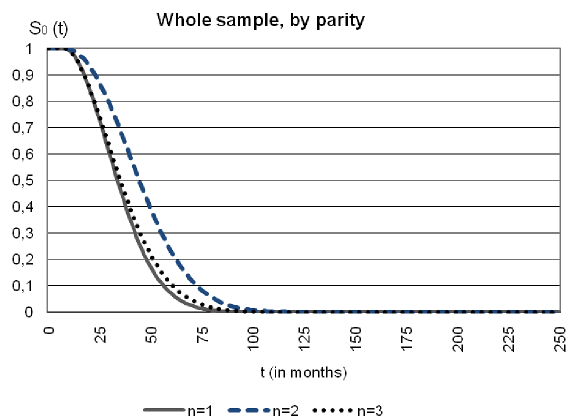
We also consider $\Big(1 - S(24 \mid X, \nu, \hat{\theta})\Big)$. Again, we compute those quantities for the median individual (setting $\nu = 0$) and compare them to the ones obtained with the Cox estimation.

Appendix D : Additional estimates of the fully parametric model

Figure D.1: Fully parametric estimation of baseline survival function



Estimates	λ	a
Whole sample	0.03***	1.93***
Co-residing wives	0.04***	1.95***



Estimates	λ	a
n=1	0.03***	1.90***
n=2	0.02***	2.10***
n=3	0.03***	1.80***

Significance levels : * p<0.10, ** p<0.05, *** p<0.01.

The graphs plot the baseline survival function $S_0(t) = \exp(-(\lambda(t-8))^a)$ for different estimated values of λ and a.

Appendix E : Robustness tests

TABLE E.1: Placebo tests on the raw difference in difference

Birth intervals (in months)	Age (1)	Work (2)	Broken union (3)	Prev. child (4)	Nb cur. child (5)	Dead child (6)	Rivals (7)
Girls	2.482	0.614	1.205	0.501	5.169	-1.565	1.624
	(7.356)	(1.600)	(1.397)	(1.331)	(3.562)	(1.553)	(1.441)
Placebo	0.747***	0.383	2.347	7.358	-0.642	2.852	2.247
	(0.150)	(1.713)	(3.928)	(5.274)	(0.560)	(1.847)	(1.989)
Girls × *Placebo*	-0.059	-0.411	-5.221	-9.101	-1.309	6.777**	-7.740**
	(0.241)	(2.784)	(5.652)	(7.490)	(0.850)	(3.055)	(3.420)
Constant	12.596***	35.114***	35.225***	35.176***	37.909***	34.707***	34.902***
	(4.635)	(0.962)	(0.838)	(0.803)	(2.371)	(0.945)	(0.885)
Observations	1024	1012	972	1024	1024	974	1024

Proportion of short interval	Age (1)	Work (2)	Broken union (3)	Prev. child (4)	Nb cur. child (5)	Dead child (6)	Rivals (7)
Girls	-0.079	0.021	0.022	0.036	-0.044	0.068*	-0.006
	(0.180)	(0.039)	(0.033)	(0.032)	(0.086)	(0.037)	(0.035)
Placebo	-0.009**	0.029	-0.092	-0.037	0.028**	-0.044	-0.032
	(0.004)	(0.041)	(0.094)	(0.127)	(0.013)	(0.044)	(0.048)
Girls × *Placebo*	0.003	0.024	0.026	-0.102	0.021	-0.138*	0.218***
	(0.006)	(0.067)	(0.135)	(0.180)	(0.020)	(0.074)	(0.082)
Constant	0.654***	0.359***	0.378***	0.370***	0.256***	0.385***	0.376***
	(0.113)	(0.023)	(0.020)	(0.019)	(0.057)	(0.023)	(0.021)
Observations	1024	1012	972	1024	1024	974	1024

Sample of co-residing wives. OLS regression on non-censored durations, no controls. Standard errors are in parentheses. Significance levels :*** Significant at the 1 percent level. **Significant at the 5 percent level. * Significant at the 10 percent level. The last column reports our baseline result using *Rivals* (cf. Table 4). The test consists of replacing this variable with another variable (*Placebo*) : age in column (1), working dummy in column (2), broken union dummy in column (3), an indicator for having children from previous unions in column (4), number of children in the current union in column (5) and an indicator for having one dead child from the current union in column (6).

Table E.2: Effect detectable after the first child : Cox estimation (Parity #1)

Hazard ratios $= e^{\beta_k}$	Whole sample (1)	Co-residing wives (2)	Poorest half (3)	Richest half (4)
G	0.98	1.03	1.00	1.13
	(.095)	(.119)	(.167)	(.195)
Rivals	0.92	0.87	0.52**	1.17
	(.158)	(.180)	(.169)	(.332)
$G \times$ *Rivals*	1.35	1.72*	2.95***	1.22
	(.323)	(.491)	(1.242)	(.503)
Controls	Yes	Yes	Yes	Yes
Log-likelihood	-1889	-1295	-552	-624
Observations	737	500	252	248
Obs in cell : *Girls* $\times$ *Rivals* $= 1$	57	38	22	16

Standard errors of e^{β_k} are in parentheses. Significance levels (for $e^{\beta} \neq 1$) : ***Significant at the 1 percent level. **Significant at the 5 percent level. * Significant at the 10 percent level. G : first child is a girl ; *Rivals* : there is at least one rival. Column (2) : sample restricted to wives living with their husband ; this sample is then split on the median income into the poorest half in column (3) and the richest half in column (4). Controls include characteristics of the woman : age at preceding birth, place of residence, level of education, age at first marriage, being or not in her first marriage, living or not with the husband, having at least one dead child, having at least one child from previous unions, rank of preceding birth. In columns (2) to (4), we include additional controls related to the husband : sector of activity of the husband, income of the husband, age difference with the husband as well as the occupation of the wife. Exact marginal-likelihood method to handle ties among non-censored durations.

Table E.3: Cox estimation (separated by parity, whole sample)

Hazard ratios $= e^{\beta_k}$	Parity #1	Parity #2	Parity #3
Girls	0.98	1.07	0.79
	(.095)	(.149)	(.204)
Rivals	0.92	1.30	0.82
	(.158)	(.211)	(.161)
Girls $\times$ *Rivals*	1.35	0.90	3.20**
	(.323)	(.300)	(1.548)
Controls	Yes	Yes	Yes
Log-likelihood	-1889	-1199	-645
Observations	737	528	324
Obs in cell : *Girls* $\times$ *Rivals* $= 1$	57	26	10

Standard errors of e^{β_k} are in parentheses. Significance levels (for $e^{\beta} \neq 1$) : *** Significant at the 1 percent level. **Significant at the 5 percent level. * Significant at the 10 percent level. *Girls* : first n children are girls ; *Rivals* : there is at least one rival. Controls include characteristics of the mother : age at birth n, place of residence, level of education, age at first marriage, being or not in her first marriage, living or not with the husband, having at least one dead child, having at least one child from previous unions. Exact marginal-likelihood method to handle ties among non-censored durations.

Table E.4: Fully parametric estimation (parameters specific to each parity, whole sample)

Hazard ratios= e^{β_k}	Parity #1	Parity #2	Parity #3
Girls	1.02	1.03	0.86
	(.130)	(.179)	(.263)
Rivals	1.00	1.38	0.85
	(.217)	(.286)	(.207)
Girls × *Rivals*	1.47	0.82	7.46***
	(.447)	(.344)	(4.767)
Controls †		Yes	
Log-likelihood		-2935	
Observations		761	

Standard errors of e^{β_k} are in parentheses. Significance levels (for $e^{\beta} \neq 1$) : *** Significant at the 1 percent level. **Significant at the 5 percent level. * Significant at the 10 percent level. *Girls* : first n children are girls; *Rivals* : there is at least one rival. † In the logit, we keep the strongest predictors of the probability to have another child : place of residence, being in first marriage, co-residing. In the survival function, we keep our variables of interest and the following controls : age at birth n, level of education, age at first marriage, co-residing, having at least one dead child, having at least one child from previous unions.

Table E.5: Impact of the number of sons : Cox estimation (pooled, co-residing wives)

Hazard ratios = e^{β_k}	$s > 0$	Breakdown by s
$s > 0$	0.93	
	(.076)	
$s = 1$		0.95
		(.080)
$s = 2$		0.82
		(.113)
$s = 3$		1.25
		(.349)
Rivals	1.44*	1.45*
	(.276)	(.277)
Rivals × $(s > 0)$	0.65**	
	(.142)	
Rivals × $(s = 1)$		0.62**
		(.142)
Rivals × $(s = 2)$		0.75
		(.202)
Rivals × $(s = 3)$		0.62
		(.336)
Controls	Yes	Yes
Observations	1117	1115

Robust standard errors of e^{β_k} are in parentheses (clustered at the woman level) - Significance levels (for $e^{\beta} \neq 1$) :*** Significant at the 1 percent level. **Significant at the 5 percent level. * Significant at the 10 percent level. s : number of sons; *Rivals* : there is at least one rival. Controls include characteristics of the woman : age at preceding birth, place of residence, level of education, age at first marriage, being or not in her first marriage, having at least one dead child, having at least one child from previous unions, occupation, sector of activity of the husband, income of the husband, age difference with the husband and rank of preceding birth. Breslow method to handle ties among non-censored durations.

Table E.6: Common intrinsic son preference : Cox estimation (wives with at least one son)

Hazard ratios= e^{β_k}	Parities #2 and #3
OneBoy	1.07
	(.149)
Rivals	1.00
	(.197)
OneBoy × *Rivals*	0.94
	(.245)
Controls	Yes
Log-likelihood	-1594
Observations	481

Sample : co-residing wives with at least one son, parities #2 and #3. Standard errors of e^{β_k} are in parentheses. Significance levels (for $e^{\beta} \neq 1$) : *** Significant at the 1 percent level. **Significant at the 5 percent level. * Significant at the 10 percent level. *OneBoy* : exactly one boy among the first children ; *Rivals* : there is at least one rival. Controls include characteristics of the mother : age at preceding birth, place of residence, level of education, age at first marriage, being or not in her first marriage, having at least one dead child, having at least one child from previous unions, rank of preceding birth, sector of activity of the husband, income of the husband, age difference with the husband, occupation of the wife. Breslow method to handle ties among non-censored durations.

Table E.7: Main result broken down by husband's status

Hazard ratios = e^{β_k}	Baseline (1)	Household Head (2)	Not Household Head (3)
Girls	1.07	1.10	0.94
	(0.088)	(0.109)	(0.141)
Rivals	0.94	0.90	1.35
	(0.112)	(0.121)	(0.311)
Girls × *Rivals*	1.54**	1.61*	0.83
	(0.334)	(0.398)	(0.460)
Controls	Yes	Yes	Yes
Observations	1117	740	377

Robust standard errors of e^{β_k} are in parentheses (clustered at the woman level) - Significance levels (for $e^{\beta} \neq 1$) :*** Significant at the 1 percent level. **Significant at the 5 percent level. * Significant at the 10 percent level. *Girls* : first children are only girls ; *Rivals* : there is at least one rival. Sample : co-residing wives ; this sample is then split on the husband's status : household head in column (2) and not household head in column (3). Controls include characteristics of the woman : age at preceding birth, place of residence, level of education, age at first marriage, being or not in her first marriage, having at least one dead child, having at least one child from previous unions, rank of preceding birth, sector of activity of the husband, income of the husband, age difference with the husband, occupation of the wife. Breslow method to handle ties among non-censored durations.

Table E.8: Impact of the gender of rivals : Cox estimation (pooled, co-residing wives)

Hazard ratios $= e^{\beta_k}$	At least 1 male (1)	At least 1 female (2)	Both (3)
Girls	1.09	1.11	1.09
	(.087)	(.089)	(.088)
Rivals_male	0.81		0.79
	(.117)		(.124)
Rivals_female		0.97	1.07
		(.142)	(.161)
Girls × *Rivals_male*	1.58*		1.47
	(.404)		(.484)
Girls × *Rivals_female*		1.33	1.09
		(.343)	(.360)
Controls	Yes	Yes	Yes
Log-likelihood	-4607	-4609	-4607
Observations	1117	1117	1117
Clusters	515	515	515
Obs in cell : *Girls* × *Rivals_male* $= 1$	48	na	48
Obs in cell : *Girls* × *Rivals_female* $= 1$	na	40	40

Robust standard errors of e^{β_k} are in parentheses (clustered at the woman level). Significance levels (for $e^{\beta} \neq 1$) : ***Significant at the 1 percent level. **Significant at the 5 percent level. * Significant at the 10 percent level. *Girls* : first children are only girls ; *Rivals_male* : there is at least one male rival ; *Rivals_female* : there is at least one female rival. Controls include characteristics of the woman : age at preceding birth, place of residence, level of education, age at first marriage, being or not in her first marriage, having at least one dead child, having at least one child from previous unions, occupation, sector of activity of the husband, income of the husband, age difference with the husband and rank of preceding birth. Breslow method to handle ties among non-censored durations.

Table E.9: Impact of the number of rivals : Cox estimation (pooled, co-residing wives)

Hazard ratios $= e^{\beta_k}$	Nb rivals ≥ 1 (1)	Nb rivals ≥ 2 (2)	Nb rivals $= 1$ (3)	Nb rivals $= 1$, by gender (4)
Girls	1.11	1.10	1.05	1.06
	(.088)	(.091)	(.087)	(.089)
Nb_shares	0.94	0.96	0.87	
	(.056)	(.056)	(.152)	
Nb_male				0.75
				(.149)
Nb_female				1.30
				(.208)
Girls $\times$ *Nb_shares*	1.11	1.09	1.90*	
	(.087)	(.082)	(.697)	
Girls $\times$ *Nb_male*				2.20**
				(.866)
Girls $\times$ *Nb_female*				0.99
				(.388)
Controls	Yes	Yes	Yes	Yes
Log-likelihood	-4608	-4039	-4097	-4095
Observations	1117	1005	1023	1023
Clusters	515	464	474	474
Obs in cell : *Girls* $\times$ *Nb_shares* > 0	64	38	26	na
Obs in cell : *Girls* $\times$ *Nb_male* > 0	na	na	na	21
Obs in cell : *Girls* $\times$ *Nb_female* > 0	na	na	na	5

Robust standard errors of e^{β_k} are in parentheses (clustered at the woman level). Significance levels (for $e^{\beta} \neq 1$) : ***Significant at the 1 percent level. **Significant at the 5 percent level. * Significant at the 10 percent level. *Girls* : first children are only girls ; *Nb_shares* : number of rivals in inheritance shares ; *Nb_male* : number of male rivals ; *Nb_female* : number of female rivals. The sample of co-residing wives in column (1) is split on the number of rivals : column (2) includes only women with 2 and more rivals whereas columns (3) and (4) include only women with exactly 1 rival. Controls include characteristics of the woman : age at preceding birth, place of residence, level of education, age at first marriage, being or not in her first marriage, having at least one dead child, having at least one child from previous unions, occupation, sector of activity of the husband, income of the husband, age difference with the husband and rank of preceding birth. Breslow method to handle ties among non-censored durations.

Table E.10: No preference for diversity : Cox estimation (pooled, co-residing wives)

Hazard ratios= e^{β_k}	Co-residing wives
Boys	0.93
	(.082)
Rivals	1.11
	(.147)
Boys $\times$ *Rivals*	0.90
	(.190)
Controls	Yes
Observations	1117

Robust standard errors of e^{β_k} are in parentheses (clustered at the woman level). Significance levels (for $e^{\beta} \neq 1$) : *** Significant at the 1 percent level. **Significant at the 5 percent level. * Significant at the 10 percent level. Controls include characteristics of the woman : age at preceding birth, place of residence, level of education, age at first marriage, being or not in her first marriage, having at least one dead child, having at least one child from previous unions, occupation, sector of activity of the husband, income of the husband, age difference with the husband and rank of preceding birth. Breslow method to handle ties among non-censored durations.

Table E.11: Robustness : Cox estimation (pooled, co-residing wives)

Hazard ratios = e^{β_k}	Baseline specification	All parities	Ethnic groups
	(1)	(2)	(3)
Girls	1.07	1.12	1.08
	(.088)	(.089)	(.088)
Rivals	0.94	0.93	0.94
	(.112)	(.092)	(.114)
Girls $\times$ *Rivals*	1.54**	1.49**	1.55**
	(.334)	(.308)	(.334)
Parities	$n \leq 3$	all n	$n \leq 3$
Controls	Yes	Yes	Yes
Ethnic group	No	No	Yes
Log-likelihood	-4607	-5633	-4604
Observations	1117	1271	1117
Clusters	515	518	515

Robust standard errors of e^{β_k} are in parentheses (clustered at the woman level). Significance levels (for $e^{\beta} \neq 1$) : ***Significant at the 1 percent level. **Significant at the 5 percent level. * Significant at the 10 percent level. *Girls* : first children are only girls ; *Rivals* : there is at least one rival. Sample : co-residing wives ; column (1) : baseline specification ($n \leq 3$, no control for ethnic groups) ; column (2) : we include all parities ; column (3) : we control for ethnic groups. Controls include characteristics of the woman : age at preceding birth, place of residence, level of education, age at first marriage, being or not in her first marriage, having at least one dead child, having at least one child from previous unions, occupation, sector of activity of the husband, income of the husband, age difference with the husband and rank of preceding birth. Breslow method to handle ties among non-censored durations.