The Pennsylvania State University

The Graduate School

College of the Liberal Arts

THE BIRTH INTERVAL AND THE ODDS OF A MALE BIRTH IN SUB-SAHARAN AFRICA:

IMPLICATIONS FOR THE SEX RATIO

A Thesis in

Sociology and Demography

by

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Submitted in Partial Fulfillment
of the Requirements
for the Degree of

Master of Arts

December 2017

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ABSTRACT

Many studies seek to understand the social and biological determinants of the sex ratio at birth, but the literature has neglected to explore how rates of fetal loss might affect a population's sex ratio at birth due to a difficulty in measuring early fetal loss at a population level. Fetal loss is important for understanding sex ratios at birth because it occurs more often for male than female fetuses (Pelletier 1998; Caselli et al. 2006; Kraemer 2000, Carlo di Renzo et al. 2007; Byrne and Warburton 1987), and poor maternal well-being is linked to higher levels of fetal loss (Kim et al. 2012; Nepomnaschy et al. 2006; Agarwal et al. 1998; Norsker et al. 2012). To address this gap in the literature, I use birth intervals as a proxy of fetal loss after controlling for other determinants of birth intervals such as contraceptive use. Using data from the Demographic and Health Surveys (DHS) in sub-Saharan Africa (SSA), I find a negative relationship between the length of birth interval and the odds of a male birth. I suggest that this relationship can be explained by repeated early fetal loss which both decreases the odds of a male birth and lengthens the birth interval.

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INTRODUCTION

Sex ratios are an important population characteristic. They shape marriage markets, labor markets and even crime rates (South et al. 2014; Angrist 2002). The sex ratio of any particular cohort changes with different mortality rates by sex, but it is largely bounded by the cohort's sex ratio at birth. A population's sex ratio at birth is determined by individual women's odds of having a male birth in that population.

Demographers consider a sex ratio of 105 males to 100 females (51:49 odds of a male) to be a "biologically normal" sex ratio at birth, but there exists considerable variation around this sex ratio (Bongaarts and Guilmoto 2015; Clarke 2000). Sex ratios at birth have reached as high as 134:100 in some regions of China (Ding Jian and Hesketh 2006) and as low as 98:100 in Botswana (Garenne 2002; Kaba 2008). Some of these deviations from a sex ratio of 105:100—such as the high sex ratios at birth in China—can be attributed to deliberate sex-selective behaviors such as sex-selective abortions and unequal parental investments in infant health. Even in the absence of sex-selective behaviors, however, there remains a substantial portion of unexplained variation in the sex ratio at birth (Pergament et al. 2002). For example, why—in the absence of sex-preference—is the sex ratio at birth 107:100 in Ireland and Jamaica, but only 98:100 in Iceland and South Africa (UN Population Division 2015)?

In the absence of sex-preference, early fetal loss can lower the odds of having a male birth and therefore, can reduce the sex ratio at birth in a population. Clinical studies show that male fetuses are more frail than female fetuses and are disproportionately lost (Byrne and Warburton 1987; Caselli et al 2006; Carlo di Renzo et al. 2007; Kellokumpu-Lehtinen and Pelliniemei 1984; Kraemer 2000; Pelletier 1998; Vatten and Skjaerven 2004). Anything which increases the incidence of fetal loss therefore decreases the odds of having a male birth (Kraemer 2000; Carlo di Renzo et al. 2007).

Measuring early fetal loss at a population level, however, is difficult, and the demographic literature has relied on exogenous shocks to the likelihood of fetal loss to examine the relationship between fetal loss and the sex ratio. At the population level, exogenous shocks that increase the likelihood of fetal loss—such as a famines, maternal stress, or exposure to environmental toxins — are associated with a decrease in the sex ratio (Cai and Feng 2005; Catalano 2003; Catalano 2006; Schnettler and Klusner 2014; Torche 2011; Valente 2015; Mazmuder and Seeskin 2015; Ruckstuhl et al. 2010; Andersson and Bergstrom 1998; Wallner et al. 2012; Hernande-Julian et al. 2014; Hamoudi and Nobles 2014). However, because the literature has relied on exogenous shocks to track the levels of fetal loss, the literature has neglected the relationship between populations' static risk of fetal loss and the odds of having a male.

This study uses birth intervals to infer fetal loss. Net of other determinants of the birth interval, fetal loss lengthens the birth interval, and this strategy has been used in the past by demographers to measure fetal loss (Sheps 1964; Potter et al. 1965; Stoeckel and Choudhury 1974; Pullum and Wililams 1977; Hebert et al. 1986; Kallan 1992; Weinstein and Wood 1993; Youssef 2005).

An increase in the rate of fetal loss therefore does two things: it decreases the odds that male fetuses survive, and it increases the length of the birth interval. In this paper, I explore whether early fetal loss—as measured by the length of birth intervals—might be an important mechanism determining women's odds of having a male birth and therefore for a population's sex ratio at birth. This is the first study to harness the relationship between birth intervals and fetal loss to examine the odds of having a male birth.

This study uses sub-Saharan Africa (SSA) as the site for investigating the relationship between birth intervals and the probability of having a male. SSA has relatively large and non-selective groups of women who are not intentionally manipulating their birth intervals. Additionally, previous studies have found no evidence of sex-selective fertility behaviors in SSA (Filmer et al. 2009; Rossi and Rouanet

2015; Fuse 2010; Basu and de Jong 2007). Using SSA as the site for the study therefore limits the confounding effects of women intentionally lengthening birth intervals as well as couples intentionally manipulating their odds of having male birth.

There is also substantive reason to focus on SSA. While a biologically natural sex ratios at birth is around 105-106 males for every 100 females, only 103 males are born for every 100 females in SSA with many sub-populations dipping even lower (Garenne 2002; Kaba 2008; UN Population Division 2015). Although a difference between 106 and 103 may seem small, small differences in sex ratios translate into large differences in absolute numbers of births. Given the number of births in SSA every year, the difference between a sex ratio at birth of 103:100 and a sex ratio of 106:100 is the difference between 250,000 male births every year. To use the language of the sex-selective abortion literature: every year, there are 250,000 "missing boys" from SSA. Because males have higher mortality rates throughout life, these skewed sex ratios at birth have the potential of leading to reductions in marriage and could alter family formation patterns in SSA. Even more importantly, low sex ratios at birth may serve as a signal of an underlying population health problem in SSA—elevated rates of fetal loss.

BACKGROUND

Fetal loss and the sex ratio at birth

In general, male fetuses suffer higher rates of fetal loss than females, just as they are more likely to suffer poor health and mortality outcomes after birth. Males are more likely to have extremely low birth weights, to be born preterm, or to be still-born (James 2008; Stevenson et al. 2000). These disadvantages do not appear instantaneously at the moment of birth, but are the cumulative result of intrauterine conditions. Male fetuses are also more likely to suffer from fetal growth retardation, to have chromosomal abnormalities, to suffer from perinatal brain damage, cerebral palsy, and from congenital deformities, (Bekedam et al.2002; Kraemer 2000; Di Renzo et al. 2007; Orzack et al. 2015). Concordantly, male fetuses are more fragile and are more likely to be lost (Byrne and Warburton 1987; Caselli, Vallin, and Wunsch 2006; Carlo di Renzo et al. 2007; Kellokumpu-Lehtinen and Pelliniemei 1984; Kraemer 2000; Pelletier 1998; Vatten and Skjaerven 2004).

Evolutionary scientists theorize that the male pre-natal and infant disadvantage are a population-level response to risky or dangerous environments. In particular, the Trivers-Willard hypothesis speculates that daughters have a reproductive advantage (a higher likelihood of producing grandchildren) in poor environmental conditions, but that sons have a reproductive advantage in good conditions. Because male reproductive success varies more than females' (a female can give birth to a maximum of approximately twenty children, while a male can potentially exclude other males and beget hundreds of children). In terms of reproductive "pay-off", sons are high risk, but potentially high reward. Daughters are less risk, but lower reward. Accordingly, the expectation is that a reproductively-successful mother would invest more (whether intentionally or not) in bearing sons when in an optimal reproductive environment. When faced with low resources and risky conditions, however, a reproductively successful mother should prefer daughters (Trivers-Willard 1973). Pongou (2013) finds support for this idea with research showing that the degree of male disadvantage in utero is not fixed, but depends partially on environment. Sex ratios tend to be lower (with fewer boys) in poor or adverse physical environments (Sander and Stoecker 2015; Terrell, Hartnet and Marcus 2011; Tragaki and Lasaridi 2009). The disproportionate effects of fetal loss on male fetuses fits within this evolutionary framework.

Birth Intervals and Fetal Loss

As noted earlier, measuring early fetal loss, particularly at a population level is difficult. Most fetal loss occurs in the first six weeks of pregnancy—well before most mothers notice or confirm their pregnancies (Copper et al. 1996; Nepomnaschy et al. 2006; Jauniaux and Burton 2005). An early fetal loss is much more likely to be interpreted as delayed menstruation rather than to be recognized and confirmed as a fetal loss.[1] Rates of self-reported fetal loss in representative surveys range from 3-15% of pregnancies (Casterline 1989), while estimates of rates of fetal loss from clinical studies range from 12-70% of all pregnancies (Graziella, Vallin, and Wunsch 2005; Macklon, Geraedts,and Fauser 2002; Wood 1994). Some estimates conclude that over 90% of pregnancy losses occur without the knowledge of the mother (Edmonds, et al. 1982,). I therefore do not rely on self-reported measures of fetal loss, but rather I use a strategy that has been used in the past by demographers. That is, I infer fetal loss from the duration of birth intervals (Sheps 1964; Potter et al. 1965; Stoeckel and Choudhury 1974; Pullum and Wililams 1977; Hebert et al. 1986; Kallan 1992; Weinstein and Wood 1993; Youssef 2005). In reproductive-aged women, fetal loss lengthens the birth interval, net of other determinants.

Fetal loss—in conjunction with postpartum amenorrhea, time until conception, and gestation—make up the easily measured birth interval (see Figure 1) (Potter 1963; Bongaarts 1981). Not every birth interval contains all four components, but the average birth interval is equal to the sum of the average duration of the four components at the population level. Determining the length contributed by any three of the components of the birth interval reveals the time contributed by the fourth. It has been estimated that without deliberate lengthening of the birth interval, the average birth interval for married/sexually active women lasts a minimum of 20 months, as follows (Bongaarts 1978):

Minimum Birth interval (20 months) =

post-partum amenorrhea (minimum 1.5 months) +

time until conception (7.5 months) +

fetal loss (2 months) +

gestation (9 months).

The duration of gestation is considered constant, but the other components of the birth intervals vary between women and between populations. Post-partum amenorrhea—the time of infertility following each pregnancy—varies by breast feeding practices and with infant mortality. Breastfeeding increases the duration of post-partum amenorrhea, and the death of a breastfeeding infant prematurely terminates breast feeding.

Time to conception is determined by fecundability (the monthly probability of conception in the absence of contraception (Gini 1924)), coital frequency, and contraceptive use. Fecundability varies between women, and there is an inverse relationship between fecundability and the birth interval (Weinstein, Wood, and Chang 1994; Preston, Heuveline, and Guillot 2000).

Each fetal loss increases the birth interval. The added time associated with each fetal loss is a combination of the length of gestation and the period of infecundability following the pregnancy before the return of ovulation. The length of time added by fetal loss is a function of the length of gestation before the loss, which in turn determines the length of the amenorrheic period following the loss (Sheps 1963). Each fetal loss resets the clock on time until conception, so for women with lower fecundability, fetal losses will have a larger effect on the length of the birth interval than in populations with shorter waiting times to conception (Preston, Heuveline, and Guillot 2000).

Perfectly measuring any three of the four components of birth intervals (post-partum amenorrhea, waiting time to conception, fetal loss, and gestation) allows the inference of the fourth. For instance, if post-

[1] Throughout this paper, I refer to all losses of conceptions before birth as "fetal loss" even though some of the losses occur in the embryonic stage rather than during the fetal stage.

partum amenorrhea, time to conception, and gestation time are fixed, but the average birth interval increases, we can assume the increase in the birth interval came from an increase in the frequency of fetal loss in the population. Birth intervals are a reliable marker of fetal loss after accounting for the other components of the birth interval.

Conceptual Framework and Hypothesis

Fetal loss is related to both the length of the birth interval and the sex of the live birth. Women with high levels of fetal loss will have both long birth intervals and lower odds of having a male birth. Women with repeated fetal loss will have long birth intervals since each fetal loss resets the clock on amenorrhea, time to conception, and gestation. Additionally, when women with repeated fetal loss do successfully carry a fetus to term, it is more likely that they will be successful in carrying a fetus with a lower risk of fetal loss (aka a female fetus). I therefore expect women with high odds of fetal loss to have long birth intervals and lower odds of having a male birth.

Conversely, women with minimal fetal loss will have shorter birth intervals and will give birth to more males. Without fetal loss adding time to their birth interval, women with minimal fetal loss will have shorter birth intervals. Additionally, in the absence of fetal loss, the odds of giving birth to male should match the odds of conceiving a male. The literature suggests that the odds of conceiving a male are much higher than one-to-one (Kellokumpu-Lehtinen and Pelliniemi 1984; Pelletier 1998) I therefore expect that women with low odds of fetal loss will have both short birth intervals and higher odds of a male birth.

Lack of Preferential Sex Behaviors in Sub-Saharan Africa

In order to clearly attribute the odds of having a male birth to fetal loss, this study looked for a population whose odds of having a male are unobscured by sex-preference behaviors. Sex-preference is an important reproductive behavior, but it is a distinct process from spontaneous fetal loss. Both can occur in a population at the same time, but sex-preference behaviors would make it difficult to determine whether different odds in having a male birth are the result of a woman's conscious choice or are the result of spontaneous fetal loss.

Cross-regional studies that look for sex-preferential fertility behavior find evidence of son-preference in Asia, the Middle East, and Northern Africa, but not in SSA. Basu and DeJong (2010) used the Demographic and Health Surveys (DHS) to reconstruct birth histories for women and found significant evidence of differential stopping behavior (DSB) in South Asia, Southeast Asia, and in Northern Africa. DSB is characterized by continuing childbearing until a couple reaches their ideal number of sons (Basu and DeJong 2010). In sub-Saharan Africa, however, Basu and DeJong found no support of DSB. Another study used 158 DHS surveys to examine sex-composition and fertility behavior patterns globally. Filmer, Friedman, and Schady (2009) found statistical evidence of DSB in Central Asia, South Asia, Southeast Asia, the Middle East, and in Northern Africa. Latin America, the Caribbean and SSA, however, did not exhibit any evidence of differential stopping.

These finding are consistent with studies that look for son-preference fertility behaviors in Africa alone. Hypothesizing that women under pressure to have a son would exhibit shorter birth intervals after the birth of a daughter than after the birth of a son, Rossi and Rouanet (2015) used birth intervals as a measure of son-preference in fertility behavior. Within the continent of Africa, this measure of son preference indicates the existence of son-preference in Northern Africa, but not in SSA. Concordantly, breastfeeding behaviors display no significant indication of son preference in SSA (Chakravarty 2012; Garenne 203), and examinations of survey questions on gender preference in the DHS confirm that "balance preference" or "no preference" dominates in SSA (Fuse 2010).

MATERIAL AND METHODS

To examine the relationship between the odds of a male birth and fetal loss (proxied by birth interval), I use the IPUMS Demographic and Health Surveys (DHS) surveys in SSA. The DHS is a cross-sectional survey that collects information on fertility and health, including a complete fertility history among all sample women ages 15-49, and I use the most recent survey for all countries in SSA that include key variables. The study therefore uses mother and birth information from the following DHS surveys: Benin (2011), Mali (2012), Mozambique (2011), Niger (2012), Nigeria (2013), Rwanda (2010, Zimbabwe (2010), Uganda (2011), Burkina Faso (2010), and Zambia (2013). IPUMS created harmonized data files for all mothers, children, and I use the mother- and children-files for this study. I merged the mother data sets with the child-level data sets. The women's data set includes a full birth history up to twenty births and the child data set includes detailed information on all births in the last five years preceding the survey. I then rotated the file so that births reported by the mothers are the individual unit of analysis.

The DHS is a cross-sectional data set, and I therefore restrict my data set to women whose birth occurred within six months of the interview date. This was done to ensure that the characteristics collected about the women in the survey were as close as possible to their characteristics during the interval between births. For instance, although the DHS asks if a woman is a second or order wife, it only asks this of women at the time of the interview. No retrospective information on this variable is available, and I therefore cannot use such a variable to accurately assume wife-rank status 10 years ago. After limiting my sample to all births in the past six months, my sample includes 19,463 births.

Because the model uses birth intervals as a proxy for fetal loss, it is important to account for other components of birth intervals. Most importantly, I restrict my sample to women who have never intentionally tried to manipulate their birth intervals. Women's intentions and actions to space childbearing have a substantial effect on the time to conception, and I therefore use a variety of variables to eliminate variation in birth intervals caused by conscious spacing behavior. I excluded women who say they ever used family planning. I exclude not only users of modern methods of contraception, but also women who say they used breastfeeding, abstinence, withdrawal, or any traditional or folkloric method to plan childbearing. After these exclusions, the sample includes 14,622 births.

I model the sex of the second birth as a function of the interval between the first and second birth. I do not model the sex of the first birth using birth intervals because it is hard to determine the beginning of the first birth interval (i.e., the beginning of exposure to regular intercourse) in the sub-Saharan context. Both first intercourse and first union are problematic measures of the commencement of regular intercourse in the sub-Saharan context (Larsen 1997; Ngalinda 1998). The distribution of birth interval length in this sample differed substantially between second and third births, suggesting a strong parity effect on fecundity. Additionally, limiting the analysis to only one parity holds constant the effects of parity on fecundity. It also allows me to hold constant the parity effects on the odds of having a male birth and facilitates more straightforward interpretations of the mother's age and fecundity absent of parity effects (Juntunen, Kvist and Kauppila 1997; James 1987).

When modeling the sex of the second birth, I include only women who have had both a first *and* a second birth, eliminating left-censoring bias. Limiting the sample in this way does not unduly bias the generalizability of the sample. Among non-contraceptive users in sub-Saharan Africa, most have at least two births. Of sexually active, non-contraceptive users who have at least one birth, less than 1% of women do not progress to parity two by age 45-49. After limiting my sample to second births that occurred within six-months of the interview, the sample includes 2,880 births. I then excluded births that occurred more than 100 months (8 years and 4 months) after the first birth.[2] This excluded 2.4% of the sample. The final sample includes 2,599 births of the same parity to women who have never intentionally manipulated the timing of their births.

To address item non-response, I used multiple imputation to handle missing data assuming multivariate normal distributions for missing variables in Stata 14. I generated 5 imputations.

Dependent variable

The dependent variable is the sex of the second birth (male = 1, female = 0).

Independent variables

The main independent variable is the time between the first and second birth, measured in months. Most second births occurred two and a half to three and a half years after the first birth, but the distribution is right skewed.

The model attempts to control for two other variable components of the birth interval besides fetal loss: post-partum amenorrhea and time to conception (there is no need to control for the 9 month period of gestation given the invariability of this component).

The duration of post-partum amenorrhea is highly related to breastfeeding practices. Unfortunately, the DHS data does not have a measure of how long the first child was breastfed. Lactational amenorrhea can add significant time to a birth interval, and this could be considered a serious limitation. In the absence of a direct measure of the duration of post-partum infecundability, however,I controlled for a dummy variable indicating whether or not the first child died in the first two years of life. The death of the child necessarily requires cessation of breastfeeding and has been shown to significantly influence a population's average birth interval through shortening the amenorrheic period (Preston 1978; Grummer-Strawn, Stupp, and Mei 1998). Also, the DHS asks a question on whether or not she is still breastfeeding her current (second) child. Most women still are, given that I restricted the sample to women who had given birth in the last six months. Whether or not the second child died explains 56% of the variation in whether or not women are still breastfeeding their second child.[3] This gives me some assurance that I am capturing at least some of the variation in the duration of breastfeeding with the dummy variable of whether the first child had died. I rely on this to capture individual women's patterns in breastfeeding in conjunction with country, union status, socioeconomic position, urbanicity, and religion to capture social patterns in the duration and intensity of breastfeeding practices (Brown 2006; Page and Lesthaeghe 1981)

To control for variation in time to conception, I include indicators of biological fecundity and coital frequency. As an attempt to measure a woman's fecundity, the model includes the DHS's measure of her fecundity as well as the time between her first intercourse and her first birth measured in months. The DHS categorizes women as infecund if they are not menopausal, postpartum amenorrheic, or pregnant, but have had no birth in the five years preceding the survey despite regular intercourse without contraception or if she says she cannot get pregnant, or has had a hysterectomy. To capture coital frequency, the model includes several dichotomous variables for union status: for being in a monogamous marriage, a polygamous marriage, a non-marital union, or no union at all. I tested interactions with these union statuses by whether the woman is currently living with her spouse or partner, but these interactions were not significant and therefore excluded from the model. I also included a measure of the length of time the woman has been in the union, because past research has found a significant negative relationship

[3] I also tested a more complex variable accounting for not only if the child died before the age of 2, but how old the child was at death. The variable was not significant and results from AIC and BIC models showed the more complex model to not be a useful addition.

between frequency of intercourse and the length of the union (Lesthaeghe, et al. 1981). Although the DHS variable "time since last sex" is a useful measure of coital frequency for some fertility studies in sub-Saharan Africa (Hindin and Muntifering 2011; Brown 2000), I do not use this variable because all women in my sample are recently post-partum, and coital frequency post-partum cannot be generalized to coital frequency before conception. I did look at the time since last intercourse variable to see if the women in the sample are practicing post-partum sexual abstinence as this practice would lengthen birth intervals; over 80% of the sample has had intercourse since having given birth less than 6 months ago. Nor does the duration of the birth interval vary between women who have had sex since their second birth and women who have had not. Moreover, when women adhere to post-partum sexual abstinence in SSA, the practice overwhelmingly aligns with polygamy, for which I control (Schoenmaeckers et al. 1981).[4]

Even though the evidence suggests there is no sex-preference fertility behavior in sub-Saharan Africa, the model also controls for the woman's ideal proportion of sons as well as the sex of the first birth. I used ideal proportion of sons rather than ideal number of sons, because ideal number of sons was highly correlated with ideal number of daughters. In other words, women who want many sons want many children in general. Ideal proportion of sons measures family sex preference rather than family size preference.

The model includes demographic controls including: a continuous variable of the age of the mother at the time of the second birth,[5] the woman's wealth quintile, educational attainment, religion, and urbanicity. The wealth quintiles rank each household into five measures of comparative wealth based on a cumulative living standard from selected assets in the respondent's household. Educational attainment is measured in categories: no schooling, primary schooling, and secondary or higher. I keep the DHS' original religion coding for Muslims, Catholics, and "no religion," but I collapse the DHS categories of Protestant, Anglican, Methodist, and "other Protestant" into one Protestant category. I also collapse responses of Pentacostals, Charismatics, Apostolic Sect, Seventh Day Adventist, "Other Christian", Ethiopian Orthodox, Celestial Church of Christ, and African Zionist into "Other Christian". I combine Traditional/spiritual, Traditional, Animist, and Voodoo into "Other." I use the DHS variable for urbanicity which classifies large cities (capital cities and cities with over 1 million population), small cities (population over 50,000), and towns as urban. It defines rural as countryside.

Plan of Analysis

First, I examined the bivariate relationship between the birth interval and the sex of the birth with a cross tabulation of the two variable and by examining a kernel density plot. I then estimate a logistic regression predicting the probability of a male birth by the length of the birth interval (Model 1). In subsequent models, I add controls for coital frequency, fecundity, and family preference (Model 2). Model 3 adds demographic controls of the mother, and the last model (Model 4) adds controls for country. I then calculated the predicted odds of having a male birth at different birth intervals comparing the base model (Model 1) and the full model (Model 4). Then, to illustrate the implications of the results for the population-level sex ratios at birth, I translate the predicted odds of having a male birth into predicted sex ratios at birth (pr(male)/pr(female))*100. Lastly, I perform robustness checks with different functional forms of the main independent variable and with different subsamples.

[4] There is no significant variation in time since last intercourse between polygamous and non-polygamous women in this post-partum sample.
[5] I tried the model with age as a continuous variable as well as a 5-year age groups; there were no significant differences between the models.

RESULTS

Table 2 shows the basic bivariate relationship between the time to the second birth and the sex of the second birth. The basic bivariate relationship shows a general relationship between the sex of the birth and the preceding birth interval. Although the interval between 2.5 and 3 years deviates a little from the trend, the earliest births show high ratios of males to females born (139.35 and 108.77, respectively). The longest birth intervals show low ratios of males to females (89.90 and 101.65). After running a basic logistic regression predicting the sex of the birth with the birth interval alone without covariates (Model 1), the predicted values shows a clear negative relationship between the odds of a male birth and the duration of the birth interval. These results are consistent with the expectation that longer birth intervals are associated with higher rates of fetal loss and therefore with a lower ability to bring a male fetus to term.

Table 3 presents the results for the full model (Model 4) estimating the probability of a male birth; longer birth intervals are significantly associated with lower odds of a male birth. The birth interval is the most significant variable in the model, and the magnitude of the effect is relatively large. Each additional month between births is associated with 1.4% lower odds of having a male birth, or each year is associated with 15.34% lower odds of having a male birth. The results are consistent across all the models and the effect of the birth interval increases in magnitude and significance across models as covariates are included.

Net of other confounders, moving from shorter to longer birth intervals is significantly associated with decreased odds of a male birth. The predicted odds of a woman having a son who had a birth interval of 12 months (1 year) are 60:40 while the predicted odds of a woman having a son who had a birth interval of 60 months (5 years) are 46:54 (see Table 3). Women who have shorter birth intervals have much higher odds of successfully giving birth to a male than a female, controlling for everything else, and women with long birth intervals are more likely to have a daughter than a son.

The sex-preference variables show evidence of women enacting sex-preference. Although the coefficient is not significant, births to women who want more daughters than sons show lower odds of being male. Women who say they want significantly more sons than daughters have 8% higher odds of having a son. Interestingly, however, this is only a very small proportion of the sample (2.6% of women), and more women in the sample say they want more daughters than say they want more sons. The modal woman in the sample, however, wants sex-balance in her family (see Table 1).

The demographic controls show mixed results. Being in the fourth wealth quintile is significantly associated with lower odds of having a male birth compared to the poorest wealth quintile, but there is no clear pattern between wealth and the odds of having a male birth. There is a positive relationship between education and the odds of having a male birth, although none of the coefficients are significant.[6] Religion does not significantly predict the sex of the birth other than for Catholics. Catholics show significantly lower odds of having a male birth in Models 2 and 3, but these results lose significance after adding controls for country. Considering the sex of an infant is considered a mostly random variable, it is not surprising that few of the other coefficients are significant.

Translating the odds of a male birth to a sex ratio at birth shows the dramatic effects different odds in having a male birth can have on a population. Short birth intervals have high predicted sex ratios, and longer birth intervals have much lower predicted sex ratios at birth (see Figure 2). Translating the odds of a male birth born 12 months after the mother's first birth to a sex ratio produces a sex ratio of 147 males to 100 females. Although this sex ratio is extremely high, the rarity of births born after such a short interval prevents the average sex ratio in the population from being too high. The predicted sex ratio at birth for births born at the mean birth interval almost perfectly matches the actual sex ratio at birth for the

[6] Wealth and education are correlated (r=0.36), but excluding one of them does not make the other significant.

sample (103:100). The predicted sex ratio at birth associated with birth intervals longer than 4 or more years, however, shows low odds of having a male birth. In fact, it is more likely for births to be female than male. The predicted sex ratio at birth associated with a birth interval of four years would be 98:100, and the predicted sex ratio at birth for births born eight years after the first is even more exaggerated at 57:100. The small decrease of 1.4% lower odds of a male per month quickly accumulates to produce dramatically lower predicted sex ratios with longer birth intervals.

To test the robustness of the findings, I first test several functional forms of the birth interval. Results from a log-Likelihood Ratio test shows that adding quadratic terms do not improve the model fit compared to the simple linear measure of the duration of the birth interval. I also took the log of the birth interval to normalize its distribution and ran the model. I additionally ran the model having recoded the time between the first and second birth from a continuous variable to a categorical variable with the following groupings: less than 18 months, between 18 and 42 months, between 42 and 60 months, and greater than 60 months. I selected these intervals based on literature which suggests birth outcomes vary by very early intervals (less than 18 months) and very long intervals (greater than 60 months) (Kangatharan 2016; Conde-Agudelo, Rosas-Bermudez, Kafury-Goeta 2006; Rutstein 2005; Kozuki et al. 2013). I also created categories based on standard deviations away from the mean. Results from BIC and AIC tests showed that the simple continuous measure yielded a better model fit than the logged transformation or either of the categorical variables. Even though a linear continuous measure of the birth interval proves to be the best model fit, a significant and negative relationship between the birth interval and the odds of having a male birth persists regardless of the functional form of the birth interval.

To assess whether using the subsample is driving the results, I also estimated the models for alternative subsamples, including: (1) all second births in the last 6 months to women who have used any method to intentionally spaced their births without limiting birth intervals longer than 100 months (N=2,880) (2) all second births in the last six months to both non-users and contraceptive users (N= 12,224), (3) for all second births to non-users of family planning, with no restrictions on when the second birth happened (N=62,902). When the sample includes birth intervals longer than 100 months, the relationship remains highly significant and the main story (that longer birth intervals are associated with lower odds of a male birth remains). The functional form of the relationship between the birth interval and the odds of having a male birth, however, becomes cubic rather than linear. When the sample includes women who use family planning, the relationship between the birth interval and the odds of having a male birth disappears. The coefficient for birth interval becomes perfectly neutral (OR=1.000) and statistically insignificant. When I expand the sample to include all women who have ever had a second birth, there is still a significant and negative relationship between the length of the birth interval and the odds of having a male birth, although the birth interval becomes slightly less significant. The loss of significance with the expanded sample suggests that the control variables meaningfully help explain the variation in the sex ratio at birth, and that they lose their explanatory power in this sample with the expanded time horizon. The pervasiveness of a negative relationship, however, suggests a genuine relationship between the birth interval and the odds of having a male birth.

DISCUSSION

The findings show a significant association between the birth interval and the odds of giving birth to a male. The existence of this relationship after controlling for behaviors suggests that early fetal loss plays a major role in the odds of having a male birth. None of the women in the sample are using modern contraception or any method of purposefully spacing their family. The selective nature of the sample in this study does not limit the applicability of the findings to those in the sample. Even though intentional birth spacing obscures the relationship between birth intervals and the probability of a male birth, this is the result of my measure—birth intervals—and not because fetal loss acts differently on the sex ratio among women who intentionally space their births.

These findings of this study can be understood within the context of a process in which many more males are conceived than females, and women with lower rates of fetal loss successfully bear these males to term. This study's estimates of the sex ratio at birth to women with short birth intervals match the estimates from the literature of the sex ratio at conception. This study estimates the sex ratio at birth following a birth interval of only 12 months to be 149:100. This sex ratio closely approximates estimates of the sex ratio at conception which vary from 169-140 males per 100 females (Kellokumpu-Lehtinen and Pelliniemi 1984; Pelletier 1998). The very short birth interval (12 months) suggests that these women have almost no fetal loss. In the absence any of fetal loss, the sex ratio at birth would perfectly match the sex ratio at conception. The findings in this study align with the hypothesis that women with minimal fetal loss will have short birth intervals and higher odds of having a male birth.

The findings from this study are consistent with the hypothesis that women who have higher rates of fetal loss have longer birth intervals. Even though each fetal loss "resets the coin toss" so to speak, women with repeated early fetal loss are more likely to successfully carry a fetus to term if the fetus has a low risk of fetal loss (is a female). This study found a significant relationship between having long birth intervals and lower odds of a male birth. A sizeable portion of the women in this study had birth intervals longer than 5 years (13.97%) (see Table 1), and thereby have very low predicted odds of having a male birth. These women with long birth intervals tend to be slightly older, less educated, and poorer than the rest of the women in the sample. A birth interval longer than five years still seems inordinately long to attribute to fetal loss in a context with regular intercourse without contraception. Yet, evidence suggests that low fecundity and early fetal loss often coincide (Hakim, Gray, and Zacu 1995), and the interaction of a low probability conception with repeated early fetal loss could easily result in very long birth intervals. The findings of this study support the hypothesis that women with repeated early fetal loss have both long birth intervals and low odds of having a male birth.

The findings in this study are also consistent with previous research on birth intervals. First, adverse birth outcomes are associated with longer birth intervals. Specifically, birth intervals longer than 4 or 5 years are significantly associated with low birth weight, preterm birth, and pre-eclampsia (Conde-Agudelo, Rosas-Bermudez, and Kafury-Goeta 2007), and most likely, fetal loss. And, longer birth intervals are associated with the sex ratio at birth. For example, a previous study of historical French Canadians in a natural fertility setting found a negative relationship between their birth intervals and the odds of having a male birth (Nonaka et al. 1998).

This study's findings on sex-preference may warrant future research. Even though the women who say they want mostly sons are a minority in this sample (2.6%), this study found that these preferences are being translated into higher odds of having a son. Future research should examine *how* these women are enacting their son-preference. It may be, however, that part of this observed association between son-preference and having a son is simply post-facto rationalization. The women are asked their family composition preference *after* the second birth. The association between women who say they want more than half their children to be male and the odds of actually having a male birth may be driven by women whose two births are sons, who say they want the two sons they already have, and that they want fewer than four children. If this is the case, however, post-facto rationalization should work the same for daughters as well, and although wanting more daughter than sons is associated with higher odds of having a daughter, these results are not significant.

The current paper has several limitations. This study uses imperfect measures to infer the magnitude of the other contributors of the birth interval. The survey did not have a direct measure of the length of post-partum amenorrhea following the first birth. However, there is no literature to suggest that the duration of breastfeeding influences the sex of the next birth. Given the existence of a relationship between post-partum amenorrhea and the birth interval but the absence of a relationship between post-partum amenorrhea and the sex of the next birth, the study's lack of a good measure of post-partum amenorrhea does not nullify the study's findings. It simply dilutes the relationship between the birth interval and the odds of having a male birth. A better measure of post-partum amenorrhea should intensify the observed relationship between the birth interval and the odds of having a male birth.

The study does not have a direct measure of coital frequency before the conception of the second child. There are very distinct differences in coital frequency among sexually active women in sub-Saharan Africa (Brown 2000; Hindin and Muntifering 2011, Westoff 2007). Differences in coital frequency compose only a small source of differences in fecundability, however (Weinstein, Wood, and Chang 1994), and the model does control for type of union and the length of the union.

Another potential problem with this study is that different women may have different levels of fecundity which the controls did not pick up. For example, some women may ovulate more regularly, or have partners with higher sperm counts. This could mean that some women have longer birth intervals because they are less fecund and not because of higher rates of fetal loss. However, there is a stronger link in the literature, between fetal loss and the sex ratio than there is between infecundity and the sex of the resulting child. Recent research also suggests that a large portion of what appears to be infecundity results not from the inability to conceive, but rather from high rates of repeated early fetal loss (Macklon, Geraedts, and Fauser 2002).

Given the existence of a relationship between fetal loss and the sex ratio at birth, further research ought to explore whether populations with persistently low sex ratios at birth have consistently higher rates of fetal loss. If some populations have higher rates of fetal loss than others, this warrants a health intervention. Demography has paid attention to the relationship between fetal loss and the sex ratio at birth within a population by studying changes in the sex ratio following exogenous shocks to a population's likelihood of fetal loss. This literature which studies within-population change over time has concluded that sex ratios at birth are a "sentinel health indicator." Differences in levels of fetal loss between populations deserve the same attention given to fluctuations within populations. Moreover, such a finding would contradict the idea that variations in sex ratios between populations are the result of unmalleable biological factors.

For instance, African Americans in the United States have consistently lower sex ratios at birth compared to their white contemporaries. The demographic and sociological literature has largely left this phenomenon untouched. Biologists, however, have noted it and attribute it to some genetic factor of African Americans (Ruder 1986). Yet race in the United States is largely a social construction. The African American experience in the United States is distinct from other racial experiences, and African Americans in the United States face a host of distinct obstacles that worsen their health. African Americans also have consistently lower birth weights and neo-natal mortality, both of which the social science literature has studied rigorously. Given that fetal loss is considered the extreme end of poor natal health, future research could examine a link between the incidence of fetal loss and the sex ratio at birth of African Americans.

Despite the limitations, this is the first study to harness the relationship between birth intervals and fetal loss to examine the odds of having a male birth. The paper shows a robust relationship between the birth interval and the odds of having a male birth, net of other confounders. I attribute the negative relationship between the birth interval and the odds of having a male birth to repeated early fetal loss. This finding warrants further research and suggests that variation in odds of having a male birth between populations may be the result of differences in maternal and prenatal health between populations.

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APPENDIX: FIGURES AND TABLES

Figure 1: Components of a birth interval

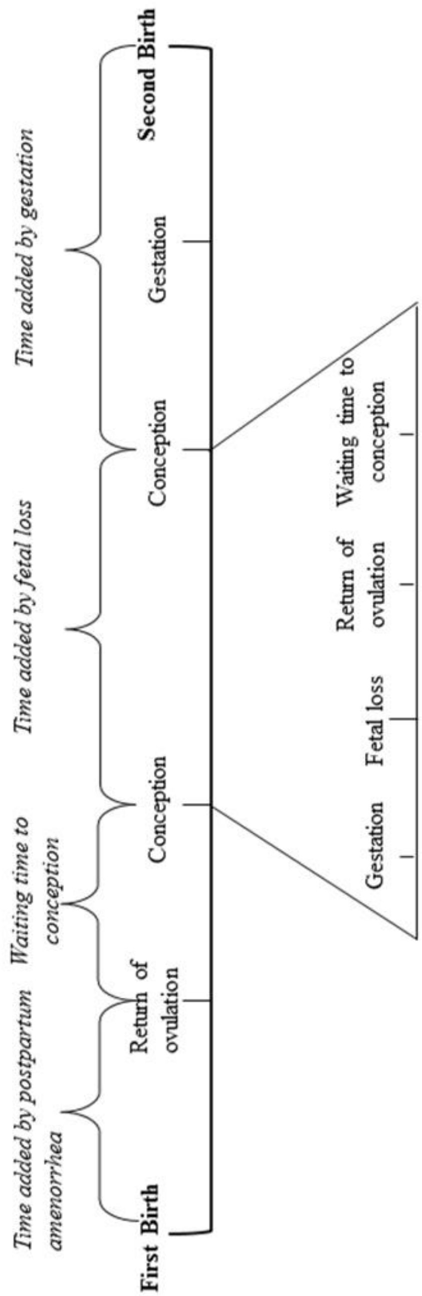

Table 1: Descriptives of women in the sample: women who have had a second birth in the last 6 months and have not used any method of birth spacing

Measure	Mean or Percentage	SD
Birth Interval (between 1st and 2nd birth)	43.66 months	15.66
Less than 2 years	5.62	
2 - 2.5 years	13.74	
2.5 - 3 years	16.04	
3 - 3.5 years	18.12	
3.5 - 4 years	13.01	
4 - 5 years	19.51	
Greater than 5 years	13.97	
Age of mother		
15-19	13.37	
20-24	53.23	
25-29	25.07	
30-34	6.42	
35+	1.91	
Household Wealth Index in Quintiles		
Poorest	22.67	
Poorer	20.52	
Middle	20.10	
Richer	19.44	
Richest	17.26	
Highest Education Level		
No Education	60.21	
Primary	22.43	
Secondary or Higher	17.36	
Religion		
Muslim	45.24	
Catholic	17.22	
Protestant	11.42	
Other Christian	17.85	
Other	5.03	
No Religion	3.23	
Urban	25.80	
Rural	74.20	
Union Status		
Monogamous Marriage	62.81	
Polygamous Marriage	21.74	
Non Marital Union	11.60	
No Union (Widowed, Divorced, Separated)	3.85	
Never Union	0.00	
Mean years in current union	4.71	2.88
Fecund	97.60	
Sex of First Birth		
Male	50.97	
Female	49.03	
Desired proportion of sons		
0	19.90	
Less than 1/2	13.13	
One half	40.80	
More than 1/2	19.65	
More than 3/4	2.60	
Non-specific answer	3.92	
Country		
Benin	19.97	
Mali	9.69	
Mozambique	6.67	
Niger	25.35	
Nigeria	16.67	
Rwanda	3.68	
Zimbabwe	1.11	
Uganda	3.58	
Burkina Faso	8.92	
Zambia	4.38	

Data Source: DHS: Benin (2011), Mali (2012), Mozambique (2011), Niger (2012), Nigeria (2013), Rwanda (2010, Zimbabwe (2010), Uganda (2011), Burkina Faso (2010), and Zambia (2013).

Sample: Second births that occurred within 6 months of interview to women who have not used any method of birth spacing N=2,599

Table 2: Bivariate relationship between sex of birth and birth interval

Length of birth interval	Females	Males	Sex Ratio	N
Less than 2 years	41.78	58.22	139.35	146
2 - 2.5 years	47.90	52.10	108.77	357
2.5 - 3 years	51.56	48.44	93.95	417
3 - 3.5 years	47.56	52.44	110.26	471
3.5 - 4 years	47.93	52.07	108.64	338
4 - 5 years	52.66	47.34	89.90	507
Greater than 5 years	49.59	50.41	101.65	363
Total	**49.25**	**50.75**	**103.05**	**2,599**

Data source: DHS: Benin (2011), Mali (2012), Mozambique (2011), Niger (2012), Nigeria (2013), Rwanda (2010, Zimbabwe (2010), Uganda (2011), Burkina Faso (2010), and Zambia (2013). Sample: second births that occurred within 6 months of interview to women who have not used any method of birth spacing N=2,599

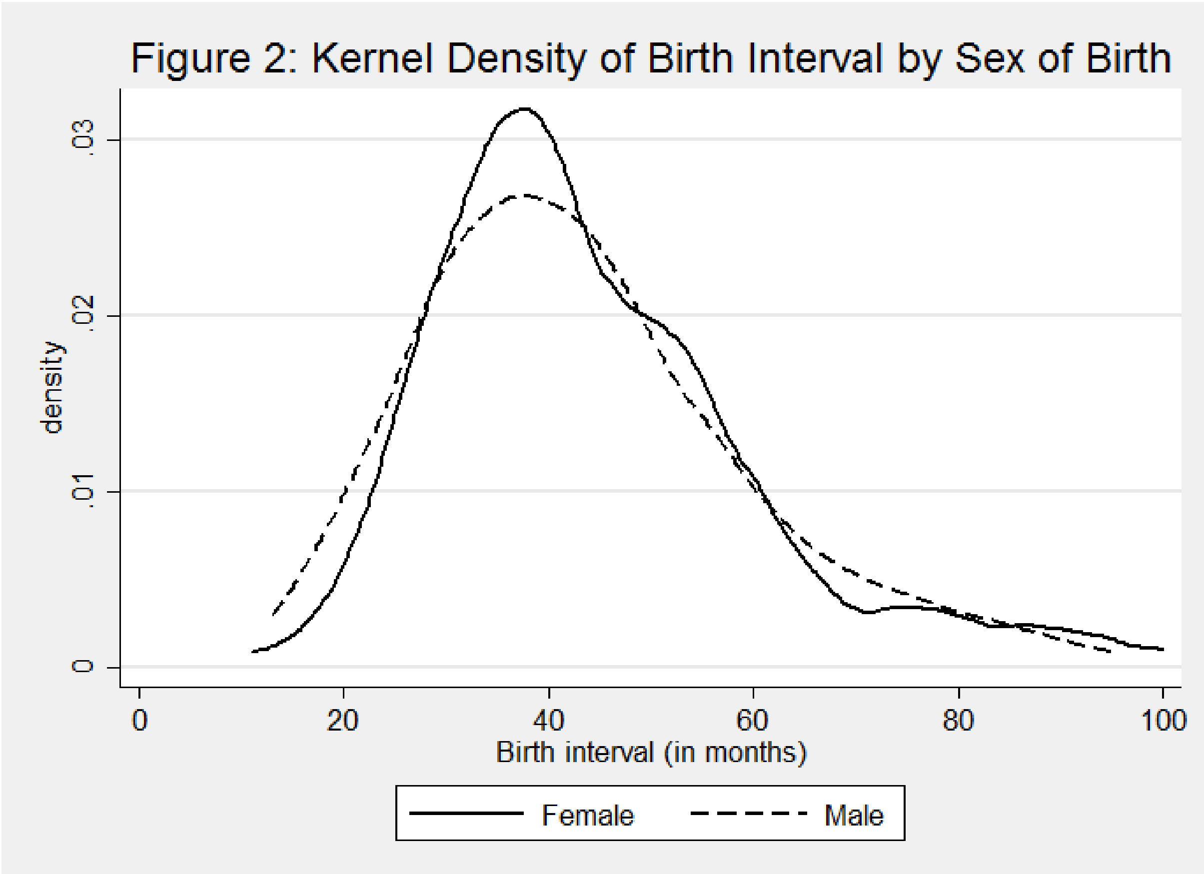
Figure 2: Kernel Density of Birth Interval by Sex of Birth
density
0
.01
.02
.03
0
20
40
60
80
100
Birth interval (in months)
Female
Male

Table 3: Results from logistic regression predicting the odds of a male birth by length of birth interval

	MODEL 1		MODEL 2		MODEL 3		MODEL 4	
	Odds Ratio	SE	Odds Ratio	SE	Odds Ratio	SE	Odds Ratio	SE
Birth Interval	0.992 ***	0.00	0.989 ***	0.00	0.987 ***	0.00	0.986 ***	0.00
First birth died early			1.06	0.15	1.06	0.15	1.04	0.15
Fecund			0.91	0.23	0.92	0.24	0.94	0.24
Time to First Birth			1.00	0.00	1.00	0.00	1.00	0.00
Marital Status (ommitted category = monogomous marriage)								
Polygamous Marriage			1.00	0.10	1.03	0.10	1.04	0.11
Non marital Union			1.10	0.14	1.12	0.15	1.11	0.16
No Union			0.84	0.19	0.81	0.18	0.81	0.19
Length of Union			1.02	0.03	1.03	0.03	1.04	0.03
Ideal Proportion of Sons (reference = one half)								
Zero			0.88	0.10	0.88	0.10	0.90	0.10
Less than 1/2			0.81	0.10	0.82	0.10	0.81	0.10
More than 1/2			1.06	0.11	1.06	0.12	1.06	0.12
More than 3/4			1.82 *	0.50	1.89 *	0.53	1.91 *	0.53
Non-numeric			1.28	0.27	1.22	0.26	1.23	0.26
Sex of First Birth			0.91	0.07	0.91	0.07	0.91	0.07
Age at the Birth					1.00	0.01	1.00	0.01
Wealth Quintile (reference = poorest)								
Poorer					1.12	0.14	1.11	0.14
Middle					0.87	0.11	0.88	0.11
Richer					0.74 *	0.10	0.75 *	0.10
Richest					0.75	0.12	0.78	0.13
Education (reference = no education)								
Primary					1.05	0.11	1.00	0.12
Secondary or higher					1.16	0.15	1.10	0.16
Religion (reference = Muslim)								
Catholic					0.78	0.09	0.81	0.10
Protestant					1.03	0.14	1.08	0.18
Other Christian					0.98	0.12	1.01	0.13
Other Christian					0.92	0.17	0.97	0.19
No Religion					1.17	0.28	1.22	0.30
Urban					0.89	0.11	0.91	0.11
Country Variables (ommitted country = Mali)								
Benin							0.88	0.16
Niger							0.83	0.13
Mozambique							0.98	0.22
Nigeria							1.00	0.17
Rwanda							0.98	0.29
Zimbabwe							0.81	0.35
Uganda							0.86	0.23
BurkinaFaso							0.94	0.18
Zambia							0.93	0.26
Constant	1.45 ***	0.17	1.48	0.44	1.88	0.89	2.13	1.06

*** p<.001 ** p<.01 *p<.05

Data source: DHS: Benin (2011), Mali (2012), Mozambique (2011), Niger (2012), Nigeria (2013), Rwanda (2010, Zimbabwe (2010), Uganda (2011), Burkina Faso (2010), and Zambia (2013).

Sample: second births that occurred within 6 months of interview to women who have not used any method of birth spacing N=2,599

Figure 3: Predicted sex-ratio at birth by birth interval