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Fertility, Agricultural Labor Supply, and Production

Instrumental Variable Evidence from Uganda

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ABSTRACT

Human fertility is likely to affect agricultural production through its effect on the supply of agricultural labor. Using the fact that in traditional, patriarchal societies sons are often preferred to daughters, we isolated exogenous variation in the number of children born to a mother and related it to agricultural labor supply and production outcomes in Uganda—a country that combines a dominant agricultural sector with one of the highest fertility rates in the world. We found that fertility has a sizable negative effect on household labor allocation to subsistence agriculture. Households with lower fertility devote significantly more time to land preparation and weeding, while larger households grow less matooke and sweet potatoes. We found no significant effect on agricultural productivity as measured in terms of yield per land area.

Keywords: fertility, labor supply, instrumental variables, boy preference

1. INTRODUCTION

At the most basic level, subsistence farmers in rural Africa combine natural with human resources to make a living. They use mainly household labor on their own small plots to produce food for their own consumption. Accordingly, the quantity of family members available for agricultural labor is an important determinant of wellbeing. More children means mothers, and to a lesser extent fathers, will need to spend more time caring for children, meaning less time will be available to spend on agricultural activities. Since women are known to supply most of the agricultural labor, the loss in time due care for offspring and additional rest needed during pregnancies may hurt subsistence households disproportionately.

Uganda has one of the highest fertility rates in the world. Even in the context of large reductions in child mortality rates, total fertility rates remain stubbornly high. On average, Ugandan women in rural areas bear 6.8 children over the course of their reproductive lives (UBOS 2012). At the same time, a substantial part of the population lives in rural areas making a living out of semi-subsistence agriculture. Ugandan agriculture accounts for about 35 percent of gross domestic product and employs about 73 percent of the active labor force. Virtually all households that reside in rural areas engage in farming, and about 80 percent are small-scale, semi-subsistence farmers. The question of how fertility affects wellbeing through its effect on household labor supply and agricultural production is therefore relevant. For example, knowledge of how fertility affects time allocation by different categories within the household is important to gender-stream efforts related to crop intensification and commercialization.

In this study, we investigated the effect of fertility on agricultural production at the household level. In particular, we investigated the effect of the number of biological children on household member labor input in agriculture (further categorized as land preparation, weeding, input application, and harvesting). We also looked at the effect of fertility on crop portfolio, area cultivated, production, and productivity for the five most important crops. However, fertility is a choice variable to agricultural households. For instance, mothers who work long hours in the field may try to avoid becoming pregnant because this would only increase their hardship. If fertility, agricultural labor allocation and agricultural production were jointly determined, just looking at correlations would be misleading, and so we needed to find a way to separate the exogenous variation from the part that is jointly determined.

Our identification strategy was a simple but powerful quasi-experimental approach inspired by the work of Angrist and Evans (1998). We used the fact that in conservative, patriarchal societies such as Uganda's, male off-spring are generally preferred to female. This preference and the random nature of the newborns' sex determination gives rise to particular fertility patterns. For example, households that have a girl as the firstborn are likely to have more children (Jayachandran and Kuziemko 2011). In other words, we used the sex of off-spring as an instrumental variable (IV) to determine the exogenous component of fertility at the household level. We expected that such a two-stage least squares (2SLS) approach would yield consistent estimates for the causal effect of fertility on agricultural labor supply and associated agricultural production within the household.

There is an active debate among labor economists on the relationship between fertility and labor supply. Angrist and Evans (1998) used the fact that American couples prefer to have children of different sexes, and that they are likely to keep trying if the first two are of the same sex, as a source of exogenous variation. We argue that in a developing country context, the sex of the firstborn makes more sense as an instrument. Indeed, this instrumentation strategy has been used in such a context. Gupta and Dubey (2006) used the sex of the first two children to predict exogenous variation in fertility in India and its effect on wellbeing. We thought it was too ambitious to relate fertility directly to poverty and related measures of consumption because the sex of the first two children may directly affect consumption, violating the exclusion restriction. We restricted ourselves to the agricultural labor allocation of adult household members, area planted, and production. In addition, most studies that look at the effects of fertility on labor allocation in a developing country context use data from Asian countries. The high incidence of selective abortion in these countries may mean the sex of the first child or children becomes endogenous as well.

This was likely to be much less of a problem in our application, which is, to our knowledge, the first such application to an African country.

We found that the sex of the firstborn, the sex of the first two children born, and the percentage of girls as a share of the total number of children all significantly explain observed fertility, measured as the gap between actual children born and a theoretical maximal fecundity for each age cohort. Fertility has a strong negative effect on the number of days the mother works in the field. We also find some evidence of a negative effect on the father, but the size of the effect is only half of that on the mother. Households with lower fertility devote significantly more time to land preparation and weeding. We also found that smaller households produce more matooke. This effect holds to a lesser extent for sweet potatoes. We found no impact on yields.

This article is organized as follows. The next section gives a brief overview of the most prominent papers that are related to our study. Then, we make our case for the use of the sex of the firstborn as an instrument, using literature that documents child gender and reproductive behavior. We then present the data we used in our application, and describe the main variables we will used in the analysis. Next, we present the results. In this section, we first take a close look at the first stage regression. We then look at the effect of fertility on household labor supply, considering differential effects depending on specific agricultural labor activities. We then turn to the effect of household size on aspects of agricultural production and productivity. The final section concludes.

2. RELATED STUDIES

Fertility and the related concept of household size affects household wellbeing through consumption and production. Lanjouw and Ravallion (1995) focused on the consumption side effects of household size in a developing country context. They noted the contradiction between widely-held views that larger households are often poorer (due to increased competition for food) and scale economies in consumption. They found that, if economies of scale are accounted for within households, the negative correlation between household size and consumption expenditure disappears. On the production side of the farm household, the effect of household size is equally ambiguous. Some may argue that larger households have more labor available within the household. The additional advantage of this labor is that it is not subject to the moral hazard effects often attributed to hired labor[1]. But at the same time, more dependents within the household means more time needs to be allocated to caring for them. Also, agricultural labor and agricultural production may be subject to diminishing returns.

The relationship between fertility and household labor supply has been studied most carefully in the field of labor economics. Since this literature is so extensive, we only mention two of the most influential works here. The first is Angrist and Evans (1998), who attempted to quantify the effect of fertility on labor supply in the US. They dealt with the endogeneity of the number of a woman's children by exploiting the fact that Americans tend to prefer two sibling genders in their households. They argued that parents of same-sex siblings are significantly and substantially more likely to go on to have an additional child. They found that more children does indeed reduce the women's participation in the labor force, but that the effect is less pronounced than previous studies had suggested. They found no effect on the labor force participation by the fathers.

Another paper that tried to answer the same question is Rosenzweig and Wolpin (1980a). In this paper, the exogenous variation in the number of children was obtained by using the occurrence of multiple births (twins) at first birth as an instrument. The authors argued that the comparison between women who gave birth to a singleton at first birth and women who gave birth to twins at first birth allowed them to identify the causal effect of an extra child on an outcome (in their case labor supply). Since the occurrence of twins is exogenous, there was no danger that heterogeneity in women's preferences contaminated the estimated coefficients. The study found that household size reduces female labor supply, but that the effect is only temporary[2].

Gupta and Dubey (2006) used the sex of the first two children as a natural experiment and found that household size increased poverty in India. They used essentially the same argument as we make in the next section. However, welfare, and the related concept of poverty, rely on consumption per capita. Consumption per capita as the independent variable is likely to be problematic in a two-stage least squares setting. There is a real danger that the instrumental variable will affect the outcome variable directly, instead of only through its influence on family size. For instance, if boys consume on average more than girls, the exclusion restriction would be violated. There is also some evidence from Indonesia. Kim et al. (2009) looked at the relation between consumption and fertility. Kim and Aassve (2006), related fertility to the allocation of labor within households. However, they moved away from the direct instrumental variable approach that is standard and instead estimated a reproduction function taking into account endogenous contraceptive choice.

All the above studies employed data from Asia. It is well known that gender at birth is already skewed in many Asian countries. For instance India, from which Gupta and Dubey (2006) drew their sample, is particularly known for selective abortion of girls (Jha et al. 2011). This non-random distribution of sex of children opens the door to potential correlation between the instrument and the error term of the

[1]For instance, Feder (1985) argues this may the reason why small farms appear to be more efficient than large farms

[2]In addition to studies that investigate the causal effect of household size on (female) labor supply, there are also a range of papers that test Becker's quantity-quality fertility model (Becker 1960; Becker and Lewis 1973). Many of these articles also used twins (Rosenzweig and Wolpin 1980b; Black et al. 2005) and/or sibling sex composition (Conley 2000; Angrist et al. 2010) as instruments.

structural equation. One example would be that less educated, poorer households that depend heavily on agriculture engage more in the abortion of female fetuses. In the context of weak instruments, such correlation can seriously bias estimates (Bound et al. 1995).

In this paper, we will try to address some of the above challenges. We used the sex of the firstborn and variations thereof as our instrumental variables. We will relate fertility to agricultural labor supply and agricultural production, since there is a direct link between these three variables. We will also concentrate on Africa. Here, while there is a boy preference, reproduction rate norms are high and the cost of raising children is low. This means that selective abortion is much less of a concern.

3. BOY PREFERENCE AND FERTILITY

There is quite a bit of evidence that parents prefer boys over girls in many developing countries[3]. For instance there is a large body of literature that looks at correlations between sex and variables related to wellbeing or quality of children. Significant differences in these outcomes are then considered proof of sex bias. Das Gupta (1987) and Sen (1990) looked at excess mortality among female infants in India. Chen et al. (1981) and Pande (2003) investigated differential access to health, in Bangladesh and India respectively. Behrman (1988) and Hazarika (2000) found a correlation between sex and nutrition and Behrman et al. (1982), Davies and Zhang (1995), and Alderman and King (1998) all investigated correlations between the gender of children and education.

However, at a more basic level, boy preference is already revealed by parents who, if asked in surveys, for example, often state clearly that they prefer boys to girls. Such preferences lead to a particular decision rule with respect to fertility, where the likelihood that children are added to the household is positively correlated to the number of surviving girls in the household. The preference for boys over girls results in what Jayachandran and Kuziemko (2011) refer to as the "stop-after-a-son" fertility pattern. There are indeed many studies that show empirically that in settings characterized by son preference, a couple that has just had a son is more likely to stop having children (Das 1987) or wait longer to have the next child (Trussell et al. 1985; Arnold et al. 1998; Clark 2000; Drèze and Murthi 2001).

Jayachandran and Kuziemko (2011) argued that son preference leads mothers to breastfeed daughters and children without brothers for a relatively shorter time. Since breastfeeding is an effective birth control method, this observed behavior also explains why couples with a son seem to wait longer before they have the next child. In addition, this underlying consequence of sex bias may partly explain a range of outcomes observed in the area of health, mortality and possibly even educational attainment. The model that Jayachandran and Kuziemko (2011) develop shows that even when parents want both boys and girls to have the same health and education, disparities can arise passively because of fertility preferences. The model shows that a "try until you have a boy" fertility rule results in girls having on average more siblings, leading to more competition for resources within the household.

The occurrence of boy preference is explained by various cultural and economic factors documented in the anthropological and demographic literature. In countries where no formal, risk-free old age insurance (such as pensions) is available, parents may choose to invest more in children who will have a higher chance of being able to support them in old age (Behrman et al. 1982). Anthropological and demographic evidence emphasize the dominant role of males in traditional patrilineal societies where descent and inheritance are transmitted through the male line. Furthermore, male children strengthen the relationship between the wife and her husband's kin (by guaranteeing the continuation of his lineage) and secure the mother's access to inheritance and a place to live upon the husband's death. Older women have power through their sons and rule over their daughters-in-law (Kandiyoti 1988). The spread of primary schooling in Africa south of the Sahara has also affected fertility patterns (Lloyd et al. 2000). Since boys are more likely to be sent (and kept) in school than girls, the extra cost associated with primary schooling will be higher in families with more boys. This, in turn, will encourage families who already have boys to reduce fertility.

Most of the evidence on the existence of boy preference comes from Asian countries. There are relatively few inquiries into sex preference in Africa south of the Sahara. Even more, it is often assumed that gender preferences are much lower or even absent there. This is surprising, since many of the cultural and economic factors that are observed in Asia equally apply to Africa. One study that documents significant gender bias in Africa is Anderson and Ray (2010), who found skewed sex ratios at older age in favor of men. Another study of a small community in Nigeria found that almost 90 percent of surveyed respondents reported male sex preference (Eguavoen et al. 2007). What is different from the Asian

[3]In developed countries, there is a preference for a mix of sexes among children, as shown in, for example Angrist and Evans (1998).

context, though, is that the female gender deficit is not only present birth, but throughout the entire age spectrum[4]. Milazzo (2014) argued that gender bias is likely not to be found at birth in the African context, where high fertility is culturally valued and less costly for families that still rely on support from the extended family system. In Uganda, preference for boys has been extensively documented in Beyeza-Kashesya et al. (2010).

Even in Western societies, preference for firstborn sons, rather than daughters, has been observed. For example, Marleau and Saucier (2002) reported an extensive list of studies that found men and/or women prefer a boy as their firstborn. Even in the United States, Angrist and Evans (1998) found some evidence of an association between having a male child and reduced childbearing at higher parities - in addition to the mixed-child preference. Accordingly, we felt that the sex of the firstborn (or closely related indicators) would provide a valid instrument for fertility at the household level in Uganda.

[4]The large effects documented in Anderson and Ray (2010) have recently been attenuated in Klasen and Vollmer (2013), who confirmed that only young adult women were missing within households.

4. DATA AND DESCRIPTIVE STATISTICS

We used the Uganda National Household Survey (UNHS) 2005–2006, obtained directly from the Uganda Bureau of Statistics (UBOS). Although it is somewhat dated, we chose this survey because it has much more information about agriculture than the more recent UNHS of 2009–2012, or that of 2012–2013. The 2005–2006 UNHS we chose was structured with a standard Living Standards Measurement Study -Integrated Survey on Agriculture (LSMS-ISA) in mind- it collected detailed information on a sample of almost 43,000 people in 7,500 households in Uganda.

Ideally, we would have liked to use a sample of households where all desired children were born. The fact that we were working with a cross-section of households in which women were at different stages of their reproductive lives, created some problems. Assume a couple that has just formed and is entering their reproductive stage. In our sample, such households showed up with a smaller-than-average household size. Now, if the firstborn happened to be a girl, this could have mistakenly been interpreted as running against our hypothesis that households where the firstborn was a girl would have higher fertility. On the other hand, if the first child was a boy, this could have led us to put too much confidence in our hypothesis, as the smaller household size was not only due to the fact that the firstborn was a son, but also to the fact that the household had only just entered its reproductive stage. The fact that we were working with a cross-section of households rather than historical data on all births by women who had reached the end of their reproductive lives was also reflected in the average number of children. This was only 3.13 children, while women bear almost 7 children over their entire reproductive period.

To deal with this problem, we worked with the difference in the actual number of children reported and the maximum reproductive capacity for a woman at a certain age, rather than simply working with the number of children in the household[5]. We referred to this measure of fertility as the *fertility gap* or *shortfall in fertility*. To get this fertility gap, we would have needed to estimate the average age at menarche within the population, and then simply divide into age the pregnancy period plus post pregnancy lactation in-fecund period. In addition, we would also have needed to incorporate the maternal mortality ratio for "censoring" the lives of women who have had too many children and thus increased mortality rates (and exit from the sample). Instead, we took the 95th percentile of total fertility rates per age from the Demographic and Health Survey of Uganda done in 2011 (UBOS 2012). This is probably a good approximation of the upper bound by age-of-fertility in the population.

The selection of children was based on the household roster of the UNHS 2005–2006. In particular, we selected individuals that were indicated as son or daughter of the household head. There is another potential problem when using a cross sectional survey such as the UNHS that only looks at reported dependents currently living within the household to calculate the difference between actual and theoretical fertility. Older women may have been living in households where some of the older children had already left the household. Thus, at around the age of thirty, the gap between reported children in the household and theoretical fertility would have started to increase more rapidly because of children growing old enough to start households of their own. More troubling, the reported gender of the oldest son or daughter living in the household may not have been the gender of firstborn. To overcome this problem, we restricted our sample to households where the mother was between 16 and 32. We chose the cut off age of 32 because at this age, the mother's firstborn will turns 16, which is our entry age into the sample of mothers. Restricting our sample in this way had a second advantage. For some of the indirect outcome variables we used, such as productivity, there was a risk that the gender of the firstborn had a direct effect on the outcome, instead of only through fertility[6]. Restricting our sample to households with only young mothers meant that the children were also likely to be younger, and thus less likely to engage in agricultural production, making a violation of the exclusion restriction less likely.

[5] Alternatively, one could use the number of children within the household and control for the age of the mother. We have also run the analysis using this strategy and came to virtually the same conclusions.

[6] For example, boys may have a different effect on productivity than girls due to their physical differences.

Looking at the sex of the firstborn is only one possible strategy. One may argue that the sex of the first born is not very relevant in a context where women bear on average almost 7 children. Indeed, it is likely that households will get a second child irrespective of the sex of the first. This is supported by Jayachandran and Kuziemko (2011), who found that the difference in breastfeeding duration between boys and girls is largest near target family size, when gender is most predictive of subsequent fertility. Therefore, we not only used the sex of the firstborn child as an instrument for fertility, but also experimented with alternative instruments such as an indicator that the first two children would be girls or a variable that expresses the share of female children in the total family size. The next section presents some preliminary statistics that suggest how gender patterns in the household are related to fertility.

Gender of Offspring and Fertility

This section makes a case for the different instruments used in the analysis. While the next section runs a first stage regression, this one presents some simple descriptive statistics to show that gender of the first few children, as well as the share of male children in a women's total number of children, affects fertility. Table 4.1 summarizes our findings. The first two columns in the top panel check if households that have a daughter as firstborn are more likely to have extra children. We simply calculated the percentage of households that had more than one child conditional on their first born being a son or a girl. In other words, we calculated the probability that a household had at least one additional child (prop +1). We found that in the sub-sample where the first child is a boy, about 37.46 percent of households would have at least one more child. However, if the first child happened to be a girl, almost 40 percent of households would have at least one more child. This confirmed our hypothesis that households have a higher chance of adding children if the firstborn is a girl.

Table 4.1 Gender and fertility

Variable	Prob +1		Prob +1		Average fertility
1st=boy	0.375	1st=boy, 2nd=boy	0.132	% daughters<0.5	2.78
1st=girl	0.393	1st=girl, 2nd=boy	0.134	% daughters>0.5	2.88
		1st=girl, 2nd=girl	0.139		
	gap		gap		gap
1st=boy	2.46	1st=boy, 2nd=boy	2.46	% daughters<0.5	2.41
1st=girl	2.26	1st=girl, 2nd=boy	2.38	% daughters>0.5	2.32
		1st=girl, 2nd=girl	2.06		

Source: Author's calculations based on the UNHS 2005–2006 data.

We also looked at the effect of the firstborn's gender on the shortfall that exists when actual fertility is compared with theoretical fertility. The first two columns in the bottom panel report this fertility gap for these two groups of households. We found that households that have a boy as the firstborn child have an average fertility gap of about 2.46 children. Consistent with the proposition that households with a firstborn girl are likely to have more children conditional on age, we found that the gap is smaller when the firstborn is a girl (2.26). In other words, households where the firstborn is a girl are closer to the theoretical maximal fertility than those households whose firstborn is a boy. The difference in the fertility gap between the group of households with a firstborn boy and the group with a firstborn girl is significant (p=0.003).

The third and fourth column present the same statistics, but now looks at the sex of the first two children. We looked at three possible scenarios. If the first two children were both boys, we expected that the chance that the household would have extra children would be lowest. We found the probability of adding children in this case to be just over 13 percent (top panel). If the first was a girl and the second was also a girl, we expected the probability that the household would have additional children to be highest. In this case, there was indeed an almost 14 percent chance that a couple would add at least one child to the household. For those households that had a girl first but whose second child was a boy, we expected the probability of increasing household size to lie between the two, which indeed turned out to be the case. The lower panel shows that the gap between actual and potential household size is also largest when the first two children are boys. The gap is smallest when the first two children are both girls. All this again confirms our proposition that boy preference affects fertility.

Finally, columns 5 and 6 propose the share of female children in the total number of a household's children as a potential instrument. As already stated above, because of boy preference, female children are likely to live in larger families and so we expected a positive correlation between this measure and household size. For the time being, we simply divided the sample in two, conditional on whether more or less than half of the children were female. We found that the average number of children was indeed smaller in the sub-sample where less than half of the children were girls as opposed to the sub-sample where the majority were girls (top panel, 2.78 children as opposed to 2.88). We also found a difference in the fertility gap that was significant with an associated p-value of 0.021 (bottom panel). Households where girls were in a majority were closer to the theoretical maximal household size.

Agricultural Labor Supply

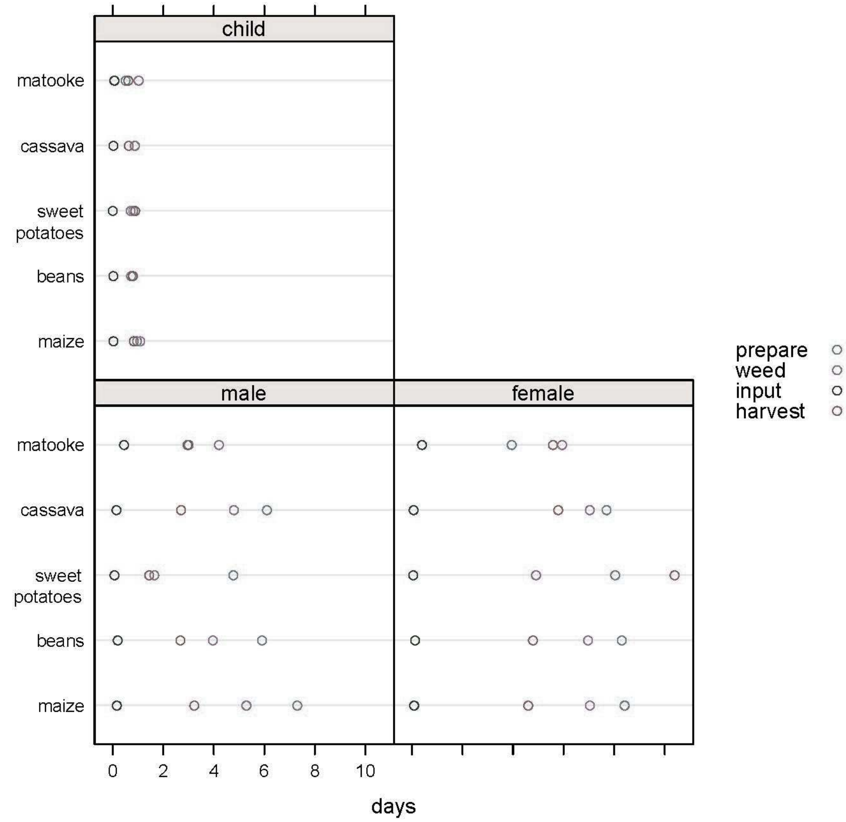
This section looks briefly at some descriptive statistics on labor supply in agriculture, one of the prime pathways through which fertility is likely to affect productivity and wellbeing. Most of Uganda has two cropping seasons. The first runs from January to June. The second is from July to December. The UNHS 2005–2006 interviewed households twice over the course of one year to capture this feature. It visited households in the beginning of 2005 to capture the second 2004 cropping season (which runs from July to December 2004). Enumerators revisited the households at the end of 2005 to record information from the first 2005 cropping season (which runs from January to June 2005). Our study only considers the 2004 July to December cropping season, as data for labor allocation in agriculture was unavailable for the 2005 cropping season.

Figure 4.1 shows time reported in the fields along different dimensions[7]. Women seemed to do most of the work, and reported child labor was negligible[8]. This already indicated that the tradeoff between the time lost by the mother because of rearing the children and the time gained by extra hands was likely to work against agricultural production. Typical for Uganda is the short amount of time spent on applying inputs. Farmers in Uganda use very limited amounts of fertilizer and other inputs, so also the time spent on applying them is short. There is also some heterogeneity in the time spent on different crops. For instance, matooke is allocated less time than maize, both for men and for women. However, there are also differences between the sexes. For instance, women spend much more time cultivating sweet potatoes, and to a lesser extent beans, than men.

[7]The dimensions are mother, father or child; land preparation, input application, weeding and harvesting; and crop. The crops are the five most widely grown crops in Uganda.

[8]While the relationship between fertility and child labor is an important research question, we do not consider this in the present study. Reported child labor occurrence indeed seems limited in Uganda. More importantly, the instruments we propose in this study (gender of first born child/children and sex composition of children within the household) are likely to directly influence the number of days the children worked in agriculture, instead of only through fertility, risking to violate the exclusion restriction.

Figure 4.1 Average number of days worked



Source: Author's creation.

Production

We also investigated fertility's effect on production of some of the most important crops. In particular, we looked at fertility's effect on the likelihood that a household cultivate each of the five most important crops. The first row in Table 4.2 reports on the percentage of households that grew each of the crops. Over 50 percent reported growing maize, beans and cassava. We also looked at the impact on area cultivated, measured in acres. Households on average allocated about half an acre to maize, while the least space was reserved for sweet potatoes. We also expressed area cultivated as a share of total area under cultivation. We found that about 17 percent of total land area was allocated to maize, while only 8 percent was allocated to sweet potatoes. The next line reports average production in kilograms at the household level. This may seem low, but this is because households that reported that they did not produce the crop were also part of the average. We also divided by household size. Finally, Table 4.2 reports yields for the five crops, defined as the amount of each crop harvested per unit of land (per acre).

We also aggregated the different corps by weighting them by average prices. We used prices from FoodNet[9]. In particular, we averaged prices observed in Kampala's Nakawa market over the July-to-December period in 2004. Doing so, we found that the average total value derived from these five crops was about UGX98,500, which translates to about UGX45,000 per capita[10]. About 40 percent of the households in their reproductive age did not cultivate any of these five crops. On average, about 0.69 acres was allocated to these five crops. The yield per acre was about UGX220,220.

[9] http://www.foodnet.cgiar.org/

[10] UGX stands for Ugandan Shillings, the national currency. At the time of the survey, USD1 = UGX1,780.

Table 4.2 Descriptive statistics for crop production

Variable	Maize	Beans	Sweet potatoes	Cassava	Matooke
Growing crop (% of households)	59.1	52.6	38.8	51.8	43.1
Crop area (acre)	0.473	0.263	0.165	0.301	0.264
Crop area (% of total area)	17.1	12.6	8.2	12.0	10.2
Production (kg)	38.7	12.8	85.1	83.2	348.9
Production per capita (kg)	18.1	6.4	38.7	38.4	168.6
Yield (kg per acre)	358.2	128.5	1096.3	1030.8	2067.3

Source: Author's calculations based on the UNHS 2005–2006 data.
Note: kg = kilogram

5. RESULTS

This section presents the results of our two-stage least squares estimates that looked at the causal impact of fertility on various agriculture related outcomes. The section starts by presenting the first stage regression of our proposed instruments on the fertility gap. It then gives a detailed description of the second stage regression that focuses on fertility's effect on agricultural labor supply. This section also explores fertility's effect on area planted, production, and productivity.

The First Stage Regression

Table 5.1 reports the results for the first stage regression, which linked the sex of first child/children to fertility. The dependent variable, as explained above, was the difference between the maximum number of children of a typical woman at her age and the actual number of children the women bore[11]. We referred to this as the fertility gap (*fgap*) or fertility shortfall. This is actually the reverse of fertility, as the bigger the gap, the lower the number of children in the household in a given age cohort. Apart from the exogenous variable that was excluded from the second stage regression elaborated in the next sections, we included a series of control variables that were clearly exogenous to fertility in all four specifications of the first stage. The first exogenous control variable, *femhead,* was an indicator variable that took the value of 1 if the household head was female. The second, *urban*, was an indicator variable that took the value of 1 if the household resided in an urban area[12]. Next, we included three dummies to account for the education level of the mother. The first, *mprim*, took the value of 1 if the mother had completed primary education. The second, *msec,* was the additional effect of having completed secondary education. The third, *mthird,* was the additional effect of the mother having completed tertiary education. The comparison category was therefore households where the mother did not complete at least primary education. We also added two community variables that were likely to influence household size. These wew *school* which was a dummy variable that took the value of 1 if there was a school in the village, and *health,* which was a dummy that took the value of 1 if there was a public health center or clinic in the community. Finally, we also added an indicator (*cdied*) that took the value of one if a son or daughter of the mother had died in the past.

We experimented with four different possible excluded instruments. Model (1) used an indicator that took the value of 1 if the firstborn in the household was a girl as an excluded instrument (*oldestgirl*). The coefficient was significant at the 1 percent level and had the expected sign. Having a girl as the firstborn offspring reduced the fertility gap by about 0.2 children. In other words, households that had a girl as a firstborn tended to be closer to maximal fecundity. For the controls, we found that households where females were the head had a significantly larger fertility gap. The effect was very large, suggesting that such mothers had more than 1 child less than the maximum. Also, in urban areas, households seemed to have significantly fewer children. Schooling of the mother seemed to reduce the number of children only at the secondary and tertiary level. There seemed to be some indication that mothers who had completed primary education had a slightly smaller fertility gap than mothers who had not completed even primary education. The community variables did not seem to have an effect on the fertility gap. Finally, having lost a child in the past left a significant additional fertility gap compared with households that had never lost a child. However, the additional gap was much lower than 1, suggesting a substantial replacement effect in Ugandan fertility.

[11]The maximum number of children has been estimated from the DHS and is actually the 95th percentile.

[12]In some specifications where we expect regional variation in the outcome variable to be important, such as for production and yields for certain crops, we also include dummies for the four regions in both the first and second stage equations. This addition did not significantly change other estimated parameters in the first stage.

Table 5.1 First stage regression: OLS estimation of fertility gap

Variable	(1)	(2)	(3)	(4)
ldestgirl	-0.203**			
	(0.067)			
2oldestgirls		-0.190*		
		(0.082)		
3oldestgirls			-0.147	
			(0.117)	
percentfmales				-0.278**
				(0.094)
femhead	1.186**	1.168**	1.201**	1.186**
	(0.098)	(0.105)	(0.118)	(0.098)
urban	0.322**	0.273**	0.077	0.325**
	(0.083)	(0.097)	(0.115)	(0.083)
mprim	-0.155*	-0.025	-0.009	-0.159*
	(0.075)	(0.082)	(0.094)	(0.075)
msec	0.259*	0.220+	0.193	0.257*
	(0.101)	(0.121)	(0.150)	(0.101)
mthird	1.058**	0.914**	0.755+	1.060**
	(0.192)	(0.250)	(0.403)	(0.192)
health	0.095	0.107	0.151	0.090
	(0.124)	(0.146)	(0.171)	(0.124)
school	0.040	-0.005	-0.118	0.043
	(0.070)	(0.078)	(0.090)	(0.070)
cdied	0.284**	0.204+	0.117	0.285**
	(0.100)	(0.108)	(0.127)	(0.100)
cons	2.172**	1.946**	1.782**	2.209**
	(0.074)	(0.075)	(0.079)	(0.081)
r2	0.091	0.075	0.065	0.091
N	2656	2036	1391	2656

Source: Author's calculations based on the UNHS 2005–2006 data.
Notes: OLS = Ordinary least square. Huber-White standard errors in parentheses, +, * and ** denote significance at the 10, 5 and 1 percent level respectively.

Model (2) used an indicator that equaled 1 if the first two children born to the mother in the household were both girls as excluded instrument (*2oldestgirls*). Using this instrument only made sense if we confined ourselves to households that had at least two children, hence the reduction in the sample size. As in (1) the parameter on the excluded instrument was significantly negative, in line with our hypothesis. The control variables were very close to what they had been in model (1). Model (3) went one step further and considered the first three children. In this case, the indicator, *3oldestgirls* was one only if the 3 first children were all girls. This again only made sense for households that had at least three children, further reducing the sample size. The coefficient estimate was again negative, but this time it was not significant anymore. We assumed that the reduced sample size in this model might have reduced the power of the t-test too much.

Model (4) used a continuous variable as an excluded instrument (*percentfemales*). We calculated the share of girls among children as a share of the total number of children in the household. Again, the coefficient on this instrument had the expected sign. A higher share of females within the households was associated with a smaller fertility gap. This was consistent with Jayachandran and Kuziemko (2011), who observed that the "try until you have a boy" fertility rule leads to an outcome where larger households have on average more girls. Again, the other variables were similar to the previous models. We found that a daughters-only household (*percentfemales* = 1) would be on average 0.28 children larger than a sons-only household (*percentfemales* = 0).

While most of our instruments were significant and had the expected sign, they explained only a small part of the variance in the outcome. When all exogenous controls were included, the R-squared was indeed rather low. If we ran partial regressions, regressing the excluded instruments one by one on the dependent variable, the R-square dropped below 1 percent. The F-value of a regression with only excluded instruments —another important indicator of the strength of the instruments according to Bound et al. (1995)—also dropped to 9.46[13]. In other words, we had serious concerns that our instruments were weak. We therefore used inference that was robust to weak instruments. In particular, we relied on the Anderson-Rubin test statistic to gauge the significance of the endogenous variable in all subsequent regressions (Staiger and Stock 1997).

Household Labor Supply

This section turns to fertility's effect on total household adult labor supply (Table 5.2). It also looks at labor supply separately for the mother and the father (Table 5.3), and by activity (Table 5.4).

Table 5.2 investigates the effect of our main variable of interest, the fertility shortfall, on the number of days worked in agriculture (land preparation, input application, weeding and harvesting)[14]. The first column of the table reports the result without taking into account endogeneity of number of children. It reports ordinary least squares (OLS) estimates that explain the number of days that adults reported to have worked on the household farm in the 2004 agricultural season. Agricultural work was defined as work related to land preparation, input application, weeding and harvesting. We saw that there was no significant correlation between the number of days worked and fertility as measured by the fertility gap. We did find significant and negative effects when the household was being headed by a female (*femhead*) and the household was located in an urban area (*urban*). Primary and secondary education of mothers (*mprim* and *msec*) did not seem to systematically relate to the number of days worked in agriculture, but mothers who had finished tertiary education (*mthird*) appeared to work less in agriculture. The OLS estimates also showed positive correlations between a school in the community (*school*) and days worked in agriculture and between a deceased child in the past (*cdead*) and days worked by the parents. There is also some indication of a positive correlation between health centers in the community (*health*) and days worked by both parents.

[13]As a rule of thumb, it is often stated that one has weak instruments if this F-statistic is smaller than 10.

[14]We have also done this analysis using days worked per acre of land held by the household. However, since average land holdings are about 1.1 acre and there seems to be no systematic relationship between farm size and labor supply, the results were very similar.

Table 5.2 Effect of fertility on total time worked in agriculture

Variable	(1) OLS	(2) 2SLS	(3) 2SLS	(4) 2SLS	(5) LIML
fgap	-0.122	46.799+	64.297+	69.227**	66.349**
	(1.245)	(32.112)	(48.865)	(35.208)	(36.319)
femhead	-38.213**	-95.268*	-116.211+	-122.540**	-119.040**
	(4.793)	(39.623)	(59.496)	(44.167)	(45.268)
urban	-23.411**	-35.338**	-40.541**	-41.038**	-40.307**
	(5.780)	(11.352)	(13.796)	(13.622)	(13.516)
mprim	1.392	8.132	6.058	11.354	10.940
	(4.897)	(7.802)	(8.092)	(9.128)	(9.065)
msec	-6.121	-10.595	-7.207	-12.734	-12.459
	(6.516)	(8.899)	(11.617)	(11.167)	(10.850)
mthird	-20.792*	-70.642+	-83.864	-94.470*	-91.412*
	(9.946)	(39.814)	(54.795)	(45.224)	(46.122)
health	-15.118*	-21.809*	-18.573	-25.008+	-24.597+
	(7.694)	(11.099)	(14.923)	(13.928)	(13.544)
school	12.570**	7.633	9.887	5.274	5.576
	(4.793)	(6.914)	(9.145)	(7.881)	(7.879)
cdead	13.573*	3.161	1.904	-1.816	-1.177
	(6.273)	(10.788)	(12.672)	(12.954)	(12.801)
cons	87.193**	-7.527	-29.950	-52.803	-46.992
	(4.718)	(64.960)	(90.557)	(70.927)	(73.206)
N	2016	2016	1620	2016	2016
instrument:	-	1st = girl	1st & 2nd = girl	% girl	1st = girl & % girl

Source: Author's calculations based on the UNHS 2005–2006 data.

Notes: OLS = Ordinary least square; 2SLS = Two-stage least squares; LIML = Limited information maximum likelihood. Huber-White standard errors in parentheses, +, * and ** denote significance at the 10, 5 and 1 percent level respectively.

Models (2) to (4) estimated the same models, but instrumented the fertility gap with a single excluded instrument. In model (2), the instrument was an indicator taking the value of 1 if the firstborn was a girl. The coefficient on the fertility gap then became positive and significant at a 10 percent level, implying that higher fertility (and hence a shrinking fertility gap) caused a reduction in the number of days the parents worked on the family fields. Model (3) used the sex of the first and second born as instruments for the fertility gap. Model (4) used the share of daughters as an instrument. The estimate of the fertility effect became higher, and is now significant even at a 1 percent significance level.

Finally, model (5) used both the gender of the firstborn and the share of daughters as the excluded instruments[15]. According to the Hansen-J statistic, our model that used multiple instruments was valid (Hansen-J=0.849; p-val=0.357). We thus assumed this was the preferred specification. Each additional child caused a reduction of about 66 days of labor in agriculture by the parents. With respect to the other variables in the regression, we found some signs that households where the mother had finished tertiary education appeared to be less engaged in agriculture.

Table 5.3 differentiates between work done by the mother and by the father. For the sake of space, it only shows the coefficient on the fertility gap, but we also added the exogenous control variables that

[15] We used limited information maximum likelihood (LIML), as this is known to have better small sample properties than 2SLS in overidentified models with weak instruments (Angrist and Krueger 2001).

were also included in the first stage regression. Full results can be found in the Appendix (Table A.1 and Table A.2). The top panel in Table 5.3 shows the effect of fertility on time worked in agriculture by the mother. The OLS estimate is not significant (model (1)). Accounting for endogeneity of fertility using the exogenous variation caused by the sex of the firstborn rendered the fertility gap significant at a 5 percent level (model (2)). An increase in the fertility gap per age cohort by one child lead a mother to work almost 30 days more in subsistence farming. Cycling through the results with the alternative instruments, the results change little with respect to significance. In all, an additional child seemed to reduce the number of days the mother worked in agricultural production by about 40 days. Full results are reported in Table A.1.

Table 5.3 2SLS estimates of household labor supply

	(1) OLS	(2) 2SLS	(3) 2SLS	(4) 2SLS	(5) LIML
			days worked mother		
	0.533	29.890*	54.070**	40.841**	38.773**
	(0.735)	(17.769)	(33.364)	(19.353)	(19.085)
			days worked father		
	-0.620	10.928	5.915	22.327*	20.076+
	(0.668)	(13.983)	(25.595)	(13.580)	(14.756)
N	2016	2016	1620	2016	2016
instrument:	-	1st = girl	1st & 2nd = girl	% girl	1st = girl & % girl

Source: Author's calculations based on the UNHS 2005–2006 data.
Notse: OLS = Ordinary least square; 2SLS = Two-stage least squares; LIML = Limited information maximum likelihood. Huber-White standard errors in parentheses, +, * and ** denote significance at the 10, 5 and 1 percent level respectively.

The first column of the second panel reports the same OLS regression but with the number of days adult males worked as the dependent variable. As was the case with female labor, the fertility gap did not seem to correlate to male labor supply when we did not take into account the endogenous nature of fertility choices. Table A.2 in the appendix gives full results and the OLS results are in the first column. We found that living in urban areas led to farmers reporting fewer days worked in the field. In households headed by females, we also found a large negative effect on the number of days that men worked in agriculture-related activities. This is because, in most cases, households are headed by females because the male head is missing, leading to fewer days reported in the field. We also found some evidence of males working less if the mother had higher education. This is most likely because men with higher education choose women with higher education to marry and the other way around.

Judged by the instrumental variable models from (2) to (5), the effect of fertility on labor supply by the father was less clear cut. When we used the sex of the first born (model 2) and the sex of the first two children born (model 3) as instruments, the coefficient was positive but not significantly different from zero. If we instrumented the fertility gap using the percentage of females, we found some indication that more children might reduce time allocated to working in the field by the father. The effect, however, is only half the size of the reductions we found for women. The over-identified model in model (5) showed a significant effect at the 10 percent level only. These findings were similar to what others have found. For instance, in their study on labor supply response to fertility in the United States, Angrist and Evans (1998) also found that women work less while men did not alter their labor supply in response to having more children. Kim and Aassve (2006) found that Indonesian women reduced their working days in response to the higher fecundity in both rural and urban areas.

Table 5.4 looks at reported labor by activity instead of by sex. Again, the results in Table 5.4 only show the coefficients on the fertility gap. Full results are in the appendix. Model (1) in the top panel presents OLS results for number of days worked on land preparation. There were no effects from fertility in this specification. Again, as expected, households living in urban areas spent significantly less time preparing land. Female-headed households also allocated less time to land preparation. There was also some indication that women who had tertiary education were less engaged in land preparation.

Table 5.4 2SLS estimates of household labor allocation

	(1) OLS	(2) 2SLS	(3) 2SLS	(4) 2SLS	(5) LIML
			time allocated to land preparation		
	-0.048	15.876	41.127**	26.028**	25.278*
	(0.528)	(12.737)	(25.884)	(13.720)	(14.761)
			time allocated to input application		
	0.051	1.812	0.849	2.372	2.224
	(0.136)	(1.632)	(1.686)	(1.829)	(1.741)
			time allocated to weeding		
	0.026	17.708*	25.959*	21.385*	20.492*
	(0.468)	(11.724)	(17.541)	(11.730)	(11.253)
			time allocated to harvesting		
	-0.082	5.315	-3.591	13.852+	12.485+
	(0.468)	(10.597)	(20.563)	(8.987)	(10.379)
N	2015	2015	1619	2015	2015
instrument:	-	1st = girl	1st & 2nd = girl	% girl	1st = girl & % girl

Source: Author's calculations based on the UNHS 2005–2006.
Note: OLS = Ordinary least square; 2SLS = Two-stage least squares; LIML = Limited information maximum likelihood. Huber-White standard errors in parentheses, +, * and ** denote significance at the 10, 5 and 1 percent level respectively.

Model (2) presents the same model, but instruments fertility with the indicator for the first child being a girl. When we ran this model, the fertility effect became positive, but was not significantly different from zero. In model (3), which looks at the sex of the first two children, the fertility gap effect became significant. The effect remained significant when we instrumented the fertility shortfall by the percentage of females born (model (4)) and in the over-identified model (5), but the effect size shrank. An additional child reduced time allocated to land preparation by about 25 days.

The second panel repeated the same five models, but used days spent on input application as the dependent variable. In none of the five specifications did fertility seem to have a significant impact. Overall, time spent on input application was very limited anyway, as can be seen in Figure 4.1. In all, households spent only about one day applying inputs (including planting). The only significant effect we found is that households where the mother has at least primary education allocated more time to input application (Table A.4).

The third panel presents results for time spent on weeding. The results were similar to the ones for land preparation, but the effects are smaller. Each extra child reduced time allocated to weeding by about 20 days. Full results in the appendix (Table A.5) show significant negative effects for female-headed households and for households in urban areas. We also found that households in communities that have a health center spent fewer days on weeding. Finally, the last panel looks at the effects of fertility on days worked for harvesting. There was no significant positive association between the number of children in the family and the number of days spent harvesting if we used only our binary instrument. We found a positive effect if we instrumented the fertility gap by the share of girls among siblings, but the effect was small compared to the other effects.

The above suggests that fertility affects time allocated to land preparation and weeding in a negative way. Harvesting seems to be less related to family size. Probably, when crops are ready to be harvested, farmers are more likely to put in the extra effort. This seems to be less evident for work that has an uncertain payoff in the future, such as weeding. The reductions of time allocated to land preparation and weeding may reduce both area planted and agricultural productivity. The next section looks at this question.

Area Planted, Production, and Productivity

This section looks at fertility's effect on production and productivity. It looks at productivity defined as kilograms harvested per acre of the five most important products separately. It also looks at the value of total production, the value of production per acre, and the value of production per capita.

Table 5.5 reports on the second stage regressions of different aspects of production for the five most important crops. The table only reports the results for the coefficient on the fertility gap for the instrumental variable regression that uses the share of girls as excluded instrument. The regressions include the same control variables as in the previous sections. However, we now also added regional dummies, as some crops are grown more in some regions than in others. When the dependent variable was binary or censored, we estimated a tobit or probit using the methods described in Newey (1987).

Table 5.5 2SLS estimates of effect of fertility on crop mix, area, production, and yield

Variable	Maize	Beans	Sweet potato	Cassava	Matooke
growing	0.431	-0.193	0.556+	-0.003	0.622+
	(0.359)	(0.332)	(0.401)	(0.302)	(0.428)
total area	0.432	-0.119	0.187	0.105	0.603+
	(0.425)	(0.199)	(0.254)	(0.289)	(0.425)
area share	0.028	-0.085	0.169+	-0.027	0.043
	(0.087)	(0.080)	(0.115)	(0.081)	(0.081)
production	56.799	-4.677	182.878	29.714	1931.785+
	(71.795)	(22.600)	(176.010)	(189.684)	(1255.565)
production per capita	26.593	-1.079	55.706	-11.725	2026.613*
	(38.247)	(12.727)	(87.191)	(103.129)	(1135.716)
yield	-22.060	41.329	69.834	-480.421	-383.889
	(169.179)	(54.371)	(694.248)	(728.459)	(882.517)

Source: Author's calculations based on the UNHS 2005–2006 data.

Notes: Huber-White standard errors in parentheses, +, * and ** denote significance at the 10, 5 and 1 percent level respectively. All regressions use the share of female children in total number of children as instrument.

The first row looks at the probability that a household grows the respective crop. For instance, the first entry in the first row tells us that the fertility gap did not affect the probability that households cultivate maize. The third entry, however, shows that households that had a higher fertility gap were more likely to grow beans. Similarly, we found that higher fertility significantly reduced the probability that matooke would be grown. The next row looks at the total area reported to be used to grow each crop, measured in acres. We found a positive effect of the fertility gap on the area used to grow matooke. Fertility seemed to be unrelated to the area used to grow any of the other crops. However, smaller households that grew more matooke might simply have had larger land holdings. Therefore, it was useful to also relate fertility to the share of each crop in total in terms of land size. This gave an idea of the relative importance of each crop within the household. This is presented in the next row. In this case, it seems that households with more children allocated less land as a share of total land to sweet potatoes. The next row looks at the value of production in kilograms. Only for matooke, larger households seemed to obtain a significantly lower quantity of matooke. The next row looks at production per capita. The lower production of matooke persisted if we accounted for household size. Finally, the fertility gap had no significant effect on yield for any of the products.

Finally, Table 5.6 presents results for total production and productivity, using the prices for the different crops. Again we used five different models. The first one was again a regression that did not take endogeneity into account. While in the previous regressions this was typically OLS, this might now have changed to a probit or tobit regression, depending on the nature of the dependent variable. The second regression instrumented the fertility gap with the sex of the firstborn. The third model used the sex of the first two children born to the mother and the fourth used the share of girls among the children. As before, the fourth model instrumented the fertility difference by two instruments: the sex of the firstborn and the share of girls among the children.

Table 5.6 2SLS estimates of total production

Variable	**(1) OLS**	**(2) 2SLS**	**(3) 2SLS**	**(4) 2SLS**	**(5) LIML**
production	-3.411	21.056	-20.275	16.988	19.334
(x UGX1000)	(2.697)	(46.911)	(57.977)	(48.493)	(44.172)
production/capita	4.133**	10.501	-9.251	14.518	12.340
(x UGX1000)	(1.527)	(26.009)	(24.107)	(27.183)	(24.605)
area	-0.033	0.120	0.042	0.145	0.132
	(0.023)	(0.343)	(0.443)	(0.359)	(0.325)
yield	-1.697	43.088	-96.861	0.013	8.818
	(2.862)	(81.451)	(140.870)	(61.118)	(69.674)
instrument:	-	1st = girl	1st & 2nd = girl	% girl	1st = girl & % girl

Source: Author's calculations based on the UNHS 2005–2006 data.

Notes: OLS = Ordinary least square; 2SLS = Two-stage least squares; LIML = Limited information maximum likelihood; UGX = Ugandan shilling. Huber-White standard errors in parentheses, +, * and ** denote significance at the 10, 5 and 1 percent level respectively.

The first row gives results for the change in production. There seemed to be no detectable effect from fertility on the total value of the production of the five crops. The second row expresses this production in per capita terms. The OLS estimates showed a positive effect of an increase in the fertility gap. However, if we confined ourselves to the exogenous part of fertility in the IV regressions, the effect disappeared. The next row looks at a change in the total area allocated to the five crops. It showed again no significant effect from fertility. The final row, which looked at productivity defined as the total value of the five crops divided by the total area allocated to these five crops, also showed no causal impact from family size.

6. CONCLUSION

We looked at the effect of fertility, defined as the number of biological children born to a mother, on agricultural production and its determinants. One of the most evident determinants was household agricultural labor. The identification strategy we used relies on the premise that, in partilineal societies, boys are preferred to girls in terms of offspring. Households that have a girl as the first child will have a higher propensity to add more children to the household. The fact that the sex of the first child is exogenous can be used to identify the causal impact of additional children on other variables such as labor supply and productivity. Similarly, the fertility rule whereby one is more likely to stop having children after a boy means that, on average, larger households consist of more girls. Therefore, the share of females in the total number of children can also be used as an instrument.

Our first stage regression performed reasonably well. We found a significant negative effect of an indicator variable that the firstborn was a girl on a variable that measures the shortfall from fecundity. We equally found a negative effect of an indicator that the first two children were female. Finally, we also found that households with a relatively higher share of girls were negatively related to the fertility gap. While our instruments were significant and had the correct sign, explanatory power as measured by the partial R-squared was low. We therefore used inference methods in the second stage that were robust to weak instruments.

In the second stage regression, we found that fertility affects the time both women and men allocated to agricultural production. However, most of the labor time lost as a consequence of an exogenous increase in children was borne by the woman. Land preparation and weeding, especially, were activities that seemed to suffer from excessive fertility. When we looked at crops, we found that only matooke and sweet potatoes were significantly affected by fertility.

Matooke is the most important stable crop in Uganda, providing 18 percent of caloric intake (Haggblade and Dewina 2010). The finding that young households that have higher fertility were reducing the most important source of calories suggests that higher fertility also causes under-nutrition. Sweet potatoes are also a typical food security crop, with a low return but also low risk (Dercon 1996). It is also a crop that is mostly under the control of the women, who do much of the work on the field.

That said, the fact that we relied on a cross-section of households also limits extent to which our conclusions can be generalized. It may well be that couples that have more children profit much more from larger household size at a later stage in life. For instance, in households where the mother has reduced fertility, she may have more time to work in agricultural activities. In addition, children may provide cheap and flexible labor at a later age. Therefore, we want to stress that our results only hold for the subset of "young" households, where the woman is between 16 and 32.

There are different ways in which the negative effect of fertility on labor and production can be influenced. First, our analysis reconfirms the need for fertility-reducing policies. Apart from known fertility-reducing policies such as women's education and improved maternal health care, the most promising policies should try to work on the root cause of increased fertility. This should be done by reducing households' propensity to have higher fertility if the firstborn is a girl. We can think of a host of policies that would do this by pushing against the patrilineal nature of these societies. For example, Uganda may consider changing its land act to make it similar to what Kenya recently did and give equal inheritance rights to both girls and boys.

Policy response involves addressing cultural issues related to high fertility, some of which may face considerable resistance. Changing a set of cultural values is likely to be a very slow process. Meanwhile, the government of Uganda should support the nutritional needs of young families. It should also consider introducing agricultural technologies that save on agricultural labor, especially for women.

APPENDIX: SUPPLEMENTARY TABLES

Table A.1 Effect on days worked by mother (full results)

Variable	(1) OLS	(2) 2SLS	(3) 2SLS	(4) 2SLS	(5) LIML
fgap	0.533	29.890*	54.070**	40.841**	38.773**
	(0.735)	(17.769)	(33.364)	(19.353)	(19.085)
femhead	-7.906*	-43.577*	-71.806+	-56.883*	-54.370*
	(3.715)	(21.981)	(40.831)	(24.354)	(23.883)
urban	-13.317**	-20.758**	-25.334*	-23.534**	-23.009**
	(3.536)	(6.910)	(10.247)	(8.041)	(7.811)
mprim	0.023	4.177	3.472	5.726	5.434
	(2.754)	(4.566)	(5.864)	(5.194)	(5.085)
msec	1.115	-1.624	2.608	-2.646	-2.453
	(4.706)	(5.869)	(9.177)	(6.913)	(6.676)
mthird	-15.001*	-46.192*	-70.087+	-57.827*	-55.630*
	(6.961)	(23.251)	(38.987)	(25.869)	(25.530)
health	-7.274	-11.634+	-9.296	-13.260	-12.953
	(4.607)	(6.967)	(11.406)	(8.326)	(8.044)
school	6.740*	3.705	3.073	2.573	2.786
	(2.703)	(3.837)	(6.100)	(4.461)	(4.334)
cdied	6.810+	0.209	-2.949	-2.253	-1.788
	(3.874)	(6.592)	(9.328)	(7.510)	(7.333)
cons	49.054**	-10.209	-48.668	-32.314	-28.140
	(2.815)	(36.011)	(61.945)	(39.017)	(38.525)
N	2016	2016	1620	2016	2016
Instrument:	-	1st = girl	1st & 2nd = girl	% girl	1st = girl & % girl

Source: Author's calculations based on the UNHS 2005–2006 data.

Notes: OLS = Ordinary least square; 2SLS = Two-stage least squares; LIML = Limited information maximum likelihood. Huber-White standard errors in parentheses, +, * and ** denote significance at the 10, 5 and 1 percent level respectively.

Table A.2 Effect on days worked by father (full results)

Variable	(1) OLS	(2) 2SLS	(3) 2SLS	(4) 2SLS	(5) LIML
fgap	-0.620	10.928	5.915	22.327*	20.076+
	(0.668)	(13.983)	(25.595)	(13.580)	(14.756)
femhead	-30.349**	-44.381**	-39.124	-58.231**	-55.496**
	(2.216)	(16.903)	(30.436)	(16.952)	(18.155)
urban	-10.103**	-13.030**	-14.318**	-15.919**	-15.349**
	(2.902)	(4.347)	(5.249)	(5.056)	(5.016)
mprim	1.374	3.008	2.241	4.620	4.302
	(2.717)	(3.271)	(2.967)	(3.769)	(3.687)
msec	-7.239*	-8.317*	-9.670*	-9.381*	-9.171*
	(3.120)	(3.454)	(3.770)	(4.404)	(4.175)
mthird	-5.829	-18.098	-9.522	-30.209+	-27.817
	(4.926)	(16.387)	(26.183)	(17.049)	(17.973)
health	-7.849*	-9.564*	-9.534*	-11.257*	-10.922*
	(3.589)	(4.045)	(4.444)	(5.168)	(4.913)
school	5.826*	4.632	7.310	3.454	3.687
	(2.843)	(3.630)	(4.782)	(3.620)	(3.742)
cdied	6.756*	4.159	5.314	1.596	2.102
	(3.272)	(4.374)	(5.030)	(5.209)	(5.089)
cons	38.067**	14.756	26.774	-8.254	-3.711
	(2.377)	(28.289)	(47.345)	(27.361)	(29.750)
N	2016	2016	1620	2016	2016
instrument:	-	1st = girl	1st & 2nd = girl	% girl	1st = girl & % girl

Source: Author's calculations based on the UNHS 2005–2006 data.

Notes: OLS = Ordinary least square; 2SLS = Two-stage least squares; LIML = Limited information maximum likelihood. Huber-White standard errors in parentheses, +, * and ** denote significance at the 10, 5 and 1 percent level respectively.

Table A.3 Effect on days spend on preparing fields (full results)

Variable	(1) OLS	(2) 2SLS	(3) 2SLS	(4) 2SLS	(5) LIML
fgap	-0.048	15.876	41.127**	26.028**	25.278*
	(0.528)	(12.737)	(25.884)	(13.720)	(14.761)
femhead	-15.173**	-34.568*	-64.891*	-46.932**	-46.018*
	(2.023)	(15.704)	(31.846)	(17.259)	(18.412)
urban	-6.940**	-11.013*	-17.002*	-13.610*	-13.418*
	(2.623)	(4.481)	(7.827)	(5.417)	(5.493)
mprim 2	-2.251	-0.069	-0.597	1.322	1.219
	(2.209)	(3.307)	(4.513)	(3.585)	(3.703)
msec	-2.980	-4.482	-2.924	-5.439	-5.368
	(2.832)	(3.537)	(6.729)	(4.377)	(4.336)
mthird	-11.214**	-28.140+	-51.405+	-38.931*	-38.133*
	(3.993)	(15.545)	(29.529)	(17.591)	(18.553)
health	-3.914	-6.215	-6.296	-7.682	-7.573
	(3.233)	(4.293)	(8.266)	(5.446)	(5.379)
school	2.919	1.343	0.511	0.338	0.412
	(1.932)	(2.546)	(4.654)	(3.079)	(3.073)
cdead	5.262+	1.632	-2.547	-0.682	-0.511
	(2.719)	(4.291)	(7.133)	(5.142)	(5.203)
cons	35.817**	3.741	-39.561	-16.708	-15.197
	(2.261)	(26.072)	(47.770)	(27.523)	(29.771)
N	2015	2015	1619	2015	2015
instrument:	-	1st = girl	1st & 2nd = girl	% girl	1st = girl & % girl

Source: Author's calculations based on the UNHS 2005–2006 data.

Notes: OLS = Ordinary least square; 2SLS = Two-stage least squares; LIML = Limited information maximum likelihood. Huber-White standard errors in parentheses, +, * and ** denote significance at the 10, 5 and 1 percent level respectively.

Table A.4 Effect on days spend on input application (full results)

Variable	(1) OLS	(2) 2SLS	(3) 2SLS	(4) 2SLS	(5) LIML
fgap	0.051	1.812	0.849	2.372	2.224
	(0.136)	(1.632)	(1.686)	(1.829)	(1.741)
femhead	0.349	-1.795	-0.493	-2.478	-2.297
	(0.747)	(1.611)	(1.756)	(1.717)	(1.628)
urban	-0.514	-0.964	-0.674	-1.107	-1.069
	(0.386)	(0.713)	(0.615)	(0.788)	(0.763)
mprim	0.466*	0.707*	0.657**	0.784*	0.764*
	(0.182)	(0.325)	(0.246)	(0.376)	(0.358)
msec	0.978	0.812	1.158	0.759	0.773
	(0.881)	(0.814)	(1.126)	(0.802)	(0.804)
mthird	-0.684	-2.555	-1.897	-3.151	-2.993
	(1.202)	(2.575)	(2.550)	(2.835)	(2.745)
health	0.198	-0.057	-0.088	-0.138	-0.116
	(0.489)	(0.611)	(0.613)	(0.687)	(0.663)
school	-0.045	-0.219	-0.108	-0.275	-0.260
	(0.289)	(0.302)	(0.339)	(0.302)	(0.300)
cdead	-0.142	-0.543	-0.235	-0.671	-0.637
	(0.222)	(0.513)	(0.401)	(0.591)	(0.564)
cons	0.526	-3.021	-0.935	-4.150	-3.851
	(0.448)	(3.370)	(3.238)	(3.793)	(3.609)
N	2015	2015	1619	2015	2015
instrument:	-	1st = girl	1st & 2nd = girl	% girl	1st = girl & % girl

Source: Author's calculations based on the UNHS 2005–2006 data.

Notes: OLS = Ordinary least square; 2SLS = Two-stage least squares; LIML = Limited information maximum likelihood. Huber-White standard errors in parentheses, +, * and ** denote significance at the 10, 5 and 1 percent level respectively.

Table A.5 Effect on days spend on weeding (full results)

Variable	(1) OLS	(2) 2SLS	(3) 2SLS	(4) 2SLS	(5) LIML
fgap	0.026	17.708*	25.959*	21.385*	20.492*
	(0.468)	(11.724)	(17.541)	(11.730)	(11.253)
femhead	-13.623**	-35.158*	-45.639*	-39.637**	-38.550**
	(1.728)	(14.648)	(21.645)	(14.779)	(14.180)
urban	-9.339**	-13.862**	-15.649**	-14.802**	-14.574**
	(2.340)	(4.495)	(5.509)	(4.756)	(4.613)
mprim	-0.489	1.934	1.010	2.438	2.315
	(1.772)	(2.772)	(3.137)	(2.968)	(2.880)
msec	-2.274	-3.942	-2.715	-4.289	-4.205
	(2.594)	(3.488)	(4.834)	(3.904)	(3.775)
mthird	-5.350	-24.143	-30.226	-28.052+	-27.103+
	(4.259)	(14.960)	(20.546)	(15.193)	(14.766)
health	-8.375**	-10.930**	-10.259+	-11.461**	-11.332**
	(2.059)	(3.844)	(5.363)	(4.270)	(4.136)
school	4.598**	2.848	3.290	2.484	2.572
	(1.751)	(2.336)	(3.165)	(2.470)	(2.417)
cdead	5.040*	1.010	0.656	0.172	0.375
	(2.323)	(4.315)	(5.183)	(4.525)	(4.412)
cons	29.534**	-6.083	-17.387	-13.490	-11.692
	(1.783)	(23.464)	(32.493)	(23.516)	(22.524)
N	2015	2015	1619	2015	2015
instrument:	-	1st = girl	1st & 2nd = girl	% girl	1st = girl & % girl

Source: Author's calculations based on the UNHS 2005–2006 data.

Notes: OLS = Ordinary least square; 2SLS = Two-stage least squares; LIML = Limited information maximum likelihood. Huber-White standard errors in parentheses, +, * and ** denote significance at the 10, 5 and 1 percent level respectively.

Table A.6 Effect on days spend on harvesting (full results)

Variable	(1) OLS	(2) 2SLS	(3) 2SLS	(4) 2SLS	(5) LIML
fgap	-0.082	5.315	-3.591	13.852+	12.485+
	(0.468)	(10.597)	(20.563)	(8.987)	(10.379)
femhead	-10.650**	-17.223	-6.270	-27.620*	-25.955*
	(1.771)	(12.609)	(24.287)	(11.128)	(12.624)
urban	-6.876**	-8.257**	-8.124*	-10.440**	-10.091**
	(1.873)	(2.977)	(3.852)	(3.362)	(3.402)
mprim	2.741	3.481	3.569+	4.650+	4.463+
	(1.906)	(2.131)	(2.042)	(2.511)	(2.448)
msec	-1.835	-2.344	-2.761	-3.149	-3.020
	(2.343)	(2.335)	(2.849)	(2.969)	(2.822)
mthird	-3.720	-9.457	-0.912	-18.530+	-17.078
	(3.366)	(11.938)	(20.598)	(11.097)	(12.290)
health	-2.340	-3.120	-0.215	-4.353	-4.156
	(3.167)	(3.283)	(4.032)	(4.004)	(3.853)
school	5.394*	4.860+	6.889+	4.015	4.150
	(2.150)	(2.803)	(3.685)	(2.604)	(2.760)
cdead	3.501	2.271	3.443	0.325	0.637
	(2.351)	(2.914)	(3.565)	(3.384)	(3.345)
cons	22.551**	11.680	29.707	-5.516	-2.763
	(1.469)	(21.379)	(38.106)	(18.133)	(20.921)
N	2015	2015	1619	2015	2015
instrument:	-	1st = girl	1st & 2nd = girl	% girl	1st = girl & % girl

Source: Author's calculations based on the UNHS 2005–2006 data.

Notes: OLS = Ordinary least square; 2SLS = Two-stage least squares; LIML = Limited information maximum likelihood. Huber-White standard errors in parentheses, +, * and ** denote significance at the 10, 5 and 1 percent level respectively.

Table A.7 Total production (full tobit results)

Variable	(1) OLS	(2) 2SLS	(3) 2SLS	(4) 2SLS	(5) LIML
fgap	-3.411	21.056	-20.275	16.988	19.334
	(2.697)	(46.911)	(57.977)	(48.493)	(44.172)
femhead	-69.276**	-97.608+	-41.725	-92.870	-95.599+
	(13.700)	(55.994)	(67.508)	(57.696)	(52.888)
urban	-191.662**	-199.111**	-183.271**	-197.891**	-198.597**
	(13.527)	(18.843)	(19.978)	(19.209)	(18.219)
mprim	31.542**	35.247**	40.222**	34.604**	34.970**
	(9.652)	(12.321)	(11.070)	(12.382)	(12.050)
msec	-14.660	-20.572	-17.281	-19.563	-20.150
	(14.527)	(17.768)	(20.052)	(17.924)	(17.324)
mthird	29.624	2.888	78.249	7.286	4.749
	(35.753)	(58.152)	(63.402)	(59.720)	(55.538)
health	-12.603	-15.259	-10.523	-14.869	-15.089
	(15.952)	(16.802)	(19.526)	(16.816)	(16.688)
school	37.105**	36.348**	39.394**	36.483**	36.411**
	(9.238)	(9.316)	(10.397)	(9.273)	(9.279)
cdied	15.566	8.472	9.039	9.642	8.970
	(12.878)	(18.682)	(18.569)	(18.990)	(18.093)
cons	153.527**	100.065	197.728	108.964	103.831
	(12.388)	(103.164)	(121.588)	(106.568)	(97.211)
N	2637	2637	2020	2637	2637
instrument:	-	1st = girl	1st & 2nd = girl	% girl	1st = girl & % girl

Source: Author's calculations based on the UNHS 2005–2006 data.

Notes: OLS = Ordinary least square; 2SLS = Two-stage least squares; LIML = Limited information maximum likelihood. Huber-White standard errors in parentheses, +, * and ** denote significance at the 10, 5 and 1 percent level respectively.

Table A.8 Total production per capita (full tobit results)

Variable	(1) OLS	(2) 2SLS	(3) 2SLS	(4) 2SLS	(5) LIML
fgap	4.133**	10.501	-9.251	14.517	12.340
	(1.527)	(26.009)	(24.107)	(27.183)	(24.605)
femhead	-39.851**	-47.223	-10.658	-51.859	-49.346+
	(7.938)	(31.039)	(28.068)	(32.335)	(29.453)
substrat	-98.207**	-100.147**	-67.933**	-101.377**	-100.709**
	(8.560)	(10.451)	(8.290)	(10.764)	(10.150)
mprim	20.087**	21.052**	17.671**	21.646**	21.324**
	(5.220)	(6.835)	(4.606)	(6.946)	(6.716)
msec	-3.727	-5.265	-4.161	-6.221	-5.707
	(8.632)	(9.846)	(8.330)	(10.040)	(9.644)
mthird	20.268	13.311	55.245*	8.899	11.294
	(24.284)	(32.189)	(26.298)	(33.422)	(30.880)
health	-8.756	-9.447	-8.607	-9.909	-9.653
	(8.925)	(9.326)	(8.131)	(9.440)	(9.308)
school	12.227*	12.030*	11.697**	11.910*	11.976*
	(4.969)	(5.167)	(4.325)	(5.203)	(5.172)
cdied	5.932	4.085	2.685	2.916	3.551
	(7.121)	(10.360)	(7.726)	(10.649)	(10.082)
cons	49.910**	35.996	68.721	27.226	31.979
	(6.365)	(57.199)	(50.558)	(59.739)	(54.151)
N	2637	2637	2020	2637	2637
instrument:	-	1st = girl	1st & 2nd = girl	% girl	1st = girl & % girl

Source: Author's calculations based on the UNHS 2005–2006 data.

Notes: OLS = Ordinary least square; 2SLS = Two-stage least squares; LIML = Limited information maximum likelihood. Huber-White standard errors in parentheses, +, * and ** denote significance at the 10, 5 and 1 percent level respectively.

Table A.9 Total area (full tobit results)

Variable	(1) OLS	(2) 2SLS	(3) 2SLS	(4) 2SLS	(5) LIML
fgap	-0.033	0.120	0.042	0.145	0.132
	(0.023)	(0.343)	(0.443)	(0.359)	(0.325)
femhead	-0.511**	-0.689+	-0.530	-0.718+	-0.703+
	(0.097)	(0.410)	(0.516)	(0.427)	(0.389)
urban	-1.351**	-1.398**	-1.345**	-1.406**	-1.401**
	(0.111)	(0.138)	(0.153)	(0.142)	(0.134)
mprim	0.172*	0.196*	0.250**	0.199*	0.197*
	(0.071)	(0.090)	(0.085)	(0.091)	(0.089)
msec	-0.175	-0.212	-0.215	-0.218	-0.215+
	(0.111)	(0.130)	(0.153)	(0.133)	(0.127)
mthird	-0.035	-0.202	0.088	-0.230	-0.215
	(0.210)	(0.428)	(0.487)	(0.443)	(0.411)
health	-0.133	-0.150	-0.158	-0.153	-0.151
	(0.108)	(0.123)	(0.150)	(0.125)	(0.123)
school	0.310**	0.306**	0.331**	0.305**	0.305**
	(0.070)	(0.068)	(0.080)	(0.069)	(0.068)
cdied	0.103	0.058	0.012	0.051	0.055
	(0.093)	(0.137)	(0.142)	(0.140)	(0.133)
cons	0.793**	0.459	0.630	0.403	0.432
	(0.080)	(0.755)	(0.930)	(0.788)	(0.715)
N	2637	2637	2020	2637	2637
instrument:	-	1st = girl	1st & 2nd = girl	% girl	1st = girl & % girl

Source: Author’s calculations based on the UNHS 2005–2006 data.

Notes: OLS = Ordinary least square; 2SLS = Two-stage least squares; LIML = Limited information maximum likelihood. Huber-White standard errors in parentheses, +, * and ** denote significance at the 10, 5 and 1 percent level respectively.

Table A.10 Total yield (x1000 UGX per acre)

Variable	(1) OLS	(2) 2SLS	(3) 2SLS	(4) 2SLS	(5) LIML
fgap	-1.697	43.088	-96.861	0.013	8.818
	(2.862)	(81.451)	(140.870)	(61.118)	(69.674)
femhead	-32.827*	-82.434	62.869	-34.721	-44.474
	(13.620)	(91.141)	(149.672)	(69.314)	(78.576)
urban	-19.120	-31.562	-17.771	-19.596	-22.042
	(17.498)	(29.340)	(36.155)	(24.900)	(26.586)
mprim	9.668	12.719	17.577	9.785	10.384
	(11.889)	(14.153)	(21.954)	(12.962)	(13.267)
msec	34.954+	32.322	24.785	34.854+	34.336+
	(20.106)	(22.005)	(27.654)	(20.229)	(20.478)
mthird	11.594	-41.631	115.467	9.562	-0.903
	(38.115)	(99.945)	(156.775)	(78.774)	(87.257)
health	-9.314	-18.619	0.558	-9.669	-11.498
	(17.760)	(26.018)	(27.844)	(21.578)	(22.827)
school	10.851	5.932	28.583	10.663	9.696
	(11.524)	(15.622)	(23.309)	(13.193)	(13.852)
cdied	1.373	-14.487	22.702	0.767	-2.351
	(16.065)	(30.019)	(38.702)	(25.321)	(27.246)
cons	178.153**	154.719	449.143	244.882+	226.452
	(16.668)	(171.814)	(283.600)	(130.707)	(148.357)
N	1567	1567	1278	1567	1567
instrument:	-	1st = girl	1st & 2nd = girl	% girl	1st = girl & % girl

Source: Author's calculations based on the UNHS 2005–2006 data.

Notes: OLS = Ordinary least square; 2SLS = Two-stage least squares; LIML = Limited information maximum likelihood. Huber-White standard errors in parentheses, +, * and ** denote significance at the 10, 5 and 1 percent level respectively.

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