Son Preference, Fertility Decline and the Non-Missing Girls of Turkey

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Abstract

Couples in Turkey exhibit son preference through son-biased differential stopping behavior that does not cause a sex ratio imbalance in the population. Demand for sons leads to lower (higher) ratios of boys to girls in large (small) families. Girls are born earlier than their male siblings. Son-biased fertility behavior is persistent in response to decline in fertility over time and across households with parents from different backgrounds. Parents use contraceptive methods to halt fertility following a male birth. The sibling sex composition is associated with gender disparities in health. Among children who were born in the third parity or later, female infant mortality is 1.5 percentage points lower if the previous sibling is male. The female survival advantage, however, disappears if the previous sibling is female. Having an older female sibling shifts the gender gap in infant mortality rate by 2 percentage points in favor of males.

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1 Introduction

"A manly man shall have a son, a manly one."

Turkish proverb

In human populations where there is no prenatal intervention, the ratio of males to females at birth tends to be constant (Hesketh and Xing, 2006). Moreover, if parents have no gender preference, the sex of children within a family is expected to follow a binomial distribution. There is, however, a substantial body of literature showing that parents with a son preference skew the sex composition of their children via gender discrimination in relative care and fertility stopping rules.

The case of "missing women," a phenomenon brought to the public's attention by Sen (1990), leads to a substantial deficit of girls in the population due to sex-selective abortion and excess female mortality. Every year, two million girls worldwide under the age of five are estimated to be missing. Of these, 70 percent were never born (World Bank, 2011). The implications of persistent, abnormally high sex ratios in South Asia and elsewhere have been studied extensively.[1]

Differential stopping behavior (henceforth DSB), on the other hand, implies that parents with a preference for sons would continue to bear children until they reach a desired number of boys (Basu and De Jong, 2010). Without prenatal manipulation, DSB alone does not alter the population sex ratio or the sex ratio across birth parities at the aggregate level. But, assuming that parents can have a finite number of children,[2] then as a result of DSB, females have a greater number of siblings and are born in relatively earlier parities than their

[1] For recent discussions, Chung and Gupta (2007) and Edlund and Lee (2013)) for South Korea, Qian (2008) for China, Jayachandran (2014) for India, and Guilmoto and Duthé (2013) for Armenia, Azerbaijan, and Georgia.

[2] See Yamaguchi (1989) for the implications with no upper bound on the number of children.

male siblings.

In this paper, I focus on family composition in Turkey, a patriarchal society with strong son preference and Muslim identity but without any history of surplus males in the population. I provide strong evidence of male-biased, differential stopping fertility behavior in the absence of prenatal sex selection. By using population data and birth statistics, I show that the long-term trend of sex ratio at birth hovers around the natural level in Turkey and that, like in most parts of the world, child mortality is slightly higher for males than for females. Family-level data show that the sex ratios are also balanced across birth parities. As predicted by the DSB model, however, the sex ratio at last birth is highly skewed in favor of males and males are more likely to grow up in smaller families.

Next, I exploit the first child's sex outcome —a purely random process in the absence of prenatal sex selection— to identify the causal effects of son preference on fertility behaviors. Parents have fewer children if the first child is male than if the first child is female. The number of children in families with first-born daughters is, on average, 6.7 percent larger than families with first-born sons. I show that contraceptive use is the only mechanism through which couples halt fertility after a male birth. Quantile regression results indicate that despite the lower fertility predicted by more schooling, higher age at first birth, and urbanization along with other characteristics, the strong response to the absence of sons is persistent.

Sibling sex composition is associated with significant health disparities between boys and girls. I argue that parents are more likely to proceed to the next parity after a female birth and favor sons in health investment if the older sibling is female. Among children who were born in the third parity or later, while the overall infant mortality[3] rate is higher for males than females, the female survival advantage disappears if the previous sibling is female. Girls with an older male sibling are 1.5 percentage points more likely to survive up to age one, whereas the gender difference in mortality completely vanishes among those with an older female sibling.

The results of this study suggest that DSB causes important early-life disparities through allocative preference in favor of sons among families who are seeking a boy. Importantly, DSB is common in countries that are geographically close and culturally similar to Turkey, notably in Central Asia and North Africa (Filmer et al., 2009; Yount et al., 2000; Basu and De Jong, 2010). Thus, the results documented here have the potential to inform health policy not just in Turkey, but in other countries as well.

2 A Simple Model

Consider a simple illustration of DSB with a three-period model in which there are N couples, each of which has a target of one son, and the maximum number of children per couple is three. Assuming that, without prenatal sex selection, sex distribution at birth is binomial with equal probabilities, $\frac{N}{2}$ couples will have a boy as their first child and the other $\frac{N}{2}$ will have a girl. Those who bear a first-born son would discontinue childbearing, as their target has been met. As a result, $\frac{N}{2}$ families will end up having a family composition of a single boy (B). The remaining $\frac{N}{2}$ couples will have a second child, of which $\frac{N}{4}$ will have a first-born girl and a second-born boy (GB). At this point, these families will also stop childbearing since they have reached their target. The remaining $\frac{N}{4}$ families will have a third child, of which $\frac{N}{8}$ will end up with a first-born girl, a second-born girl, and a third-born boy (GGB)

[3]Death of a child under one year of age.

while $\frac{N}{8}$ will end up with a first-born girl, a second-born girl, and a third-born girl (GGG). In this hypothetical society, the family composition will be as follows: $\frac{N}{2}$ families will have B, $\frac{N}{4}$ GB, $\frac{N}{8}$ GGB and another $\frac{N}{8}$ GGG.

The theoretical implications of such a stopping rule on family composition is shown in Table 1. First and foremost, the population sex ratio is balanced. There are equal numbers of males and females born in the population, $\frac{7N}{8}$ children of each sex. The sex ratio is also perfectly balanced within birth parities. There are $\frac{N}{2}$ males and females born in the first parity, $\frac{N}{4}$ in the second, and $\frac{N}{8}$ in the third. DSB, however, changes the sex composition within families. For example, as shown in Table 1, single child families are composed exclusively of males. The number of males and females are equal in families with two children, and the sex ratio is 0.20 in families with three children. Accordingly, the sex ratio at last birth (henceforth SRLB) is highly male-skewed. In families with one or two children, the last birth is always male. The SRLB is only balanced among families with 3 children, a mechanical result of the three-children maximum.

Basu and De Jong (2010) provide the simulated effects of DSB on family composition with different combinations of maximum and desired numbers of boys, and show similar results. Seidl (1995) and Jensen (2003) each use slightly different models but their implications are also similar: the desire for boys leads to lower (higher) ratios of boys to girls in large (small) families, the SRLB is male-skewed, and girls are born earlier than their male siblings.

3 Relevant Literature on Stopping Rules

There is a rich empirical literature that documents the effects of DSB in countries where son preference has historically been strong. Among others, Park (1983) and Park and Cho (1995) show that the sex ratio of siblings in small families in Korea is skewed in favor of boys and sex at the last birth is highly correlated with the decision of having an additional child. In India, where smaller families have a higher proportion of boys, son targeting is especially pronounced in rural areas and exhibits substantial regional variation (Clark, 2000; Basu and De Jong, 2010). The same patterns have been observed in Vietnam (Pham et al., 2012). In these countries, stopping behavior interacts with the common practice of sex-selective abortion. In China and India, Ebenstein (2007) demonstrates that the gender imbalance is almost exclusively concentrated among couples who are seeking a boy. That is, women continue conceiving until they bear sons, but an excess number of girls conceived in between are missing. Hesketh and Xing (2006) show that in 1992 in South Korea, at the peak of the gender discrimination at birth, the sex ratios were 1.13, 1.96, 2.29 for the second, third and fourth birth parity, respectively.

Fertility stopping rules are also prevalent in countries with balanced sex ratios. Filmer et al. (2009) show strong DSB patterns revealed in Central Asia. In rural Menoufia, Egypt, Yount et al. (2000) find that son-biased family planning translates into fewer births among families with living sons. Basu and De Jong (2010) confirm this finding at the country level in Egypt. In a more striking study, Dahl and Moretti (2008) show that in the United States, parents with at least two children follow a son-biased fertility stopping rule to reach a son when the previous children are all girls. The same study shows qualitatively similar fertility behaviors in Mexico, Colombia, and Kenya.

4 Data and Descriptive Analysis

4.1 Data

The aggregate data come from several different sources. Population sex ratios are calculated from the Population Censuses and Address Based Population Registration System (henceforth ABPRS), a register-based census that collects demographic data based on the place of usual residence. Both sources of data are provided by the Turkish Statistical Institute (henceforth, TurkStat) and include the entire population. Population estimates by sex and 5-year age groups are available in the 1985, 1990 and 2000 Population Censuses while ABPRS provides population estimates for the period from 2008 to 2013 on an annual basis. In addition, TurkStat provides yearly birth statistics collected by the Central Population Administrative System (MERNIS) from 2001 to 2013. The data include all of the births in Turkey that were registered with each district population office.

Household-level analysis is based on the 1993, 1998, 2003 and 2008 waves of the Turkish Demographic and Health Survey (TDHS). This is a nationally representative survey, and the pooled data contain 28,151 married or previously married women, aged 15 to 49, including their complete fertility histories, family planning prevalence and demographic characteristics. The analysis sample includes 25,600 women, all of whom have given birth at least once.

4.2 Population Sex Ratios

In an effort to document the sex ratio trends at birth and among children under age five, I calculate the number of boys per girl for each year that the data are available. Figure 1 shows the estimated sex ratios from 1985 to 2013, with the y-axis scaled to the commonly accepted natural sex ratio range at birth (1.02-1.08 boys per girl). The population sex ratios are strikingly balanced over the last 28 years in Turkey. The sex ratio for children under five years old varies between 1.05 and 1.065. Correspondingly, birth statistics follow a similar trend. From 2001 to 2013, for every female, 1.055 to 1.057 males were born. In comparison, from 1962 to 1980 in 24 European countries, the aggregate ratio of male to female births was between 1.05 and 1.07 (Coale, 1991).

Figure 1 includes the sex ratios at birth from each survey year in TDHS as well. In order to investigate the differential gender mortality, I also calculated the sex ratios for those who survived until the age of five. Overall, TDHS does a good job of replicating the sex ratios calculated from the censuses. The point estimates are not statistically different from the population sex ratios. The consistency of reported sex ratios in TDHS relative to the population data speaks to the accuracy of reporting in the survey. Importantly, the sex ratio under 5 is below the sex ratio at birth for each survey year indicating a lower male-female ratio for the survivors. Like in most countries, this is a natural result of higher child mortality for boys compared to girls.[4] In the pooled TDHS data, 83 out of every 1,000 females die before the age of 5 compared to 90 males indicating a significant difference in child mortality.

Altogether, the aggregate data show no evidence of sex-selective abortion or excess female infant mortality for the study period during which abortion was legal for up to 10 weeks of gestation.[5]

4.3 Family Sex Ratios

To explore the role of DSB in sibling sex composition, I use family-level data from the TDHS and start by disaggregating the sex ratio analysis by sibship size. Sibship size refers to the

[4] Females are less likely to die from infections and respiratory ailments due to their stronger immune system (Drevenstedt et al., 2008).

[5] The abortion law was passed in 1983 and remains with slight modifications up to the present time.

number of children who are alive[6] and sex ratio is the average number of boys per girl within a family. In the presence of a son-biased stopping rule, parents tend to halt fertility after a male birth. Therefore sex ratios should be biased in favor of boys in small families and gradually decrease with the number of siblings.

The TDHS spans a period in which Turkey witnessed both a leap in economic development and a dramatic decline in fertility.[7] The decline in fertility coupled with rapid economic development might have changed both the gender preference and the ability to satisfy such preference. In the interest of capturing the time trend in fertility choices, the results are reported separately for each survey year.

Table 2 shows the sex ratios by total number of living children for each survey year. As predicted by the DSB hypothesis, males are more likely to be in single-child or two-children families. Despite the consistent decrease in average family size from 1993 to 2008, the sex ratio imbalance conditional on number of children remains persistent. For example, the pooled estimates show that, on average, there are 1.2 boys per girl in families with fewer than three children (column 5, upper panel). The sex ratio is 1.11 in three-children families and still in favor of boys, although to a lesser extent. On the other hand, families with more than three children are dominated by females: the ratio of boys to girls plunges to 0.92 in families with five or more children. The female surplus in large families brings down the sex ratio to 1.04 at the aggregate level. The overall sex ratio is balanced for each survey year as well. Strikingly, skewed sex ratios, conditional on sibship size, are similar in different survey

[6] 83.5 percent of deaths in the sample occurred within the first year after birth, hence sex ratios for children who are alive seem to be accurate approximations of the actual sibling sex composition.

[7] The annual average GDP per capita growth was around 2.71% between 1993 and 2008, which corresponds with an increase in real GDP per capita from 5435 to 7730 in constant 2005 U.S. dollars. The World Bank estimates that the total fertility rate declined from 2.8 births per woman in 1993, to 2.1 in 2008, corresponding to a 25 percent decline in the total fertility rate (http://data.worldbank.org - last accessed July 7, 2015).

years, showing a consistency in male-biased reproductive behavior between 1993 and 2008 (columns 1-4, upper panel).

In the lower panel of Table 2, the sample is restricted to women aged 35 to 49 in order to observe the sex ratios among the couples who have most likely finished childbearing. The sex ratio imbalance is even greater in nearly-completed families. In small families (number of children ≤ 3), the average sex ratio is 1.21, and falls to 0.94 among those with more than three children (column 5, lower panel).

The sex ratio at last birth (SRLB) is another measure to test the presence of son-targeting fertility behavior. If parents are more likely to cease childbearing after a male birth, the SRLB should be male-skewed. Table 3 shows the average sex ratios by total number of births and birth order, with the SRLB depicted in bold. The upper panel contains calculations for the full sample and the lower panel is restricted to women aged 35 to 49. In both panels, independent of the mother's birth history, the last birth is consistently male-skewed, families seek boys at all birth parities. In the upper panel, the number of males per female is slightly above 1.20 in the last birth parity, even among very large families. For example, the SRLB among couples with six births is 1.23, while the same families' earlier parities are highly female-skewed. This may indicate either an unusually strong persistence in seeking a boy, or the "gambler's fallacy" in son targeting. If parents believe that the sex of the next child is contingent on the existing sibling sex composition, they are less likely to stop childbearing after a girl than couples who are aware of the fact that each child's sex is an independent event.

The lower panel of Table 3 shows that son preference is revealed more strongly among nearly-completed families. Families with three children or fewer exhibit abnormal sex ratios in favor of boys at all parities (columns 1-3, lower panel), whereas, for those with more than three children, only the SRLB is male-skewed. Earlier birth parities in large families are highly female-skewed since couples continue childbearing after a female birth. For example, in families with four births, the sex ratio ranges between 0.90 and 0.94 in the first three parities, while the SRLB is 1.31, a clear indicator of families stopping once they reach a son (row 4, lower panel).

As a robustness check for the prenatal sex selection in higher birth parities, I conducted several tests. The sex ratio for second-born children conditional on a first-born daughter is 1.04, and not skewed. The sex ratio for third-born children after two females is 1.02. Lastly, without conditioning on the sex composition of previous births, the sex ratios in the second-, third-, and fourth parities are 1.05, 1.02, and 1.05, respectively. In countries with prenatal sex selection, the likelihood of sex-selective abortion is substantially higher if the children are all females, and the sex ratios become more imbalanced in higher birth parities. This is not the case in Turkey, where parents apply male-biased stopping rules but do not practice sex-selective abortion.

To summarize, DSB is the only mechanism by which couples in Turkey pursue son preference, and prenatal sex selection is not a common practice. As documented in the existing literature, the skewed sex ratio distribution conditional on family size is persistent despite the economic development and fertility decline. The next section offers an empirical strategy to identify the changes in fertility behavior that have led to the patterns shown above.

5 Empirical Analysis

5.1 Identification Strategy

Without prenatal manipulation, the sex of the first-born child is a random draw. Parents with a son preference, however, respond differently to this exogenous shock. This makes it possible to exploit the first child's sex as a source of an exogenous variation in order to identify the causal effects of son preference on several fertility decisions. The reduced form equation in this context is:

$$y_{irt} = \alpha + X_i'\Gamma + \tau Z_{irt} + \theta_r + \delta_t + \omega_{rt} + u_{irt} \tag{1}$$

where y_{irt} is the fertility outcome (number of pregnancies, number of children born, number of children alive and indicators for current contraceptive use and having any pregnancy termination in the past)[8] for mother i, who is living in region r, and was interviewed in survey year t. Z is an indicator of a first-born female, X is a vector of family background covariates (mother's age, age at first birth, years of education, ethnicity, husband's age and years of education, rural residence, co-residence of husband's parents, and dummies for whether the marriage was arranged and husband's family or husband paid bride price), θ_r and δ_t control for survey-year and region fixed effects while ω_{rt} captures the region specific year effects. Importantly, adjusting for these control variables in equation (1) does not affect the estimated parameter τ given that Z is random. It does, however, improve precision.[9]

The parameter τ reflects the effect of a first-born daughter compared to a first-born son on couple's fertility decisions. As mentioned earlier, male infant mortality is higher due to

[8] In the survey, pregnancy termination is defined as having a miscarriage, abortion, or still birth.

[9] Online Appendix Table 1 documents the results from the OLS regressions with and without adjustment for covariates.

purely biological reasons; therefore Z might affect y through both differential mortality rate for males and son targeting. For example, a woman will be more likely to have another pregnancy if the first child dies and the mortality risk is higher among male children. In order to isolate the effect of son preference from the male differential mortality, equation (1) controls for the survival status of the first child. The regression sample is restricted to the women with a singleton first birth who represent 99.1 percent of the total sample. Although these adjustments make no statistical difference in the estimation results, they avoid a potential confusion in the interpretation of τ.

For causal inference, the error term in equation (1) should be uncorrelated with Z. This is a major concern in countries with abnormal sex ratios at birth because the child's sex is a prenatal choice due to the common practice of sex-selective abortion. In such cases, children's sex is likely to be correlated with unobserved family characteristics. There is no fully robust test to validate the exogeneity assumption, yet comparing the families with first-born sons and first-born daughters helps. Observed family characteristics can altogether explain more than 50 percent of the variation in sibship size. Thus, despite not being perfect, the comparison is highly informative regarding the validity of the random assignment assumption. As a further examination, I estimate the following regression:

$$Z_{irt} = \gamma + X_i'\Phi + \theta_r + \delta_t + \epsilon_{irt} \tag{2}$$

using a logit model and report the joint χ^2-test result for the null hypothesis that all the estimated coefficients in the right-hand side of equation (2) are jointly equal to zero.

When estimating equation (1), I use OLS as well as the Poisson likelihood function when the response variable is a count, i.e. number of pregnancies, number of children born, and number of children alive. There are two reasons to go beyond the standard linear model: first, the functional form in the Poisson model ensures a positive predicted value for each family, second, Poisson estimates show the percentage change in sibship size induced by a female birth, an alternative indicator that shows the change in fertility level with respect to the baseline fertility preference. The effect of a first-born female on family size depends on couples' competency at fertility control, hence the deviation from the baseline fertility level might be a better indicator when comparing families with different backgrounds since it takes into account the overall family planning behavior.

5.2 Estimation Results on Fertility

I present the family background characteristics by first child's sex in Table 4. There are no statistically significant differences between families with first-born sons and those with first-born daughters with regards to any of the sample characteristics. The p-value for the overall χ^2-statistic from the regression in equation (2) is 0.53 with an extremely low pseudo-R^2. Strictly speaking, the coefficient vectors Φ, δ_t and θ_r in equation (2) are jointly equal to zero.[10] Given the large sample size, the data strongly support the assumption that the sex outcome of the first child is not manipulated. Additionally, the overall sex ratio is balanced among higher parities independent of the first-child's sex. The sex ratio of subsequent siblings is 1.04, both in families with first-born sons, and those with first-born daughters.

DSB sharply affects the average number of siblings. In Table 5, the pooled sample OLS estimates show that women with first-born daughters have about 0.20 more pregnancies, 0.19 more births and 0.18 more living children than women with first-born sons (columns 1-3, panel A). The maximum likelihood estimate from the Poisson model reveals that this corresponds to a 6.7 percent increase in number of living children (column 3, panel A).

[10] See Online Appendix Table 2 for the full set of individual coefficients.

Results in panels (B) through (D) are based on separate regressions for each age group. The estimated DSB effects on family size are small for the youngest cohort, and similarly large for the older age categories. The increase in sibship size induced by a first-born female for the youngest mother cohort aged 15 to 29 is 0.06 children, or 3.4 percent (column 3, panel B). The estimated family size effects are much higher for the older cohorts. If the first child is female, women aged 30 to 39 have around 0.25 more children, corresponding to an 8.3 percent increase in number of living children (column 3, panel C). The results are quantitatively similar for the oldest cohort (column 3, panel D). The discrepancy of the estimates between young and old cohorts is due to the fact that some of the young women have not had, and are still pursuing, a son. The change in contraceptive use behavior among young couples confirms this argument. Women aged 15 to 29 with first-born daughters are 2.6 percentage points less likely to use either a traditional or modern contraceptive method than those with first-born sons (column 4, panel A).[11] The difference is weaker in older cohorts (column 4, panels C and D).

Irrespective of age category, the probability of pregnancy termination is unrelated to the first child's sex, suggesting that families do not use abortion for reaching the desired sex composition (column 5, panels A-D). Nevertheless, the results must be interpreted with caution because pregnancy termination is self-reported and the survey question does not allow to identify whether or not it was a health-related procedure[12]. Underreporting of abortion cases would bias the estimated coefficient towards zero.

[11]Traditional methods include coitus interruptus, periodic abstinence, and vaginal douche while modern methods include the pill, injections, female or male condom, intrauterine device, and sterilization.

[12]Specifically, the survey question asked: "has the respondent had ever had a pregnancy that was terminated by a miscarriage, abortion, or still birth, i.e., did not result in a live birth".

The change in fertility behaviors induced by the first-child's sex reveals two important findings: first, son preference has a sizeable impact on family size through DSB, second, women are more likely to use contraceptive methods when the first-born child is male. In other words, contraceptives are used as a tool for stopping fertility after a son.

5.3 Heteregenous Effects on Fertility

Pooled-sample estimates might mask heterogeneous effects on families with different backgrounds. A common way to reveal treatment heterogeneity is to interact the treatment indicator, in this case the first-born female indicator, with family characteristics. The results from the interaction effects are included in the Online Appendix Tables 3-6. Overall, the effect of a first-born female on number of living children is similar across survey years, suggesting that son targeting endures despite the decline in fertility over time. A first-born female significantly increases the sibship size for all the subgroups, categorized by parents' education level, type of marriage, or residential status. With the exception of educated women, the percentage changes in number of children are statistically indistinguishable. The relative family size effect among women with secondary or higher education is significantly smaller.

A recent alternative to interaction effects, borrowed from the clinical literature, is to use a set of covariates to predict outcomes among the untreated group. The regression coefficients from the sample of untreated group are then used to predict outcomes for the full sample. After stratifying the predicted values into quantiles, the treatment effects are estimated within each quantile. This procedure thus creates an index of predicted outcomes by using all the relevant covariates instead of interacting each one with the treatment dummy. Abadie et al. (2013), however, show that the OLS estimator is severely biased in finite samples due to overfitting and recommend using either leave-one-out (LOO) or repeated split sample (RSS) estimators. The leave-one-out estimator avoids overfitting simply by excluding each observation when estimating the coefficients used to calculate its own predicted value. Alternatively, the repeated-split sample estimator divides prediction sample into two groups and uses only one of them for prediction. When this is repeated many times and averaged over the number of repetitions, the small sample bias vanishes.[13] .

I use the families with first-born sons as the "control" group and use endogenous determinants of fertility level (mother's age at first birth, father' and mother's years of education, region and rural residency) to predict number of siblings. Duflo (2012) notes that decrease in fertility and increase in age at first birth are highly correlated with higher income and education. Urbanization and migration from agricultural to industrial regions are also associated with economic growth and prosperity. Note that this prediction step simply involves dividing the sample into quantiles and is not concerned with causality. The key assumption for the causal identification is that within each fertility quantile, the sex of the first child is random.

Table 6 reports both adjusted and unadjusted differences for each fertility quintile using LOO and RSS algorithms[14]. Unadjusted differences are simple differences in the average number of children among families with first-born females and first-born males for the corresponding quintile (columns 1 and 3). As before, adjusted differences control for the full set of covariates (columns 2 and 4). The similarity of the unadjusted and adjusted results speaks to the exogeneity of Z_i, and the type of estimator used does not make a statistical difference in the estimated quantile treatment effects.

[13]See Abadie et al. (2013) for the detailed description of the methodology.

[14]Jeremy Ferwerda provides a Stata routine at https://ideas.repec.org/c/boc/bocode/s457801.html - Last accessed July 7, 2015.

Stratification reveals high variation in number of children across fertility quintiles. At the lowest predicted fertility quintile, families with first-born sons bear on average 1.69 children compared to 4.41 children for the highest quintile (column 5, $\hat{\tau}_1$ and $\hat{\tau}_5$). The number of additional children induced by a female first birth also declines in response to lower fertility, but the relative change is strongest at the median level (column 6, $\hat{\tau}_3$). If the first child is female, number of children increases by 0.077 children (4.6 percent) among women in the lowest predicted fertility quintile (columns 4 and 6, $\hat{\tau}_1$). The change in number of siblings is 0.23 children (9.4 percent) at the middle quintile and 0.27 children (6 percent) for the highest quintile. Considering the significant family size differences between predicted fertility quintiles, DSB shows a relatively flat response to decline in fertility. Son preference is significantly prevalent at each fertility level, even among families with 1.69 children, and causes similar changes in the number of siblings. In other words, the lower fertility predicted by better education, more income and urbanization neither eliminate nor drastically change the son-biased fertility preference.

5.4 Health Effects on Children

In addition to changing fertility behavior and causing differences in sibling sex composition, DSB might also give rise to health disparities between boys and girls. Rosenblum (2013) develops an economic model in which sibling sex composition leads to a differential allocation of family resources among boys and girls. In this framework, sons provide a future economic gain to parents while daughters come with a future economic burden. The economic gain from an extra son is larger if the existing proportion of sons is relatively small in the family; therefore, the smaller the proportion of boys, the greater the incentive for households to favor boys in health investment.

I use a difference-in-differences approach similar to Rosenblum (2013) to test the sibling sex composition hypothesis. In the absence of prenatal sex selection, the sex of the child is random at any birth parity. If the previous sibling is a girl, however, families have an incentive to invest more in boys in the next parity. Therefore, in the case of male-biased allocative preference, the gender difference in health should lead to a relative male advantage if the previous sibling is female.

Table 7 compares the observable characteristics of parents by each child's sex for each of the first four birth parities. Family backgrounds of first-, second-, and third-born boys are identical to first-, second-, and third-born girls, respectively. Mother's age and age at first birth, which are expected to be correlated, are somewhat statistically higher for mothers who have a fourth-born boy than mothers who have a fourth-born girl, but the differences are very small. Like earlier, I use a logit model to test if all the differences in the means reported in the table are jointly equal to zero. The p-values for the joint χ^2-tests are indicated under the observable characteristics for each birth order. None of the p-values indicate a significant difference between families of boys and girls.

The exogenous variation in children's sex, however, causes significant changes in the sibling sex composition and family size. Families who had a female in the previous parity are more likely to have another birth and on average have a higher fertility. For example, mothers who have a second-born female have 0.16 more births and are 4.8 percentage points more likely to have additional children, a clear indication of stopping after a male birth (panel 2). Consistent with the previous findings, the differential stopping behavior is clear across all birth parities.

In an effort to investigate how the previous sibling's sex changes the gender health gap in the next parity, I use several different versions of the following difference-in-differences estimator:

$$y_i = \mu_0 + \mu_1 Z_{i1} + \mu_2 Z_{i2} + \mu_3 (Z_{i1} \times Z_{i2}) + \eta_i \tag{3}$$

where y_i is the early-life health outcome (infant mortality and nutrition) and Z_{i1} is a female indicator for child i while Z_{2i} is a dummy variable and equals one if the previous sibling is female. For example, when the sample is restricted to second-born children and the outcome is the infant mortality, μ_1 shows the girl-boy difference in infant mortality if the first child is male whereas $\mu_1 + \mu_3$ shows the same difference if the first child is female. Thus μ_3 is the difference-in-differences estimator and is expected to be positive if a previous female sibling causes the boys to be more valuable and to shift the infant mortality gender gap in favor of males.

I begin with presenting the regression results for the second-born children. This provides the most generalizable estimates because first, the sex of the first-born child is random and second, most of the families in Turkey have at least two children. Although the sex of children is random at any parity, parents who proceed to the next parity after a female birth might be different from parents who proceed to the next parity after a female birth. Therefore, restricting the sample to second-born children attempts to address the self-selection of families into higher birth orders since families typically have at least two children. In other words, Z_{i1} and Z_{i2} in equation (3) are both plausibly exogenous.

The difference-in-differences estimator, nevertheless, is still informative in regards to resource allocation by gender, even for higher birth orders. For example, assume that parents who have a second-born girl are identical to parents who have a second-born boy but those who have a third child after a second-born girl are wealthier than those who have a third child after a second-born boy. In this case, independent of their gender, children who have an older male sibling would have poorer health than children who have an older female sibling due to wealth differences between parents. To put this another way, Z_{1i} in equation (3) is still purely exogenous while Z_{2i} is not. However, in this example, μ_2 captures the wealth differences triggered by the previous sibling's sex whereas the interaction term μ_3 shows the gender differential effect of having an older female sibling. If only boys are better off by having an older female sibling, which is an evidence of a treatment heterogeneity, this might be an indicator of a male-biased allocative preference. It is important to note that the results for children in higher birth orders are less generaralizable since the sample is restricted to high fertility households. But they reflect potentially important behavioral responses to the gender composition of a family.

In TDHS, the retrospective birth history includes mortality information covering all births by the same mother. The nutrition outcomes are only available for children under the age of five. Anthropometric measurements are constructed by taking the height and weight of children and a child's immunization is self-reported if an official immunization record is missing. Following the definitions of the World Health Organization, I create two dichotomous outcome variables that reflect the child's nutritional status: stunting, which refers to being less than two standard deviations below the age- and gender-normalized median height for the reference population, and being underweight, which refers to being less than two standard deviations below the age- and gender-normalized median weight for the reference population.

Table 8 presents the regression results for the second-born children. Panel (1) shows the infant mortality rates for the second-born children by the first- and the second-born siblings' sex compositions. Independent of the first-child's sex, second-born girls have a lower mortality rate than second-born boys, albeit the estimated difference is insignificant in all cases. Overall, there is no indication of improvement in male mortality compared to female mortality after a female birth. These findings hold for the nutrition outcomes as well. Although the estimated probabilities of stunting and being underweight are slightly higher for girls with an older female sibling, the difference-in-differences estimators are not statistically significant. The regression results in Table 8 are robust to covariate adjustment[15].

Gender disparities in health, however, emerge among children who were born in the third-parity and later. Panel (1) in Table 9 shows the mean infant mortality rates for boys and girls who were born in the third parity or later. Female infant mortality is significantly lower than the male infant mortality by 1.5 percentage points if the previous sibling is male. The biological female advantage, however, disappears if the previous sibling is female. The difference-in-differences estimate shows a statistically significant 2 percentage points shift in female-male mortality gap induced by having an older female sibling. The point estimates in the lower panel are identical after controlling for birth order and other covariates. The nutrition estimates suggest a similar pattern but the statistical inference is weaker due to the small sample sizes. In all the regressions, standard errors are clustered by mother in order to capture any correlations in the health outcomes of siblings.

Importantly, Table 9 shows that the difference-in-differences estimator is mostly driven by the improvement in male mortality. Infant mortality rates are similar among girls who have an older male sibling and who have an older female sibling. Male infant mortality rate is, however, 1.4 percentage points lower for males who have an older female sibling compared to males who have an older male sibling. The difference is highly significant in both adjusted

[15] Adjusted regressions control for the year of survey, region, year of survey and region interactions, mother's age, age at first birth, years of education, ethnicity, rural residence, husband's age and years of education, patrilocal residence, whether the marriage was arranged and husband paid a bride price plus indicator variables for missing husband's age, husband's years of education, arranged marriage and bride price payment.

and unadjusted regressions[16].

I further investigate the gender differences in immunization outcomes for BCG (Bacillus Calmette-Guerin), DPT (diphtheria, pertussis, tetanus), Polio, and MMR (measles-mumps-rubella) vaccinations. Overall, independent of the previous sibling's sex, the differences in vaccination rates between males and females are small and not statistically significant[17]. This is not surprising, since immunization rates are high in Turkey: child vaccination is free of charge, and is part of routine procedure in public hospitals.

6 Discussion

In Turkey, the trend in the sex ratio at birth fluctuates around the commonly accepted natural sex ratio and there is no evidence or documented history of sex-selective abortion. On the other hand, couples exhibit strong son preference through family planning and are more likely to halt fertility after a male birth. My analysis reveals that contraceptive use after a male birth is a contributing factor to an abnormal sex ratio distribution conditional on sibship size. I provide additional evidence that abortion is not a common practice for reaching the desired sex composition.

Still, the demand for sons is persistent in response to decline in fertility over time and across households with parents from different backgrounds. These findings are consistent with Yount et al. (2000), who show that the dramatic increase in modern contraceptive use in Egypt from the 1980s to the early 1990s resulted in a decline in fertility but had no

[16] The estimated coefficients for $\hat{\mu}_2$ from equation (3) are not shown, but available upon request from the author.

[17] See Online Appendix Table 7 for the results.

impact on son preference. My findings further suggest that families that exceed their ideal family size as a result of a previous female birth favor males in health investment in the next parity. Similar to the findings in India documented in Rosenblum (2013), boys are better off if they have older female siblings. In other words, the empirical evidence suggests that families who are seeking a boy show allocative preference in favor of them once they are born.

The significant changes in family structures and health discrepancies reported here raise the question of what would happen if parents did not persistently seek sons. One prediction is that girls and boys would have similar a number of siblings, which would improve equality in intra-household resource allocation (Becker and Lewis, 1974). Lee (2007) provides empirical evidence on quantity-quality trade-off in Korea, by showing that the exogenous increase in family size caused by a first-born female decreases the investment in education for each sibling. Another prediction is that boys would be less likely to be born in the last parity. Black et al. (2005) show that the negative impact of the higher birth order is largest for last born children. One possibility therefore, is that the absence of son-biased fertility behavior could further improve gender equality at birth, though by favoring males. More importantly, girls in households with a high proportion of girls would be better off since the gender discrimination in health investment emerges predominantly in these large households.

References

Abadie, A., M. M. Chingos, and M. R. West (2013): "Endogenous stratification in randomized experiments," *NBER Working Paper No. 19742*.

Basu, D. and R. De Jong (2010): "Son targeting fertility behavior: some consequences and determinants," *Demography*, 47, 521–536.

Becker, G. S. and H. G. Lewis (1974): "Interaction between quantity and quality of children," in *Economics of the family: marriage, children, and human capital*, UMI, 81–90.

Black, S. E., P. J. Devereux, and K. G. Salvanes (2005): "The more the merrier? The effect of family size and birth order on children's education," *The Quarterly Journal of Economics*, 669–700.

Chung, W. and M. D. Gupta (2007): "The decline of son preference in South Korea: the roles of development and public policy," *Population and Development Review*, 33, 757–783.

Clark, S. (2000): "Son preference and sex composition of children: evidence from India," *Demography*, 37, 95–108.

Coale, A. J. (1991): "Excess female mortality and the balance of the sexes in the population: an estimate of the number of "missing females"," *The Population and Development Review*, 517–523.

Dahl, G. B. and E. Moretti (2008): "The demand for sons," *The Review of Economic Studies*, 75, 1085–1120.

Drevenstedt, G. L., E. M. Crimmins, S. Vasunilashorn, and C. E. Finch (2008): "The rise and fall of excess male infant mortality," *Proceedings of the National Academy of Sciences*, 105, 5016–5021.

Duflo, E. (2012): "Women Empowerment and Economic Development," *Journal of Economic Literature*, 50, 1051–79.

Ebenstein, A. (2007): "Fertility choices and sex selection in Asia: analysis and policy," *Mimeo, Princeton University.*

Edlund, L. and C. Lee (2013): "Son preference, sex selection and economic development: the case of South Korea," *NBER Working Paper No. 18679.*

Filmer, D., J. Friedman, and N. Schady (2009): "Development, modernization, and childbearing: The role of family sex composition," *The World Bank Economic Review*, lhp009.

Guilmoto, C. and G. Duthé (2013): "Masculinization of birth in Eastern Europe," *Population and Societies*, 1–4.

Hesketh, T. and Z. W. Xing (2006): "Abnormal sex ratios in human populations: causes and consequences," *Proceedings of the National Academy of Sciences*, 103, 13271–13275.

Jayachandran, S. (2014): "Fertility decline and missing women," *NBER Working Paper No. 20272.*

Jensen, R. T. (2003): "Equal treatment, unequal outcomes? Generating sex inequality through fertility behaviour," *Mimeo, Harvard University.*

Lee, J. (2007): "Sibling size and investment in childrens education: an Asian instrument," *Journal of Population Economics*, 21, 855–875.

Park, C. B. (1983): "Preference for sons, family size, and sex ratio: an empirical study in Korea," *Demography*, 20, 333–352.

Park, C. B. and N.-H. Cho (1995): "Consequences of son preference in a low-fertility society: imbalance of the sex ratio at birth in Korea," *Population and Development Review*, 59–84.

Pham, B. N., T. Adair, P. S. Hill, and C. Rao (2012): "The impact of the stopping rule on sex ratio of last births in Vietnam," *Journal of Biosocial Science*, 44, 181–196.

Qian, N. (2008): "Missing women and the price of tea in China: the effect of sex-specific earnings on sex imbalance," *The Quarterly Journal of Economics*, 123, 1251–1285.

Rosenblum, D. (2013): "The effect of fertility decisions on excess female mortality in India," *Journal of Population Economics*, 26, 147–180.

Seidl, C. (1995): "The desire for a son is the father of many daughters," *Journal of Population Economics*, 8, 185–203.

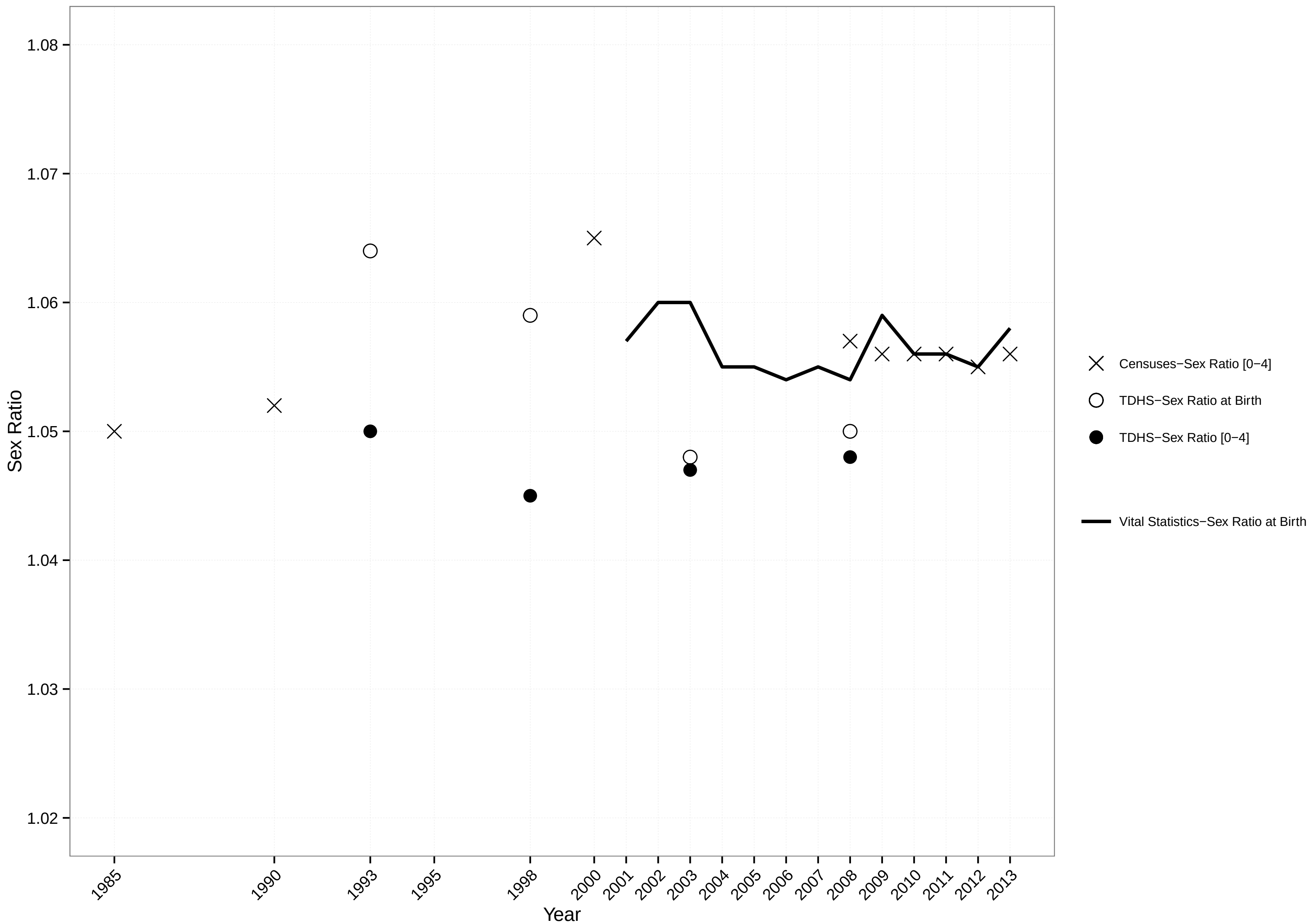
Sen, A. (1990): "More than 100 million women are missing," *The New York Review of Books*.

World Bank (2011): *World development report 2012: gender equality and development*, World Bank Publications.

Yamaguchi, K. (1989): "A formal theory for male-preferring stopping rules of childbearing: sex differences in birth order and in the number of siblings," *Demography*, 26, 451–465.

Yount, K. M., R. Langsten, and K. Hill (2000): "The effect of gender preference on contraceptive use and fertility in rural Egypt," *Studies in Family Planning*, 31, 290–300.

Figure 1. Sex Ratio Trends



Note: Table shows the estimated sex ratios at birth and under age five from different data sources. Census estimates show the sex ratio under age 5 and are gathered from Population Censuses (1985, 1990, and 2000) and Address Based Population Registration System (2008-2013). Vital statistics show the sex ratio at birth and are gathered from Central Population Administrative System (2001-2013). Turkish Demographic and Health Survey (TDHS) estimates are from author's calculations, with *y*-axis scaled to the commonly accepted natural sex ratio range at birth (1.02-1.08 boys per girl), and *x*-axis labels are only shown for the years that the data were available.

Table 1. Implications of a Simple Son-Biased Differential Stopping Rule on Sibling Sex Composition

	Sibling Sex Composition				
Birth Parity	B	GB	GGB	GGG	Sex Ratio
First	$\frac{N}{2}$ Boys	$\frac{N}{4}$ Girls	$\frac{N}{8}$ Girls	$\frac{N}{8}$ Girls	1.00
Second		$\frac{N}{4}$ Boys	$\frac{N}{8}$ Girls	$\frac{N}{8}$ Girls	1.00
Third			$\frac{N}{8}$ Boys	$\frac{N}{8}$ Girls	1.00
Family Size	1 Child	2 Children	3 Children		
Sex Ratio	Only male	1.00	0.20		
Sex Ratio at Last Birth	Only male	Only male	1.00		
Aggregate Number of Children	Boys			Girls	
	$\frac{7N}{8}$		=	$\frac{7N}{8}$	

Note: This table shows the sibling sex composition from a three-period model, in which there are N couples, each of which has a target of one son and the maximum number of children per couples is three. In the upper panel, each cell reports the number of children born by sex, birth parity, and sibling sex composition. The last column shows the sex ratio by birth parity. The middle panel reports the overall sex ratio and the sex ratio at last birth by family size. The lower panel shows the aggregate number of males and females. Sex ratio refers to the number of males per female. "B" indicates a single boy, "GB" indicates a first-born girl and a second-born boy, "GGB" indicates a first-born girl, a second-born girl, and a third-born boy, while "GGG" indicates three girls.

Table 2. Sex Ratios by Total Number of Living Children and Year of Survey

Women Aged 15 to 49

	Survey Year				
Total Number of Living Children	1993 (1)	1998 (2)	2003 (3)	2008 (4)	Pooled (5)
1	1.17	1.16	1.22	1.22	1.20
2	1.21	1.19	1.21	1.21	1.20
3	1.13	1.09	1.07	1.14	1.11
4	0.97	1.01	0.91	0.90	0.94
5+	0.93	0.92	0.94	0.89	0.92
Overall	1.05	1.04	1.04	1.05	1.04
Total Number of Children					
Born	3.34	3.19	3.05	2.92	3.11
Still Alive	2.94	2.87	2.80	2.72	2.83
N	5923	5578	7360	6739	25600

Women Aged 35-49 (Pooled Sample)

Family Size	Sex Ratio	*N*	Family Size	Percentage	Sex Ratio
1	1.24	857	Small ($n \leq 3$)	60.8%	1.21
2	1.31	3506			
3	1.15	3049			
4	0.96	1913	Large ($n > 3$)	39.2%	0.94
5+	0.93	2859			
Overall	1.04	12184		100%	1.04

Note: In the upper panel, each cell shows the sex ratio of siblings by total number of living children in the family and survey year for women aged 15 to 49. The average number of children born per family and the average number of children alive are reported by survey year at the bottom of the upper panel. In the lower panel, column (1) shows the sex ratio by family size for the pooled sample of women aged 35 to 49. Column (4) reports the percentage of small ($n \leq 3$) and large ($n > 3$) families for the same age group, where n indicates the number of living children in the family. Column (5) reports the sex ratio of siblings within small ($n \leq 3$) and large ($n > 3$) families for women aged 35 to 49. The populationl sex ratio is reported at the bottom of the lower panel. Sex ratio refers to average number of males per female. Family size 5+ indicates families with 5 children or more. Sample includes women with at least one birth history and sample sizes are shown with N.

Table 3. Sex Ratios by Birth Order

Women Age 15-49 (Pooled Sample)

Number of Births	Birth Order 1	2	3	4	5	6	7
1	**1.21**						
2	1.19	**1.19**					
3	1.08	1.06	**1.26**				
4	0.92	0.89	0.93	**1.20**			
5	0.98	0.89	0.94	1.07	**1.23**		
6	0.84	0.98	0.89	0.91	1.00	**1.23**	
7+	0.97	0.94	0.81	0.90	0.94	0.98	0.98
Average birth order		*Boys* = 2.75		*Girls* = 2.78			

Women Age 35-49 (Pooled Sample)

Number of Births	Birth Order 1	2	3	4	5	6	7
1	**1.24**						
2	1.30	**1.31**					
3	1.11	1.07	**1.27**				
4	0.94	0.90	0.92	**1.31**			
5	0.99	0.89	1.00	1.11	**1.21**		
6	0.85	1.04	0.87	0.92	1.03	**1.25**	
7+	0.98	0.93	0.84	0.90	0.94	0.99	0.98
Average birth order		*Boys* = 3.18		*Girls* = 3.20			

Note: In both panels, each cell shows the sex ratio by birth order and total number of births. The sex ratio at last birth (SRLB) is depicted in bold. The average birth order by sex is reported at the bottom of each panel. In the the upper panel, the sample includes 79,674 births from women aged 15-49. In the lower panel, the sample includes 48,340 births from women aged 35 to 49. Sex ratio refers to the average number of males per female, and 7+ indicates the children with a birth order 7 or more.

Table 4. Baseline Characteristics of Families by First Child's Sex

Family Characteristics	First child's sex			*t*-test	
	Boy	Girl	Difference	*p*-value	*N*
Mother					
Age	34.07	34.13	-0.053	0.61	25366
Age at first birth	20.66	20.59	0.067	0.17	25366
Years of education	4.93	4.99	-0.062	0.19	25366
Non-Turkish	0.20	0.19	0.005	0.32	25366
Residential					
West	0.27	0.27	0.002	0.76	25366
South	0.16	0.16	-0.003	0.48	25366
Central	0.20	0.20	0.001	0.82	25366
North	0.13	0.13	0.004	0.31	25366
East	0.23	0.23	-0.004	0.45	25366
Rural	0.30	0.30	0.003	0.61	25366
Husband					
Age	38.61	38.72	-0.115	0.33	23140
Years of education	7.02	7.07	-0.047	0.33	25269
Patrilocal residence	0.12	0.12	-0.005	0.21	25366
Marriage					
Arranged by families	0.61	0.61	0.005	0.44	25355
Paid bride price	0.23	0.24	-0.005	0.38	24956
			p-value, joint χ^2-test = 0.53		
			N=25366 pseudo-R^2=0.0006		

Note: This tables compares the families with first-born sons and first-born daughters. The first column reports the indicated covariate mean for families with first-born sons, the second column reports the indicated covariate mean for families with first-born daughters, the third column reports the difference between the first and the second columns, the fourth column shows the *p*-values, which are based on a two-sample *t*-test of difference in means assuming equal variances. The last column shows the number of non-missing observations for each covariate. At the bottom, the *p*-value from the joint χ^2-test is shown. The joint χ^2-test is based on a logit regression of first child's sex (equals 0 if a boy and 1 if a girl) on all variables in the table, survey year dummies plus indicator variables for missing husband's age, husband's years of education, arranged marriage, and bride price payment. The null hypothesis is that all slope coefficients are jointly equal to zero. Regression sample size and pseudo-R^2 are shown at the bottom.

Table 5. Effect of First Child's Sex on Parents' Fertility Behavior

		Outcomes				
		Number of Pregnancies (1)	Number of Births (2)	Number of Living Children (3)	Contraceptive Use (4)	Pregnancy Termination (5)
(A) Age 15-49	$\hat{\tau}^{OLS}$	0.204***	0.189***	0.184***	-0.016***	-0.001
		(0.023)	(0.017)	(0.015)	(0.005)	(0.005)
	$\hat{\tau}^{MLE}$	0.053***	0.062***	0.067***		
		(0.006)	(0.005)	(0.005)		
	$\bar{y}\|Z_i = 0$	3.82	3.02	2.73	0.70	0.26
	N			25366		
(B) Age 15-29	$\hat{\tau}^{OLS}$	0.087***	0.058***	0.061***	-0.026***	-0.001
		(0.022)	(0.016)	(0.015)	(0.010)	(0.007)
	$\hat{\tau}^{MLE}$	0.039***	0.031***	0.034***		
		(0.010)	(0.008)	(0.008)		
	$\bar{y}\|Z_i = 0$	2.29	1.93	1.82	0.70	0.12
	N			8301		
(C) Age 30-39	$\hat{\tau}^{OLS}$	0.254***	0.267***	0.250***	-0.015*	-0.012
		(0.036)	(0.026)	(0.023)	(0.008)	(0.009)
	$\hat{\tau}^{MLE}$	0.060***	0.080***	0.083***		
		(0.008)	(0.008)	(0.007)		
	$\bar{y}\|Z_i = 0$	3.96	3.11	2.85	0.78	0.29
	N			9657		
(D) Age 40-49	$\hat{\tau}^{OLS}$	0.273***	0.234***	0.233***	-0.008	0.016
		(0.056)	(0.040)	(0.034)	(0.010)	(0.011)
	$\hat{\tau}^{MLE}$	0.053***	0.059***	0.065***		
		(0.010)	(0.009)	(0.009)		
	$\bar{y}\|Z_i = 0$	5.37	4.13	3.60	0.58	0.37
	N			7408		

Note: Each column shows the effect of a first-born female on the number of pregnancies, number of births, number of living children, current contraceptive use (includes withdrawal, periodic abstinence, vaginal douche, the pill, injections, female or male condom, intrauterine device, or sterilization), and any pregnancy termination in the past (includes miscarriages, abortions or still births). In each of the panels (A) through (D), for women in the indicated age group, $\hat{\tau}^{OLS}$ shows the OLS estimate and $\hat{\tau}^{MLE}$ shows the maximum likelihood estimate assuming a Poisson process. Mean outcomes for families with first-born males are shown with $\bar{y}|Z_i = 0$. N refers to the number of observations in the regression. All regressions control for the first born's survival, year of survey, region, year of survey and region interactions, mother's age, age at first birth, years of education, ethnicity, rural residence, husband's age and years of education, patrilocal residence, whether the marriage was arranged and husband paid a bride price plus indicator variables for missing husband's age, husband's years of education, arranged marriage and bride price payment. Heteroskedasticity-consistent standard errors are in parentheses. Significance levels are indicated by * $<$ .10, ** $<$ .05, *** $<$ .01.

Table 6. Endogenous Stratification Results on the Number of Living Children

	Repeated Split Sample		*Leave-One-Out*				
Quantile	Unadjusted	Adjusted	Unadjusted	Adjusted	$\bar{y}_k \vert Z_i = 0$	%Δ	N_k
	(1)	(2)	(3)	(4)	(5)	(6)	(7)
$\hat{\tau}_1$	0.096***	0.076**	0.095***	0.077***	1.69	0.046	5073
	(0.021)	(0.019)	(0.021)	(0.019)			
$\hat{\tau}_2$	0.152***	0.144***	0.137***	0.128***	2.12	0.060	5067
	(0.027)	(0.022)	(0.029)	(0.024)			
$\hat{\tau}_3$	0.234***	0.218***	0.256***	0.229***	2.44	0.094	5081
	(0.034)	(0.026)	(0.039)	(0.028)			
$\hat{\tau}_4$	0.213***	0.215***	0.209***	0.219***	2.99	0.073	5073
	(0.046)	(0.031)	(0.047)	(0.035)			
$\hat{\tau}_5$	0.283***	0.259***	0.295***	0.265***	4.41	0.060	5072
	(0.071)	(0.044)	(0.071)	(0.044)			

Note: This table shows the effects of a first-born daughter on the number of living children for each of the predicted fertility quantiles. The outcome is the number of living children in the family. Columns (1)-(4) show the treatment effects for each fertility quantile, $\hat{\tau}_k$, where $k = \{1,2,3,4,5\}$. Columns (1) and (2) are estimated with the repeated split sample estimator. Columns (3) and (4) are estimated with the leave-one-out estimator. Both estimation methods are provided in Abadie et al. (2014). Column (5) shows the mean number of children for families with a first born male, indicated with $\bar{y}_k|Z_i = 0$ for each fertility quantile. Column (6) shows the percentage change (%Δ) in family size induced by a first-born female and calculated by dividing the treatment effect in column (4) by the mean number of children in column (5). Variables that are used to predict the fertility quantiles are the mother's age at first birth, mother's and father's years of education, rural residence, and region. Adjusted regressions control for the firstborn's survival, year of survey, mother's age, age at first birth, education level, ethnicity, region, rural residence, husband's age, husband's education level, patrilocal residence, whether the marriage was arranged and husband paid a bride price plus indicator variables for missing husband's age, husband's education, arranged marriage and bride price payment. The number of repeated split sample repetitions is 100. Bootstrapped standard errors are in parentheses. Significance levels are indicated by * $<$ .10, ** $<$ .05, *** $<$ .01.

Table 7. Family Characteristics, Number of Births and the Sibling Sex Composition by Birth Parity and Child's Sex

	(1)			(2)			(3)			(4)		
Family	First-child's sex			Second-child's sex			Third-child's sex			Fourth-child's sex		
Characteristics	Boy	Girl	Diff.	Boy	Girl	Diff.	Boy	Girl	Diff.	Boy	Girl	Diff.
Mother												
Age	34.07	34.13	-0.053	35.82	35.74	0.081	37.39	37.52	-0.128	38.85	39.49	0.357**
Age at first birth	20.66	20.59	0.067	20.10	20.14	-0.043	19.36	19.38	-0.017	18.90	18.74	0.161**
Years of education	4.93	4.99	-0.062	4.46	4.40	0.066	3.39	3.31	0.079	2.45	2.47	-0.015
Non-Turkish	0.20	0.19	0.005	0.21	0.21	-0.005	0.27	0.27	-0.006	0.35	0.36	-0.008
Husband												
Age	38.61	38.72	-0.115	40.34	40.23	0.111	42.04	42.08	-0.043	45.53	43.27	0.253
Years of education	7.02	7.07	-0.047	6.72	6.66	0.062	5.89	5.91	-0.021	5.21	5.17	0.043
Residence												
Rural	0.30	0.30	0.003	0.32	0.32	-0.001	0.36	0.37	-0.009	0.40	0.41	-0.010
Patrilocal	0.12	0.12	-0.005	0.09	0.10	-0.006	0.09	0.08	0.008*	0.07	0.07	0.006
Marriage												
Arranged by families	0.61	0.61	0.005	0.65	0.65	-0.003	0.71	0.72	-0.003	0.75	0.75	0.001
Paid a bride price	0.23	0.24	-0.005	0.26	0.27	-0.001	0.36	0.35	0.005	0.44	0.46	-0.019*
p-value, joint χ^2-test	0.57			0.92			0.12			0.32		
Differential Stopping												
Total Number of births	3.02	3.20	-0.181***	3.54	3.70	-0.160***	4.47	4.74	-0.269***	5.56	5.74	-0.179***
Boys (%)	0.67	0.35		0.65	0.37		0.61	0.40		0.58	0.40	
Girls (%)	0.33	0.65		0.35	0.63		0.39	0.60		0.42	0.60	
P (Having More Children)	0.79	0.81	-0.020***	0.60	0.65	-0.048***	0.56	0.65	-0.081***	0.59	0.65	-0.053***
Number of observations	25,366			20,397			12,701			7,676		

Note: This table compares the children in families by sex of the child at each parity. Panel (1) compares the families with first-born sons and first born-daughters. Panel (2) compares the families with second-born sons and second-born daughters. Panel (3) compares the families with third-born sons and third-born daughters. Panel (4) compares the families with fourth-born sons and fourth-born daughters. The reported differences are from a two-sample *t*-test of difference in means assuming equal variances. At the bottom of family characteristics, the *p*-value from the joint χ^2-test is shown. The joint χ^2-test is based on a logit regression of the child's sex at each birth parity (equals 0 if a boy and 1 if a girl) on all variables in the table, and survey year dummies plus indicator variables for missing husband's age, husband's years of education, arranged marriage and bride price payment. The null hypothesis is that all slope coefficients are jointly equal to zero. Significance levels are indicated by * $<$ .10, ** $<$ .05, *** $<$ .01.

Table 8. The Effects of the First-born Sibling's Sex on the Second-born Child

	(1) Infant Mortality				(2) Stunting				(3) Underweight			
	First-born				First-born				First-born			
	Boy		Girl		Boy		Girl		Boy		Girl	
	Second-born		Second-Born		Second-born		Second-born		Second-born		Second-born	
	Boy	Girl	Boy	Girl	Boy	Girl	Boy	Girl	Boy	Girl	Boy	Girl
Infant Mortality												
Mean	0.069	0.066	0.067	0.062	0.126	0.131	0.118	0.130	0.049	0.049	0.038	0.056
Standard Deviation	[0.25]	[0.25]	[0.25]	[0.24]	[0.33]	[0.34]	[0.32]	[0.34]	[0.22]	[0.22]	[0.19]	[0.23]
Girl-Boy difference	-0.003		-0.005		0.006		0.013		-0.000		0.018*	
	(0.005)		(0.005)		(0.016)		(0.016)		(0.009)		(0.011)	
Difference-in-differences	-0.002				0.007				0.018			
	(0.007)				(0.023)				(0.015)			
Covariate Adjusted Difference-in-differences	-0.004				0.005				0.017			
	(0.007)				(0.022)				(0.014)			
N	20,397				3,399				3,399			

Note: This table compares the health outcomes of the second-born children by first-born sibling's sex. Panel (1) compares the infant mortality rates, Panel (2) compares the probability of stunting, and Panel (3) compares the probability of being underweight. Infant mortality is defined as the death of a child under the age of one. Stunting is defined as being less than two standard deviations below the age- and gender-normalized median height for the reference population. Being underweight is defined as being less than two standard deviations below the age- and gender-normalized median weight for the reference population. Height and weight regressions only include children under the age of five at the time of the interview. Girl-boy difference estimator shows the gender difference in infant mortality by previous sibling's sex. Difference-in-difference estimator shows the difference in girl-boy differences between children who has a previous female sibling and children who has a previous male sibling. The covariate adjusted results are from the regressions that control for the year of survey, region, year of survey and region interactions, mother's age, age at first birth, years of education, ethnicity, rural residence, husband's age and years of education, patrilocal residence, whether the marriage was arranged and bride's family received a bride price plus indicator variables for missing husband's age, husband's years of education, arranged marriage and bride price payment. Heteroskedasticity-consistent standard errors are in parentheses. Significance levels are indicated by * $<$ .10, ** $<$ .05, *** $<$ .01.

Table 9. The Effect of Previous Sibling's Sex on Next Parity

Pooled sample estimates ($n \geq 3$)	(1) Infant Mortality Birth order, $n-1$				(2) Stunting Birth order, $n-1$				(3) Underweight Birth order, $n-1$			
	Boy		Girl		Boy		Girl		Boy		Girl	
	Birth order, n		Birth order, n		Birth order, n		Birth order, n		Birth order, n		Birth order, n	
	Boy	Girl	Boy	Girl	Boy	Girl	Boy	Girl	Boy	Girl	Boy	Girl
Outcome												
Mean	0.094	0.080	0.080	0.085	0.216	0.231	0.201	0.229	0.089	0.090	0.087	0.103
Standard Deviation	[0.29]	[0.27]	[0.27]	[0.28]	[0.41]	[0.42]	[0.40]	[0.42]	[0.29]	[0.29]	[0.28]	[0.30]
Girl-Boy difference	-0.015***		0.005		0.015		0.028*		0.001		0.016	
	(0.005)		(0.004)		(0.017)		(0.016)		(0.012)		(0.011)	
Difference-in-differences	0.020***				0.013				0.015			
	(0.006)				(0.023)				(0.016)			
Covariate-adjusted Difference-in-differences	0.020***				0.018				0.022			
	(0.006)				(0.022)				(0.016)			
N	33,039				5,064				5,064			

Note: This table compares the infant mortality rates, probability of stunting and being underweight between boys and girls by previous sibling's sex. Regression samples are restricted to children who were born in the third birth parity or later. Panel (1) compares the infant mortality rates, Panel (2) compares the probability of stunting, and Panel (3) compares the probability of being underweight. Infant mortality is defined as the death of a child under the age of one. Stunting is defined as being less than two standard deviations below the age- and gender-normalized median height for the reference population. Being underweight is defined as being less than two standard deviations below the age- and gender-normalized median weight for the reference population. Height and weight regression only include children under the age of five at the time of the interview. Girl-boy difference estimator shows the gender difference in health outcomes by previous sibling's sex. Difference-in-difference estimator shows the difference in girl-boy differences between children who has a previous female sibling and children who has a previous male sibling. The lower panel shows the same results from the regressions that control for the child's birth order, year of survey, region, year of survey and region interactions, mother's age, age at first birth, years of education, ethnicity, rural residence, husband's age and years of education, patrilocal residence, whether the marriage was arranged and husbad paid a bride price plus indicator variables for missing husband's age, husband's years of education, arranged marriage and bride price payment. Standard errors are in parentheses and clustered by mother. Significance levels are indicated by * $<$ .10, ** $<$ .05, *** $<$ .01.