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

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Abstract

Son preference is usually revealed by both gender discrimination in relative care and son-targeting fertility stopping rules. This article shows that couples in Turkey exhibit strong son preference without causing a gender imbalance in the population. Estimation results reveal that a first-born daughter increases the average sibship size by 6.6 percent through male-biased differential stopping fertility behavior. Contraceptive use is the primary tool to halt fertility following a male birth among young couples. Families with a highly educated mother are much less likely to seek sons, while father's education has no association with the degree of son preference. The differential demand for sons is persistent despite economic development and decline in fertility predicted by more schooling, higher age at first birth and urbanization along with other endogenous determinants. The relationship between degree of son preference and fertility follows an inverse U-shaped path reaching a peak at the medium fertility level.

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1 Introduction

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

Turkish proverb

In the absence of manipulation, both the sex ratio at birth and the population sex ratio are remarkably constant in human populations (Hesketh and Xing, 2006). Moreover, the gender of children within a family is expected to follow a binomial distribution if parents do not have a sex preference. However, there is a substantial body of literature suggesting that in some regions of the world, parents may skew the sex composition of their children via gender discrimination in relative care and son targeting fertility stopping rules.

The former, a phenomenon brought to public attention by Sen (1990) as the case of "missing women" leads to a substantial deficit of girls in the population due to sex selective abortion and excess female mortality. World Bank (2011) estimates around 2 million girls under the age of five are missing every year and of these, 70 percent were never born. The implications of persistent, abnormal high sex ratios in South Asia[1] and elsewhere[2] have been studied extensively.

The latter, differential stopping behavior (DSB), implies that parents with a preference for sons would continue to bear children until they reach a desired number of boys, given the household resource constraints (Basu and De Jong, 2010). In the absence of sex selective abortion or infanticide, DSB does not alter the population sex ratio. However, it does affect sibling sex composition within families and causes important gender disparities even when no sex bias in intra-family resource allocation exists. Assuming that the maximum number

[1] For recent discussions in this research area, see for example Chung and Gupta (2007) and Edlund and Lee (2013)) for South Korea, Qian (2008) for China and Jayachandran (2014a) for India.

[2] See Guilmoto and Duthé (2013) for Armenia, Azerbaijan, and Georgia.

of children parents can have is finite,[3] DSB will result in females having a greater number of siblings and being born earlier than boys in relatively large families.

A simple illustration of DSB involves a two-period fertility decision model in which every couple has a target of having one son, with a maximum of two children per family. Assuming that sex distribution at birth is binomial with equal probabilities, half of the couples will have a boy as their first child and the other half will have a first-born girl. Those who bear a first-born son would discontinue childbearing, as their target has been met. First-child lottery "losers" will go on to have another child and again, half of them will bear a second-born boy and the others will bear another girl. In this hypothetical society, all the girls are members of two-children families while average sibship size for boys is 4/3. Additionally, in families with mixed-gender children, all the girls are first born and all the boys are second born.

Basu and De Jong (2010) provide the simulated effects of DSB on family composition with different combinations of maximum sibship size and desired number of boys. Seidl (1995) and Jensen (2003) use a slightly different model but the implications of these models are also similar: the desire for boys leads to lower (higher) ratio of boys to girls in large (small) families. As mentioned above, stopping rules skew neither the sex ratio nor the birth order at the population level. In the example above, the total number of girls equals the total number of boys, and the average birth order for each gender is 4/3 in the population.

The implications of DSB have been demonstrated empirically in countries where son preference has historically been strong. Park (1983) and Park and Cho (1995) found that the sex ratio of siblings in small families in Korea is skewed in favor of boys and gender at the

[3] Yamaguchi (1989) for the implications with no limit on the maximum number of children.

last birth is highly correlated with the decision of having an additional child. Clark (2000) found that smaller Indian families have a higher proportion of boys while Basu and De Jong (2010) indicate that son targeting is higher in rural India and exhibits regional variation. Similar findings are reported for Vietnam (Pham et al., 2012).

While these patterns are striking, establishing causality from DSB to family size and sibling sex composition is challenging because the relationship is confounded with the effects of sex selective abortion and discrimination in care practices for girls. In China and India, Ebenstein (2007) demonstrates that the sex ratio among the first child is around the natural rate (approximately 106 boys for 100 girls). However, sex ratios become increasingly imbalanced with the birth order and the gender imbalance is almost exclusively concentrated among couples who are seeking a boy (Ebenstein, 2007). That is, women continue conceiving until they bear sons, but an excess number of girls conceived in between are missing. Similarly, Ebenstein (2010) shows that in regions of China where the One Child Policy is controlled more strictly the gender imbalance increases because son preference manifests in an interaction of missing girls and DSB. Empirically, it is difficult, if not impossible, to isolate DSB from differential treatment of male and female fetuses during pregnancy.

In this paper, I focus on family composition in Turkey — a patriarchal society with strong son preference and Muslim identity but without large cohorts of surplus males. The study is the first to provide causal evidence on strong male-biased, differential stopping fertility behavior in the absence of imbalanced sex ratios at birth or differential female mortality. By using the Population Censuses and birth statistics from The Central Population Administrative System (MERNIS), I provide evidence that the long-term trend of sex ratio at birth hovers around the natural level in Turkey and that child mortality is slightly higher for males than for females, like in most parts of the world. However, family-level data from the Demographic and Health Survey show that the sex ratio at last birth is highly skewed in favor of males and that males are more likely to grow up in relatively smaller families due to DSB. A flip-of-the-coin approach in son targeting leads to a “gambler’s fallacy” behavior in fertility: parents who had girls in earlier parities are more likely to continue childbearing as if they are bound to have a son in the next parity. The sex ratio at last birth is highly male skewed even in very large families. Consequently, girls are more likely to be the older children in the family.

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 DSB on family structure and fertility behaviors. If the first child is a male, parents are more likely to stop childbearing than if the first child is a female. The number of children in families who have a first-born daughter is 6.6 percent greater compared to families with a first-born son. Contraceptive use is the primary mechanism through which young couples halt the fertility after a male birth and the extent of son preference varies little by family background. Educated mothers appear to be an exception and exhibit weaker son preference while father’s education is unrelated to the degree of son preference. Turkish data do not support the cultural explanations that have been proposed in previous literature.

Borrowing from the randomized experiment literature, I use endogenous stratification to investigate the relationship between son preference and family size. The results indicate that the son preference is persistent in response to the fertility decline predicted by more schooling, higher age at first birth and urbanization along with other characteristics. Similar to countries with missing women, favoring sons through DSB is relentless despite economic development.

There are several candidate explanations that could justify son preference without missing women. First, dowry is a common source of income received by the bride's family therefore the marriage puts a strain on the finances of the families with male children. As a result, daughters are not perceived as a future financial burden at birth. Second, Islam explicitly forbids all types of infanticide[4] and does not tolerate abortion as a family planning method. Infanticide was a recognized practice in the Arabian peninsula before the rise of Islam; it was denounced and totally rejected by the Prophet Muhammad (Giladi, 1990). In addition, the majority of Muslim scholars agree that a necessary abortion performed for a health related reason escapes being labeled a sin and is a religiously neutral action (Bowen, 1997). Third, considering the middle income level and the universal health care system, families in Turkey are much less likely to face a severe crisis that would require sacrificing the welfare of girls in favor of boys, as shown in Rose (1999).

These results are important from a methodological perspective. This paper shows that the sibling sex composition is a family-level decision and therefore aggregate indicators might not fully reflect the degree of son preference. Couples can reach the desired sex composition through sex selective abortion, infanticide or contraceptive use and gender-neutral abortion following a male birth. While all reflect the same fertility preference, only the "missing" girls can be empirically observed at the aggregate level. Therefore, when investigating the roots of son preference, comparing sex ratios across countries or regions might not gauge the true variation in male-preferring fertility behavior. As shown in the remainder of this paper, Turkey offers a vivid example of such a case.

[4]See for example Qur'an Surah Al-Isra, Chapter 17, Verse 31 (17:31).

2 Data and Desctiptive Analysis

2.1 Data

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 The Turkish Statistical Institute (henceforth, TurkStat) and cover the entire population. Population estimates by gender and 5-year age groups are available in the 1985, 1990 and 2000 Population Censuses while ABPRS provides population estimates for the period 2008-2013 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. Registration is mandatory within one month after birth and parents who fail to comply must pay punitive fines.[5] A national identity card is required for access to the universal health care system in Turkey, which forms another important incentive for birth registration.

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 of 28,151 ever-married women, aged 15-49, including their complete fertility histories, family planning prevalence and demographic characteristics.

[5]http://www.nvi.gov.tr (last accessed October 27, 2014.)

2.2 Population Sex Ratios

In an effort to document the sex ratio trends at birth and among children aged under five, I calculate the number of boys per girl for each year that the data are available. In Figure 1, census data are illustrated by black geometric shapes. Sex ratios indicate a strikingly consistent gender balance over the last 28 years in Turkey. The sex ratio for children under 5 years old varies between 1.05-1.065 from 1985 to 2013. Correspondingly, birth statistics follow a similar trend. The black line reveals that 1.055-1.057 males were born for every female between 2001 and 2013. In comparison, for aggregate births from 1962 to 1980 in 24 countries in Europe, the ratio of the number of male to female births ranged from 1.05 to 1.07; the median ratio was 1.059 (Coale, 1991). Note that sex ratio estimates for children come from four different data sources thus have higher variance and y-axis is scaled to the commonly accepted natural sex ratio range at birth (1.02-1.08 boys per girl).

In Figure 1, red points indicate the reported sex ratios at birth from each survey year in TDHS. In order to investigate the differential gender mortality, I also report the sex ratio of those who survived until the age of five. These estimates are shown with blue points. 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 lack of misreporting in the survey. Additionally, the blue points are below the red points 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. Because females have more vigorous immune responses and greater resistance to infection, female infants have lower mortality from infections and respiratory ailments (Drevenstedt et al., 2008).

Altogether, the data show no evidence of sex-selective abortion or excess female infant mortality for the study period during which abortion up to 10 weeks of gestation was legal and funded by the government[6].

2.3 Family Sex Ratios

To explore the role of DSB in sibling composition, I use the family-level data and start with disaggregating the sex ratio analysis by sibship size. Sibship size refers to the number of children who are alive[7] 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 the fertility after a male birth. Therefore sex ratios should be biased in favor of boys in small families and gradually increase with the number of siblings.

In total, the sample includes 25,600 women who have at least one birth record. TDHS span a period in which Turkey has witnessed a leap in economic development[8] and a dramatic decline in fertility[9]. This has potentially important implications on the sex composition. The decline in fertility and economic development might have changed both the gender preference and the ability to satisfy such preference. Accordingly, the results are reported separately for each survey year in the interest of capturing the time trend in fertility choices.

Table 1 shows the first set of descriptive results. As predicted by the DSB hypothesis,

[6] The abortion law was passed in 1983 and preserved with slight modifications up to the present time. In 2004, the requirement for spousal consent was rescinded.

[7] 83.5 percent of deaths in the sample occurred within the first year after birth hence sex ratios for children who are alive seem accurate approximations of the actual sibling sex composition.

[8] Annual average GDP per capita growth was around 2.71% between 1993-2008, which corresponds an increase in real GDP per capita from 5435 to 7730 in constant 2005 U.S. dollars (http://data.worldbank.org - last accessed October 27, 2014.)

[9] World Bank estimates that total fertility rate declined from 2.8 births per woman in 1993 to 2.1 in 2008 corresponding to a 25 percent decline in total fertility rate (http://data.worldbank.org - last accessed October 27, 2014.).

males are more likely to be in a single child or two children families. Despite the consistent decrease in average family size from 1993 to 2008, gender imbalance in small families remains persistent. In the last column of Table 1, the pooled estimates show that on average, there are 1.2 boys per girl in families with less than 3 children. The sex ratio is still in favor of boys in three children families although to a lesser extent. Notwithstanding, families with more than three children are dominated by females: the ratio of boys to girls plunges to 0.92 in families with 5 or more children in the pooled sample. The female surplus in large families brings down the sex ratio to a natural rate ($\approx$1.05) at the aggregate level. The overall sex ratio is balanced for each survey year as well. Strikingly, skewed sex ratios are similar in different survey years, showing a consistency in male-biased reproductive behavior between 1993 and 2008.

In the lower panel of Table 1, I restrict the sample to women aged 35-49, an arbitrary choice of age interval, to observe the sex ratios among the couples who most likely ended childbearing. The gender imbalance is even greater in nearly completed families. In small families (i.e. number of children $\leq$ 3), the average sex ratio is 1.21 and falls to 0.94 among those with more than three children.

A simple calculation helps to put these results in perspective. As shown in the lower panel of Table 1, roughly 43 percent of the boys grow up in a small family compared with 36 percent of girls. In the absence of skewed sex composition, both should be around 40 percent. This corresponds to a sizeable gender population exchange ($\approx$ 6 percent), i.e., girls who were to be born in small families are replaced by boys who would have been born in large families in the absence of gender preference.

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 2 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 35-49. In both panels, independent of the mother's birth history, the last birth is consistently male-skewed, i.e., families seek boys at all family sizes. In the upper panel, on average, the number of males per female is slightly above 1.2 in the last birth parity, even among very large families. For example, the SRLB among couples with six children 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 gender of the next child is contingent on the existing sibling sex composition, they are less likely to stop childbearing after a girl compared to the case in which families are aware of the fact that each child's gender is an independent event.

The average birth order for boys and girls (last row of the upper panel in Table 2) is very similar without conditioning on family size since the sex of each child is independent of the birth order at the population level. However, girls are typically born in earlier parities and grow up in large families.

As expected, the lower panel of Table 2 shows son preference is revealed more strongly among nearly completed families. Small families with three children or less exhibit abnormal sex ratios in favor of boys at all parities, whereas only SRLB is male-skewed for those with more than three children. Earlier birth parities in large families are highly female-skewed since couples continue childbearing after a female birth.

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, skewed sex ratio distribution conditional on family size is persistent despite economic development and fertility decline.

The underlying interpretations come with two caveats. First, the negative relationship between the sex ratio and family size does not directly imply causal evindence of DSB. Bennett (1983) argues that a woman might stop childbearing not because she is satisfied with her current family composition, but because she fears having a child of the undesired sex. Similarly, proceeding to the next parity does not necessarily imply displeasure with the current sex composition of children but rather reflects a desire for more children, regardless of the sex. The challenge is therefore to measure son preference without conditioning on family size. Second, it is hard to identify the mechanisms through which families skew the sex ratio and to explain the persistence in son targeting with the sex composition tables. The next section offers an empirical strategy to address the causality issue and attempts to shed light on the demographic and cultural determinants of son preference.

3 Empirical Analysis

3.1 Identification Strategy

Without prenatal manipulation, the gender of the first-born child is a random draw. Jensen (2003) is the first study to use the first child's sex as an instrument for sibship size. Lee (2008) follows a similar approach in Korea, while Rosenblum (2013) explains the sex differential in child mortality in India by the first-born child's gender. In the current setting, I use the firstling sex as a source of random variation in order to identify the causal effects of son preference on several fertility outcomes. The reduced form equation in this context is:

$$y_i = \alpha + X_i'\gamma + M_i'\theta + \tau Z_i + D_i'\delta + u_i \tag{1}$$

where y_i is the fertility outcome for mother i (number of pregnancies, number of children ever born, number of children alive and indicators for current contraceptive use and having any pregnancy termination[10] in the past), Z_i is an indicator of a female first-born and X_i is a vector of demographic pre-treatment covariates (mother's age, age at first birth, years of education, ethnicity, as well as husband's age and years of education). While current residential characteristics (region, rural residence and co-residence of husband's parents) are post-treatment covariates, the gender of the first-born is not correlated with any of these factors in the data and controlling for these variables in equation (1) does not affect the estimated causal parameter τ. However, these variables are useful for identifying the heterogeneity in son preference and thus are included in X_i. M_i is a vector of marriage-specific characteristics (whether the marriage was arranged and husband's family or husband paid dowry), D_i indicates a vector of separate intercepts for each survey year and u_i captures the random fluctuations in fertility.

The parameter τ reflects the differential effect of a first-born daughter compared to a first-born son on fertility outcomes. As mentioned earlier, male infant mortality is higher due to purely biological reasons; therefore Z_i might affect y_i 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 women with a singleton first birth who represent 99.1 percent of the total sample. Although these

[10]Miscarriage, abortion or still birth.

adjustments make no statistical difference in the estimation results, they avoid a potential confusion in the interpretation of τ.

For causal inference, the disturbance term u_i in equation (1) should be uncorrelated with Z_i. This is a major concern in countries with abnormal sex ratios at birth because the child's gender is a prenatal choice due to the common practice of sex selective abortion. In such cases, Z_i is not random and is likely to be correlated with unobserved family characteristics. Another potential confounding factor is maternal mortality: since boy preference increases the number of births, it also increases the maternal mortality risk and therefore can change the composition of women observed in the sample in the case of high maternal mortality.

While there is no fully robust test to validate the unconfoundedness assumption, inspecting the observable family characteristics by gender of the first child may help. I examine whether Z_i is related to observed characteristics in X_i and M_i and estimate the following regression:

$$Z_i = \mu_0 + X_i'\mu_1 + M_i'\mu_2 + D_i'\delta + \epsilon_i \tag{2}$$

using a logit model and report the joint χ^2-test result for the null hypothesis $\mu_1 = \mu_2 = 0$. X_i and M_i can altogether explain more than 50 percent of the variation in the family size. Despite not perfect, comparing the fertility related observable characteristics between families with a first-born female and first-born male is informative about the validity of the random assignment assumption.

Equation (1) is estimated using OLS and the Poisson likelihood function when the response variable is a count. 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, comparing the degree of son preference across different demographic groups requires a common scale of inference. For example, more educated women have fewer children than the less educated; hence the estimated effect of Z_i on the conditional mean will tend to be small for this subgroup. However, relative change in the sibship size induced by a female birth is a better indicator since it shows the change in fertility preference with respect to the baseline level. Difference in exponential conditional means yields to this desired semi-elasticity:

$$E[y_i|X_i, M_i, D_i, Z_i = 1] - E[y_i|X_i, M_i, D_i, Z_i = 0] = \tau \ E[y_i|X_i, M_i, D_i, Z_i] \quad (3)$$

where τ measures the relative change in the conditional expected sibship size induced by a first-born daughter. Figure 2 shows the frequency distribution for count response variables along with their sample mean and the variance. A graphical examination of unconditional distributions suggests that a count model might be a good candidate to mimic the data generating process especially for sibship size. The empirical distribution of the number of children per family is right-skewed and exhibits little dispersion between the sample mean ($\hat{\mu} = 2.82$) and the variance ($\delta^2 = 3.10$). Note that the estimated standard errors are heteroscedasticity-consistent in all the regressions and thus control for overdispersion in the case of estimated count models.

3.2 Summary Statistics

I present family background characteristics by first child's gender in Table 3. There are no statistically significant differences between families with a first-born son and daughter on any of the sample characteristics. The p-value for the overall χ^2-statistic from the regression[11]

[11] See Appendix Table 1 for the full set of individual coefficients.

in equation (2) is 0.42 with an extremely low pseudo-R^2. Strictly speaking, the coefficient vectors μ_1 and μ_2 in equation (2) are jointly equal to zero. Given the large sample size, the data strongly support the assumption that the sex outcome of the first child is random and not manipulated.

All the demographic and marriage-specific variables reported in Table 3 have no impact on the estimated parameter τ given that Z_i is uncorrelated with the family background [12], however, they improve precision.

Table 3 also reveals some of the Turkish cultural attitudes towards family structure in which economic calculus has a role. Patrilocal residence remains a strong social preference representing 12 percent of families in the data. From an economic perspective, this likely represents a traditional pension system in which families invest in sons who later provide old-age support. Ebenstein (2014) argues that strong patrilocal norms can explain the male-preferring fertility behavior in East Asia, South Asia and the South Caucasus including the Muslims in Azerbaijan. One may thus expect son preference to be stronger among couples who live with the husband's parents. To put this another way, τ in equation (3) is expected to be higher among couples with patrilocal residence.

Jayachandran (2014b) brings forward an equally convincing argument in her review of gender inequality. The desire to protect female safety and "purity" may contribute to the relative deprivation of women because constrained physical mobility reduces the economic returns of having female children. Female sexual inexperience is also an important aspect of marriage decision for most men in Turkey (Sakalh-Uğurlu and Glick, 2003). Fox (1975)

[12] See Appendix Table 2 for the results from the OLS regressions with and without adjustment for covariates in Table 3.

points out that in Turco-Islamic culture, arranged marriage is a major tool to preserve the purity of women because modern forms of matching constitute *prima facie* evidence of contact with other men before the husband. Data confirm that parents meet parents more often before couples meet each other with 61 percent of the marriages in the sample arranged by parents.

Another interesting feature of marriage arrangements between families is that in almost one quarter of the marriages, the bride's family receives a dowry. As argued by Sambrani et al. (1983), bride-price might either reflect the compensation for bride's contribution to the husband's household income-generation or a price for the bride's "purity". In both cases, it is an indicator of a traditional marriage which is perceived as an economic transaction between couple's families.

3.3 Estimation Results

DSB affects the average sibship size sharply. In Table 4, the pooled sample OLS estimates ($\tau^{\hat{OLS}}$) reported in columns (1)-(3) of panel (A) show that women who had a first-born daughter had about 0.20 more pregnancies, 0.19 more births and 0.18 more living children than women who had a first-born son. In column (3) of panel (A), maximum likelihood estimates ($\tau^{\hat{MLE}}$) from equation (3) show that on average a first-born daughter increases the family size by around 6.6 percent for the full sample.

Results in panels (B)-(D) are based on separate regressions for each age group. The estimated DSB effects on sibship size are small for the youngest cohort and similarly large for the older age categories. Column (3) of panel (B) indicates a 3.4 percent increase in family size induced by a first-born female for the youngest mother cohort aged 15-29. Estimated average effect of a first-born female is around two times higher for the older cohorts. This is due to the fact that some of the young women have not had, and are still pursuing, a son. The results in column (4) confirm this argument. Women aged 15-29 with a first-born daughter are less likely to use either a traditional[13] or modern[14] contraceptive method compared to those with a first-born son. OLS results in column (4) of panel (B) uncover a 2.5 percentage point increase in current contraceptive use following a male birth among the youngest women in the sample. The difference is almost non-existent 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 gender, suggesting that families do not use abortion for reaching the desired sex composition. Nevertheless, the results shown in column (5) must be interpreted with caution because pregnancy termination is self-reported and the survey question does not allow one to identify whether it was a health-related procedure or the couple's independent decision for abortion. The survey question asked whether the respondent ever had a pregnancy that was terminated by a miscarriage, abortion, or still birth, i.e., did not result in a live birth, without asking the specific reason. Underreporting of abortion cases would bias the estimated coefficient towards zero.

Results presented in Table 4 reveal two important findings. First, son preference has a sizeable impact on family size through DSB. Second, it shows that young women are more likely to use contraceptive methods when the first-born child is male; that is, contraceptives are used as a tool for stopping fertility after a son.

[13]Coitus interruptus, periodic abstinence and vaginal douche.

[14]Pill, injections, female or male condom, intrauterine device and sterilization.

3.4 Heterogeneity in Son Preference

Next, I investigate the differential effect of a first-born daughter on family size by interacting Z_i with demographics and marriage-specific variables and using sibship size as the outcome.

Table 5 and Table 6 show the main effect for Z_i plus its interaction with each category of the variable of interest. I report the OLS estimate of τ for each interacted category, the average sibship size for families with a first-born son ($\bar{y}|Z_i = 0$) and the percentage change in sibship size ($\%\Delta$) estimated by the Poisson maximum likelihood model. All specifications control for the full set of covariates as employed in Table 3 except the years of education for which I use a categorical variable for the level of the last school that each parent attended[15]. Finally, I test if the estimated treatment effects differ by the category of interest. In Table 5, the test result is indicated by the p-value for the χ^2-test of joint significance for the interaction terms. In Table 6, all the interacted variables are binary, therefore I report the difference in the estimated effects on family size between two categories and the standard error of the difference.

The first panel in Table 5 shows that the increase in family size associated with a first-born daughter is similar in each survey year. OLS results indicate that in 1993, families with a first-born daughter have 0.16 more children than families with a first-born son, corresponding to a 5.7 percent increase in family size. The average number of children per family decreases from 2.86 in 1993 to 2.62 in 2008, while the effect of a first born daughter remains much the same. The χ^2-test of joint significance indicates that the degree of son preference is not different across survey years.

[15] Results from various classifications of the educational attainment are similar and are available from the author by request.

In Table 5, panels (2) and (3) examine the relationship between son preference and parental education. The number of siblings starkly decreases in more educated families, but son targeting endures. Families are more likely to stop childbearing after a boy at all education levels. DSB, however, manifests differently among educated women: those with a secondary school or higher education are much less likely to seek sons. Meanwhile, education of fathers has no discernible association with the degree of son preference. A first-born daughter increases the family size by 0.06 children or 3.6 percent among women with secondary or higher education, an impact that is less than half of the estimated increase in family size among women with primary school education. The low p-value for the test of joint significance confirms that women with higher education exhibit significantly weaker preference for sons. The percentage change in family size as a result of a first-born daughter is similar for fathers of different educational backgrounds. The difference in family size between families with a first-born female and a first-born male is 4.9 percent if the father is uneducated and 6.7 percent if the father has a secondary or higher education.

One may find it puzzling to observe no improvement in gender-differential fertility behavior over time since economic growth should have also increased women's education. The main underlying cause of persistence in son preference is that the relationship between education and sex-biased fertility practice is not exactly linear: DSB peaks at the primary education level for women which could be explained by needing a minimum knowledge of the efficient contraceptive use.

From 1993 to 2008, the average years of education for women increased from 4.26 to 5.42, but the share of women with only primary education remained around 55 percent due to a substantial decrease in the number of women with no formal education and a similar increase in the proportion of secondary education. Consequently, although women are more educated, the proportion with the strongest DSB remained the same. As this process continues, however, the relationship between economic growth and gender bias in fertility behavior might be different in the future.

If the male-biased fertility behavior is a result of intra-household bargaining, differences in parental education may matter. I also examined whether families with equally-educated parents differ from families with parents of differing educational levels and found no difference.[16] The mother's education alone seems to be the driving factor behind decreasing son preference.

Table 6 investigates other cultural and economic explanations for son preference that were introduced in earlier studies. In panel (1), the estimated treatment effect is slightly higher for families living in a patrilocal residence with respect to the rest but the difference is not statistically significant. Similarly, panels (2) and (3) show that the gender semi-elasticity of demand for children is nearly the same among different marriage arrangements. A first-born daughter increases family size by around 6.6 percent irrespective of whether the marriage was arranged or the bride's family received a dowry. Altogether, the results in Table 6 do not support theories that rely on the intuition that cultural practices make girls economically less desirable.

Regional variation, rural residence and ethnicity have little role in explaining son preference either[17]. Unlike South Asian countries or China, couples in Turkey seems more homogeneous in sex-biased fertility behavior with the exception of educated mothers for whom the excess demand for sons is most rare.

[16] Results are not reported, available upon request from the author.
[17] Results are not reported, available upon request from the author.

3.5 Endogenous Stratification

The relationship between economic development and son preference is not obvious, since fertility decline might lead to a stronger male-preferring fertility behavior. Without a change in sex composition preference, a decrease in the maximum number of children that couples could bear might lead to a more pronounced manipulation of the sibling sex composition because of two reasons. First, education leads to a more efficient contraceptive use. For example, Dinçer et al. (2013) estimate that a change in compulsory schooling in Turkey from 5 to 8 years in 1997 raised the proportion of women using modern family planning methods by eight to nine percent. Second, higher age at first birth implies a shorter fertility age interval. Kirdar et al. (2012) show that the new compulsory schooling law also increased the average age at first birth substantially. Couples are therefore less likely to reach the desired number of sons without additional effort to bias the sibling sex ratio. In the meantime, as shown in the previous section, education itself leads to a more neutral preference of sex composition. The relationship between family size and son preference is therefore ambiguous since any factor that influences fertility level will most likely change the preference for sex composition, and vice versa.

The random assignment of the first-born child's gender allows the use of endogenous stratification to unfold the relationship between the preference for fertility level and sex composition. Endogenous stratification is widely used in randomized experiments[18] to estimate treatment effect on subgroups. Typically, subgroups are defined by terciles or quintiles of the predicted outcome without treatment. In this context, I quantify the effect of a first-born daughter on family size for different fertility levels predicted by endogenous covariates that

[18]See for example Goldrick-Rab et al. (2012).

evolve as a result of economic development.

In a recent paper, Abadie et al. (2013) show that leave-one-out and repeated split sample estimators are consistent and have well-behaved small sample properties when stratifying predicted outcomes. While their study is based on randomized experiments, the implications are the same for any observational study in which the assignment to treatment is unconfounded.

Using the notation of Abadie et al. (2013), let the treatment effect be the difference in the number of children between families with a first-born daughter and a first-born son, which can be computed as comparing the sample average outcomes:

$$\hat{\tau} = \frac{\sum_{i=1}^{N} y_i Z_i}{\sum_{i=1}^{N} Z_i} - \frac{\sum_{i=1}^{N} y_i (1 - Z_i)}{\sum_{i=1}^{N} (1 - Z_i)} \tag{4}$$

where as before, y_i is the sibship size and Z_i is a female indicator for the first-born child in the family i. The first step is to regress y_i on W_i, a set of predictors for the fertility level using the sample of women with a first-born son, i.e., $Z_i = 0$ and estimate the following regression coefficient vector:

$$\hat{\pi} = \sum_{i=1}^{N} \left(W_i (1 - Z_i) W_i^{'} \right)^{-1} \sum_{i=1}^{N} W_i (1 - Z_i) y_i \tag{5}$$

The next step is predicting a sibship size, $W_i'\hat{\pi}$, for the full sample of women. The treatment effect for each quintile $k = \{1, 2, 3, 4, 5\}$ of the empirical distribution of $W_i'\hat{\pi}$ then follows:

$$\hat{\tau}_k = \frac{\sum_{i=1}^{N} y_i I_{\left[Z_i=1, c_{k-1} < W_i'\hat{\pi} \leq c_k\right]}}{\sum_{i=1}^{N} I_{\left[Z_i=1, c_{k-1} < W_i'\hat{\pi} \leq c_k\right]}} - \frac{\sum_{i=1}^{N} y_i I_{\left[Z_i=0, c_{k-1} < W_i'\hat{\pi} \leq c_k\right]}}{\sum_{i=1}^{N} I_{\left[Z_i=0, c_{k-1} < W_i'\hat{\pi} \leq c_k\right]}} \tag{6}$$

where I is an indicator function and equals one if the condition (shown in square brackets) is satisfied. The lower and upper limits for the quintile k is indicated by c_{k-1} and c_k respectively. While $\hat{\tau}_k$ is consistent, Abadie et al. (2013) show that it is severely biased in finite samples due to overfitting and recommend using either leave-one-out ($\hat{\tau}_k^{LOO}$) or repeated split sample ($\hat{\tau}_k^{RSS}$) estimators, instead of $\hat{\tau}_k$ in equation (6). They caution that the bias increases with the number of regressors in W_i while other covariates can be incorporated when estimating the treatment effects in the second step[19] .

I use the endogenous determinants of fertility level (mother's age at first birth, father's and mother's years of education, region and rural residency) to predict the number of children for each family. Duflo (2012) notes that fertility decreases (and age at first birth increases) with income and education. Urbanization and migration from agricultural to industrial regions are also highly correlated with economic growth and prosperity. A naïve regression of sibship size on these predictors confirm this observational pattern: endogenous predictors have expected signs and can explain up to 30 percent of the variation in family size[20]. Note that

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

[20]Regression results are not shown, available from the author upon request.

the first step of endogenous stratification simply involves dividing the sample into quintiles based on the predicted fertility and is not concerned with causality. The key assumption for the causal identification is that within each quintile, the gender of the first-born child is as good as random.

Table 7 reports both adjusted and unadjusted differences estimated using $\hat{\tau}_k^{LOO}$ and $\hat{\tau}_k^{RSS}$ for quintiles and OLS estimator $\hat{\tau}$ from equation (4) for the full sample. Unadjusted differences shown in Table 7 are simple differences in the average number of children between families with a first-born female and a first-born male for the corresponding quintile. Adjusted differences control for the full set of covariates as employed in Table 3. The consistency 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 quintile treatment effects.

Column (5) in Table 7 reveals the high variation in family size across quintiles of predicted fertility. At the lowest fertility quintile, families with a first-born son bear on average 1.69 children compared to 4.41 children for the highest fertility quintile. Columns (1)-(4) shows that 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) provides the percentage change in family size (% Δ) as a result of DSB. If the first child is a female, the family size increases by 4.6 percent among the women in the lowest predicted fertility quintile. The estimated treatment effect reaches a peak at the median fertility quintile with 9.4 percent and declines to 6 percent for the highest fertility level. In conclusion, DSB shows a relatively flat response to decline in fertility. The relationship follows an inverse U-shaped path reaching a peak at the medium fertility level.

To put these results in perspective, the average fertility difference is 0.68 children between two quintiles and very large in magnitude compared to the decline in family size (≈ 0.21 children) during the 15 years that span the study period. Given the persistence in DSB effects across quintiles, one can conclude that couples in Turkey strongly desire at least one son for all levels of predicted fertility, including the most educated couples living in wealthy urban areas. Table 7 shows no evidence that the relative devaluation of a female birth is likely to disappear with declining fertility in the near future. More broadly, economic development that comes with fertility decline predicted by better education, more income and urbanization does not necessarily eliminate the gender-biased fertility preference.

4 Conclusion

In Turkey, the trend in the sex ratio at birth is remarkably constant around the commonly accepted natural sex ratio and there is no evidence or documented history of gender discrimination in relative care. 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 by young couples after a male birth is a contributing factor to an abnormal sex ratio distribution conditional on family size. I provide additional evidence that abortion is not a common practice for reaching the desired sex composition.

Unlike countries with “missing women”, son preference in Turkey manifests little spatial or demographic heterogeneity with the exception of educated women. Son preference is much weaker among women with secondary or higher education whereas father’s education does not reduce the sex bias in fertility preference.

Cultural norms are usually at the heart of the son preference discussion and its implications on the differential treatment of boys and girls are of great importance. This study shows that couples with patrilocal residence or traditional marriage arrangements do not differentiate in male-preferred fertility behavior compared to nuclear families with modern marriage practices.

I provide consistent estimates of son preference for different quintiles of fertility predicted by the endogenous sample characteristics. Quintile treatment effects indicate that the fertility decline does not necessarily translate into a gender-neutral pattern in fertility behavior.

While this opens room for policy action, and gender equality at birth can be an important policy goal on its own, this study does not go beyond the first stage effects of male-biased fertility stopping rules. In the absence of intrahousehold allocation of resources in favor of sons, DSB alone might not lead to negative long-term effects for girls. There is little causal evidence in the economic literature that supports the child quantity and quality trade-off predicted by Becker and Lewis (1974). However, birth order effects on children's education and other related later life economic outcomes are proven to be strong in internally valid settings[21]. DSB favors females in birth order. As shown in Black et al. (2005), if the negative impact of the higher birth order is largest for the last born children, DSB might controversially lead to better later life outcomes for females. Given the strong effects of DSB on family size and birth order, this provides an intriguing direction for future work.

[21] See for example Kristensen and Bjerkedal (2007).

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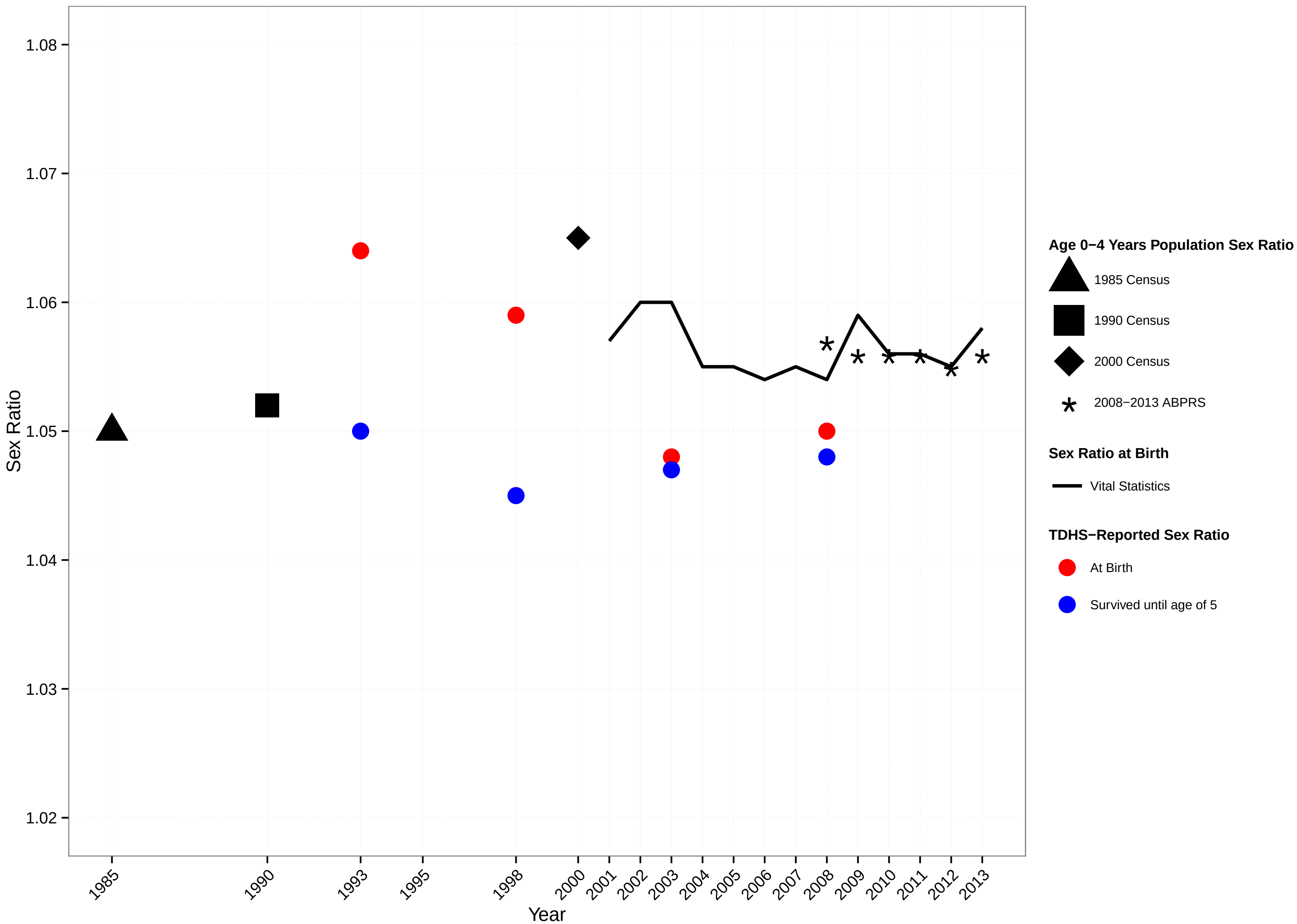
Figure 1. Sex Ratio Trends



Note: Table shows the number of boys per girl for each data source. Estimates from census data were shown with geometric shapes. Black line show estimates from the vital statistics. Blue and red points are estimates from the TDHS. *y*-axis is scaled to the commonly accepted natural sex ratio range at birth (1.02-1.08 boys per girl) and *x*-axis labels are shown for the period that the data were available.

Figure 2. Empirical Distribution of Count Variable Outcomes

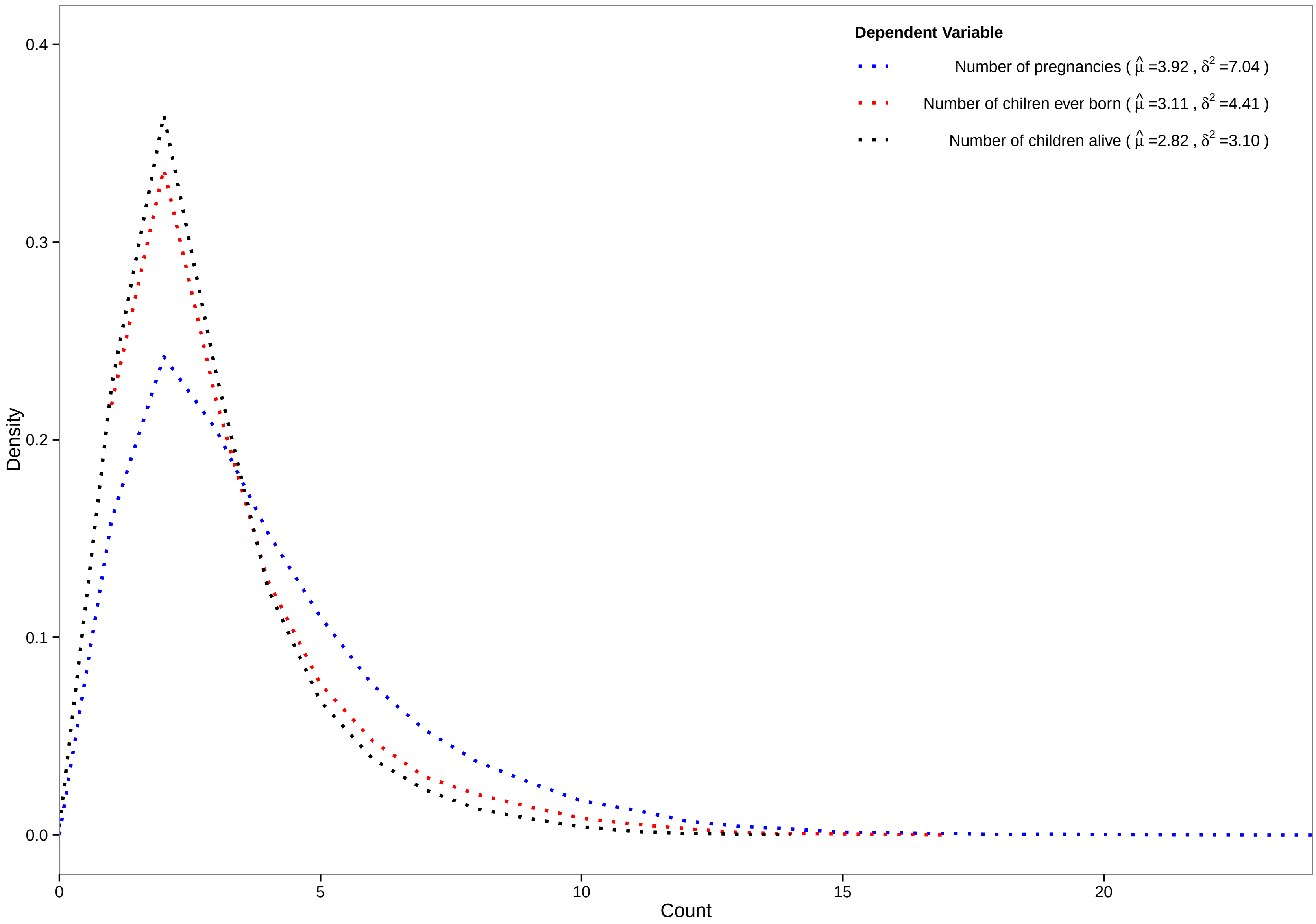


Table 1. Sex Ratios by Sibship Size and Year of Survey

Women Aged 15-49

Sibship Size	Sex Ratio 1993	1998	2003	2008	Pooled
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
Number of Children					
Ever 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)

Sibship 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: Sex ratio refers to the number of males per female, n indicates the number of living children in a family and 5+ indicates families with more than 5 children. Sample includes ever married women with at least one birth history and sample size is shown with N. In the lower panel, the column percentage refers to the percentage of small and large families in the sample.

Table 2. Sex Ratios by Birth Order

Women Age 15-49 (Pooled Sample)

Total Number of Births	Birth Order 1	2	3	4	5	6	7	*N*
1	**1.21**							4969
2	1.19	**1.19**						15584
3	1.08	1.06	**1.26**					15210
4	0.92	0.89	0.93	**1.20**				11864
5	0.98	0.89	0.94	1.07	**1.23**			8800
6	0.84	0.98	0.89	0.91	1.00	**1.23**		6588
7+	0.97	0.94	0.81	0.90	0.94	0.98	0.98	16659
				Average Birth Order				
Overall		*Boys* = 2.75				*Girls* = 2.78		79674

Women Age 35-49 (Pooled Sample)

Total Number of Births	Birth Order 1	2	3	4	5	6	7	*N*
1	**1.24**							761
2	1.30	**1.31**						6184
3	1.11	1.07	**1.27**					8217
4	0.94	0.90	0.92	**1.31**				7536
5	0.99	0.89	1.00	1.11	**1.21**			6145
6	0.85	1.04	0.87	0.92	1.03	**1.25**		5046
7+	0.98	0.93	0.84	0.90	0.94	0.99	0.98	14451
				Average Birth Order				
Overall		*Boys* = 3.18				*Girls* = 3.20		48340

Note: Sex ratio refers to the number of males per female, 7+ indicates the children whose mother had more than 7 birth histories. Sample includes all the birth records of the ever married women and sample size is shown with *N*. The sex ratio at last birth (SRLB) is depicted in bold.

Table 3. Baseline Characteristics of Families by First Child's Gender

Family Characteristics	First child's gender		Difference	t-test	N
	Boy	Girl		p-value	
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
Dowry paid to bride's family	0.23	0.24	-0.005	0.38	24956

p-value, joint χ^2-test = 0.42

N=25366 Pseudo-R^2=0.0006

Note: The p-values are based on a two-sample t-test of difference in means assuming equal variances. The joint F-test is based on a logit regression of first child's gender (equals 0 if boy and 1 if girl) on all variables shown in the table plus indicator variables for missing husband's age, husband's years of education, arranged marriage and dowry. Separate dummies for each survey year are included in the logit regression but not in the joint χ^2-test. Regression sample size and Pseudo-R^2 are shown at the bottom.

Table 4. Estimation Results by Age Category

		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***	0.188***	0.183***	-0.016***	-0.000
		(0.023)	(0.017)	(0.015)	(0.005)	(0.005)
	$\bar{y}\|Z_i = 0$	3.82	3.02	2.73	0.70	0.26
	N			25366		
	R^2	0.52	0.57	0.54	0.13	0.12
	$\hat{\tau}^{MLE}$	0.053***	0.062***	0.066***		
		(0.006)	(0.005)	(0.005)		
(B) Age 15-29	$\hat{\tau}^{OLS}$	0.087***	0.058***	0.060***	-0.025***	-0.001
		(0.022)	(0.016)	(0.015)	(0.010)	(0.007)
	$\bar{y}\|Z_i = 0$	2.29	1.93	1.82	0.70	0.12
	N			8301		
	R^2	0.50	0.59	0.58	0.14	0.08
	$\hat{\tau}^{MLE}$	0.039***	0.031***	0.034***		
		(0.010)	(0.008)	(0.008)		
(C) Age 30-39	$\hat{\tau}^{OLS}$	0.250***	0.263***	0.247***	-0.014*	-0.012
		(0.035)	(0.026)	(0.023)	(0.008)	(0.009)
	$\bar{y}\|Z_i = 0$	3.96	3.11	2.85	0.78	0.29
	N			9657		
	R^2	0.45	0.55	0.52	0.14	0.08
	$\hat{\tau}^{MLE}$	0.060***	0.079***	0.082***		
		(0.008)	(0.008)	(0.007)		
(D) Age 40-49	$\hat{\tau}^{OLS}$	0.272***	0.234***	0.234***	-0.008	0.016
		(0.056)	(0.040)	(0.034)	(0.010)	(0.011)
	$\bar{y}\|Z_i = 0$	5.37	4.13	3.60	0.58	0.37
	N			7408		
	R^2	0.40	0.54	0.50	0.19	0.06
	$\hat{\tau}^{MLE}$	0.052***	0.058***	0.065***		
		(0.010)	(0.009)	(0.009)		

Note: Estimated with OLS ($\hat{\tau}^{OLS}$) and maximum likelihood method ($\hat{\tau}^{MLE}$) assuming Poisson process. Coefficients reported are for the indicated age category. All regressions control for first born's survival, year of survey, mother's age, age at first birth, years of education, ethnicity, region, rural residence, husband's age, husband's years of education, patrilocal residence, whether the marriage was arranged and bride's family received a dowry plus indicator variables for missing husband's age, husband's years of education, arranged marriage and received dowry. Heteroskedasticity-consistent standard errors are in parentheses. Significance levels are indicated by * $< .10$, ** $< .05$, *** $< .01$. Mean outcomes for families with a first born male are shown with $\bar{y}|Z_i = 0$. N refers to number of observations in the regression for each age category.

Table 5. Interaction Effects on the Sibship Size

	Survey Year (1)				*Mother's Education* (2)				*Father's Education* (3)		
Category	OLS	$\bar{y}\|Z_i=0$	% Δ	Category	OLS	$\bar{y}\|Z_i=0$	% Δ	Category	OLS	$\bar{y}\|Z_i=0$	% Δ
1993	0.162*** (0.031)	2.86	0.057*** (0.011)	No Education	0.255*** (0.046)	4.19	0.057*** (0.010)	No Education	0.236** (0.099)	4.73	0.049** (0.019)
1998	0.151*** (0.032)	2.79	0.054*** (0.011)	Primary	0.206*** (0.018)	2.51	0.082*** (0.007)	Primary	0.212*** (0.023)	2.97	0.069*** (0.007)
2003	0.211*** (0.028)	2.68	0.075*** (0.010)	Secondary $\geq$	0.060*** (0.020)	1.79	0.036*** (0.010)	Secondary $\geq$	0.143*** (0.017)	2.17	0.067*** (0.007)
2008	0.203*** (0.028)	2.62	0.076*** (0.010)								
p (joint χ^2)	0.40		0.27	p (joint χ^2)	< 0.001		< 0.001	p (joint χ^2)	0.05		0.62
N	25366		25366		25366		25366		25283		25283
R^2	0.55				0.55				0.55		

Note: Outcome is the number of living children per woman. Estimated with OLS and maximum likelihood method assuming Poisson process. Coefficients reported are for the indicated category and are estimated by the interaction of that category with the first child's gender ($Z_i = 1$ if female) plus the main effect. The reported p-values are from χ^2 tests that the interactions with Z_i are jointly equal to zero. All regressions control for first born'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 bride's family received a dowry plus indicator variables for missing husband's age, husband's education, arranged marriage and received dowry. Heteroskedasticity-consistent standard errors are in parentheses. Significance levels are indicated by * < .10, ** < .05, *** < .01. Mean number of children for families with a first born male are shown with $\bar{y}|Z_i = 0$. N refers to number of observations in each regression. The percentage change in the mean outcome induced by a first born female is indicated by % Δ and estimated with maximum likelihood.

Table 6. Interaction Effects on the Sibship Size

	Patrilocal Residence (1)				*Arranged Marriage* (2)				*Dowry Paid* (3)		
Category	OLS	$\bar{y}\|Z_i = 0$	% Δ	Category	OLS	$\bar{y}\|Z_i = 0$	% Δ	Category	OLS	$\bar{y}\|Z_i = 0$	% Δ
No	0.185***	2.81	0.065***	No	0.147***	2.28	0.065***	No	0.163***	2.39	0.066***
	(0.016)		(0.005)		(0.021)		(0.008)		(0.015)		(0.006)
Yes	0.179***	2.15	0.080***	Yes	0.208***	3.02	0.067***	Yes	0.250***	3.83	0.067***
	(0.038)		(0.016)		(0.020)		(0.006)		(0.042)		(0.010)
Difference	-0.006		-0.015	Difference	0.061**		-0.003	Difference	0.087*		0.001
	(0.042)		(0.017)		(0.029)		(0.011)		(0.044)		(0.012)
N	25366		25366		25355		25355		24956		24956
R^2	0.55				0.55				0.55		

Note: Outcome is the number of living children per woman. Estimated with OLS and maximum likelihood method assuming Poisson process. Coefficients reported are for the indicated category and are estimated by the interaction of that category with the first child's gender ($Z_i = 1$ if female) plus the main effect. The reported *p*-values are from χ^2 tests that the interactions with Z_i are jointly equal to zero. All regressions control for 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 bride's family received a dowry plus indicator variables for missing husband's age, husband's education, arranged marriage and received dowry. Heteroskedasticity-consistent standard errors are in parentheses. Significance levels are indicated by * $< .10$, ** $< .05$, *** $< .01$. Mean number of children for families with a first born male are shown with $\bar{y}|Z_i = 0$. *N* refers to number of observations in each regression. The percentage change in the mean outcome induced by a first born female is indicated by % Δ and estimated with maximum likelihood.

Table 7. Endogenous Stratification Results on Sibship Size

	Repeated Split Sample		*Leave-One-Out*				
Quantile	Unadjusted	Adjusted	Unadjusted	Adjusted	$\bar{y}_k \mid Z_i = 0$	% Δ	N_k
	(1)	(2)	(3)	(4)	(5)	(6)	(7)
$\hat{\tau}_1$	0.096*** (0.024)	0.076** (0.017)	0.095*** (0.025)	0.077*** (0.017)	1.69	0.046	5073
$\hat{\tau}_2$	0.147*** (0.027)	0.141*** (0.025)	0.137*** (0.030)	0.128*** (0.026)	2.12	0.060	5067
$\hat{\tau}_3$	0.234*** (0.034)	0.219*** (0.027)	0.256*** (0.039)	0.229*** (0.031)	2.44	0.094	5081
$\hat{\tau}_4$	0.209*** (0.042)	0.219*** (0.039)	0.209*** (0.045)	0.219*** (0.041)	2.99	0.073	5073
$\hat{\tau}_5$	0.290*** (0.061)	0.257*** (0.044)	0.295*** (0.063)	0.265*** (0.045)	4.41	0.060	5072
			Ordinary Least Squares				
$\hat{\tau}$			0.196*** (0.022)	0.183*** (0.015)	2.73	0.067	25366

Note 1: Outcome is the number of living children per woman. Quantile treatment effects are estimated by repeated split sample and leave-one-out estimators provided in Abadie et al. (2014). Variables that are used to predict the quantiles are mother's age at first birth, mother's and father's years of education, rural residence, region. Adjusted regressions control for 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 bride's family received a dowry plus indicator variables for missing husband's age, husband's education, arranged marriage and received dowry. Number of repeated split sample repetitions is 100. Bootstrapped standard errors are in parenthesis. Significance levels are indicated by * $< .10$, ** $< .05$, *** $< .01$. Mean number of children for families with a first born male are shown with $\bar{y}_k|Z_i = 0$ for each quantile $k = \{1,2,3,4,5\}$. N_k is the number of observations in each quantile. The percentage change in the mean outcome induced by a first born female is indicated by % Δ and calculated by dividing adjusted $\hat{\tau}_k^{LOO}$ by $(\bar{y}_k|Z_i = 0)$.

Note 2: Full sample treatment effect ($\hat{\tau}$) is estimated via OLS using heteroskedasticity-consistent standard errors. Adjusted regression controls for the same full set of coefficients as listed above. Mean number of children for families with a first born male are shown with $\bar{y}|Z_i = 0$. The percentage change in the mean outcome induced by a first born female is indicated by % Δ and calculated by dividing adjusted $\hat{\tau}^{OLS}$ by $(\bar{y}|Z_i = 0)$.

Appendix Table 1

Full Set of Coefficients for the Logit Regression in Equation (2)

Variable	Coefficient	Standard Error	95 % Confidence Interval	
Mother's age	0.0014	0.0029	-0.0042	0.0070
Mother's age at 1^{st} birth	-0.0067	0.0037	-0.0139	0.0005
Mother's years of education	0.0065	0.0047	-0.0027	0.0157
Mother Non-Turkish	-0.0629	0.0404	-0.1421	0.0162
West	-0.0539	0.0418	-0.1359	0.0280
South	-0.0127	0.0448	-0.1005	0.0750
Central	-0.0597	0.0439	-0.1458	0.0263
North	-0.0876	0.0492	-0.1840	0.0088
Rural	-0.0128	0.0294	-0.0704	0.0447
Patrilocal Family	0.0808	0.0434	-0.0043	0.1659
Father's age	0.0015	0.0026	-0.0036	0.0066
Father's age missing	-0.0167	0.0477	-0.1103	0.0769
Father's years of education	0.0013	0.0042	-0.0068	0.0095
Father's education missing	0.0749	0.2041	-0.3251	0.4750
Arranged marriage	-0.0248	0.0280	-0.0798	0.0301
Arranged marriage missing	-0.4056	0.6322	-1.6447	0.8334
Dowry received	0.0432	0.0341	-0.0237	0.1101
Dowry received missing	-0.0906	0.1053	-0.2970	0.1157
Survey year=1998	0.0151	0.0393	-0.0620	0.0922
Survey year=2003	0.0158	0.0378	-0.0583	0.0899
Survey year=2008	-0.0106	0.0389	-0.0868	0.0657
Constant	-0.0433	0.1090	-0.2569	0.1702

Pseudo-$R^2 = 0.0006$
Number of Observations = 25366
χ^2-test statistic = 18.62
Prob [$\chi^2 > 18.62$]= 0.4154

Appendix Table 2

OLS Regression Coefficients for Fertility Outcomes (Age 15-49) , Table 4

	Number of Pregnancies		Number of Births		Number of Living Children		Contraceptive Use		Ever Had Abortion	
Variable	Coef.	Std. Err.	Coef.	Std. Err.	Coef.	Std. Err.	Coef.	Std. Err.	Coef.	Std. Err.
$\hat{\tau}^{OLS}$ (Unadjusted)	0.202	0.033	0.181	0.026	0.196	0.022	-0.014	0.006	0.003	0.005
$\hat{\tau}^{OLS}$ (Adjusted)	0.204	0.023	0.188	0.017	0.183	0.015	-0.016	0.005	0.000	0.005
Firstborn is still alive	-1.061	0.053	-0.954	0.041	0.324	0.034	0.041	0.011	0.018	0.010
Mother's age	0.170	0.003	0.125	0.002	0.109	0.002	-0.006	0.001	0.015	0.001
Mother's age at 1^{st} birth	-0.188	0.004	-0.153	0.003	-0.132	0.002	-0.004	0.001	-0.018	0.001
Mother's years of education	-0.060	0.004	-0.061	0.003	-0.055	0.003	0.007	0.001	0.007	0.001
Mother Non-Turkish	0.730	0.039	0.811	0.032	0.759	0.028	-0.124	0.009	-0.055	0.008
West	-0.955	0.038	-0.985	0.029	-0.862	0.025	0.104	0.009	0.062	0.009
South	-0.517	0.041	-0.512	0.032	-0.420	0.028	0.055	0.010	0.021	0.009
Central	-0.595	0.040	-0.641	0.030	-0.582	0.026	0.062	0.010	0.029	0.009
North	-0.571	0.045	-0.573	0.033	-0.476	0.029	0.101	0.011	0.023	0.010
Rural	0.109	0.028	0.281	0.022	0.217	0.019	-0.025	0.007	-0.062	0.006
Patrilocal Family	-0.264	0.033	-0.262	0.027	-0.212	0.024	0.032	0.010	-0.006	0.007
Father's age	-0.004	0.003	-0.007	0.002	-0.007	0.002	0.001	0.001	0.001	0.001
Father's age missing	-0.459	0.047	-0.367	0.035	-0.363	0.030	-0.434	0.010	-0.015	0.010
Father's years of education	-0.040	0.004	-0.043	0.003	-0.031	0.002	0.007	0.001	0.005	0.001
Father's education missing	0.247	0.196	0.082	0.173	0.059	0.144	-0.094	0.043	0.060	0.042
Arranged marriage	-0.061	0.025	-0.009	0.019	0.032	0.016	0.022	0.006	-0.021	0.006
Arranged marriage missing	-0.215	0.741	-0.280	0.619	-0.176	0.555	-0.051	0.117	0.110	0.123
Dowry received	0.555	0.036	0.433	0.028	0.332	0.024	-0.034	0.008	0.027	0.007
Dowry received missing	0.326	0.111	0.207	0.077	0.201	0.065	-0.006	0.024	0.035	0.024
Survey year=1998	-0.259	0.039	-0.210	0.028	-0.170	0.024	-0.074	0.009	-0.026	0.009
Survey year=2003	-0.544	0.036	-0.354	0.027	-0.268	0.023	-0.008	0.008	-0.075	0.008
Survey year=2008	-0.788	0.036	-0.471	0.027	-0.350	0.023	0.011	0.008	-0.111	0.008
Constant	4.338	0.108	4.201	0.081	2.640	0.070	0.838	0.026	0.076	0.023
Number of Observations	25366		25366		25366		25366		25366	
R^2	0.5186		0.5723		0.5371		0.1308		0.1154	