Transition of Son Preference: Child Gender and Parental Inputs in Korea

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Abstract

Sex ratio at birth remains highly skewed in many Asian countries due to son preference. The ratio in South Korea, however, has declined from 1990 and reached natural levels in 2007. This paper studies the transition of son preference in Korea by testing whether parent's time, educational, and health inputs differ by the sex of the child. Following Dahl and Moretti (2008), our empirical strategy exploits randomness of the first child's sex to overcome potential bias from endogenous fertility decisions. We find that relative to girls, boys have mothers who work fewer hours, spend less time on household chores, and receive more academic private education. These gender differences, however, have declined substantially over the past two decades. No child gender effects are found in the likelihood or duration of breastfeeding. Altogether, evidence suggests the existence, and weakening of, son preference in Korea.

JEL Codes: J13, J16, O12
Keywords: child gender, parental inputs, son preference, economic and cultural transition

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1 Introduction

Son preference has persisted across many generations, particularly in patriarchal Asian societies. This is well represented by the grossly skewed sex ratio at birth: per 100 girls, 115 boys are born in China, 112 in India and Hong Kong, and 111 in Vietnam.[1] Advanced technologies that allow prenatal sex selection (e.g. ultra-sound) and the increasing desire for smaller families have induced more parents to opt for their preferred—male—child.

*** Figure 1 here ***

However, South Korea (hereafter, Korea), which shares many of the traditional norms with nearby countries, seems to be heading towards gender neutrality. Sex ratio at birth surpassed 116 in 1990, but has declined steadily until reaching 107 in 2007 and remaining at natural levels since then (see Figure 1). Although there is still some bias among higher order births, Korea is no longer a country with thousands of "missing women" (Sen, 1992; Edlund and Lee, 2013).

Does the decline in sex ratio at birth imply a weakening of son preference, or do parents in Korea still treat boys and girls differently? This paper studies parent's time, monetary, and health inputs by child gender during the period of decreasing sex ratio at birth. Although natural sex ratio at birth means that discrimination has diminished in the "extensive" margin of whether or not to have the child, it may linger in the "intensive" margin after birth in the form of differential parental treatment. Such parental behavior could result in non-trivial differences in child's human capital accumulation and career choices later on.[2] Hence, we test whether there are boy-girl differences in mother's labor supply, child's household work time, expenditures on private out-of-school education, duration and likelihood of breastfeeding, and further investigate how these measures changed during the past two decades.

Gender differences in parental investments has been documented in various countries. Previous studies look at the effect of child gender on fertility (Abrevaya, 2009; Almond et al., 2013), parents' marriage probability and stability (Dahl and Moretti, 2008), parents' labor supply (Lundberg and Rose, 2002; Choi et al., 2008), health investments (Jayachandran and Kuziemko, 2011; Barcellos et al., 2014), and educational inputs (Baker and Milligan, 2013). Significant boy-girl differences are found in both developing and developed countries, with the magnitude being larger in developing countries.[3]

[1]Numbers are 2015 estimates available from the CIA World Factbook. For more details, see `https://www.cia.gov/library/publications/the-world-factbook/fields/2018.html` (accessed May 5, 2016).

[2]Almond and Currie (2011) reviews the growing literature on the long-term impacts of early childhood influences.

[3]See Lundberg (2005) and Bharadwaj et al. (2015) for an overview of the literature on child gender and family behavior.

However, previous work cannot show whether and how these child gender effects evolve over time. Korea is an extremely interesting case to study in this sense, as it is the only Asian country that has escaped the imbalanced sex ratio at birth.[4] The transitional experience could bridge the gap between existing studies on developing and developed countries and have important implications for other countries where son preference remains strong to this day. Although there has been research on trends in sex ratio at birth (e.g. Chung and Das Gupta, 2007; Lee and Paik, 2006) and the existence of gender discrimination in childcare in Korea (e.g. Choi and Hwang, 2015; Edlund and Lee, 2013; Kang, 2011), extensive evidence on over-time changes of gender effects is very limited.

The findings would also help understand the seemingly incongruent gap in women's educational attainment and their labor market outcomes in Korea and other developed Asian countries. Like the US, Korea's college entrance rate is now higher among girls than boys.[5] However, it still lags behind other OECD countries in female labor force participation, women's relative wage (to men's), and other measures of gender equality.[6] Investigating how much parents invest in their sons versus daughters, and interpreting why some gender effects exist (and are diminishing), would shed light on policies necessary to close down the gender gap.

To study parental inputs on various dimensions, we use data from several sources, including Korean Labor and Income Panel Survey (KLIPS), Korean Time Use Survey (KTUS), Korean Education Longitudinal Study (KELS), Private Education Expenditures Survey (PEES), and National Survey on Fertility, Family Health and Welfare in Korea (NSF).

Our empirical strategy exploits randomness of the first child's sex to overcome potential bias from endogenous fertility decisions, as in Dahl and Moretti (2008). Throughout the sample period of the past two decades, sex ratio at birth for first-born children shows no evidence of sex-selective abortions. Also, our data shows no difference in various observable characteristics between parents whose first child is male versus female. Even in the absence

[4]In Taiwan and Hong Kong, sex ratio at birth had been increasing until very recently and suddenly dropped to around 107 during the past couple years. It is yet too early to determine whether sex imbalance has improved in these countries. In Hong Kong, for example, the estimated sex ratio at birth for jumped back to 113 in 2014. Korea is unique in that the ratio of males to females at birth has been declining continuously for over two decades after soaring in the 1980s and reaching a peak in 1990. The unique experience of Korea is also pointed out by *The Economist* in `http://www.economist.com/node/15606229` (accessed May 5, 2016). Consequently, in 2008, the Korean Constitutional Court overturned the longstanding ban against the disclosure of a fetus' gender. Sex ratio at birth has been remaining around the natural ratio of 106 without the ban for the past several years and reached 105.3 in 2013.

[5]In Korea, girls' college entrance rate surpassed boys' in 2009 at 82.4 percent versus 81.6 percent (Statistical Yearbook of Education).

[6]According to OECD Employment Outlook 2013, employment to population ratio among women aged 15–64 in Korea ranks 25th out of 34 OECD countries (at 53.5 percent), and the gender earnings gap remains the largest (at 37 percent).

of sex-selective abortion, however, child gender in higher order births may be correlated with parents' preference on child gender if there are parents employing son-biased fertility stopping rules. Thus, boy-girl difference estimates from analyses including children of all birth orders may be biased.

Findings from KLIPS and KTUS provide evidence of important differences in time allocations—both parent's and child's—by child gender. Mothers of girls are more likely to be working and work longer hours compared to mothers of boys, a few years after first childbirth. Such gender gaps are not observed among women who gave birth after 2007, however. As for the time use of children, we find that girls on average spend twice as much time as boys (of the same age) in housework activities such as meal preparation and cleaning the house. The gender gap in housework time and participation rate halves from 1999 to 2009, however, from 1 hour per week to half-an-hour and from 18 percentage points to 9, respectively. Even at young ages, stereotypical gender roles arise in the household, but decreasingly so in recent years.

The effects of offspring gender on parents' educational inputs are studied using PEES and KELS. We focus on expenditures on children's private out-of-school education as they consist a major component of childrearing expenses in Korea.[7] Parents are 3 percentage points more likely to make expenditures and spend 23 dollars more per month on private out-of-school education in core academic subjects for their first-born boys than for first-born girls at middle-school age.[8] Higher expenditures on boys seems to reflect higher expectations regarding their academic achievement and career choices. Further analysis on monthly private education spending in a broader range of academic subjects for children at all grade levels indicates that the boy-girl difference has narrowed down from 10 dollars per month in 2007 to 1 dollar in 2012, however.

Unlike results on time or educational inputs, we find no evidence on differential treatment of boys and girls in terms of breastfeeding initiation or duration. Our analysis on 1990–2009 birth cohorts of infants using the NSF data shows that male first-borns are neither more likely to be breastfed nor breastfed for longer than female first-borns. Also, the boy effect is uniformly insignificant across all 20 birth years.

How can we interpret these findings? We explore three potential explanations for gender differential childcare: 1) cost of raising girls are higher than boys, 2) parents engage in

[7] Authors' calculation using PEES shows that about 79 percent of K-12 students received some kind of private out-of-school education, such as private tutoring, group tutoring, cram schools, or online courses, in 2007–2012. A major differences in educational investment arises from expenditures from private out-of-school education because the vast majority of primary schools are public and the majority of public and private secondary schools are quite homogeneous in terms of curriculum and tuition rates.

[8] Core academic subjects include Korean, Math, and English.

compensatory behavior, or 3) parents have son preference. The first hypothesis predicts that parents with first-born girls would be less likely to have additional children than those with first-born boys. This is rejected by the existence of son-biased fertility stopping rules. The compensatory behavior hypothesis predicts that parents would spend more resources on the child with greater needs—on average, male. Additional analyses using educational data, however, suggest the opposite: private education spending is positively correlated with child's previous academic performance. Son preference is the most consistent explanation of all the results in this paper. Parents invest more in boys than girls because they gain higher utility from them for economic or non-economic factors. Thus, improving gender equality at school and in the labor market seem to be both reasons for, and consequences of, weakening son preference in Korea.

The remainder of the paper is organized as follows. Section 2 describes the empirical strategy and econometric issues. Section 3 presents empirical results. Interpretations of our findings are given in Section 4. Section 5 concludes.

2 Empirical Strategy

Boy-girl differences in parental inputs can be measured using a regression model as follows:

$$y_i = \alpha + \beta Boy_i + \mathbf{X}_i'\boldsymbol{\gamma} + \varepsilon_i, \tag{1}$$

where y_i is the parental input of interest for child i, Boy_i is an indicator that equals one if the child is male and zero otherwise, $\mathbf{X}_i$ is a vector of baseline family characteristics, and ε_i is an error term. The OLS estimate of β captures the average effect of child gender on parental inputs if child gender is randomly assigned or exogenous conditional on covariates.

Although child gender can be considered being randomly determined at conception, boy-girl difference parameter in (1) may not be consistently estimated using observational data for two reasons. First, in societies where sex-selective abortions are prevalent, girls are less likely to be born into families with strong son preference. This implies that parents who are less biased against girls would be overrepresented among parents of daughters compared to the average population of parents. In this case, the effect of *Boy* on parental inputs would be underestimated due to the selection bias.

The downward bias of the OLS estimate due to sex-selective abortions relies on the argument that discrimination on the "extensive" and "intensive" margins are positively correlated. It is not obvious ex-ante whether girls would be treated better if they were born to parents less likely to engage in prenatal sex selection. Recent studies show empirical evidence that an increase in the practice of prenatal sex selection could lead to enhanced well-being, in terms of after-birth mortality rates and nutrition status, of girls in various countries. (See Lin et al. (2014) for the case of Taiwan, Hu and Schlosser (2015) for India, and Lee and Lee (2015) for Korea.)

*** Figure 2 here ***

Sex-selective abortion is still a serious issue in many Asian countries, including China and India, as reflected in the severely distorted sex ratio at birth. Korea once had a sex ratio at birth as skewed as those countries, but has started heading back to normality in the early 1990s. As a result, sex ratio at birth in Korea has been remaining within the normal range of 104–107 for first births since 1991, for second births since 2003, and for total births since 2007 (see Figure 2).

*** Table 1 here ***

As reported in Table 1, observable characteristics of parents at time of first childbirth presents very little difference by child gender.[9] Calculations from Vital Statistics show no statistically significant differences in parents' education level and the fraction of children born out-of-wedlock. Although not reported in Table 1, there are no differences in region by first child gender as well. Parents' age are shown to be slightly lower for boys than for girls but the magnitude is nearly zero (0.03 years). Although the difference is statistically significant, it seems to be merely due to the enormous size of the Vital Statistics data (in excess of 4,000,000 observations). Thus, it is reasonable to assume that first child gender is not subject to sex-selective abortions in Korea during the past two decades.

*** Table 2 here ***

Second, even in the absence of sex-selective abortions, child gender would not be exogenous if parents make fertility decisions based on gender of previous children. In many countries, parents are less likely to have additional children following a son than a daughter (i.e. son-biased fertility stopping rules).[10] Table 2 suggests that some families in Korea follow son-biased fertility stopping rules. Whether measured by the total number of children or the likelihood of having two (three) or more children, having a male first child reduces subsequent fertility.[11]

[9]Tables A2–A6 check covariate balances for each of the dataset we use in Section 3.

[10]Dahl and Moretti (2008) and Barcellos et al. (2014) provide evidence of families' fertility behavior following son-biased stopping rules in the US and India, respectively.

[11]We use KLIPS for this analysis because Vital Statistics lacks information on sibling's gender.

Both gender-neutral and son-biased couples would stop childbearing after reaching their preferred number of children—often two in Korea and many other developed countries. Suppose that the desired number of children is the same regardless of the type of parents but that son-biased parents may stop having a child before reaching the target number once they have a boy. In this case, a representative sample of children in all birth orders would result in a sample of parents in which son-biased parents are underrepresented and gender-neutral parents are overrepresented (Bharadwaj et al., 2015). With the selective sample of parents, boy-girl difference estimates would be biased downwards.

The downward selection bias arising from son-biased fertility stopping rules would become more pronounced if the regression model (1) controls for the number of children. This is because couples who have a son and stop having children before reaching their preferred number of children would be effectively left out from the estimation sample of parents sharing the same target number of children. Consequently, son-biased parents would be further underrepresented conditioning on the number of children.

On the other hand, son-biased parents may be willing to have a larger family until they have a son. For example, son-biased parents would have a third child if they have two girls, whereas gender-neutral parents stops at two children of any gender mix. In this case, son-biased parents are overrepresented and gender-neutral parents are underrepresented in a representative sample of children in all birth orders. In this case, boy-girl difference estimates would be biased upwards.

Overall, the *Boy* effect estimate would be biased in the presence of a son-biased fertility stopping rule. The selection bias could go in either direction depending on the fertility rate of son-biased and gender-neutral families.

To address the selection bias from son-biased fertility stopping rules, we exploit randomness of the first child's gender as in Dahl and Moretti (2008). Then Boy_i in the regression model (1) indicates whether the first child is male and the estimation sample would only consist of first-borns. It is reasonable to assume that first-child gender is exogenous during our sample period, because there has been no evidence of prenatal sex selection among first-borns since 1991. Son-biased stopping rules imply that first-born girls would end up having more siblings than first-born boys, however. As Bharadwaj et al. (2015) outlines, the identification strategy relying on the first child's gender thus consistently estimates the *total* effect of child gender on parental inputs, including any indirect effects through subsequent fertility choices that depend on the gender of the first-born child. The fact that first-born's gender influences subsequent fertility is part of the mechanism through which differential treatment occurs.

3 Estimation Results

3.1 Mother's Labor Supply

Labor supply adjustments are commonly observed with the beginning of parenthood as the value of the couple's home time rises after childbirth. Changes in labor market outcomes following parenthood need not be symmetrical between the couple, however, and in most cases they are not. Comparing the relative value of home time and market time, the value of women's time at home becomes greater than men's after childbirth due to both economic (e.g. gender wage gap) and physical (e.g. breastfeeding) reasons (Becker, 1991). Thus mothers usually reduce their labor supply whereas fathers may slightly increase or decrease theirs.[12]

In Korea, the gender difference in labor supply response to childbirth is particularly stark: men's response remains minimal whereas many women quit working. As a result, female labor force participation rate by age-group is M-shaped in Korea, dipping around childbearing age. Long working hours (the longest among OECD countries) and the Confucianist tradition that women are responsible for childcare and household duties (even more so than in Western countries), make balancing work and family difficult for working mothers with young children.[13] Although market substitutes for childcare are available as in the US, many believe mother's active involvement in school and after-school activities to be critical for the child's academic success.

In this context, we study whether the gender of the child has any role in mother's choice of labor supply after childbirth. We use data from the Korean Labor Income Panel Study (KLIPS), a longitudinal study of a representative sample of Korean households and individuals living in urban areas (comparable to the PSID of the US). We use all seventeen waves of KLIPS spanning 1998–2014 and construct a sample of households with mother-children pairs. To minimize the probability that some of the children might have left the household, we use information on relation to household head to identify the birth order of each child. As in Table 1, Table A2 confirms that observable characteristics of would-be parents do not differ by first child's gender.

Table 3 reports the effect of first child gender on mother's labor supply for a sample of women who were observed some time before as well as after first childbirth. Columns

[12] Lundberg and Rose (2002) discuss how the effects of children on men's labor market outcomes are ambiguous. As in Becker's work, specialization would predict that husbands focus more on the labor market. On the other hand, the value of both parents' time as inputs to childcare increase after a child is born; this effect would predict an increase in husbands' time at home.

[13] Korea's working hours has been the highest among all OECD countries from 2000 to 2007, and has been the second highest since 2008. According to 2013 statistics, yearly average working hours in Korea is 2,163 hours whereas the OECD average is 1,770 hours.

(1) and (2) look at both intensive and extensive margins (average weekly working hours) whereas columns (3) and (4) focus on the extensive margin (employment).[14] We run separate regressions for periods that span 2 and 5 years since first childbirth to see both short- and long-run responses, considering that women can take one year of maternity leave in Korea.[15] Control variables include corresponding labor supply variables prior to childbirth as well as mother's and father's demographic characteristics, year, and region. The independent variable of interest is a dummy equal to 1 if the first child is a boy.

In panel A, we find that relative to first-born girl mothers, boy mothers work about 2 hours per week less and are slightly less likely to be employed during the five years following childbirth. No statistically significant child gender effects are observed in the two-year period. To additionally investigate whether these *Boy* effects changed over time, we divide the sample into three groups according to woman's year of first childbirth in Panel B: 1999–2002, 2003–2006, and 2007–2014.[16] Roughly one-third of the sample belong to each group. All four columns consistently show that the coefficient on the first child being male is negative and large in the earliest birth-year group (1999–2002) but becomes smaller in magnitude in the second birth-year group (2003–2006) and is near zero in the most recent birth-year group (2007–2014). Among women who gave birth between 1999 and 2002, the boy effect on working hours is statistically significant even within two years after first childbirth. The difference in five-year average working hours between mothers of boys and girls declines from 7.37 hours per week, to 1.21 hours, then 0.39 hours. Similar trends are observed in the extensive margin as well, although statistically insignificant due to small sample size. Results are robust to alternative time horizons of the outcome variable and birth-year intervals.[17]

*** Table 3 here ***

Table 3 above, however, has the limitation of collapsing women's labor supply response into one measure pre- and post-childbirth rather than depicting how it rolls out over the years. Hence, we now modify equation (1) by introducing a series of dummy variables D_{it}^k

[14] KLIPS divides respondents into two groups—employed and not employed (including not in labor force). An individual is defined as employed if he/she worked at least 1 hour for pay during the survey week, worked at least 18 hours as unpaid family worker, or had a job from which he/she was temporarily absent (e.g., because of illness or vacation). Following KLIPS guidelines, working hours equal average working hours for temporary workers, the sum of regular and extra hours for permanent workers, and zero for not employed. Those with weekly working hours exceeding 168 are excluded.

[15] Women who have been employed for at least 180 days and have children under age 8 can apply for paid maternity leave in Korea. Applicants receive 40 percent of their regular monthly pay, with an upper limit of 1,000,000 KRW (roughly 1,000 USD) per month.

[16] Women who gave birth in 1998 (first wave of KLIPS) are not included in the analysis because we do not have information on their pre-childbirth labor supply.

[17] We do not present results on horizons longer than 5 years, because then there would not be observations for women who gave birth recently.

for years relative to first childbirth ($D_{it}^k = 1$ if k-th year relative to first childbirth, and 0 otherwise).

The estimation equation is:

$$y_{it} = \alpha_i + \delta_t + \sum_{k \geq -1} \lambda_k D_{it}^k + \sum_{k \geq -1} \beta_k Boy_i \times D_{it}^k + \varepsilon_{it}, \tag{2}$$

where y_{it} is the labor supply of woman i in year t. Coefficients β_k captures the causal effect of first child's gender on mother's labor supply in the k-th year relative to first childbirth (with $k = 0$ as the reference group). δ_t captures calendar year effects common to all mothers regardless of birth year, such as temporal changes in mothers' labor supply and fertility behaviors due to business cycles or work-family policy interventions. λ_k controls for relative time effects common to all women regardless of child gender. Finally, individual fixed effects α_i are included to absorb unobserved heterogeneity across individuals. Standard errors are clustered at the individual level.

Equation (2) is a generalized version of the "differences-in-differences" (DID) method (Jacobson et al., 1993). By using mothers of girls as a control group, we estimate the effect of first child gender on mother's labor supply, allowing the effect to vary by the number of years relative to childbirth. The identifying assumptions are thus analogous to the standard DID model: 1) women share common (relative) time trends in labor supply and fertility outcomes if first child gender is the same, and 2) there are no separate (calendar) time trends by first child gender.

Condition 1) is likely to be satisfied because first child gender is random in all years during the sample period. Another way of examining condition 1) is to test $H_0 : \beta_{-1} = 0$. Table A2 confirms that pre-childbirth observable characteristics are balanced across first child's gender. In order for condition 2) to hold, there should not have been any shocks or interventions during the analysis period that treated mothers differently by first child's gender. We are unaware of any relevant events that are child gender-specific, however, and the fact that estimates of β_k are not sensitive to adding or dropping calendar year fixed effects δ_t also reassures the validity of our approach.[18]

*** Figure 3 here ***

Figure 3 plots the estimated coefficients β_k for women's working hours and employment, respectively. Confirming condition 1), β_k is statistically indifferent from zero at $k = -1$.

[18] Of course, even when a policy is neutral to child gender in design, there could be different behavioral responses among mothers with male first child versus female fist child. This would not threaten the internal validity of the DID design, however, but concerns the interpretation of our findings. See Section 4 for a discussion on potential explanations.

The trend is downward-sloping thereafter. That is, the gender gap in mother's labor supply widens with years following childbirth such that mothers with first-born boys work less—both in terms of extensive and intensive margins—than those with first-born girls. The finding is also consistent with what is implied in Table 3: *Boy* effects are larger (more negative) in the five-year horizon than two-year horizon.[19]

As mentioned in Section 2, the child gender effects estimated here include any indirect effects that operates through subsequent fertility choices. The findings are thus more surprising given that families are less likely to have additional children following a son. Labor supply of mothers with first-born boys are relatively more depressed compared to those with first-born girls despite the fact that on average, the former have fewer children to take care of.

3.2 Child's Time Use at Home

Another way to analyze parental time inputs is to observe the kind of activities parents share with their children at home. We use the Korean Time Use Survey (KTUS), which reports how much time per day individuals spend on various activities, to study whether boys' and girls' time allocation differ. The study was collected by the National Statistical Office in 1999, 2004, and 2009. Each wave covers household members older than age 10 in 17,000, 12,750 and 8,100 households nationwide, respectively.

Unfortunately, KTUS does not contain detailed household member characteristics (including birth order and total number of children) nor information on which child an activity was carried out with.[20] Therefore, even if we observe one child respondent within the household we cannot be certain that the child is first-born or the only child because there may be child(ren) not currently living in the household or too young to be included in the study. It is also not possible to separate out the time parents spend reading to their son versus their daughter, for example.

With these data limitations, we examine time diaries reported directly by the children in the study and focus on the amount of time they spend doing household chores. The idea is that although we cannot directly assess the time parents spend with their first-born son versus daughter on activities that are believed to enhance a child's human capital, the time boys and girls spend on housework would provide suggestive evidence of the expectations parents have of their sons versus daughters, at least in terms of gender roles.

[19]We do not divide the sample into three groups as in Table 3 here, because the sample size at each k would be too small to make meaningful comparisons.

[20]KTUS has information on whom an activity was carried out with, but the categories are too crude for our use—alone, spouse, preschool children, other family members, and non-family members.

*** Table 4 here ***

Table 4 reports the effect of child gender on housework time and participation among respondents between age 10 and 18. Housework encompasses activities such as food preparation, washing dishes, doing the laundry, and cleaning the house. In the first two columns, we look at hours per week spent on housework (including zeros) and in columns (3) and (4), we look at whether the child participated in any housework during the survey date.

Pooling all three waves of KTUS in panel A, we find that the coefficient on child gender is negative and statistically significant. Boys spend about 1 hour per week less on housework and are 16 percentage points less likely to do any housework compared to girls. The results are robust to controlling for the child's age, mother's employment status, and other parental characteristics such as parent's age and educational attainment.

Panel B reveals another interesting pattern: gender difference in child's housework time and participation decreased in size across years. In 1999, boys spent about one hour less (per week) on housework compared to girls whereas in 2004, the gap decreases to 0.7 hours and in 2009, to 0.5 hours. Similar trend is observed in the extensive margin: the gender gap in housework participation decreases from 18 percentage points in 1999 to 13 percentage points in 2004 and to 9 percentage points in 2009. Changes are driven not only by girls reducing their load of household chores but also boys increasing theirs.[21] Although the size of the coefficients in 2009 are still non-trivial given that the mean housework time and participation rate is only about 1 hour per week and 28 percent, respectively, a child's gender role at home has become slightly more equal during the ten-year period.

One note of caution is that the coefficient on child gender cannot be interpreted causally because the KTUS sample is not restricted to first-born respondents. As discussed in Section 2, children from gender-biased parents may be over- or underrepresented given that some Korean families apply the son-biased fertility stopping rule.

3.3 Educational Inputs

Next, we investigate whether there are boy-girl differences in educational inputs in terms of money and expectations. We focus on parents' private education spending and expectations on educational attainment and career choices. Here, private education indicates private out-of-school education, such as private tutoring, cramming schools, and online courses, but do not include private school fees. Given that schools are quite homogeneous in terms of

[21]Mean of girls' housework time decreased from 1.57 to 1.13 hours per week and participation from 39.34 to 30.04 percent. Mean of boys' housework time increased from 0.55 to 0.64 hours per week and participation from 20.76 to 20.80 percent.

curriculum and tuition rates, expenditures on private out-of-school education are considered to be a main area of educational investments in Korea. We use data from the Korean Education Longitudinal Study (KELS) and the Private Education Expenditures Survey (PEES) to study educational investments.

The KELS provides data on learning experiences and transitions to work of a nationally representative sample of seventh-graders who were first surveyed in 2005. Similar to the National Education Longitudinal Study of 1988 (NELS:1988) in the US, data are collected from students, parents, teachers and school principals every year. We pool the first three waves of the KELS and construct a sample of middle school students born in 1992 or 1993.[22] The survey asks parents about monthly expenditures on private out-of-school education of Korean, Math, and English, which are key subjects of the college entrance exam. The survey also asks parents about the highest education level and occupational choices expected for their child.

The PEES is a nationwide survey specifically designed to collect data on the demand for private out-of-school education. Parents of students attending elementary, middle, or high schools are surveyed twice a year since 2007. We use the first six years of data from 2007 to 2012. The PEES contains detailed information about private education spending not just on academic subjects, but also on non-academic subjects, such as art, music, sports, and hobby activities. Academic subjects covered by the PEES not only include Korean, Math, and English, but also Science, Social Science, other foreign languages, computer programming, and critical writing. One limitation of the PEES data is that we cannot distinguish the birth order of children.

In Table 5, we find that sons receive higher educational investments on the three major subjects of the college entrance exam than daughters. Columns (1) and (2) report the boy-girl differences in monthly expenditures on private out-of-school education, which represent the *Boy* effect on both extensive and intensive margins. In Columns (3) and (4), we look at the extensive margin only by focusing on whether parents spend on their child's private out-of-school education.

*** Table 5 here ***

Column (2) in panel A presents that parents whose first child is male spend about 23 dollars more per month on their eldest child's private out-of-school education of the three key academic subjects compared to those whose first child is female. The difference is about 9 percent at the mean. Column (4) shows that parents are 3 percentage points (4 percent at

[22]We do not use later years of data because attrition rate becomes substantially higher once sample members enter high school in the fourth wave.

the mean) more likely to make expenditures on their first child's private education if the first child is a boy. Thus, the *Boy* effect on private out-of-school education spending is positive on both extensive and intensive margins.

Note that the estimated boy-girl differences in panel A are robust to controlling for baseline household characteristics, including parents' age and education, when we focus on first-born children.[23] This provides another piece of evidence that first child gender is exogenously determined without being contaminated by sex-selective abortions.

When we look at children at all birth orders in panel B, the estimated difference is slightly lower but within the standard error bounds of the estimate obtained from the sample of first-borns in panel A. The similar results in panels A and B imply the selection bias arising from sex-selective abortions in higher order births or son-biased fertility stopping rules seems small in the KELS sample. Or, it might be the case that the downward and upward biases discussed in Section 2 are cancelled out.

Panel C of Table 5 reports the estimated *Boy* effect on private education spending obtained from a PEES sample of children at all birth orders. The analysis in panel C is restricted to middle school students from the PEES 2007 in order to make the estimation sample comparable to the KELS sample in panel B. Also, private education spending is restricted to the three subjects, Korean, Math, and English, to make the outcome variable as consistent as possible with the one used in panel B. Compared to the estimates (23 dollars per month) in panel B, the effect on monthly expenditures is lower in terms of magnitude (18 dollars per month) but similar in percentage terms (9 percent at the mean). On the extensive margin, similar boy effect estimates are obtained.

Data from the PEES give us a more comprehensive understanding of expenditures on private out-of-school education. We use the full PEES sample of students attending primary and secondary schools in 2007–2012 to analyze private education spending on academic and non-academic subjects separately. The results in panel A of Table 6 show that the effect of *Boy* on private education spending is positive for academic subjects, whereas it is negative for non-academic subjects. That is, parent's private education spending pattern differs by the gender of the child. The magnitude of the boy-girl difference estimate for academic subjects is only about one ninth of that found from the KELS sample because the boy effect is nearly zero for primary schoolers (Appendix Table A7) and the effect for secondary schoolers has substantially decreased over time (Appendix Table A8).

*** Table 6 here ***

[23] The full list of control variables include parents' age, age squared, and dummies for parents' education (less than high school, high school, college or more), survey year, and urban rural classification (Seoul, large cities, small cities, rural).

In panel B of Table 6, we find that the positive effect of *Boy* on private education spending has substantially decreased over time for academic subjects. Panel B also presents that the boy-girl difference has slightly narrowed down for non-academic subjects. The gender differences and their over-time variations are mainly driven by expenditures for students at secondary schools (see panel B of Appendix Table A7 and A8).

The boy-girl difference estimates in Table 6 are likely to be biased since we cannot restrict the analysis to first-born children using the PEES data. The bias is likely to be small though, as the analysis of the KELS data in Table 5 indicates.[24]

*** Table 7 here ***

Higher expenditures on private education for boys may reflect higher expectations on their academic achievement and labor market outcomes. Our study using the KELS data finds that parents' expectations on their children's educational attainment are on average 0.24 years longer for first-born sons than for first-born daughters (See Table 7). The difference is small but statistically significant at the one percent level and exists despite the recent reversal of the gender gap in college entrance rates.

*** Figure 4 here ***

Figure 4 indicates that there is a substantial boy-girl difference in parents' aspirations for their children's career choices as well. The KELS asks parents to select two occupations that they would like their children to have in the future. We plot the fraction of parents who select each occupation by first child's gender. Parents of sons are more likely to select high-wage professions or those that require advanced degrees, such as doctor, professor, lawyer, and CEO, than parents of daughters. Parents of daughters tend to select female dominated occupations, such as nurse and fashion stylist, or the ones with a high level of job security, including teacher and pharmacist. Teacher is parents' most preferred choice for their daughter's future occupation (selected by 54 percent). Although 20 percent of parents state that they want their children to be scientists or engineers when their first child is male, the number decreases to only three percent when their first child is female. It seems that parents' expectations on sons' versus daughters' career choices are generally consistent with occupational gender stereotypes. The boy-girl difference in the fraction of parents selecting each occupation is statistically significant at the one percent level for all

[24] If downward bias from sex-selective abortions in higher order births is dominating, the boy-girl difference estimates are likely biased towards zero. As sex ratio at birth for higher parities heads toward the natural ratio, this type of selection bias would become less severe in later years. If this is true, the gender gap estimates from earlier years could be biased downwards more severely than those from more recent years. Then, the over-time decrease in the gap could have actually been more dramatic.

the listed occupations. Also, the estimated differences are robust to adding control variables on baseline household characteristics.

3.4 Breastfeeding

We also look at health inputs during early childhood. Breastfeeding duration is commonly used to measure early childhood investment in the literature. Previous studies show that in developing countries with strong son preference, breastfeeding duration is shorter for girls so that mothers try out for a son (Jayachandran and Kuziemko, 2011; Barcellos et al., 2014). This could result in gender health disparities since breast milk is a major source of nutrition for infants and thus a main determinant of child health in developing countries where food and clean water are scarce. In developed countries, infant formula is routinely used as an affordable substitute for breast milk when nursing involves substantial costs in terms of time, physical conditions, or emotional stress.[25] Given that infant formulas does not have many of the benefits that breast milk has, breastfeeding duration could partly indicate how much mothers are willing to invest on child health.

We use repeated cross-sectional data from the National Survey on Fertility, Family Health and Welfare (NSF). The NSF is designed to provide information on marriage, pregnancy, contraception, childbirth, child rearing, work-life balance, support for parents, and family health. The first survey started in the 1960s but data are available from 1991, since when the NSF has been conducted every three years. Our estimation sample is restricted to first-born children with information on breastfeeding in NSF 1994–2009 (6 trennial surveys). The NSF asks mothers about pre and postnatal care only for their youngest children born in the past three years. Thus, the estimation sample consists of first-borns less than 3 years old without siblings.

*** Table 8 here ***

Our empirical analysis find that the duration or likelihood of breastfeeding do not differ by child gender in Korea during the past two decades. Table 8 shows that boys are neither more likely to be breastfed nor breastfed for longer than girls. Rather, the signs of *Boy* effect estimates indicate that breatfeeding duration or likelihood may slightly decrease when the child is male (although the boy effect is insignificant at the 5% significance level). The results are robust to additionally controlling for a low birthweight dummy (see columns (3) and (6) in Table 8) or to different functional form assumptions (see Appendix Table A9). When we allow boy effects to vary across child's birth cohorts, we cannot reject the null

[25] Lee and Lee (2015) describes time, phisical, and psychic costs associated with breastfeedign in Korea.

hypothesis that the boy effects are uniformly zero across all birth years from 1990 to 2009 (see Figure 5).

*** Figure 5 here ***

4 Interpretations

The evidence in Section 3 reveals that mother's labor supply, child's time use at home, and parent's private education spending all still differ by child gender in Korea, although the gender gaps are decreasing. On the other hand, no differential treatment is observed in parent's health inputs throughout the sample period. How can we explain these findings? In this section, we discuss three possible explanations for child gender effects then provide further evidence to distinguish between them.

Differential cost. First, the cost of raising boys and girls may differ. That is, child gender could directly affect the household budget constraint. If marital costs are much higher for daughters than sons, for example, parents would have to save more for the former. Even if parents do not have gender bias and the total amount of parental input is the same for boys and girls, the *net* resources available for them during childhood could differ by child gender in this case.

Compensatory behavior. Second, boys may have greater needs than girls. Infant mortality is higher among boys than girls in most countries including Korea.[26] Boys are also known to have more attention difficulties and lower noncognitive skills compared to girls (see, for example, Bertrand and Pan, 2013; Jacob, 2002; Ready et al., 2005). Depending on parent's level of inequality aversion, altruistic parents may engage in compensatory behavior by investing more in the weaker—male—child (Behrman et al., 1982).

Son preference. Lastly, parents may have son preference due to gender differences in economic returns, psychic returns, or both. Despite the recent educational crossover and increases in female labor force participation, Korea's gender gap in wages remains one of the largest among OECD countries at 37 percent.[27] Parents would invest more in boys than girls if the former is more likely to reap the benefits of these investments and is more reliable for old-age support.

Son preference due to gender differences in non-economic returns is also a possibility in a country like Korea, where Confucianism dictates that only sons "carry on the family line."

[26] According to KOSIS, infant mortality rate in 2014 is 3.2 for male and 2.8 for female infants. Extensive research indicates that due to differences in genetic makeup, boys are biologically weaker and more susceptible to diseases and premature death (see for example, Naeye et al., 1971; Waldron, 1983).

[27] Source: 2013 OECD Earnings Distribution Database.

For instance, the eldest son is considered to be responsible for holding memorial services for the family's ancestors whereas married women are considered to belong to the husband's household in these ceremonies. These rituals have become less important nowadays, but they are still carried out by most families and is often cited as a reason why married couples "need to have a son."

In discussing the son preference hypothesis, however, we do not attempt to separate out these two types of returns. This is because it is very difficult, if not impossible, to isolate "pure taste" from economic elements. Both parent's preferences and child's human capital are endogenously formed and interact in reinforcing ways. If parents prefer boys to girls and hence, invest more in them, boys will inevitably yield higher economic returns than girls in the future, and would further justify parent's bias for them. Whether (psychic) taste for sons could have been formulated if not for the history of gender differences in economic returns is highly questionable.[28] Confucian traditions are also not unrelated to the fact that only men could provide for the family in the past.

Keeping this in mind, we now conduct additional exercises in an effort to distinguish between the three channels—differential cost, compensatory behavior, and son preference. If the first interpretation is true such that boys are less expensive to raise than girls, couples whose first-born is male will be more likely to have another child compared to those whose first-born is female due to an income effect (assuming that children are normal goods). This is rejected in Korea, however, by the presence of son-biased stopping rules (Table 2). Also, unlike countries like India where large dowries are common, the newlywed's home (by far the largest marriage expenditure) is usually supplied by the groom's family in Korea whereas other gifts are prepared by the bride's. Although more gender-equal exchanges have been increasing among recent marriages, differential cost is not in favor of boys in Korea.

As for compensatory behavior hypothesis, if this is true, we should observe parents spending more time and money on those children who are having difficulties. We should also find that the effect of child's deficit on parental input is stronger for boys than for girls given that boys, on average, have more health and behavioral problems than girls.

In order to test whether parents engage in such behavior, we investigate how educational inputs differ by child's academic performance from the previous year and whether the difference depends on child's gender using data from the KELS. The KELS data include individual-level scores from a standardized test on Korean, Math, and English conducted every year from 2005 to 2007. We construct a measure of child's academic performance by

[28]Studies often use gender wage gaps or female labor force participation rates as proxies for a society's gender norms because they are highly correlated (see for example, Alesina and Giuliano, 2010; Fernández and Fogli, 2009; Hwang, forthcoming). Within Korea as well, regions with lower sex ratio at birth have higher female wages (Lee, 2013)

normalizing total scores on the three subjects to have zero mean and unit variance. We estimate a modified version of equation (1) by adding last year's test score and its interaction with child gender to the regression equation.

*** Table 9 here ***

Table 9 shows that educational inputs are positively correlated with child's previous academic performance. Conditioning on child gender, parents spend more on private out-of-school education and expect a higher level of educational attainment when their children perform better in school. Such pattern of parenting behavior would reinforce ability differences among children. The positive and statistically significant coefficient estimates on last year's test score (and its square) provides evidence of reinforcing, rather than compensating, behavior in parental investments. Although boys on average receive higher educational investment compared to girls with the same level of academic achievement, we do not find an additional boy effect on the reinforcing behavior.

We conduct a similar analysis using a self-reported measure of 6th-grade academic performance available in the 2005 KELS data. Based on the self-reported measure, we classify 7th graders into three groups – low, middle, and high achievers. Table A10 reports results from regression analyses using these categorical variables of academic performance. Although estimates are not directly comparable, the results in Table A10 are qualitatively similar to those in Table 9 above.

In addition, results from Table 6 and 7 also cast doubt on the claim that boys have greater needs than girls. Parents expect longer educational years for their sons than their daughters and focus on different subjects when spending on private education – academic subjects for boys and arts for girls. If parents allocated monetary inputs unequally because boys fared worse than girls in school (particularly in core subjects), it is difficult to explain why parents then expect higher educational attainment for the former. There is no barrier against women pursuing tertiary education in Korea; college entrance rates are higher for girls than for boys since 2009. Thus parents' expectations cannot be interpreted as reflecting the status quo, either.

Lastly, we do not expect gender-specific constraints to be vastly different across developed countries and over time. Even if it is true that boys tend to have more health and behavioral problems than girls, it is difficult to argue that these problems are much larger for boys in Korea than elsewhere or in the past than nowadays. For example, girls have lower infant mortality rates and higher noncognitive skills than boys in the US as well, but studies do not agree on the direction or existence of child gender effects on parental inputs, and the magnitude is relatively small even in papers that do find gender gaps in favor of sons.[29] The boy effect is not constant over time as well, as we have shown in Tables 3, 4, and 6: gender gap in mother's labor supply, child's housework time, and private education spending decreased substantially during the past two decades.

All in all, it is difficult to explain our findings without son preference. Parents who gain higher utility from sons than daughters would spend more time and money on boys and would have higher aspirations for them regardless of their performance at school. Of course, even if these findings are due to son preference, the decline of such gender differences does not necessarily prove the weakening of son preference. Other concurrent changes could confound the link between the two. If a new policy or technology generates different behavioral responses by child gender, for example, we may observe similar trends even if son preference persists.

However, we cannot ignore the fact that as women's opportunities in higher education and the labor market improved in Korea, the gap in economic returns from sons and daughters became much smaller than before. Changes in institutions such as the abolition of the "male head of the family" (hoju) register system in 2008 symbolize a decrease in non-economic returns from sons as well. Parents now have less reason to treat children differently by gender, which enables girls' performance to improve academically and economically, which in turn reinforces parents' gender-equal treatment. This interpretation is also consistent with what happened in the "extensive" margin of discrimination: sex ratio at birth has steadily declined in Korea during the past two decades.

*** Figure 6 here ***

Finally, one straightforward evidence of the transition of son preference is the responses from a survey that asks parents whether they prefer a son or a daughter. Data from the Korean Value Survey (KVS) that asks adults "Suppose you could only have one child. Would you prefer that it be a boy or a girl?" shows that 40.4 percent say "boy," 9.8 percent "girl," and the rest "either" in 1996 (see Figure 6). By 2008, the gap diminishes to 23.5 percent answering "boy," 16.1 percent "girl," and the majority (60.5 percent) "either."[30]

[29] Child gender does seem to affect the stability of parent's relationship in the US. Dahl and Moretti (2008), for example, finds that a first-born daughter is 3.4 percent less likely to be living with her father compared to first-born son. Time allocation results are more ambiguous, however. Using PSID, Lundberg and Rose (2002) shows that both sons and daughters increase father's work hours but that there are no significant effects of child gender on mother's work hours. On the other hand, using the German Socio-Economic Panel, Choi et al. (2008) shows that a first-born son increases father's working hours by 60 hours per year but not a first-born daughter. See Lundberg (2005) for a review of the literature.

[30] According to the KVS 2008 summary report, son preference is shown to be stronger among men, individuals who are older, have lower educational attainment, and lower household income. Korean General

5 Conclusion

This paper shows that boys still receive more parental inputs than girls in Korea, but that the gender gap has been decreasing substantially over time. We use multiple micro datasets to include different measures of parental investment, and exploit the randomness of the first-born child's sex to reduce bias from sex-selective abortion and endogenous fertility decisions.

We find that women work on average 2 hours less per week during the five years following childbirth when the first born is male than when it is female. In 1999, girls on average spent an hour more per week on household work than boys of the same age, but the gap halved to 0.5 hours in 2009. The boy effect on private out-of-school education spending on academic subjects is positive on both extensive and intensive margins: parents of first-born sons spend about 23 dollars per month more and are 3 percentage points more likely to spend any money at all, than those of first-born daughters. This difference narrows down, however, from 10 dollars per month in 2007 to 1 dollar in 2012 when we including all grade levels and subjects. On the other hand, no gender differential treatment is observed in the duration or likelihood of breastfeeding.

Together with the decline of sex ratio at birth, the findings indicate that Korea is in a transitional phase. It is at a different stage from not only other developing Asian countries but also from more developed countries. Gender discrimination before or at birth has become much less common, but child gender still matters and stereotypical gender roles continue in the household. These differences are dwindling recently, however. Weaker son preference, more equal parental inputs to sons and daughters, increase in women's educational attainment and wages, seem to reinforce each other. In light of these discussions, the paper implies the importance of polices aimed at promoting gender equality. Although it is difficult to manipulate preferences, they too, are not immune to economic incentives in the long run. Reducing the gender gap in labor markets, for example, may have amplified effects by eliminating reasons for gender differential childcare in the first place.

Social Survey (KGSS) also asks about preference for sex of child in 2004, 2006, 2008, and 2012. Similar to KVS, the KGSS data shows a positive but diminishing gap in the fraction of people preferring a son versus a daughter. Interestingly, more people states they would prefer a daughter rather than a son in 2012 KGSS, which shows that the gap has overturned.

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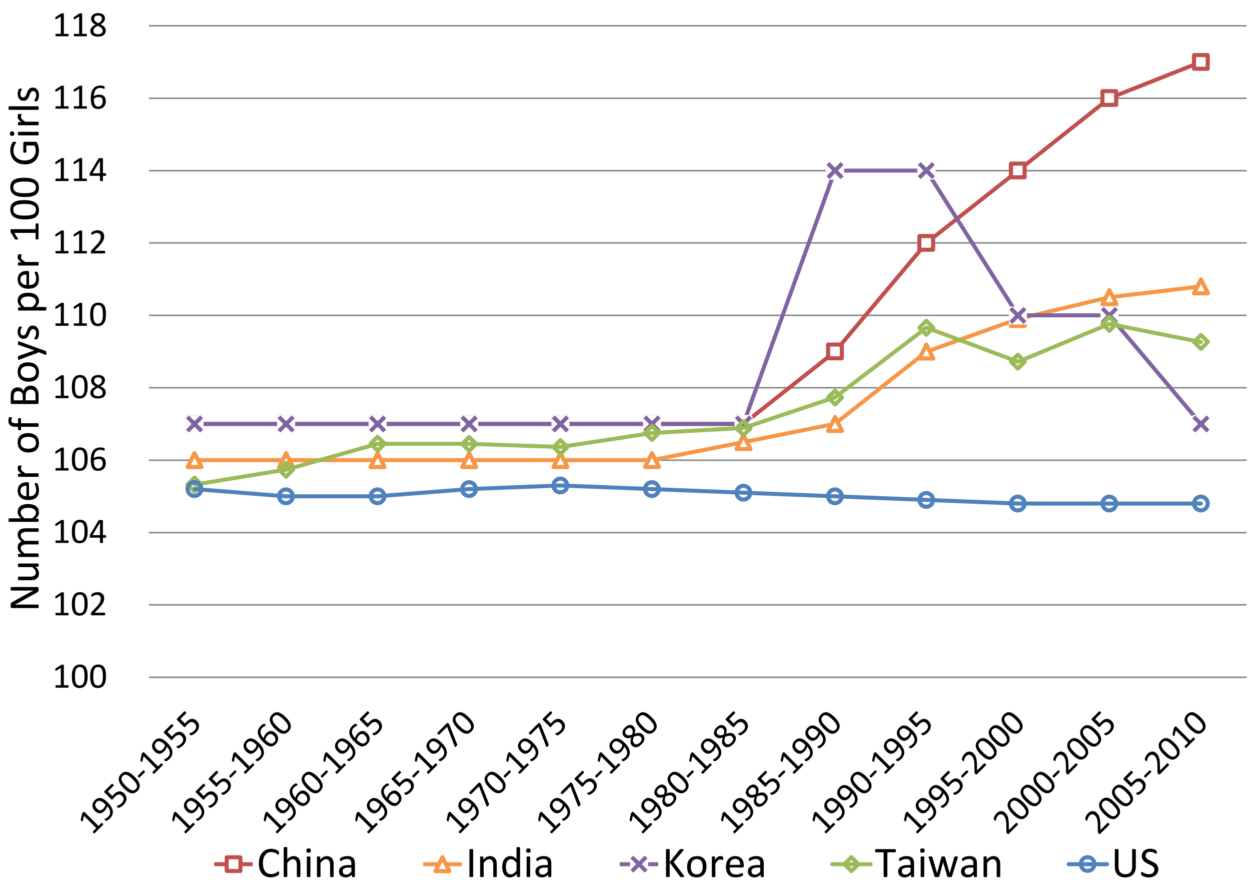


Figure 1: Sex Ratio at Birth, Selected Countries

Notes. Data for China, India, Korea, and US are from United Nations, World Population Prospects: The 2012 Revision. Data for Taiwan are from the Department of Household Registration (statistics for 1950–1979 are not available for two counties in Taiwan—Kinmen and Lienchiang).

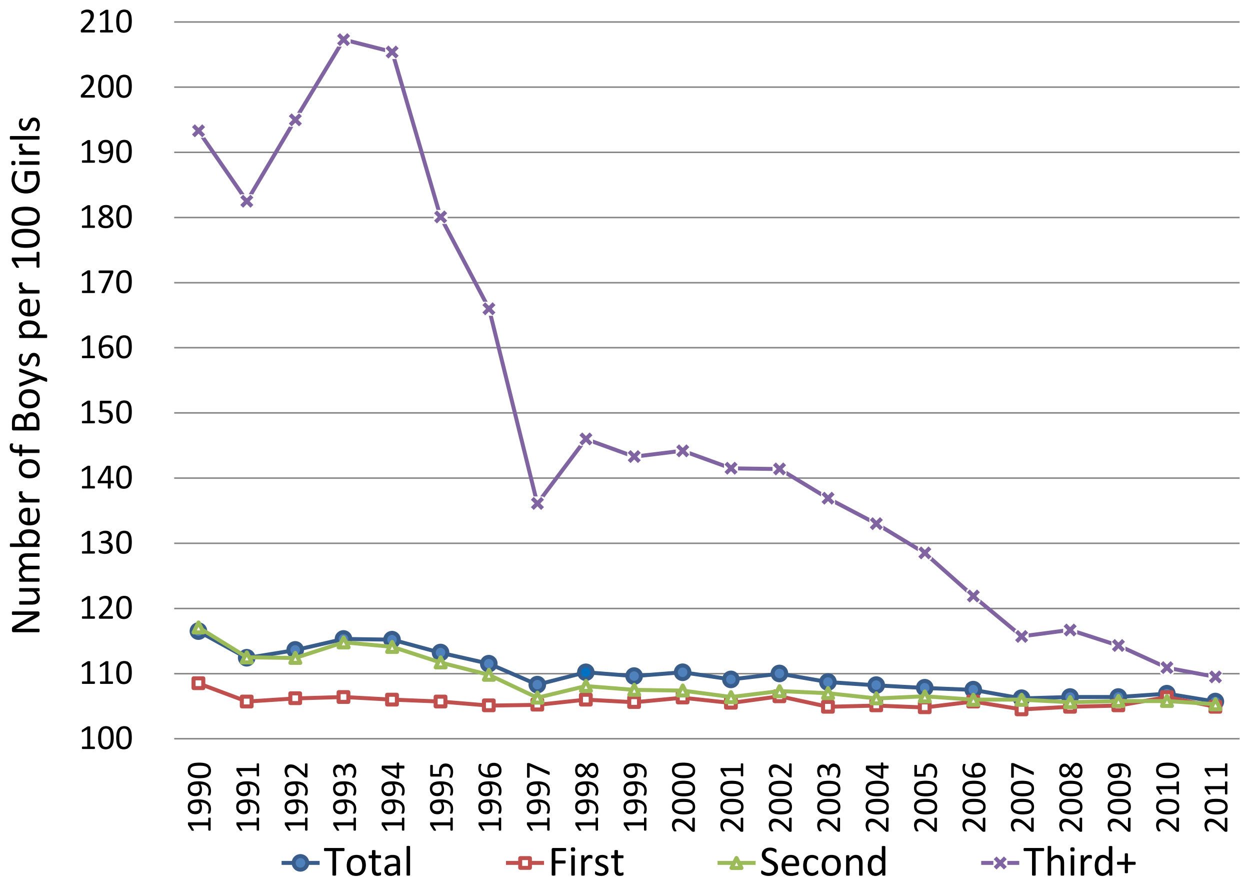


Figure 2: Sex Ratio at Birth, Korea by Birth Order

Source. KOSIS Vital Statistics.

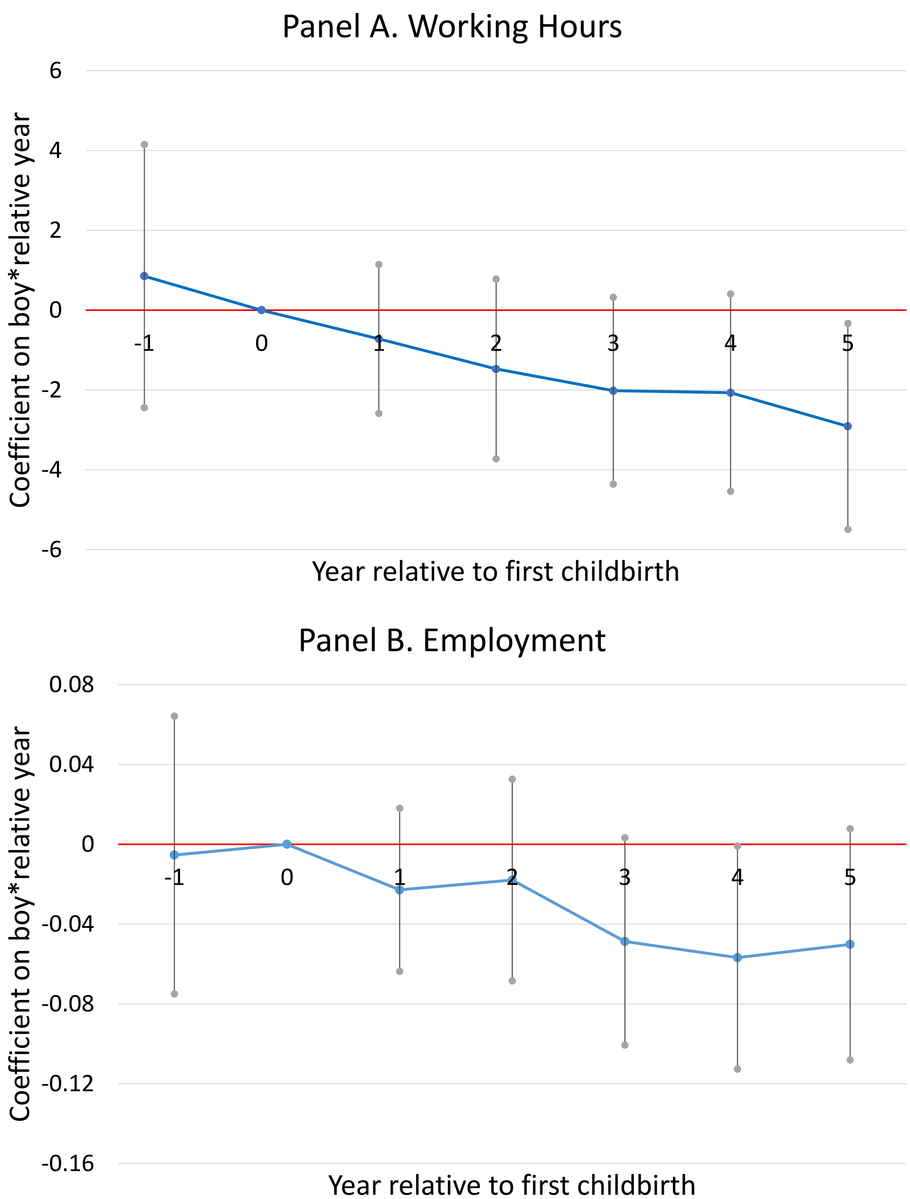


Figure 3: Mother's Labor Supply Over Time, by First Child's Gender

Notes. Estimation using KLIPS 1998–2014. The figure plots β_k, the coefficient estimate of $Boy \times k$-th year relative to first childbirth. Control variables include calendar year and relative year fixed effects as well as individuals fixed effects. Standard errors are clustered at the individual level. Vertical spikes around the point estimate represent the 95% confidence interval. Refer to equation (2).

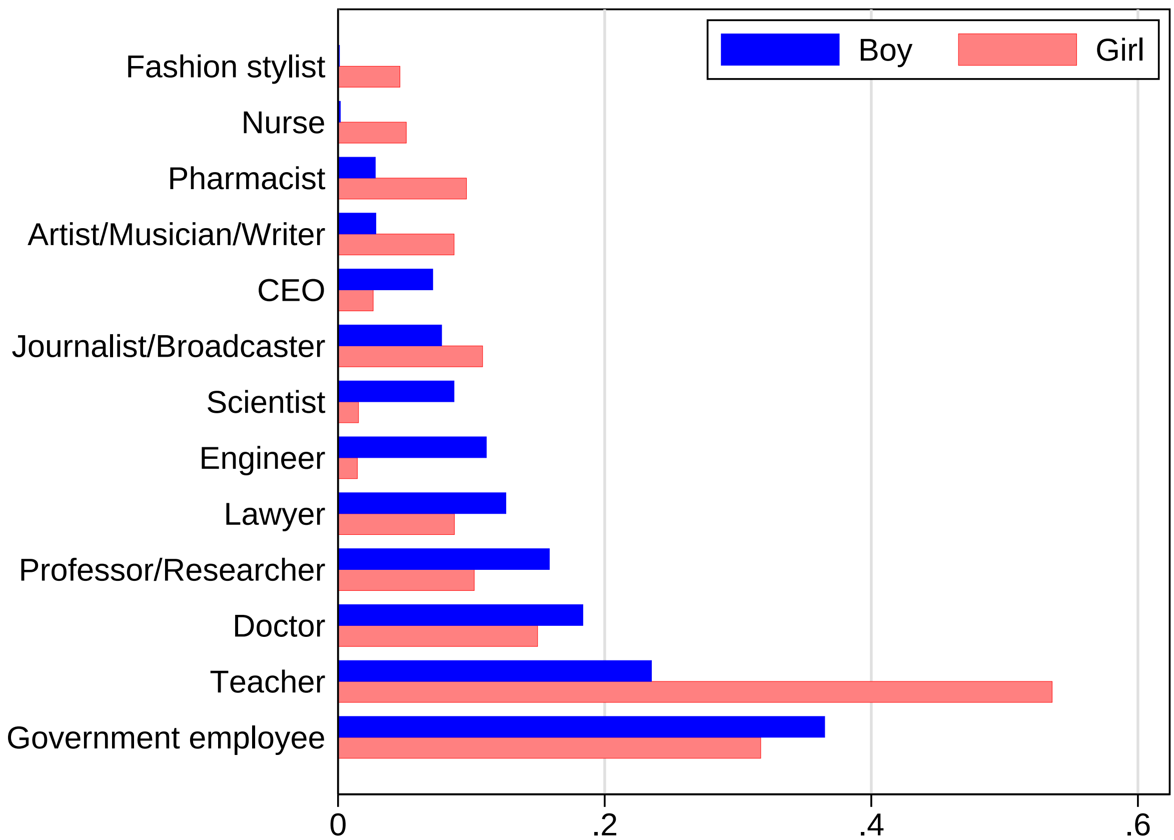
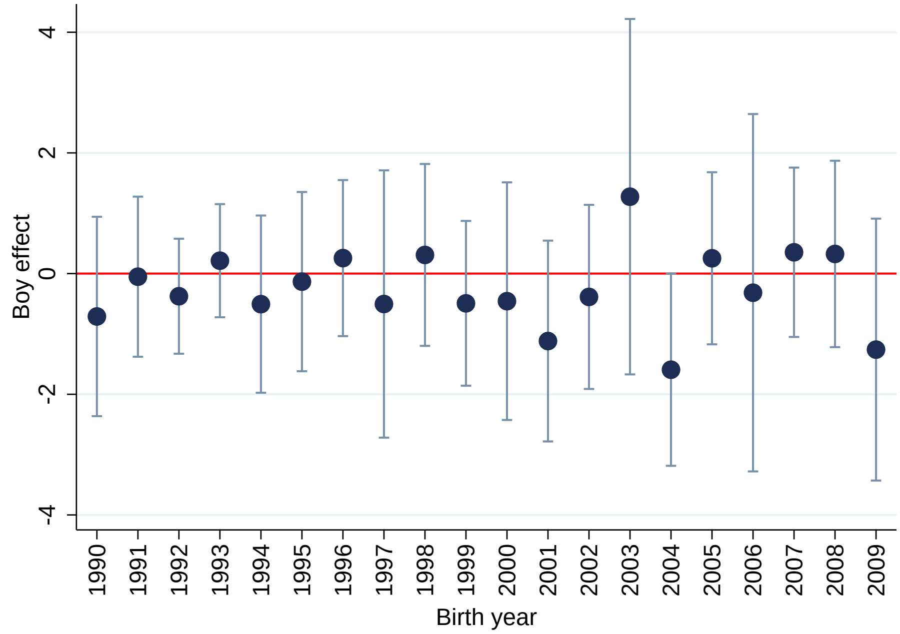


Figure 4: Parents' Expectations for Their Children's Future Occupations, by First Child's Gender

Notes. Calculations using KELS 2005–2007. Each bar represents the fraction of parents who want their eldest children to have the given occupation. The Professor/Researcher category excludes scientists. The boy-girl difference in the fraction of parents selecting each occupation is statistically significant at the 1% level for all the listed occupations. The estimated differences are robust to adding control variables listed under Table 5.

Panel A. Number of Months Breastfed



Panel B. Ever Breastfed?

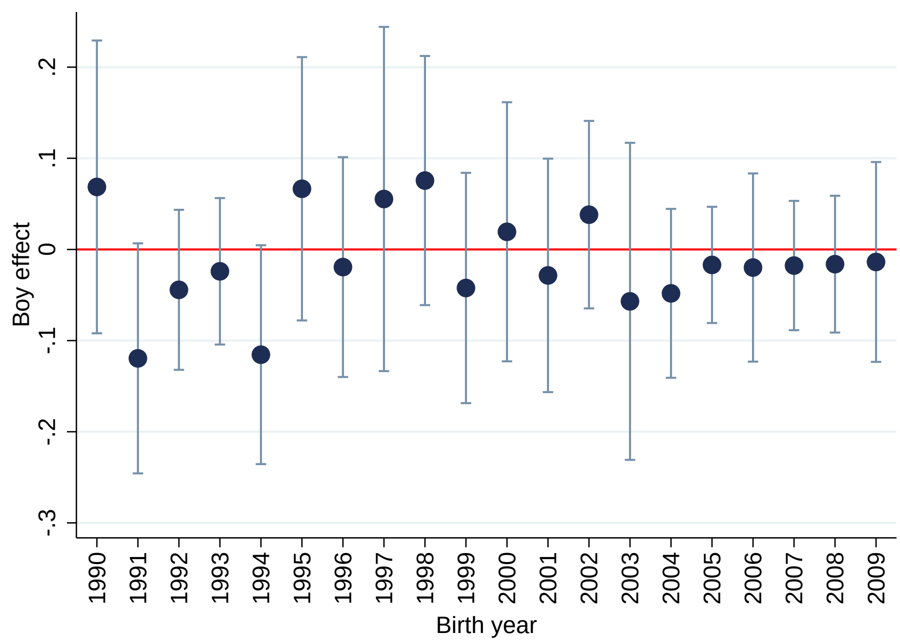


Figure 5: Breastfeeding Outcomes, by First Child's Gender

Notes. Estimation using NSF 1994–2009 (6 triennial surveys). Panel A shows censored regression results and panel B shows OLS estimates. Each dot represents the coefficient estimate on *Boy* $\times$ birth year dummy. Vertical spikes around the point estimate represent the 95% confidence interval constructed using robust standard errors. Control variables include parents' age at birth of their child, age squared, and dummies for parents' education (less than high school, high school, college or more), child's birth year (1990–2009), and birth month (January–December). Missing values in covariates are imputed with mean values and dummies for missing observations are also controlled.

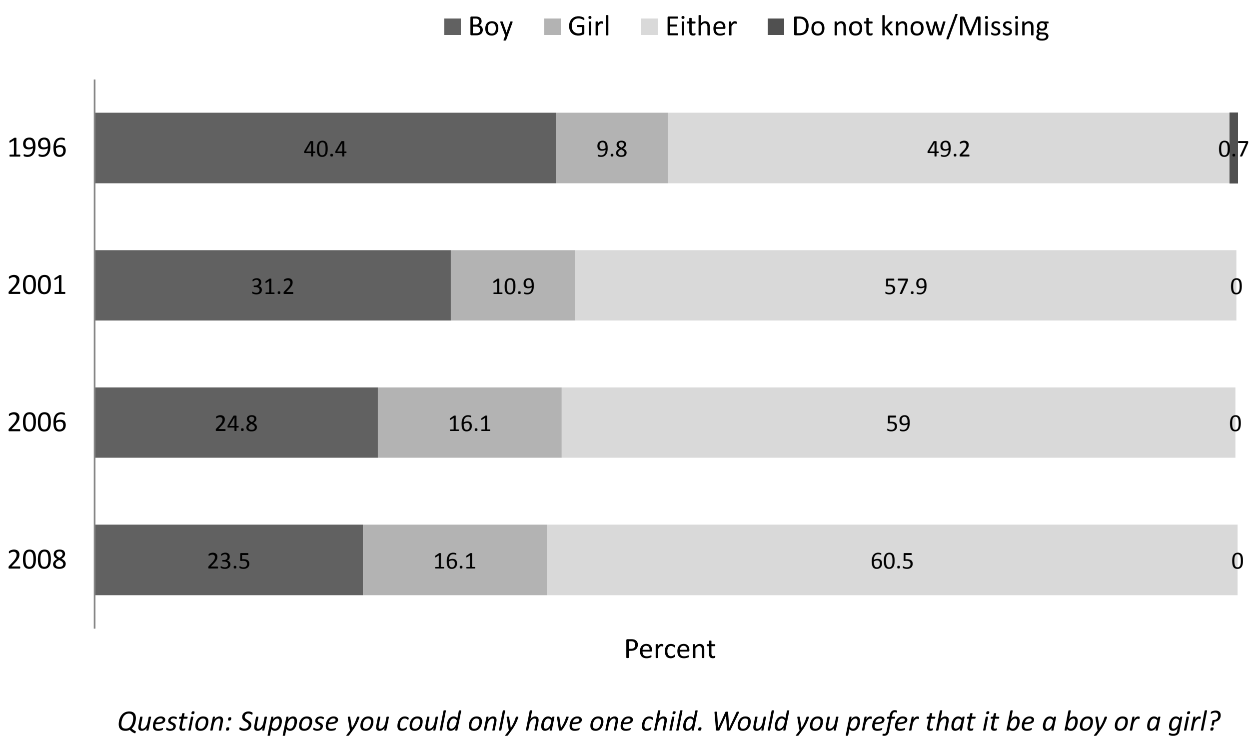


Figure 6: Preference for Boys versus Girls

Source. Korean Value Survey.

Table 1: Parent Characteristics at First Childbirth, by First Child's Gender

	First child's gender		
	Boy	**Girl**	**Difference in means**
Mother's age	28.15	28.18	-0.03***
	[3.906]	[3.930]	(0.004)
Father's age	30.94	30.98	-0.03***
	[4.226]	[4.244]	(0.004)
Mother college graduate	0.55	0.55	0.00
	[0.497]	[0.497]	(0.000)
Father college graduate	0.60	0.60	0.00
	[0.490]	[0.490]	(0.000)
Mother high school graduate	0.42	0.42	0.00
	[0.494]	[0.494]	(0.000)
Father high school graduate	0.37	0.37	0.00
	[0.482]	[0.482]	(0.000)
Out-of-wedlock	0.02	0.02	0.00
	[0.128]	[0.129]	(0.000)

Notes. Mean and standard deviations of parental characteristics at first childbirth using KOSIS Vital Statistics 1997–2012. Column (3) reports results from running separate linear regressions where each characteristic is regressed on a dummy equal to 1 if child is male. Sample size varies by variable but a total of 4,044,245 households (2,075,979 parents of boys and 1,968,266 parents of girls) have non-missing observations for all characteristics. Standard deviations in brackets. Robust standard errors in parentheses. * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$

Table 2: Number of Children, by First Child's Gender

Dependent variable:	**Number of children**	**2+ children**	**3+ children**
First child: boy	-0.16***	-0.04***	-0.12***
	(0.018)	(0.011)	(0.012)
Second child: boy			-0.15***
			(0.013)
Controls	Yes	Yes	Yes
N	4,882	4,882	3,750
Dependent variable mean	1.94	0.77	0.20

Notes. OLS estimation using KLIPS 1998–2014. Control variables include parent's age of the last wave, age squared, dummies for parent's education (less than high school, high school, college or more), year, and region. Missing values in covariates are imputed with mean values and dummies for missing observations are also controlled. Robust standard errors in parentheses. * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$

Table 3: Mother's Labor Supply, by First Child's Gender

Dependent variable:	**Mother's working hours after first childbirth**		**Mother's employment after first childbirth**	
	2 years	5 years	2 years	5 years
A. Effects for all birth years				
First child: boy	-1.44	-2.05*	0.01	-0.02
	(1.293)	(1.242)	(0.034)	(0.035)
B. Effects by birth year				
First child: boy × 1(1999 ≤ birth year ≤ 2002)	-6.03*	-7.37**	-0.07	-0.10
	(3.580)	(3.223)	(0.084)	(0.081)
First child: boy × 1(2003 ≤ birth year ≤ 2006)	-1.36	-1.21	-0.02	-0.03
	(2.456)	(2.363)	(0.065)	(0.065)
First child: boy × 1(2007 ≤ birth year ≤ 2014)	0.07	-0.39	0.06	0.02
	(1.702)	(1.662)	(0.046)	(0.047)
F-statistic testing Boy effects	1.16	1.84	1.02	0.80
identical across years	[0.32]	[0.16]	[0.36]	[0.45]
Controls	Yes	Yes	Yes	Yes
N	649	662	696	710
Dependent variable mean	13.78	13.72	0.44	0.51

Notes. OLS estimation using KLIPS 1998–2014. One observation per mother. Mother's employment after first childbirth is a dummy equal to 1 if a woman worked anytime between 0–2 years (col. 1) or 0–5 years (col. 2) after first childbirth. Mother's working hours after first childbirth is the mean of weekly working hours between 0–2 years (col. 3) or 0–5 years (col. 4) after first childbirth. Working hours equal zero when not employed. Control variables include mother's employment or average working hours one year before first childbirth, dummies for the three birth year intervals, parent's age at the time of first childbirth, age squared, dummies for parent's education (less than high school, high school, college or more), year, and region. Missing values are imputed with mean values and dummies for missing observations are controlled. Robust standard errors in parentheses. * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$

Table 4: Child's Housework Time and Participation, by Child's Gender

Dependent variable:	**Housework time**		**Any housework?**	
	(1)	(2)	(3)	(4)
A. Effects for all years				
Boy	-0.91***	-0.91***	-0.16***	-0.16***
	(0.037)	(0.036)	(0.007)	(0.007)
Controls	No	Yes	No	Yes
N	29,455	29,455	29,455	29,455
Dependent variable mean	0.98	0.98	0.28	0.28
B. Effects by year				
Boy $\times$ 1(year = 1999)	-1.03***	-1.02***	-0.19***	-0.18***
	(0.049)	(0.048)	(0.009)	(0.009)
Boy $\times$ 1(year = 2004)	-0.70***	-0.72***	-0.12***	-0.13***
	(0.055)	(0.054)	(0.010)	(0.010)
Boy $\times$ 1(year = 2009)	-0.49***	-0.48***	-0.09***	-0.09***
	(0.061)	(0.061)	(0.013)	(0.012)
F-statistic testing Boy effects identical across years	24.90	25.12	21.81	21.22
	[0.00]	[0.00]	[0.00]	[0.00]
Controls	No	Yes	No	Yes
N	29,455	29,455	29,455	29,455
Dependent variable mean	0.98	0.98	0.28	0.28

Notes. Data are from KTUS 1999, 2004, 2009. Robust standard errors in parentheses. p-values in brackets. All estimations use survey weights. Sample consists of two entries (time diaries) each of respondents between age 10 and 18. Control variables include child's age, age squared, mother's labor force participation, parent's age, age squared, dummies for parent's education (less than high school, high school, college or more), and dummies for region. Missing values in covariates are imputed with mean values and dummies for missing observations are also controlled. * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$

Table 5: Private Education Spending, by Child's Gender

Dependent variable:	**Monthly expenditures on private education**		**Any private education spending?**	
	(1)	(2)	(3)	(4)
A. First-borns (KELS 2005–2007)				
Boy	29.70***	23.34***	0.04***	0.03***
	(8.358)	(7.761)	(0.012)	(0.012)
Controls	No	Yes	No	Yes
N	8,820	8,820	8,820	8,820
Dependent variable mean	262.25	262.25	0.69	0.69
B. All birth orders (KELS 2005–2007)				
Boy	22.04***	22.81***	0.02**	0.02***
	(6.041)	(5.727)	(0.009)	(0.009)
Controls	No	Yes	No	Yes
N	18,086	18,086	18,086	18,086
Dependent variable mean	236.93	236.93	0.65	0.65
C. All birth orders (Middle school students from PEES 2007)				
Boy	13.60***	18.21***	0.02**	0.02***
	(3.429)	(3.118)	(0.007)	(0.007)
Controls	No	Yes	No	Yes
N	16,918	16,918	16,918	16,918
Dependent variable mean	198.95	198.95	0.75	0.75

Notes. Data are from KELS 2005–2007 and PEES 2007. The estimation sample includes single and two-parent families. Monthly expenditures on private education are in thousands of 2010 South Korean wons (KRWs). 1000 KRWs are worth approximately 1 USD. Private education denotes private out-of-school education of Korean, Math, and English. In panels A and B, control variables include parents' age, age squared, and dummies for parents' education (less than high school, high school, college or more), year, and urban rural classification (Seoul, large cities, small cities, rural). Panel C uses the sample of middle school students from PEES 2007. Control variables in panel C include dummies for parents' age (20-39, 40-49, 50+), parents' education (less than high school, high school, college or more), and urban rural classification (Seoul, large cities, small cities, rural). Missing values in covariates are imputed with mean values and dummies for missing observations are also controlled. Robust standard errors in parentheses. Standard errors are clustered at individual level in panel A and B. Estimation uses survey weights in panel C. * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$

Table 6: Private Education Spending at All School Levels, by Child's Gender

Dependent variable:	**Private education spending on**			
	Academic subjects		**Non-academic subjects**	
	(1)	(2)	(3)	(4)
A. Effects for all years				
Boy	1.09	2.52***	-6.09***	-5.04***
	(0.755)	(0.711)	(0.341)	(0.320)
Controls	No	Yes	No	Yes
N	481,114	481,114	481,114	481,114
Dependent variable mean	196.41	196.41	45.10	45.10
B. Effects by year				
Boy × 1(year = 2007)	8.32***	10.03***	-6.47***	-4.75***
	(1.829)	(1.704)	(0.811)	(0.763)
Boy × 1(year = 2008)	4.72**	6.03***	-7.98***	-7.06***
	(1.855)	(1.737)	(0.773)	(0.722)
Boy × 1(year = 2009)	4.68**	4.47***	-6.22***	-4.92***
	(1.844)	(1.729)	(0.873)	(0.820)
Boy × 1(year = 2010)	-5.48***	-2.91*	-7.39***	-6.21***
	(1.813)	(1.719)	(0.876)	(0.821)
Boy × 1(year = 2011)	-6.53***	-4.11**	-6.03***	-4.48***
	(1.804)	(1.720)	(0.870)	(0.817)
Boy × 1(year = 2012)	-0.51	0.75	-2.26***	-2.59***
	(1.932)	(1.846)	(0.798)	(0.763)
F-statistic testing Boy effects identical across years	10.96	10.10	6.30	4.16
	[0.00]	[0.00]	[0.00]	[0.00]
Controls	No	Yes	No	Yes
N	481,114	481,114	481,114	481,114
Dependent variable mean	196.41	196.41	45.10	45.10

Notes. Data are from PEES 2007–2012. Robust standard errors in parentheses. p-values in brackets. All estimations use survey weights. The estimation sample includes students at school levels from single and two-parent families. Estimated effects and dependent variable averages are in thousands of 2010 South Korean wons (KRWs) per month. 1000 KRWs are worth approximately 1 USD. Control variables include dummies for parents' age (20-39, 40-49, 50+), parents' education (less than high school, high school, college or more), child's school level (elementary, middle, high), year, and urban rural classification (Seoul, large cities, small cities, rural). In panel B, control variables are the ones included in panel A. In all columns of panel B, intercepts are allowed to vary across years. Missing values in covariates are imputed with mean values and dummies for missing observations are also controlled. * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$

Table 7: Expected Years of Education, by Child's Gender

Dependent variable:	**Expected years of education**			
	First borns		All birth orders	
	(1)	(2)	(3)	(4)
Boy	0.29***	0.24***	0.30***	0.30***
	(0.064)	(0.059)	(0.044)	(0.041)
Controls	No	Yes	No	Yes
N	8,986	8,986	18,541	18,541
Dependent variable mean	17.32	17.32	17.18	17.18

Notes. Data are from KELS 2005–2007. The estimation sample includes single and two-parent families. Control variables include parents' age, age squared, and dummies for parents' education (less than high school, high school, college or more), year, and urban rural classification (Seoul, large cities, small cities, rural). Missing values in covariates are imputed with mean values and dummies for missing observations are also controlled. Standard errors clustered at individual level in parentheses. * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$

Table 8: Breastfeeding, by First Child's Gender

Dependent variable:	**Months breastfed**			**Ever breastfed?**		
	Censored regression			Linear Probability Model		
	(1)	(2)	(3)	(4)	(5)	(6)
Boy	-0.31*	-0.20	-0.24	-0.02	-0.02	-0.03*
	(0.176)	(0.169)	(0.183)	(0.014)	(0.013)	(0.014)
Low birthweight			-2.26***			-0.22***
			(0.502)			(0.045)
Controls	No	Yes	Yes	No	Yes	Yes
N	3,811	3,811	3,311	3,811	3,811	3,311
Dependent variable mean	3.49	3.49	3.60	0.77	0.77	0.79

Notes. Data are from NSF 1994–2009 (6 triennial surveys). The estimation sample includes single and two-parent families. Control variables include parents' age at birth of their child, age squared, and dummies for parents' education (less than high school, high school, college or more), child's birth year (1990–2009), and birth month (January–December). Missing values in covariates are imputed with mean values and dummies for missing observations are also controlled. Robust standard errors in parentheses. * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$

Table 9: Educational Inputs, by First Child's Gender and Standardized Test Score from Previous Year

Dependent variable:	**Monthly expenditures on private education**		**Any private education spending?**		**Expected years of education**	
	(1)	(2)	(3)	(4)	(5)	(6)
Boy	39.98***	34.67***	0.05***	0.07***	0.36***	0.35***
	(10.120)	(13.444)	(0.013)	(0.018)	(0.059)	(0.078)
Test score from last year	52.73***	44.02***	0.10***	0.10***	0.63***	0.58***
	(7.318)	(7.267)	(0.010)	(0.011)	(0.043)	(0.042)
(Test score from last year)^2		22.39***		-0.01		0.13***
		(7.052)		(0.009)		(0.039)
Boy $\times$ Test score from last year	6.22	16.63	-0.01	-0.02	0.01	0.06
	(11.155)	(11.201)	(0.012)	(0.013)	(0.058)	(0.057)
Boy $\times$ (Test score from last year)^2		1.01		-0.02		-0.01
		(11.415)		(0.012)		(0.053)
F-statistic testing coefficients on terms interacted with Boy = 0		1.13		2.52		0.61
		[0.32]		[0.08]		[0.55]
Controls	Yes	Yes	Yes	Yes	Yes	Yes
N	5,836	5,836	5,836	5,836	5,872	5,872
Dependent variable mean	299.65	299.65	0.70	0.70	17.19	17.19

Notes. Data are from KELS 2006–2007. The estimation sample includes single and two-parent families. Control variables include parents' age, age squared, and dummies for parents' education (less than high school, high school, college or more), year, and urban rural classification (Seoul, large cities, small cities, rural). Missing values in covariates are imputed with mean values and dummies for missing observations are also controlled. Standard errors clustered at individual level in parentheses. * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$

Appendix

A Data Description

A.1 Korean Labor and Income Panel Survey (KLIPS)

KLIPS is a longitudinal study of a representative sample of Korean households and individuals living in urban areas. It was initiated by the Korea Labor Institute in 1998 and the most recent, 17th wave, was conducted in 2014. Household members aged 15 or older respond to the person survey. The 1st wave includes 5,000 households and the last, 17th wave includes 4,741 households. In 2009 (12th wave), 1,415 households were newly added to the original sample.

Using household ID, person ID, and relation to household head, we identify the first-born child and the child's mother. Father is defined based on spousal relationship with regards to the mother. Households with only grown-ups (no child or youngest child older than 18), with adopted children, or with a change in mother's ID anytime during the sample period (from divorce, remarriage, etc.) are excluded. Year relative to first childbirth is extrapolated using the first child's age.

A.2 Korean Time Use Survey (KTUS)

Korea's Time Use Survey is collected by the National Statistical Office every five years beginning from 1999 and provides nationally representative data of how Koreans spend their time. Household members older than age 10 record the time they spend on various activities (both main and secondary) during two survey dates.

The dependent variable, housework time, is constructed from respondents' time spent on "Household Activities" as the main activity. Household activities corresponds to the same category in the Bureau of Labor Statistics time use data and includes activities such as housework, food and drink preparation and clean-up, interior maintenance, exterior maintenance, vehicle maintenance, and household management. It does not include time spent on caring for children or other family members. Time diaries are reported in minutes. For convenience, we convert measures to weekly hours by multiplying by 7 (days) and dividing by 60 (minutes).

The data does not provide information on child's birth order, total number of children in the household, or which child a parent carried out an activity with. Thus we study child respondents' (ages 10–18) time diaries directly. Household characteristics, including parent's demographics, are matched to child entries using household ID, person ID, and relation to household head variables. Child respondents who could not be matched with parent ID were excluded from the sample. Estimation uses survey weights, which takes into account respondent's sex, age, and week(end) day of survey.

A.3 Korean Education Longitudinal Study (KELS)

The KELS provides data on learning experiences and transitions to work of a nationally representative sample of seventh-graders who were first surveyed in 2005. Similar to the

National Education Longitudinal Study of 1988 (NELS:1988) in the US, data are collected from students, parents, teachers and school principals every year. We pool the first three waves of the KELS and construct a sample of middle school students born in 1992 or 1993. We do not use later years of data because attrition rate becomes higher once sample members enter high school in the fourth wave. The sample members are supposed to be followed up every year until they become thirty years old.

A.4 Private Education Expenditures Survey (PEES)

The PEES is a nationwide survey specifically designed to collect data on the demand for private out-of-school education. Parents of students attending elementary, middle, or high schools are surveyed twice a year since 2007. We use the first six years of data from 2007 to 2012. The PEES contains detailed information about private education spending not just on academic subjects, but also on non-academic subjects, such as art, music, sports, and hobby activities. Academic subjects covered by the PEES not only include Korean, Math, and English, but also Science, Social Science, other foreign languages, computer programming, and critical writing. One limitation of the PEES data is that we cannot distinguish the birth order of children.

A.5 National Survey on Fertility, Family Health and Welfare in Korea (NSF)

We use repeated cross-sectional data from the National Survey on Fertility, Family Health and Welfare (NSF). The NSF is designed to provide information on marriage, pregnancy, contraception, childbirth, child rearing, work-life balance, support for parents, and family health. The first survey started in the 1960s but data are available from 1991, since when the NSF has been conducted every three years.

Table A1: Descriptive Statistics

	KLIPS 1998–2014	**KTUS 1999–2009**	**KELS 2005–2007**	**PEES 2007–2012**	**NSF 1994–2009**
Boy	0.52	0.54	0.52	0.53	0.54
	[0.50]	[0.50]	[0.50]	[0.50]	[0.50]
Child's age	0.00	14.05	13.82	–	1.37
	[0.00]	[2.57]	[0.90]		[0.93]
Primary school age	0.00	0.32	0.00	0.47	0.00
	[0.00]	[0.47]	[0.00]	[0.50]	[0.00]
Middle school age	0.00	0.33	1.00	0.27	0.00
	[0.00]	[0.47]	[0.00]	[0.44]	[0.00]
High school age	0.00	0.34	0.00	0.26	0.00
	[0.00]	[0.47]	[0.00]	[0.44]	[0.00]
Mother's age	29.22	40.85	40.50	–	26.92
	[3.61]	[4.06]	[3.47]		[3.57]
20 to 39	0.99	0.37	0.28	0.30	0.99
	[0.12]	[0.48]	[0.43]	[0.46]	[0.12]
40 to 49	0.01	0.57	0.69	0.62	0.00
	[0.10]	[0.50]	[0.45]	[0.49]	[0.06]
50 or over	0.00	0.03	0.03	0.05	0.00
	[0.00]	[0.16]	[0.17]	[0.22]	[0.00]
Father's age	31.78	44.15	43.72	–	29.98
	[3.82]	[4.27]	[3.37]		[4.22]
20 to 39	0.97	0.12	0.06	0.13	0.97
	[0.18]	[0.33]	[0.22]	[0.34]	[0.16]
40 to 49	0.03	0.66	0.83	0.69	0.02
	[0.17]	[0.48]	[0.35]	[0.46]	[0.14]
50 or over	0.00	0.10	0.11	0.14	0.00
	[0.05]	[0.30]	[0.28]	[0.35]	[0.06]
Mother's education					
Less than high school	0.02	0.35	0.09	0.06	0.06
	[0.15]	[0.47]	[0.28]	[0.23]	[0.24]
High school graduate	0.35	0.52	0.63	0.55	0.52
	[0.48]	[0.49]	[0.47]	[0.50]	[0.50]
College or more	0.63	0.13	0.27	0.37	0.41
	[0.48]	[0.33]	[0.44]	[0.48]	[0.49]
Father's education					
Less than high school	0.03	0.23	0.09	0.05	0.05
	[0.16]	[0.40]	[0.28]	[0.22]	[0.22]
High school graduate	0.28	0.49	0.47	0.42	0.44
	[0.45]	[0.47]	[0.49]	[0.49]	[0.50]
College or more	0.69	0.27	0.44	0.49	0.50
	[0.46]	[0.42]	[0.49]	[0.50]	[0.50]
N	1,208	29,455	18,086	481,114	3,811

Notes. Means and standard deviations at year of first childbirth for KLIPS and including all birth orders for other samples. Standard deviations in brackets. Missing values in baseline characteristics are imputed with mean values. Survey weights are used for KTUS and PEES sample.

Table A2: Parent's Characteristics One Year Prior to First Childbirth, by First Child's Gender (KLIPS)

	First child's gender		
	Boy	**Girl**	**Difference in means**
Mother's age	28.97	29.18	-0.21
	[3.45]	[3.16]	(0.249)
Father's age	32.82	32.66	0.16
	[5.58]	[5.19]	(0.406)
Mother less than HS graduate	0.02	0.00	0.02**
	[0.15]	[0.05]	(0.009)
Mother high school graduate	0.32	0.32	-0.00
	[0.47]	[0.47]	(0.035)
Mother college graduate	0.65	0.67	-0.02
	[0.48]	[0.47]	(0.036)
Father less than HS graduate	0.01	0.01	-0.00
	[0.11]	[0.12]	(0.009)
Father high school graduate	0.22	0.23	-0.01
	[0.42]	[0.42]	(0.031)
Father college graduate	0.66	0.67	-0.01
	[0.47]	[0.47]	(0.035)
Mother's employment	0.55	0.56	-0.01
	[0.50]	[0.50]	(0.037)
Mother's working hours	23.28	23.37	-0.09
	[22.65]	[22.35]	(1.811)
N	375	336	

Notes. Mean and standard deviations using KLIPS 1998–2014. Observations from families one year before first childbirth, by the gender of first child realized later in the panel. Missing values are imputed with mean values. Standard deviations in brackets. Standard errors in parentheses. * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$

Table A3: Parent's Characteristics, by Child's Gender (KTUS)

	First child's gender		
	Boy	**Girl**	**Difference in means**
Mother's age	40.88	40.81	0.07
	[4.15]	[3.95]	(0.060)
Father's age	44.19	44.10	0.09
	[4.34]	[4.20]	(0.064)
Mother less than HS graduate	0.34	0.35	-0.00
	[0.47]	[0.47]	(0.007)
Mother high school graduate	0.53	0.52	0.01*
	[0.49]	[0.49]	(0.007)
Mother college graduate	0.13	0.14	-0.01*
	[0.32]	[0.33]	(0.005)
Father less than graduate	0.23	0.23	-0.00
	[0.40]	[0.40]	(0.006)
Father high school graduate	0.50	0.49	0.01
	[0.47]	[0.47]	(0.007)
Father college graduate	0.27	0.28	-0.01
	[0.42]	[0.42]	(0.006)
N	15,314	14,141	

Notes. Mean and standard deviations using KTUS 1999–2009. Column (3) reports results from running separate linear regressions where each characteristic is regressed on a dummy equal to 1 if child is male. Missing values are imputed with mean values. Standard deviations in brackets. Standard errors in parentheses. * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$

Table A4: Parent's Characteristics, by First Child's Gender (KELS)

	First child's gender		
	Boy	**Girl**	**Difference in means**
Mother's age	40.56	40.45	0.11
	[3.28]	[3.64]	(0.117)
Father's age	43.66	43.78	-0.12
	[3.12]	[3.59]	(0.112)
Mother less than high school	0.06	0.06	-0.00
	[0.22]	[0.22]	(0.008)
Mother high school graduate	0.62	0.65	-0.03*
	[0.48]	[0.47]	(0.017)
Mother college graduate	0.33	0.30	0.03*
	[0.46]	[0.45]	(0.016)
Father less than high school	0.06	0.05	0.01
	[0.23]	[0.22]	(0.008)
Father high school graduate	0.43	0.48	-0.05***
	[0.49]	[0.49]	(0.017)
Father college graduate	0.50	0.46	0.04**
	[0.49]	[0.49]	(0.017)
N	4,344	4,476	

Notes. Mean and standard deviations using KELS 2005–2007. Column (3) reports results from running separate linear regressions where each characteristic is regressed on a dummy equal to 1 if first child is male. Missing values are imputed with mean values. Standard deviations in brackets. Standard errors clustered at individual level in parentheses. * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$

Table A5: Parent's Characteristics, by Child's Gender (PEES)

	First child's gender		
	Boy	**Girl**	**Difference in means**
Mother's age 20 to 39	0.22	0.23	-0.01***
	[0.41]	[0.42]	(0.001)
Mother's age 40 to 49	0.68	0.68	-0.00**
	[0.47]	[0.47]	(0.001)
Mother's age 50 or over	0.07	0.06	0.01***
	[0.26]	[0.24]	(0.001)
Father's age 20 to 39	0.09	0.10	-0.01***
	[0.29]	[0.30]	(0.001)
Father's age 40 to 49	0.67	0.68	-0.01***
	[0.47]	[0.47]	(0.001)
Father's age 50 or over	0.20	0.18	0.02***
	[0.40]	[0.39]	(0.001)
Mother less than HS graduate	0.06	0.06	0.00***
	[0.25]	[0.24]	(0.001)
Mother high school graduate	0.54	0.57	-0.00*
	[0.50]	[0.50]	(0.001)
Mother college graduate	0.35	0.36	-0.01***
	[0.48]	[0.48]	(0.001)
Father less than HS graduate	0.06	0.06	0.01***
	[0.24]	[0.23]	(0.001)
Father high school graduate	0.43	0.42	0.00***
	[0.49]	[0.49]	(0.001)
Father college graduate	0.47	0.48	-0.01***
	[0.50]	[0.50]	(0.001)
Child attending primary school	0.29	0.29	-0.00**
	[0.45]	[0.46]	(0.001)
Child attending middle school	0.26	0.26	0.00***
	[0.44]	[0.44]	(0.001)
Child attending high school	0.45	0.45	-0.00
	[0.50]	[0.50]	(0.001)
N	253,436	227,678	

Notes. Mean and standard deviations using PEES 2007–2012. Column (3) reports results from running separate linear regressions where each characteristic is regressed on a dummy equal to 1 if child is male. Missing values are imputed with mean values. Standard deviations in brackets. Standard errors in parentheses. * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$

Table A6: Parent's Characteristics, by First Child's Gender (NSF)

	First child's gender		
	Boy	**Girl**	**Difference in means**
Mother's age	26.91	26.92	-0.01
	[3.58]	[3.57]	(0.116)
Father's age	29.98	29.98	0.00
	[4.12]	[4.33]	(0.138)
Mother less than high school	0.06	0.07	-0.01
	[0.23]	[0.25]	(0.008)
Mother high school graduate	0.53	0.51	0.02
	[0.50]	[0.50]	(0.016)
Mother college graduate	0.41	0.42	-0.01
	[0.49]	[0.49]	(0.016)
Father less than high school	0.05	0.05	0.00
	[0.22]	[0.22]	(0.007)
Father high school graduate	0.43	0.45	-0.02
	[0.50]	[0.50]	(0.016)
Father college graduate	0.51	0.49	0.02
	[0.50]	[0.50]	(0.016)
Birthweight[a]	3.29	3.21	0.08***
	[0.44]	[0.42]	(0.015)
Low birthweight[a]	0.03	0.04	-0.00
	[0.17]	[0.19]	(0.006)
N	2,059	1,752	

Notes. Mean and standard deviations using NSF 1994–2009. Column (3) reports results from running separate linear regressions where each characteristic is regressed on a dummy equal to 1 if first child is male. Missing values are imputed with mean values. Standard deviations in brackets. Standard errors clustered at individual level in parentheses. * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$

[a] N = 1,798 for boys and N = 1,513 for girls. The sample does not include 1997 data where birthweight is unavailable.

Table A7: Private Education Spending at Primary School Level, by Child's Gender

Dependent variable:	**Private education spending on**			
	Academic subjects		**Non-academic subjects**	
	(1)	(2)	(3)	(4)
A. Effects for all years				
Boy	-1.39	-0.90	0.63	1.48***
	(1.079)	(1.033)	(0.556)	(0.530)
Controls	No	Yes	No	Yes
N	140,734	140,734	140,734	140,734
Dependent variable mean	172.67	172.67	70.39	70.39
B. Effects by year				
Boy × 1(year = 2007)	-2.19	-1.05	0.17	1.22
	(2.156)	(2.045)	(1.213)	(1.166)
Boy × 1(year = 2008)	-1.14	-1.38	-1.77	-1.18
	(2.295)	(2.175)	(1.096)	(1.043)
Boy × 1(year = 2009)	2.75	2.27	2.83**	3.55***
	(2.809)	(2.687)	(1.405)	(1.332)
Boy × 1(year = 2010)	-3.19	-2.33	1.06	2.36*
	(2.812)	(2.700)	(1.480)	(1.409)
Boy × 1(year = 2011)	-9.19***	-7.50***	-1.21	0.16
	(2.834)	(2.734)	(1.549)	(1.485)
Boy × 1(year = 2012)	4.44	4.73	2.98**	3.00**
	(3.005)	(2.934)	(1.443)	(1.396)
F-statistic testing Boy effects identical across years	2.81	2.25	2.28	2.23
	[0.02]	p0.05]	p0.04]	[0.05]
Controls	No	Yes	No	Yes
N	140,734	140,734	140,734	140,734
Dependent variable mean	172.67	172.67	70.39	70.39

Notes. Data are from PEES 2007–2012. Robust standard errors in parentheses. p-values in brackets. All estimations use survey weights. The estimation sample includes primary school students from single and two-parent families. Estimated effects and dependent variable averages are in thousands of 2010 South Korean wons (KRWs) per month. 1000 KRWs are worth approximately 1 USD. Control variables include dummies for parents' age (20-39, 40-49, 50+), parents' education (less than high school, high school, college or more), year, and urban rural classification (Seoul, large cities, small cities, rural). In all columns of panel B, intercepts are allowed to vary across years. Missing values in covariates are imputed with mean values and dummies for missing observations are also controlled. * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$

Table A8: Private Education Spending at Secondary School Level, by Child's Gender

Dependent variable:	**Private education spending on**			
	Academic subjects		**Non-academic subjects**	
	(1)	(2)	(3)	(4)
A. Effects for all years				
Boy	2.74***	5.99***	-11.40***	-10.91***
	(1.047)	(0.961)	(0.381)	(0.376)
Controls	No	Yes	No	Yes
N	340,380	340,380	340,380	340,380
Dependent variable mean	217.07	217.07	23.10	23.10
B. Effects by year				
Boy × 1(year = 2007)	18.31***	22.15***	-12.51***	-10.62***
	(2.923)	(2.676)	(0.979)	(0.972)
Boy × 1(year = 2008)	10.41***	14.71***	-14.18***	-12.83***
	(2.855)	(2.642)	(0.998)	(0.988)
Boy × 1(year = 2009)	5.02**	6.91***	-12.63***	-12.58***
	(2.429)	(2.184)	(1.002)	(0.993)
Boy × 1(year = 2010)	-8.29***	-3.36	-13.45***	-13.53***
	(2.358)	(2.149)	(0.939)	(0.929)
Boy × 1(year = 2011)	-5.58**	-1.65	-8.79***	-8.56***
	(2.313)	(2.141)	(0.850)	(0.844)
Boy × 1(year = 2012)	-3.48	-2.59	-6.86***	-7.14***
	(2.446)	(2.265)	(0.799)	(0.798)
F-statistic testing Boy effects	15.12	18.21	11.12	8.78
identical across years	0.00	0.00	0.00	0.00
Controls	No	Yes	No	Yes
N	340,380	340,380	340,380	340,380
Dependent variable mean	217.07	217.07	23.10	23.10

Notes. Data are from PEES 2007–2012. Robust standard errors in parentheses. p-values in brackets. All estimations use survey weights. The estimation sample includes secondary school students from single and two-parent families. Estimated effects and dependent variable averages are in thousands of 2010 South Korean wons (KRWs) per month. 1000 KRWs are worth approximately 1 USD. Control variables include dummies for parents' age (20-39, 40-49, 50+), parents' education (less than high school, high school, college or more), child's school level (middle, high), year, and urban rural classification (Seoul, large cities, small cities, rural). In all columns of panel B, intercepts are allowed to vary across years. Missing values in covariates are imputed with mean values and dummies for missing observations are also controlled. * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$

Table A9: Number of Months Breastfed among Those Breastfed, by First Child's Gender

Dependent variable:	**ln(Months breastfed)**			**Months breastfed**					
	Censored regression			Accelerated failure time			Proportional hazard		
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)
Boy	-0.15	-0.09	-0.09	-0.07	-0.03	-0.03	0.06	0.03	0.04
	(0.107)	(0.106)	(0.112)	(0.052)	(0.054)	(0.058)	(0.043)	(0.044)	(0.047)
Low birthweight			-0.74**			-0.49**			0.36**
			(0.366)			(0.209)			(0.157)
Controls	No	Yes	Yes	No	Yes	Yes	No	Yes	Yes
N	2,931	2,931	2,602	2,931	2,931	2,602	2,931	2,931	2,602
Dependent variable mean	0.31	0.31	0.33	4.54	4.54	4.58	4.54	4.54	4.58

Notes. Data are from NSF 1994–2009 (6 triennial surveys). The estimation sample includes single and two-parent families. Control variables include parents' age at birth of their child, age squared, and dummies for parents' education (less than high school, high school, college or more), child's birth year (1990–2009), and birth month (January–December). Missing values in covariates are imputed with mean values and dummies for missing observations are also controlled. Robust standard errors in parentheses. * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$

Table A10: Educational Inputs, by First Child's Gender and 6th Grade Achievement

Dependent variable:	**Monthly expenditures on private education**	**Any private education spending?**	**Expected years of education**
Boy	27.04	0.16**	0.85***
	(28.826)	(0.069)	(0.284)
1(6th grade achievement = middle)	1.44	0.13**	0.60***
	(23.266)	(0.053)	(0.182)
1(6th grade achievement = high)	25.93	0.18***	1.39***
	(23.691)	(0.053)	(0.193)
Boy $\times$ 1(6th grade achievement = middle)	-6.23	-0.12	-0.60**
	(30.419)	(0.074)	(0.305)
Boy $\times$ 1(6th grade achievement = high)	-6.19	-0.10	-0.46
	(30.819)	(0.073)	(0.311)
F-statistic testing coefficients on terms interacted with Boy = 0	0.02	1.31	2.03
	[0.98]	[0.27]	[0.13]
Controls	Yes	Yes	Yes
N	2,915	2,915	3,037
Dependent variable mean	190.01	0.68	17.59

Notes. Data are from KELS 2005. The estimation sample includes single and two-parent families. Control variables include parents' age, age squared, and dummies for parents' education (less than high school, high school, college or more), year, and urban rural classification (Seoul, large cities, small cities, rural). Missing values in covariates are imputed with mean values and dummies for missing observations are also controlled. Standard errors clustered at individual level in parentheses. * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$