More Missing Women, Fewer Dying Girls: The Impact of Sex-Selective Abortion on Sex at Birth and Relative Female Mortality in Taiwan

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Abstract

This study finds that the introduction of sex-selective abortion in Taiwan due to the legalization of abortion when pre-natal sex-detection technology was already available increased the fraction of males born at higher parities and changed the composition of mothers choosing to give birth. Controlling for compositional changes, we find that access to sex-selective abortion reduced relative neonatal female mortality rates for higher parity births.

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1 Introduction

The rapid rise in the fraction of males at birth during the 1980s and 1990s in countries with son preference has motivated governments such as those in China, India and South Korea to ban pre-natal sex selection, which is made possible by the combined use of technologies for abortion and pre-natal sex detection (e.g., ultrasound).[1] Such bans have received much attention and support from the international policy community.[2] There is, however, little discussion about the potential unintended consequences from such bans. Specifically, one may be concerned that there are parents with strong son preference who will neglect unwanted daughters, and this may reduce the well-being of girls who are born.[3]

We aim to address this gap in the literature and policy discussion. We investigate whether increased access to sex-selective abortion increases male-biased sex ratio at birth and reduces relative female neonatal mortality rates, where mortality is interpreted as the extreme outcome of neglect. Our study is set in Taiwan during the 1980s and early 1990s, a context where parents are known to have son preference, the data are high quality and the legalization of abortion in 1985 provides plausibly exogenous variation in access to sex-selective abortion.

The analysis exploits three sources of variation. The first is time-variation in access to sex-selective abortion caused by the legalization of abortion at a time when prenatal sex-detection technologies were already available. This allows us to avoid the potential endogeneity of access to sex-selective abortion.[4] We also exploit the cross-sectional variation in the cost of having additional children. Thus, the second and third sources of variation come from birth parity and mother's age. The more children that parents already have and the older is the mother, the fewer chances they will have to realize their preferred number of sons. This increases the willingness of parents to sex select.

To examine sex at birth, we estimate the effect of the interaction between the legalization of abortion and birth parity and the interaction of the legalization with mother's age on the sex of children at birth.

[1]For example, see Greenhalgh and Li (1995) and Parikh (1993) for discussions of these policies.

[2]For example, the *United Nations Program of Action*, an organization established partly to stop sex-selective abortion, states its objectives as "to eliminate all forms of discrimination against a girl child.. which results in harmful and unethical practices regarding female infanticide and prenatal sex selection" ((1994), Article 4.15).

[3]The logic of our hypothesis is similar to that of Donohue and Levitt (2001) for explaining the reduced crime rates of children born to mothers after abortion is legalized in the United States. They argue that abortion reduced the number of unwanted children, which in turn, increased the quality of children on average. Goodkind (1996, 1999) discusses the potential tradeoffs for policymakers.

[4]For example, if wealthier or more educated parents have better access to technologies such as Ultrasound B or safe abortions conducted by medical professionals, then a negative correlation between the fraction of males at birth and relative female child mortality rates would reflect differences in access and parental characteristics rather than the causal impact of access to pre-natal sex selection.

To examine mortality rates of girls relative to boys, we estimate the effect of the triple interaction of the sex of a child, the legalization of abortion and birth parity and the triple interaction of the sex of a child, the legalization of abortion and mother's age on the probability of the child dying.[5] Since birth parity and mother's age are highly correlated (i.e., older mothers are more likely to have higher parity births), the baseline estimates control for birth parity, mother's age and their interactions in one equation. Our data are comprised of individual-level data from birth and death registries for all individuals born in Taiwan during 1980-92.

We find that the reform increased the fraction of males born at higher parities and to older mothers. At the same time, we find that the reform reduced relative female mortality rates for higher parity births during the first month of life, when infant mortality mostly occurs in Taiwan. The relative female mortality result is only statistically significant when we control for the triple interaction of the reform, the sex of the child and mother's age. Unlike the effects on sex at birth, there is no systematic pattern in the effect of the reform on relative female mortality across mother's age.

To interpret these results, we provide a simple conceptual framework, where parents jointly decide on the sex and the total number of children. The model predicts that a reduction in the cost of selective abortion will increase the fraction of males born, especially for parents who face a higher cost of having additional children (i.e., older mothers, or households that are already near their preferred number of children). Second, the introduction of sex-selective abortion can increase fertility for mothers with high son preference, who also face a high cost of child-rearing – for example, older mothers who face a high physical cost or mothers who already have their desired number of children who face a high psychological or financial cost. Thus, the predicted effects across birth parities and mother's age capture two offsetting forces. First, there is a negative force: the increase in pre-natal sex selection for higher parities and older mothers could reduce the number of unwanted daughters born at higher parities and to older mothers, which will reduce relative female mortality rates at higher parities and for older mothers. Second, there is a positive force: if sons born to older mothers or at higher parities after the reform are better cared for and suffer lower mortality rates, then the reform will reduce relative male mortality rates and increase relative female mortality rates at higher parities and for older mothers.

The results on fertility, which show that the reform reduced the number births at higher parities and

[5] Our estimates always control for birth parity and birth year fixed effects to control for time-invariant differences across parities and changes over time that affect all parities similarly. Similarly, the estimates for mortality will always control for the full set of double interaction terms.

increased it for older mothers, are consistent with this framework. Together with the results on sex at birth and relative female mortality, they imply the following. For relative female mortality rates across birth parities, the negative forces dominate the positive ones. For relative female mortality rates across mother's age, the opposing forces offset.

More importantly, the results show that conditional on compositional changes in mothers choosing to give birth, the reform reduced relative female mortality rates for higher parity births. In other words, for parents that would give birth to higher parity children regardless of access to sex-selective abortion, the reduced cost of sex-selection due to the reform reduced the number of unwanted daughters that were born, and thereby, reduced female mortality. For such parents, the reform explains 100% of the rise in the fraction of males at birth and over 50% of the reduction in relative female neonatal mortality rates.

This paper makes two novel contributions. This study is the first to provide rigorous empirical evidence that reducing the cost of sex selection in a society with strong son-preference can lead to better outcomes for girls that are born. Second, it is the first to show that the introduction of abortion can increase fertility for some mothers by reducing the cost of the child's "quality".

Our study adds to recent works that relate fertility decisions to post-natal investment in girls. For example, a study by Hu and Schlosser (2010) finds that sex-selective abortion increases sex imbalance, but also reduces the incidence of malnutrition for the girls that are born in India. Also, Jayachandran and Kuziemko (2011) finds that the desire to give birth to sons causes daughters to be breastfed less, which results in higher female child mortality. In finding that sex-selective abortion has heterogeneous treatment effects, our work is related to a recent study by Ebenstein (2011), which finds that raising the costs of fertility and sex selection has heterogenous effects across households of different wealth and education in China.[6]

The paper is organized as follows. Section 2 describes the background. Section 3 discusses the conceptual framework and empirical strategy. Section 4 describes the data. Section 5 presents the descriptive results. Section 6 presents the regression results. Section 7 concludes.

[6]Studies in demography such as Gupta (1987), Gupta and Bhat (1997), Gu and Roy (1995), Muhri and Preston (1991), Park and Cho (1995) and Pebley and Amin (1991) have described the changes over time in sex ratios at birth by birth parity in several East Asian countries. Arnold et al. (2002) have associated access to abortion with increasing sex ratios. Abrevaya (2009) find evidence of sex-selective abortion amongst higher parity births in the United States. A recent study by Chen et al. (2013) examines the correlation between regional adoption of ultrasound B and sex selection in China during the 1980s and 90s and find that sex-selective abortion increased sex ratios at birth is consistent with our results. Two recent studies by Almond et al. (2010) and Ebenstein et al. (2013) use the same variation in region-specific ultrasound adoption to study other pre- and post-natal outcomes and female empowerment in China.

2 Background

Until the mid-1980s, induced abortion was only legal in Taiwan for a small range of medical problems as outlined by the *Eugenics Protection Law*. The prohibition against abortion was very strict, and the punishment for abortions was severe. Legal scholars believe that the rate of illegal abortions prior to the reform was very low (Hung, 2004). During the mid-1980s, increasing demand for safe abortions as a method of family planning and a growing feminist movement pushed Taiwanese legislators to make abortion legal. For our study, the relevant reform occurred on January 1st, 1985, when it became legal for women to induce an abortion for social as well as medical reasons up to the 24th week of pregnancy. The service was inexpensive and safely conducted, although it was not covered by medical insurance (Henshaw, 1990). Based on interviews that the authors of this study conducted with physicians who performed abortions during the 1980s, the cost of an abortion was on average 1% of average household income at the time. By all accounts, the reform was binding. The percentage of women who have ever had abortions increased from 23% in 1985 to approximately 27% in 1992. Abortions increased between 1985 and 1992 for all age groups, with the largest increase in women ages 29-35. The fraction of women who have ever had abortions increases over time for all education levels.[7]

During the period of our study, the most widely used technology for pre-natal sex detection in Taiwan was Ultrasound B, which can reveal the sex of the fetus beginning in the 16th week of gestation.[8] Accuracy is greatly increased by the 20th week. The procedure for revealing the sex is not invasive and the results can be easily interpreted by a trained technician. Unlike in China and India, revealing the sex of the fetus has never been prohibited in Taiwan.

The legalization of abortion combined with the use of available sex-detection technologies enabled parents to use abortion as a method of sex selection. Hence, the legalization of abortion in Taiwan can be interpreted as a decrease in the cost of sex-selective abortion.[9] Using Ultrasound B during routine pre-natal care may have also increased the quality of pre-natal care more generally. We will revisit this issue later in the section on robustness.

[7] These statistics come from the KAPS data for women age 18-44 and are discussed in detail in an earlier version of this paper (Lin et al., 2008).

[8] Amniocentesis and chorionic villus were also available, but much less common. Our estimates do not distinguish between these methods.

[9] The effect of the reform on the demand for sex-selective abortions is reflected in both anecdotal evidence and the data on the number of Ultrasound B machines, which increased by almost an order of a magnitude between 1984 and 1989. See the earlier version of our paper for more details (Lin et al., 2008).

Our empirical strategy interprets 1985-1989 as the "post" reform period.[10] We interpret relative female mortality as the outcome of the differential treatment in the care of female infants relative to male infants.[11] We focus our attention on mortality in the first month of life, which has been shown to be when most infant mortality occurs in Taiwan (Ebenstein, 2007). This is when children are most vulnerable to the lack of care and when exerting less effort can, in marginal cases, result in mortality.

3 Conceptual Framework

3.1 Conceptual Framework

The goal of the empirical exercise is to use the legalization of abortion in Taiwan to investigate whether increased access to pre-natal sex selection can reduce relative female mortality. To understand the relationship between the mechanisms underlying this effect and to motivate our empirical strategy, we present a framework that allows a reduction in the cost of pre-natal sex selection to have heterogeneous effects on pre-natal sex selection, fertility and the effort that mothers exert towards caring for children, depending on a mother's preference for sons and her cost of rearing children. The model is presented formally in the Online Appendix.

For simplicity, we assume that mothers are the only relevant decision makers in the household. Mothers who face a high cost of having additional children and have strong son preference will opt to abort female fetuses when sex-selective abortion is easily available. For example, if a mother's preferred family size is two or three children, she will be more likely to abort a female fetus if she is pregnant with her second or third child than her first. Similarly, if a mother is older in age and faces a high physical cost in bearing or rearing a child, she will be more likely to abort a female fetus than a younger mother with the same son preference.

The decision to become pregnant can also be affected by the introduction of sex-selective abortion

[10] This could cause us to underestimate the effect of the reform for two reasons. First, the implementation of the reform was phased in during 1985 and 1986 and the government made no attempt to publicize the reform (Liu, 1995). The treatment effect that we estimate include those who are not yet affected by and those who do not yet know about the reform. Second, the law was applied to all contemporaneous pregnancies but abortion was allowed only up to the 24th week (six months). The first cohort that was exposed to sex-selective abortion were not born until four months into 1985. Thus, by using January 1st, 1985 as the cutoff date, the post-reform group will include some un-treated individuals. Note that we could alternatively define each yearly cohort to be those born from April 1st - March 31st. Doing so results in very similar estimates as those reported in this paper. These alternative estimates are available upon request.

[11] For example, Banister (2004) finds no evidence of female infanticide in Taiwan, and Jayachandran and Kuziemko (2011) finds that small differences in treatment (e.g., the duration of breastfeeding) can lead to significant differential mortality rates in young children in India.

because mothers with very high son preference, who also face a very high cost of child-rearing will be more willing to become pregnant if they can increase their chances of having a son. For example, consider older mothers for whom it is physically very costly to bear children, but who strongly desire a son, or mothers who have already had their desired number of total children for whom it is psychologically or financially costly to rear additional children.[12] The introduction of sex-selective abortion will allow them to abort female fetuses in the second trimester and avoid the last few months of pregnancy and the act of giving birth. By reducing the cost of sex selection, the reform will increase fertility for women might otherwise have chosen to not try to have a son. The legalization of non-medical abortion thus *increases* the option value of pregnancy for mothers with high son preference and who face a high cost of bearing or rearing children.

Finally, access to abortion will affect the amount of effort mothers exert towards caring for female children in aggregate because only mothers who value girls sufficiently will continue to bear girls after the reform. In the extreme, low effort will result in child mortality. Thus, the reform can reduce the aggregate mortality rates of girls that are born relative to those of boys that are born.

3.2 Empirical Strategy

Our empirical strategy exploits both time variation in the cost of pre-natal sex selection caused by the reform, and cross-sectional variation from two proxies for the cost of child-rearing. The first and main proxy is birth parity, which follows from the widely held belief that most mothers (parents) have a preferred family size when they begin having children and the marginal physical, logistical or psychological "cost" of rearing a child increases as a mother approaches or exceeds her preferred family size. The second proxy is a mother's age, which is motivated by the medical evidence that it is often physically more difficult for older mothers to endure the finals months of pregnancy, the process of giving birth, or the sleep deprivation that occurs with breastfeeding.[13]

The main empirical analysis focuses on the interaction of the time variation from the reform and the cross-sectional variation from birth parity or mother's age. The conceptual framework makes unambiguous predictions for the effect of the reform on sex ratio at birth across birth parities and mother's age. To the

[12]For brevity, the formal model presented in the Online Appendix only considers the effect of the reform on fertility for older mothers who face a higher physical cost of bearing children.

[13]For example, women over the age of 35 are significantly more likely to experience pregnancy-related deaths, are more likely to experience pregnancy-related complications such as gestational diabetes, high blood pressure, placental problems, premature birth and stillbirths, and the costs for rearing a child after infancy may also be higher for older mothers. See the earlier version of this paper for references (Lin et al., 2008).

extent that these two variables proxy for the cost of bearing/rearing another child, the model predicts that lowering the cost of pre-natal sex selection will increase the fraction of boys that are born, and the increase will be larger for higher parity births and older mothers. The older the mother and the higher the birth parity, the costlier it will be for parents to try for a son again, and the more likely they are to prenatally sex select. That older mothers or parents who have already reached their desired level of fertility with strong son preferences may select to bear children because they can choose the sex prenatally reinforces the positive interaction effects between the reform and birth parity and the reform and mother's age on sex at birth.

The predictions for the effects of the reform on relative female mortality are more nuanced. The predicted effects across birth parities and mother's age capture two offsetting forces. On the one hand, the increase in pre-natal sex selection for higher parities and older mothers could reduce the number of unwanted daughters born at higher parities and to older mothers, which will reduce relative female mortality rates at higher parities and for older mothers. On the other hand, if older mothers having boys after the reform value their sons more than those having boys prior to the reform and exert more effort towards the care of sons, then relative male mortality rates will decline, and relative female mortality rates will, by definition, increase. Similarly, if the reform induces parents that have already reached their desired level of fertility to try for a son and such parents take better care of their sons, relative male mortality rates could decline for higher parity births, which will increase relative female mortality rates. Since the descriptive statistics presented in the next section shows that, in practice, the reform only increased fertility for older mothers, we expect that the negative effects of the reform on relative female mortality rates are likely to be monotonically increasing in birth parity, but are ambiguous across mother's age.

Note that the purpose of our study is not to separately identify the influences of birth parity and mother's age on the impact of the reform. As we have discussed in our conceptual framework, both are proxies of similar forces. However, since the strength of these forces may differ across birth parities and mother's age and since birth parity and mother's age are highly correlated because higher parity children are born when mothers are older, our baseline regressions will always control for the interactions of the reform with birth parity and with mother's age in one regression. Not doing so could affect the magnitude and the precision of our estimates. At the same time, we acknowledge that mother's age is an endogenous variable (i.e., mothers choose when to give birth). Thus, we will always present results with and without these controls and show that our results are qualitatively similar across the two specifications.

We will discuss the details of each specification as it becomes relevant and provide evidence for our identification assumptions after we present the main results.

4 Data

This study uses the universe of data from Taiwan's *National Birth Registries* from 1980-1992 and *Death Registries* from 1980-1996 which is comprised of approximately 4.1 million individuals.[14] The data from the two registries are linked at the individual level. The death registry reports each child's dates of birth and death, from which we calculate age at the time of birth (measured in terms of months). Our analysis focuses on death within one month, when young children are most fragile. To check that the reform indeed affects neonatal mortality, we also present results for death at six, twelve, eighteen, 24, 36 and 48 months of life for comparison. Both datasets report basic child and parental characteristics. We restrict our sample to individuals born to mothers who were 18-45 years of age at the time of birth. We take the birth and death registries at face value.[15]

Because the sample size of the individual-level data is extremely large and because government policy prevents us from using the individual-level data outside of a secured facility in Taiwan, we aggregate the individual-level data to cells according to an individual's birth year, birth parity, and mother's age at the time of birth for convenience. We retain the size of each cell and estimate all of the descriptive statistics and regressions using cell-level data weighted by cell size (i.e., population weights). The estimated means, coefficients and standard errors in our analysis are numerically identical to those from using individual data. Mother's age is arbitrarily categorized into four groups: 18 to 22, 23 to 28, 29 to 34, or 35 years of age and over.[16]

Consistent with the belief that there was little pre-natal sex selection prior to the reform, the data show that 51.7% of all children born in Taiwan prior to the reform were male, which is similar to the fraction in countries with no known son preference, such as the United States. However, after the reform, the share of males increased to 52.3%.

Our main measure of mortality is the number of deaths per 1,000 living children. In our context, mor-

[14] We were unable to obtain data prior to 1980. The birth data end in 1992 because of administrative changes in data collection that make data from later years not directly comparable.

[15] All of the children in our sample are born in hospitals. The rates of multiple births in our data is 0.84% for the entire sample, which is similar the United States for the same time period. This again supports the belief that there is little mis-measurement in our data.

[16] We group mothers who may still be in school into one group (18-22), and then define additional groups of roughly equal age intervals (23-28, 29-34), and adjust for the fact that there are fewer mothers with older ages (35+). All of our results are robust to alternative categorizations. They are available upon request.

tality rates are much higher for very young children and the highest in the first month of life, when 37% of all mortality in the first year of life occurs.[17]

5 Descriptive Evidence

Several important facts emerge from the raw data. First, Figures 1a-1d plots the fraction of males at birth by birth parity and mother's age. Our focus on the first three birth parities is motivated by the fact that during the period of our study, parents had two to three children on average (Chang, 2003; Feeney, 1991). They show that there is no difference in sex at birth across parities prior to the reform, indicated by the vertical line. After the reform, there is a small increase in the fraction of boys for second-parity births and a dramatic increase for third and higher parity births. This pattern is true for mothers of all ages, but it is most pronounced for the oldest group of mothers (note that the y-axes differ across figures and the one for the oldest group of mothers covers the largest range). These figures are consistent with the reform having heterogeneous effects in increasing the sex imbalance more for higher parity births and for older mothers. Note that the rise in the fraction of males for third and higher parity births plateaus in the early 1990s.

Next, Figures 2a-2d plot female deaths per 1,000 children during the first month of life by birth parity for each category of mother's age. The low level of mortality rates causes the means to be noisier for mortality than for sex at birth. Nevertheless, the figures show a clear drop in the mortality rates of girls for third and higher parity births for the youngest and oldest mothers in Figures 2a and 2d. The drop is the most pronounced and permanent for the oldest group of mothers in Figure 2d. The timing of the drop coincides with the reform. The fact that the drop is driven by the youngest and oldest mothers, who also experienced the largest increase in sex selection after the reform is consistent with the hypothesis that the reform reduced relative female mortality rates by allowing parents with strong son preferences to abort female fetuses.

In Figures 3a-3d, we plot the analogous figures for boys. Although there is a decline in mortality rates over time, there is no pattern across parities or mother's age groups, the decline is relatively smooth over time and does not coincide with the reform.

In summary, Figures 2a-3d show that after the reform, mortality rates for higher parity girls declined

[17] See top of Table 2. The *Death Registry* data also report the causes of death. For deaths that occur during the first month of life, we categorize the causes of death according to how likely they to be outcomes of parental neglect. When we do this, we find that higher parity births are more likely to be due to parental neglect. For brevity, we discuss this in more detail in the Online Appendix. Unfortunately for the regression analysis, the low level of mortality rates means that it is not possible to examine the causes of death by birth parity, birth year and sex in a meaningful way. Thus, our regression analysis will only examine total mortality from all causes, for which we provide more detailed descriptive statistics in the following section.

relative to those for boys. This is more succinctly illustrated by plotting the ratio of female-to-male mortality rates for each birth parity over time. To isolate the meaningful variation from the reform, we demean this variable by mothers' age and mother's age interacted with post-reform status and plot the demeaned mortality ratio over time for each parity. Figure 4 shows that the pattern is most pronounced for third and higher parity births. Prior to the reform, all else equal, female mortality rates were higher than male mortality rates. However, after the reform, female mortality rates converged towards male mortality rates (i.e., relative mortality rates declined). The timing of the change coincides with the introduction of the reform. Figure 4 is important for our study as it illustrates the variation that underlies the mortality regressions presented in the next section.

Finally, Figures 5a-5e plot the number of births by parity for all mothers and for each category of mother's age. The figures show that the number of births for all parities declined steadily over time for the youngest mothers and first- and second-parity births increased for the three groups of older mothers. These patterns are consistent with the increase in the age at first birth of mothers during this period. Furthermore, they show that total family size was declining prior to the reform as the number of third and higher parity births were declining for all age groups. Importantly, Figure 5e shows that for the oldest group of mothers who are 36 or older, the number of third and higher parity births *increases* after the reform. This is consistent with our hypothesis that the reduction in the cost of sex-selective abortion can cause older mothers who have strong son preference and face a high cost of child-rearing to have additional children.

6 Regression Evidence

6.1 The Effect on Sex at Birth

To test the hypothesis that reducing the cost of abortion will increase sex-selective abortion for mothers who face a high cost of child rearing, we regress sex at birth on the interaction of the legalization of abortion and birth parity and the interaction of the legalization of abortion and mother's age. As we discussed earlier, since mother's age and birth parity are correlated, we examine both variables in one regression. The baseline equation can be written as

$$Y_{imt} = \sum_{i=2}^{3} \beta_i (O_i \times R_t) + \sum_{m=2}^{4} \alpha_m (A_m \times R_t) + \phi_m + \gamma_i + \rho_t + \varepsilon_{it}, \tag{1}$$

where the outcome for individuals of birth parity (order) i, born to mothers that are age m at the time of birth in year t, Y_{imt}, is a function of: the interaction of birth parity dummy variables, O_i, and a dummy variable for being born in 1985 or afterwards (post reform), R_t; the interaction of dummy variables indicating mother's age, A_m, and R_t; mother's age fixed effects, ϕ_m; birth parity fixed effects, γ_i; and birth year fixed effects, ρ_t.[18] The reference groups are comprised of first-born children and children born to the youngest mothers in our sample (age 18-22). The standard errors are clustered at the birth-year and -parity levels.[19] For all regression results, we present the coefficients and p-values in the tables.

β_i is the effect of the reform for birth parity i. α_m is the effect of the reform for children whose mother's age falls within the interval that defines category m. According to the framework discussed in Section 3, if there is substitution between pre-natal and post-natal sex selection, then the reform will increase the probability of having a son and this increase will be larger in magnitude for higher birth parities and older mothers, such that $\hat{\beta}_3 > \hat{\beta}_2 > 0$ and $\hat{\alpha}_4 > \hat{\alpha}_3 > \hat{\alpha}_2 > 0$.

Table 1 presents the coefficients and p-values for sex at birth. To illustrate the influence of each set of explanatory variables, we introduce them gradually before presenting the full baseline in column (3). In column (1), we estimate the interaction effects of the reform with birth parity dummy variables. In column (2), we add controls for mother's age. In column (3), we add the interactions of the reform and mother's age. This is the baseline specification, equation (1). The estimates show that the legalization of abortion increased the fraction of boys born by 0.25 percentage-points for second-parity births and two percentage-points for third and higher parity births. These estimates are statistically significant at the 10% and 1% levels. They are also statistically different from each other.[20]

Note that the coefficients reflect the effect of the reform on higher parity births relative to first parity births, the reference group. But since the descriptive statistics in the previous section show that there is no change for first-parity births over time, the relative changes measured by the coefficients effectively reflect absolute changes. This is the case for all of the outcomes that we examine.

[18] Our estimates will always control for the main effects of birth year, birth parity and mother's age. These fixed effects will control for all changes over time that affect all children or all mothers similarly, all differences across parities that are time invariant, and the changes over a mother's life-cycle that do not change over time and have similar effects across parities. Only the interaction terms are interpreted as plausibly exogenous.

[19] There are 39 clusters.

[20] When we conduct a Wald test on whether these two coefficients are equal to each other, we can reject the null hypothesis with 99% or higher confidence (the p-value is zero).

At the same time, the reform increased the fraction of males born to mothers 36 years or older by one percentage point (relative to children born to the youngest group of mothers). This estimate is statistically significant at the 5% level. Although the interaction of the reform with mother's age is not significant for other age groups, it is interesting to note that the magnitude of the coefficient is increasing with mother's age. These results are consistent with our hypothesis that the reform increased sex selection and that the increase was larger for mothers who face a higher cost of child-rearing.

In column (4), we control for mother's educational attainment. The interaction with mother's age for mothers who are age 29-34 (the second oldest group in our sample) is now also statistically significant. In column (5), we add the interaction terms of the reform and categorical variables for mother's educational attainment.[21] Columns (4) and (5) control for the fact that mother's age at birth is correlated with education (e.g., more educated mothers may choose to give birth later), which may also affect son-bias and the extent to which mothers take up the new technology. The main results are robust to adding these controls. Interestingly, the results in column (5) show that, controlling for birth parity and mother's age, mothers with high school education are the most likely to use abortion to sex select. The estimated magnitudes are similar in size for the interactions of the reform with mothers with nine to ten years and eleven to twelve years of education, but only the former is statistically significant.

6.2 The Effect on Sex-Differential Infant Mortality

To test the hypothesis that the reform reduced relative infant mortality, we employ a strategy with a similar logic to equation (1), but which also allows the effects to vary by the sex of the child.

The baseline equation is the following:

$$\begin{aligned}
D_{imts} &= \sum_{i=2}^{3} \beta_i (O_i \times R_t \times G_s) + \sum_{m=2}^{4} \alpha_m (A_m \times R_t \times G_s) && (2)\\
&+ \sum_{i=2}^{3} \delta_i (O_i \times G_s) + \sum_{m=2}^{4} \varphi_m (A_m \times R_t) + \pi_t (G_s \times R_t) && (3)\\
&+ \sum_{i=2}^{3} \lambda_i (G_s \times O_i) + \sum_{m=2}^{4} \varpi_m (A_m \times G_s) && (4)\\
&+ \theta_s + \gamma_i + \rho_t + \phi_m + \varepsilon_{imts}. && (5)
\end{aligned}$$

[21] These categories are arbitrarily defined to contain similar numbers of children.

The outcome variable, D_{imts}, is the occurrence of a death for an individual of birth order i, born to mothers of age m in birth year t, and who are sex s. G_s is a dummy variable that equals one if the sex of child is female. All other variables are defined as before. The standard errors are clustered at the birth-parity and post-reform level. As this results in six clusters, we correct for the small number of clusters by estimating bootstrapped standard errors, which Cameron et al. (2008) recommends for correcting for small-sample bias when there are five or more clusters.[22] As before, we aggregate the individual-level data to cells for convenience and estimate the regressions with population weights. For the mortality analysis, the data are aggregated according to birth parity, birth year, the mother's age and child's sex.

Note that because we interact the impact of the reform with the dummy variable indicating that the child is female, our estimates capture the effect on relative female mortality rates.[23]

Following the discussion in Section 6, we expect that legalizing abortion will reduce relative female mortality, and that the reduction will be larger for children born to mothers who face a high cost of child-rearing. Specifically, the effects across birth parities are expected to be more negative at higher birth parities such that $\hat{\beta}_3 < \hat{\beta}_2 < 0$. However, the estimates of the interactions between post-reform, the sex of the child and mother's age categories are likely to be ambiguous given the observation that older mothers have more children after the reform.

The main results on relative mortality rates for girls from equation (2) are shown in Table 2. The outcome is the number of deaths per 1,000 children. To illustrate the influence of the different controls, Panel A first presents estimates without controlling for mother's age or its interactions. In column (1), when we examine deaths within one month of birth, the triple interaction coefficients, $O_i \times R_t \times G_s$, are negative for both second-parity births and third and higher parity births. The magnitude of the estimate is much larger for third and higher parity births. However, neither estimate is statistically significant. For mortality later in life (columns (2)-(7)), the estimates are much smaller in magnitude, which is consistent with all of the mortality effects occurring during the first month of life. These estimates are mostly statistically insignificant.

In Panel B, we present the baseline estimates where we control for mother's age and its interactions. The precision of the estimates in column (1) greatly improves relative to the analogous estimates in Panel A. The coefficients also increase in magnitude. Column (1) shows that in the first month of life, the reform

[22]We use the bootstrapped clustered residual method (Cameron et al., 2008). The clustered standard errors are smaller than robust standard errors without clustering, which implies that there is negative intra-cluster correlation.

[23]Note that in addition to the main effects of mother's age, sex, birth parity and birth year, we control for all combinations of double interaction terms between mothers' age, sex, birth parity and post-reform. Thus, our estimates control for a large range of potentially confounding influences such as secular changes in relative female mortality rates that do not vary by birth parity.

reduces relative female mortality by almost 0.19 children per 1,000 for second-parity births and by almost 0.29 per 1,000 for third and higher parity births (i.e., approximately two and three deaths per 10,000 births). We interpret the magnitude of the estimates later in Section 6.5. The estimates are statistically significant at the 1% and 5% levels.[24] The results show that conditional on mother's age and its interactions with the reform, the introduction of pre-natal sex selection reduced relative female mortality rates for higher parity births. Since we showed that there is an analogous rise in the fraction of males at births earlier in Table 1 column (3), this result is consistent with the hypothesis that sex-selective abortion reduces relative female mortality rates.

The reform has little effect on relative mortality for either parity later in life. The estimates in columns (2)-(7) show that all of the effects of the reform on mortality (for the first four years of life) occur in the first month. This is reassuring since this is the period when infant death is most likely to occur.

Several insights emerge from comparing Panels A and B. First, as we previously discussed, the inclusion of the controls for mother's age categories and their interactions is important for the precision and magnitude of the mortality results in column (1). This is most likely due to the fact that the reform affects relative female mortality rates by mother's age through several forces, which increases the variance of the estimates (when we do not control for mother's age and its interactions), and the fact that birth parity and mother's age are correlated.

A second and related point is that the triple interactions of post-reform, sex and mother's age dummy variables ($A_m \times R_t \times G_s$) in Panel B are not statistically significant for any mother's age group and the coefficients do not monotonically increase in magnitude with mother's age. This is consistent with our earlier discussion: the selection of older mothers with strong son preference into having children can reduce relative male mortality rates (and hence increase relative female mortality rates) for older mothers.[25] In other words, the opposing forces that we described in Section 3 for mother's age offset, while for birth parity, the negative force on relative female mortality rates across parities dominate the positve ones.

[24] When we conduct a Wald test on whether these two coefficients are equal to each other, we can reject the null hypothesis with 95% or higher confidence (the p-value is 0.023).

[25] Although the interaction terms are insignificant, the relative magnitudes of the coefficients suggest that after the reform, relative female mortality declined the most for the youngest mothers (the reference group) and the oldest mothers. This is consistent with the descriptive statistics shown in Figures 2a-2d. Note that the mortality results in Panel B are robust to controlling for mother's education and its interaction with post-reform and the sex of the child. These results are available upon request.

6.3 The Effect on Fertility

The results on sex ratios at birth and relative female mortality rates, together, indicate that the reform increased pre-natal sex selection and reduced relative female mortality rates for higher parity births. This is consistent with the conceptual framework, which predicts that the reform will have the most pronounced effects for parents facing higher costs of having additional children. The finding that the reform increased sex ratio at birth more for older mothers, but did not simultaneously reduce relative female mortality for older mothers, is consistent with the framework if the reform induced older mothers with strong son preference to give birth.

The estimates in Table 1 column (3) shows that the reform indeed caused a larger increase in the fraction of males born to older mothers. The estimates in Table 3 shows that the reform also increased the number of children born to older mothers. To estimate the effect of the reform on fertility, we repeat equation (1), except that the outcome variable is the number of children in each parity-year-mother's-age cell divided by the number of women in each age group in Taiwan (e.g., age 18-45).[26] The outcome variable should be interpreted as the number of births per woman. As with the results on sex at birth, we introduce the controls gradually and present our main baseline specification in column (3). It shows that the reform reduced the number of third or higher parity children by approximately 0.0102 per woman. In other words, the reform caused there to be one less child born for roughly every 100 women of child bearing age. This estimate is statistically significant at the 1% level. The fact that there is little effect for second-parity births and a significant reduction of third and higher parity births is consistent with the fact that our study takes place during a transition from a total fertility rate of above three to just below three.

The interaction effects between mother's age and post-reform are positive and monotonically *increasing* in magnitude with mother's age. It is statistically significant at the 5% for the oldest group of mothers. Relative to younger mothers, the reform caused mothers age 36 and over have on average 0.0124 child more (one child per eighty women). These results are consistent with our explanation.

We also note that in column (5), the interaction effects of mother's educational attainment and post-reform are larger in magnitude for higher levels of education, which show that more educated mothers are more likely to be induced into giving birth by access to sex-selective abortion. The interaction estimates for the two most educated groups are statistically significant at the 5% level.

[26] We interpolate the population using data from the 1980 and 1990 population censuses.

6.4 Robustness

Our preferred interpretation of the main mortality results assumes that the reform was not accompanied by other changes that would reduce relative female mortality rates for higher parity births (conditional on the baseline controls). In this section, we provide evidence for this assumption. For brevity, we focus on mortality during the first month of life.

6.4.1 Timing

We first investigate whether the changes driving the main results occur around the time of the reform and estimate two equations similar to equations (1) and (2), replacing the post-reform dummy in each with a vector of dummy variables for each year since the reform. The estimates for sex at birth and relative female mortality rates are shown in Figures 6a and 6b. There are no pre-trends and the changes are larger for third and higher parity children for both outcomes.[27] Furthermore, the timing of change coincides with the reform, which is consistent with our interpretation of the effect of the reform, but difficult to reconcile with an alternate explanation based on underlying trends.[28]

The figures also show that the increase in the fraction of males continued to grow for several years after the reform, while the reduction in relative mortality rates seems to have been realized fully immediately after the reform. This suggests that parents who were selecting post-natally prior to the reform were very sensitive to the reform, which is consistent with the belief that such parents had the strongest son preference.

6.4.2 Composition Effects

An alternative explanation for the results on mortality is that the reform changed the composition of parents having girls in such a way that relative female mortality declines due to reasons other than a decline in the neglect of female infants. For example, if the education of parents of girls relative to parents of boys increased after the reform, then the reduction in relative female mortality may be a result of girls having more educated parents who are better able to care for them, rather than a change in the average amount of effort that parents exert towards caring for girls.

[27] Note that the increase for second-parity children that we estimated in Table 1 are less apparent in this figure because of the y-axis scale. See an earlier version of the paper for more detailed figures (Lin et al., 2008). The estimates over time are noisier than those for sex at birth, which is likely due to the low level of child mortality in the data.

[28] The coefficients and p-values are shown in Online Appendix Tables A.2 and A.3. That the point estimates for the post-reform period are jointly different from those from the pre-reform period is shown by the baseline estimates presented earlier.

To investigate this, we estimate equation (2), where we replace mortality with mother's or father's educational attainment as dependent variables. Table 4 columns (1) and (2) shows that the reform had no effect on the educational attainment of parents of girls relative to parents of boys for second-parity births, and reduced it for third and higher parity births. The estimates for the latter are statistically significant at the 1% level. If educated parents are able to provide better care for their children, then our main results understate the extent to which the reduction in relative female mortality is due to an increase in the effort that the average parents of girls exert towards their care.

Similarly, column (3) shows that the average mother of a girl born after the reform was less likely to be married relative to the average mother of a girl born prior to the reform. If single mothers are not able to provide care that is as high in quality as married mothers, then this implies that our main results understate the extent to which the reduction in relative female mortality is due to an increase in the effort that the average parents of girls exert towards their care.

Another concern is that the increase in the use of Ultrasound B improved pre-natal care in such a way as to cause more boys of marginal health to be born than girls of marginal health. In this case, the reduction in relative female mortality would reflect the change in the endowment of girls relative to boys rather than the change in the effort parents exert towards caring for girls on average. We find no evidence to support this concern. In columns (4) and (5), we show that there was no effect on the relative birth status of girls in terms of whether they are part of a single birth or have low birth weight, two crude proxies for health status. This is true for both the second and third and higher parity births.[29]

6.5 Quantifying the Results

Consistent with the low level of average child mortality rates in Taiwan, the coefficients for the effect of the reform on relative female mortality rates shown in Table 2 are small in magnitude. To quantify the overall impact of the reform, we conduct two simple exercises.

First, we compare our estimated coefficients on the changes in sex at birth and relative mortality rates to the observed changes over time in the raw data. For example, sex at birth for second parity births increased 0.3 percentage-points from an average of 51.6% males during the pre-reform period to an average of 51.9%

[29]Our data provide us with only these two proxies for health. These two variables are highly correlated with each other and have both been show to be associated with poor health. After an extensive review of the medical literature, we know of no medical reason to believe that ultrasound had different effects on the pre-natal care of boys and girls, much less that such effects would vary by birth parity.

males during the post-reform period, and for third and higher parity births, it increased 1.6 percentage-points from 51.7% to 53.3%. Our baseline estimates in Table 1 column (3) show that the reforms increased the fraction of males by 0.254 percentage-points for second parity births and two percentage-points for third and higher parity births. Thus, the legalization of abortion can explain 100% of the increase in the rise in the fraction of males at birth over the period of our study.[30]

When we conduct an analogous calculation for relative mortality rates, we find that the reform explains approximately 50% of the decline in relative female mortality rates for second parity births and 100% of the decline in relative mortality rates for third and higher parity births. Assessed this way, the legalization of abortion had a large impact on sex-selection at birth and on the reduction in relative mortality rates of children after birth.

The second exercise calculates the implied number of parents who had previously neglected unwanted daughters such that the daughters died within the first month of life, and who, after the reform, switch to aborting female fetuses. We focus on second and third and higher parity births and for brevity, will only discuss the latter (columns (4) - (6) in Table 5). Column (4) row A reports the one-month female mortality rates during the pre-reform period as reported in the raw data. Column (5) shows the predicted female mortality rate during the post-reform period that is a result of the reform. The difference between the two periods shown in column (6) is the triple interaction estimate from Table 2 Panel B column (1).

Multiplying the one-month female mortality rates, for the pre- and post-reform periods (row A) and the total number of girls born during each period (row B) gives the total number of female deaths during the two periods. Row C shows that the total number of predicted female deaths were 802 and 390 during the pre- and post-reform periods, respectively. The difference between the number of girls dying in the two periods is therefore 412. Thus, the decline in the predicted number of deaths for third and higher parity births is approximately 51% of the pre-reform level mortality rate ($412/802 = 0.51$). If we conduct the analogous exercise for second parity births, we find that the predicted decline is approximately 23% ($207/(881-674)$) of the pre-reform level mortality rate.

Multiplying the fraction of missing girls for each period (row D) with the number of births each period (row E) gives the number of missing girls each period. Row F shows that 4,667 and 9,110 third and

[30]Note that the implied two percentage-point increase in the fraction of males for third and higher parity births is higher than the 1.6 percentage-point increase in the data. However, the difference is not statistically significant. Also note that in column (4), when we condition on the education of mothers and its interaction with the reform, the implied increase for third and higher parity births declines to 1.67 percentage-points.

higher parity girls were missing during the pre- and post-reform periods.[31] To calculate the fraction of parents that substituted from post- to pre-natal sex-selection, we assume that all parents that would have post-natally selected previously now select prenatally. Thus, the fraction of parents that switched from to pre-natal sex-selection is the reduction in the number of girls' deaths shown in row C column (6) divided by the total number of missing girls during the pre-reform era in row F column (4). We find that 8.8% ($412/4,667 = 0.08$) of parents of third and higher parity births switched to pre-natal selection after the reform. An analogous exercise shows that 4.3% of parents of second parity births switched after the reform.

For interpreting the two back-of-the-envelope quantification exercises, note that the baseline equation for sex at birth condition on mother's age and its interactions with the reform; for the mortality equation, it additionally conditions on the triple interactions of mother's age with the reform and the sex of the child. The coefficients thus reflect the effect of the reform holding mother's age and its interactions constant (i.e. controlling for compositional changes). This means that the estimates discussed in this section are only relevant for parents who would have had children regardless of access to sex-selective abortion and does not apply to mothers that have children only because they can now sex select pre-natally. For the mortality results, this conceptually refers to the fact that mothers induced to give birth by the reform are not switching from post-natal selection because they were not giving births prior to the reform.[32]

While it makes sense to control for the effect of the reform on fertility for older mothers, the inclusion of mother's age can also introduce endogeneity since mother's age at the time of birth is a choice variable. Thus, we also examine the implied tradeoffs from the predicted mortality rates from the regression in Table 2 Panel A, where we do not control for mother's age or its interactions. These calculations imply that legalizing abortion caused 4% of parents to switch from post- to pre-natal sex selection amongst second-parity births, and 8.5% of parents to switch amongst third and higher parity births. Thus, the implied numbers of switchers is nearly identical to those using the conditional baseline estimates.

Also note that all of the regressions include time effects, which control for secular changes over time

[31] We assume that the *natural* fraction of males at birth is 51% (the lowest observed fraction of males at birth in the United States). Hence, the natural fraction of girls is 49%. Note that because fertility declined, the number of missing girls declined even though the fraction of missing girls born declined.

[32] Note that accounting for the selection of older mother's into giving birth will most likely show that the tradeoff between pre-natal and post-natal sex selection is larger than the current quantification exercise. Conceptually, accounting for selection requires us to subtract the number of "extra" children born to older mothers after the reform. As long as the relative mortality rates for sons is lower for this group relative to children born to younger mothers, accounting for selection would cause the relative male mortality rates during the post reform period to be higher, which means that the relative female mortality rates are lower, which in turn implies that the implied tradeoff between pre- and post-natal sex selection is larger than what our current quantification exercise shows.

in fertility. Thus, the coefficients and the implied number of parents that switch to pre-natal sex-selection should be interpreted as conditional on fertility being held constant.

7 Conclusion

This paper shows that increased access to sex-selective abortion in Taiwan significantly increased the fraction of males born and reduced the relative neonatal mortality rates of girls. These effects are large in magnitude as they explain all of the increase in the fraction of males born and over half of the decline in relative female neonatal mortality rates during this period. The results make a simple point: there is a trade-off between pre- and post-natal sex selection for some parents. In the context of our study, such parents are those with strong son preference who have higher parity children regardless of access to sex-selective abortion. The findings suggest that policymakers who wish to ban sex-selective abortion should consider complementary policies that incentivize parents to invest in daughters.[33]

Studying the effects of sex-selective abortion in Taiwan has several advantages over other contexts. The quality of the data is much better than data that are available from most other countries with son-biased sex imbalances. The legislative reform provides plausibly exogenous variation in access to sex-selective abortion. Since it was legal to reveal the sex of the fetus, we can also avoid some selection issues.

It is important to note that the magnitude of our estimates are specific to the context of our study. Taiwan during the 1980s was a society with strong son preference and low child mortality rates that experienced a secular decline in the preferred number of children. Our results will likely understate the effect relative to places where son-bias is stronger, or where parents are more constrained in the number of children they can have (e.g., mainland China, which has strict family planning policies). It may also understate the effect relative to places where healthcare is generally poor such that child mortality is more sensitive to parental neglect (e.g., India).

This paper suggests several topics of future research. First, it is important to examine whether sex-selective abortion can affect more nuanced outcomes for children such as schooling or height-for-age. Second, aside from the main results, we find that older mothers actually have *more* children after the legislative reforms. The idea that the availability of abortions, by allowing parents to select the "quality" of children, may potentially induce certain parents to have more children is new and should be further explored.

[33] An example is a policy implemented in India which gives cash awards to parents who give births to daughters and also promises an award for parents when their daughters reach age 18 (Holla et al., 2007).

References

Abrevaya, Jason, “Are There Missing Girls in the United States? Evidence from Birth Data,” *American Economic Journal: Applied Economics*, September 2009, *1* (2), 1–34.

Almond, Douglas, Hongbin Li, and Lingsheng Meng, “Son Preference and Early Childhood Investments in China,” Master’s thesis, Working Paper, Tsinghua University 2010.

Arnold, Fred, Sunita Kishor, and T. K. Roy, “Sex-Selective Abortions in India,” *Population and Development Review*, 2002, *28* (4), 759–785.

Banister, Judith, “Shortage of girls in China today,” *Journal of Population Research*, 2004, *21* (1), 19–45.

Cameron, A. Colin, Jonah B. Gelbach, and Douglas L. Miller, “Bootstrap-Based Improvements for Inference with Clustered Errors,” *Review of Economics and Statistics*, July 2008, *90* (3), 414–427.

Chang, Ming-Cheng, “Demographic Transition in Taiwan,” *Journal of Population and Social Security (Population)*, 2003, *Supplement to Volume 1*, 611–628.

Chen, Yuyu, Hongbin Li, and Meng Lingsheng, “Prenatal Sex Selection and Missing Girls in China: Evidence from the Diffusion of Diagnostic Ultrasound,” *Journal of Human Resources*, 2013, *48*, 36–70.

Donohue, John J. and Steven D. Levitt, “The Impact of Legalized Abortion on Crime,” *The Quarterly Journal of Economics*, 2001, *116* (2), 379–420.

Ebenstein, Avraham, “Fertility Choices and Sex Selection in Asia: Analysis and Policy,” Master’s thesis, Working Paper, University of California at Berkeley 2007.

_ , “Estimating a Dynamic Model of Sex Selection in China,” *Demography*, 2011, *48* (2), 783–811.

_ **, Hongbin Li, and Meng Lingshen**, “The Impact of Ultrasound Technology on the Status of Women in China,” Master’s thesis, Working Paper, Hebrew University 2013.

Feeney, Griffith, “Fertility Decline in Taiwan: A Study Using Parity Progression Ratios,” *Demography*, 1991, *28* (3), 467–479.

Goodkind, Daniel, “On Substituting Sex Preference Strategies in East Asia: Does Prenatal Sex Selection Reduce Postnatal Discrimination?,” *Population and Development Review*, 1996, *22* (1), 111–125.

Greenhalgh, Susan and Jiali Li, "Engendering Reproductive Policy and Practice in Peasant China: For a Feminist Demography of Reproduction," *Signs*, 1995, *20* (3), 601–641.

Gu, B. and K Roy, "Sex ratio at birth in China, with reference to other areas in East Asia: What we know," *Asia-pacific Population Journal*, 1995, *10*, 17–42.

Gupta, Monica Das, "Selective Discrimination against Female Children in Rural Punjab, India," *Population and Development Review*, 1987, *13* (1), 77–100.

_ **and P. N. Mari Bhat**, "Fertility Decline and Increased Manifestation of Sex Bias in India," *Population Studies*, 1997, *51* (3), 307–315.

Henshaw, Stanley K., "Induced Abortion: A World Review, 1990," *International Family Planning Perspectives*, 1990, *16* (2), 59–65+76.

Holla, Alaka, Robert Jensen, and Emily Oster, "Daughters as wealth? The Effects of Cash Incentives on Sex Ratios," Master's thesis, Working Paper, Chicago Booth School 2007.

Hu, Luojia and Analia Schlosser, "Does Sex-Selective Abortion Improve Girls' Well-being: Evidence from India," Master's thesis, Working Paper, Tel Aviv University 2010.

Hung, David Sho-Chao, "Abortion Rights in the United States and Taiwan," Master's thesis, Working Paper, Kent College of Law 2004.

Jayachandran, Seema and Ilyana Kuziemko, "Why Do Mothers Breastfeed Girls Less than Boys? Evidence and Implications for Child Health in India," *The Quarterly Journal of Economics*, 2011, *126* (3), 1485–1538.

Lin, Ming-Jen, Jin-Tan Liu, and Nancy Qian, "More Women Missing, Fewer Girls Dying: The Impact of Abortion on Sex Ratios at Birth and Excess Female Mortality in Taiwan," Working Paper 14541, National Bureau of Economic Research December 2008.

Parikh, M, "Morass of sex-selective abortions: Legal inefficiencies and the continued tragedy of unwanted daughters," Technical Report, Abstracts of the 1993 Annual Meeting of the Association for Asian Studies, Ann Arbor, MI 1993.

Park, Chai Bin and Nam-Hoon Cho, "Consequences of Son Preference in a Low-Fertility Society: Imbalance of the Sex Ratio at Birth in Korea," *Population and Development Review*, 1995, *21* (1), 59–84.

Pebley, Ann and Sajeda Amin, "The impact of a public health intervention on sex differentials in childhood mortality in rural Punjab, India," *Health Transition Review*, 1991, *1*, 143–169.

tung Liu, Chung, "Female Body under Government Policy. Situation of Women in Taiwan: 1995 report, edited by Liu, Yu-Hsiu," *Chinatimes Publication*, 1995, pp. 219–254 (in Chinese).

(UN), United Nations, "Programme of Action," Technical Report, United Nations Population Conference, Cairo 1994.

Table 1: The Effect of the Reform on Sex at Birth

	Dependendent Variable: Fraction of Males				
	(1)	(2)	(3)	(4)	(5)
Dep Var Mean	**0.52**	**0.52**	**0.52**	**0.52**	**0.52**
Post x Parity = 2	0.00298	0.00277	0.00254	0.00234	0.00321
	(0.0522)	(0.0714)	(0.0991)	(0.171)	(0.0610)
Post x Parity = 3+	0.0214	0.0211	0.0200	0.0167	0.0182
	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)
Post x Mother's Age = 23-28			-0.00197	0.000743	-0.000320
			(0.117)	(0.679)	(0.902)
Post x Mother's Age = 29-34			0.00110	0.00340	0.00460
			(0.531)	(0.0862)	(0.208)
Post x Mother's Age = 35+			0.0110	0.0134	0.0131
			(0.0447)	(0.0191)	(0.0612)
Post x Mother's Years of Education = 9 -10					0.00485
					(0.0739)
Post x Mother's Years of Education = 11-12					0.00423
					(0.249)
Post x Mother's Years of Education = 12+					-0.000276
					(0.953)
Controls					
Birth Parity FE	Y	Y	Y	Y	Y
Mother's Age FE	N	Y	Y	Y	Y
Mother's Edu FE	N	N	N	Y	Y
Observations	156	156	156	156	156
R-squared	0.642	0.726	0.744	0.781	0.791

Notes: Observations are at the cell-level, which are constructed by aggregating individual-level data by birth parity, birth year and mother's age. For birth parity, individuals are categorized into groups of 1st, 2nd and 3+ births. For mother's age, individuals are categorized into groups for mothers' ages of: 18-22, 23-28, 29-34, 35+. The regressions are weighted by cell size. P-values, clustered at the birth-parity and year level, are presented in the parentheses. The raw data are reported by Taiwan's *National Birth Registries,* 1980-1992.

Table 2: The Effect of the Reform on Relative Female Mortality

Dependent Variables: The Number of Death per 1,000 Children Living in Month X of Life							
	1 Mo.	6 Mos.	12 Mos.	18 Mos.	24 Mos.	36 Mos.	48 Mos.
	(1)	(2)	(3)	(4)	(5)	(6)	(7)
Dep Var Mean	**2.251**	**0.308**	**0.127**	**0.129**	**0.090**	**0.056**	**0.045**
	A. Exclude Controls for Mother's Age and its Interactions						
Post x Girl x Parity = 2	-0.124	-0.0391	-0.0604	-0.0114	-0.0184	-0.0657	-0.018
	(0.464)	(0.468)	(0.092)	(0.860)	(0.568)	(0.036)	(0.560)
Post x Girl x Parity = 3+	-0.208	-0.00997	-0.00145	-0.0844	-0.0259	0.0175	0.0247
	(0.192)	(0.928)	(0.952)	(0.168)	(0.444)	(0.508)	(0.472)
Observations	312	312	312	312	312	312	312
R-Squared	0.589	0.161	0.189	0.175	0.229	0.098	0.095
	B. Baseline -- Include Controls for Mother's Age and its Interactions						
Post x Girl x Parity = 2	-0.186	-0.0376	-0.0695	-0.00856	-0.0226	0.0641	-0.0302
	(0.008)	(0.087)	(0.055)	(0.787)	(0.567)	0.000	(0.095)
Post x Girl x Parity = 3+	-0.287	-0.013	-0.0138	-0.0922	0.0212	0.026	0.00281
	(0.044)	(0.151)	(0.293)	(0.418)	(0.382)	(0.080)	(0.844)
Post x Girl x Mother's Age = 23-28	0.409	-0.0111	0.0843	-0.0141	0.02	0.0609	0.0546
	(0.160)	(0.959)	(0.234)	(0.648)	(0.996)	(0.566)	(0.139)
Post x Girl x Mother's Age = 29-34	0.45	-0.0372	0.0535	-0.0279	0.0271	0.0129	0.078
	(0.606)	(0.917)	(0.676)	(0.705)	(0.967)	(0.672)	(0.133)
Post x Girl x Mother's Age = 35+	0.278	0.215	0.12	0.435	-0.047	-0.166	0.0982
	(0.425)	(0.697)	(0.227)	(0.363)	(0.487)	(0.190)	(0.091)
Observations	312	312	312	312	312	312	312
R-Squared	0.668	0.205	0.224	0.201	0.278	0.146	0.13

Notes: Observations are at the cell-level, which are constructed by aggregating individual-level data by birth parity , birth year, mother's age and the sex of the child. For birth parity, individuals are categorized into groups of 1st, 2nd and 3+ births. For mother's age, individuals are categorized into groups for mothers' ages of: 18-22, 23-28, 29-34, 35+. The regressions are weighted by cell size. The regressions are weighted by cell size. Bootstrapped p-values, clustered at the birth-parity and post-reform level, are presented in the parentheses. The raw data are reported by Taiwan's *National Birth Registries, 1980-1992* and *National Death Registries*, 1980-1996. In Panel A, the controls are: post x girl, post x parity=2, post x parity=3, girl x parity=2, girl x parity=3, a dummy variable for girl, two dummy variables for birth parity and birth year fixed effects. In Panel B, the controls are the same as those in Panel A, and, in addition include: post x mother's age=23-28, post x mother's age=29-34, post x mother's age=35+, girl x mother's age=23-28, girl x mother's age=29-34, girl x mother's age =35+, and the four dummy variables for mother's age groups.

Table 3: The Effect of the Reform on Fertility

	Dependendent Variable: Number of Births per Woman				
	(1)	(2)	(3)	(4)	(5)
Dep Var Mean	**0.043**	**0.043**	**0.043**	**0.043**	**0.043**
Post x Parity=2	-0.00148	0.000864	-0.000254	0.000665	0.00361
	(0.248)	(0.458)	(0.876)	(0.788)	(0.190)
Post x Parity = 3+	-0.0125	-0.00769	-0.0102	-0.0176	0.00494
	(0.000)	(0.000121)	(0.00178)	(0.000)	(0.600)
Post x Mother's age = 23-28			0.00313	0.00875	-0.00493
			(0.347)	(0.0188)	(0.229)
Post x Mother's age = 29-34			0.00755	0.0136	-0.0133
			(0.298)	(0.0128)	(0.279)
Post x Mother's age = 35+			0.0124	0.0238	0.00255
			(0.0443)	(0.000582)	(0.819)
Post x Mother's Years of Education = 9 -10					0.0105
					(0.126)
Post x Mother's Years of Education = 11-12					0.0206
					(0.0182)
Post x Mother's Years of Education = 12+					0.0376
					(0.0307)
Controls					
Birth Parity FE	Y	Y	Y	Y	Y
Mother's Age FE	N	Y	Y	Y	Y
Mother's Edu FE	N	N	N	Y	Y
Observations	156	156	156	156	156
R-squared	0.245	0.794	0.798	0.874	0.887

Notes: Observations are at the cell-level, which are constructed by aggregating individual-level data by birth parity, birth year and mother's age. For birth parity, individuals are categorized into groups of 1st, 2nd and 3+ births. For mother's age, individuals are categorized into groups for mothers' ages of: 23-28, 29-34, 35+. The regressions are weighted by cell size. P-values, clustered at the birth-parity and year level, are presented in the parentheses. The raw data are reported by the Taiwan's National Birth Registries, 1980-1992.

Table 4: The Effect of the Reform on the Parental Characteristics of Girls and Girls' Health Outcomes

	Dependent Variables				
	Mother's Edu Years	Father's Edu Years	Parents Married	Singleton Birth	Low Birth Weight
	(1)	(2)	(3)	(4)	(5)
Dep Var Mean	**9.895**	**9.229**	**0.983**	**0.991**	**0.041**
Post x Girl x Parity = 2	0.0112	0.00245	-0.00107	0.000219	-0.00136
	(0.417)	(0.583)	(0.008)	(0.870)	(0.858)
Post x Girl x Parity = 3+	-0.102	-0.114	-0.00217	-0.000642	0.00118
	(0.009)	(0.009)	(0.009)	(0.130)	(0.267)
Post x Girl x Mother's Age = 23-28	-0.0320	-0.0163	-0.00102	0.000899	-0.00153
	(0.303)	(0.336)	(0.254)	(0.225)	(0.205)
Post x Girl x Mother's Age = 29-34	-0.0204	-0.0177	-0.00139	0.000586	-0.00105
	(0.456)	(0.523)	(0.158)	(0.123)	(0.290)
Post x Girl x Mother's Age = 35+	-0.232	-0.188	0.00458	0.00256	-0.000555
	(0.235)	(0.219)	(0.429)	(0.255)	(0.557)
Observations	312	312	312	312	312

Notes: Observations are at the cell-level, which are constructed by aggregating individual-level data by birth parity, birth year, mother's age and the sex of the child. For birth parity, individuals are categorized into groups of 1st, 2nd and 3+ births. For mother's age, individuals are categorized into groups for mothers' ages of: 18-22, 23-28, 29-34, 36+. The regressions are weighted by cell size. Bootstrapped p-values (estimated with 500 repetitions), clustered at the birth-parity and post-reform level, are presented in the parentheses. The raw data are reported by Taiwan's *National Birth Registries*, 1980-1992, and *National Death Registries,* 1980-1996. The controls are: post x girl, post x parity=2, post x parity=3, girl x parity=2, girl x parity=3, a dummy variable for girl, two dummy variables for birth parity, birth year fixed effects, addition include: post x mother's age=23-28, post x mother's age=29-34, post x mother's age=35+, girl x mother's age=23-28, girl x mother's age=29-34, girl x mother's age =35+, and the four dummy variables for mother's age groups.

Table 5: The Implied “Tradeoff” Between Pre- and Post-natal Sex-Selection

	2nd Parity			3+ Parity		
	1980-84	1985-92	Difference	1980-84	1985-92	Difference
	(1)	(2)	(3)	(4)	(5)	(6)
A. Female Mortality Rate (per 1,000)	0.00235	0.002164	-0.000186	0.00249	0.002203	-0.000287
B. Total # of Girls Born	374,870	311,642	-63,228	322,014	176,878	-145,136
C. Total # of Girls Dying = A x B	881	674	-207	802	390	-412
D. Fraction of Girls	0.484	0.481	-0.003	0.483	0.466	-0.017
E. Total Number of Births	774,524	647,904	-126,620	666,696	379,567	-287,129
F. Total # of Missing Girls = (0.49 - D) x F	4,647	5,831	1,184	4,667	9,110	4,443
G. Fraction of Parents Substituting from Post- to Pre-natal selection = (C2 - C1) / F1 or (C4 - C3) / F3		0.0444			0.0883	

Notes: Row A columns (1) and (4) are reported by the *National Birth Registry*. Row A columns (2) and (5) are calculated by substracting columns (3) and (6) from columns (1) and (4). Rows B and E are reported by the *National Birth Registry*. Other numbers are calculated by the authors. See text for more discussion.

Figure 1: Fraction of Males at Birth by Parity and Mother's Age

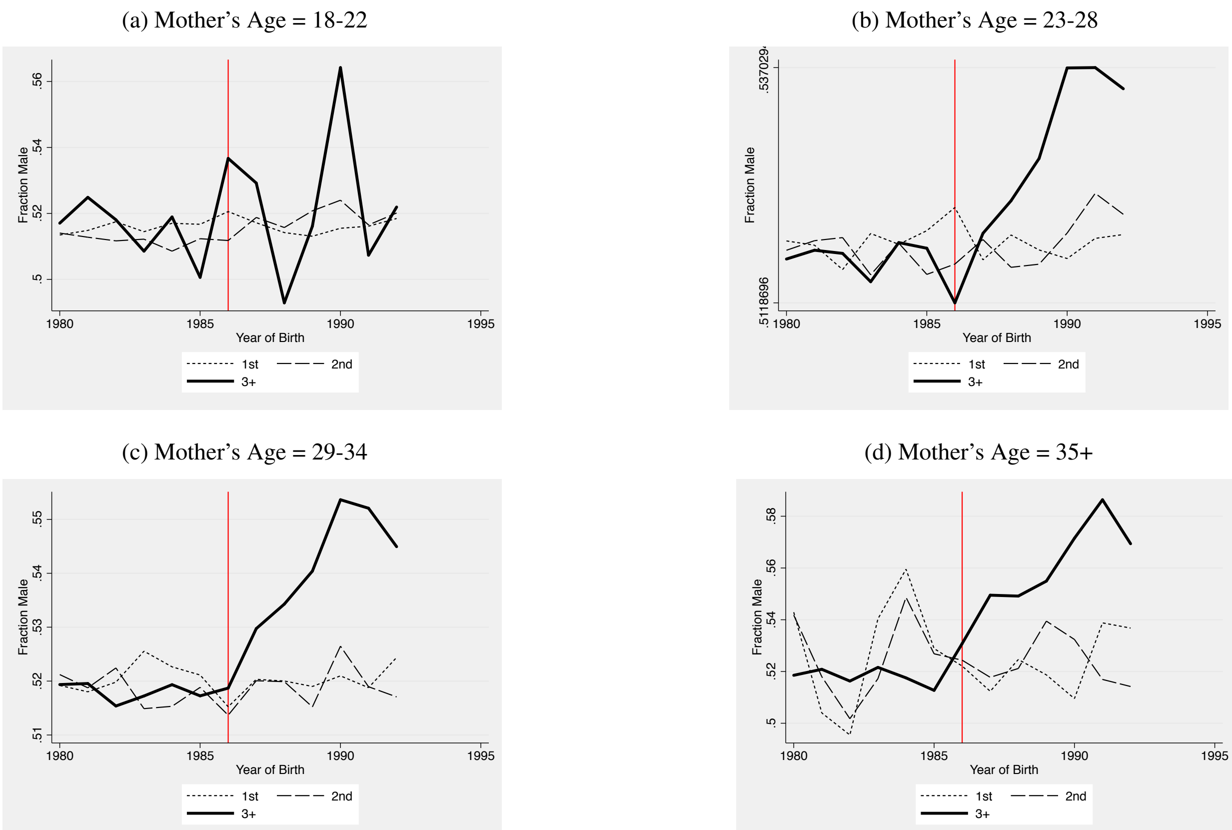
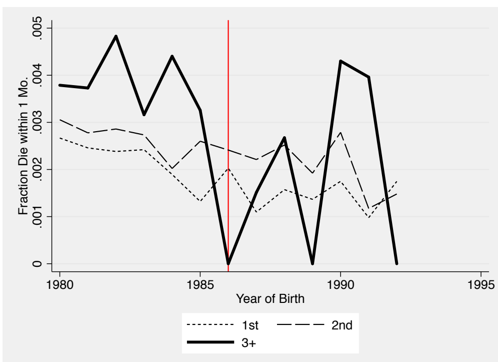
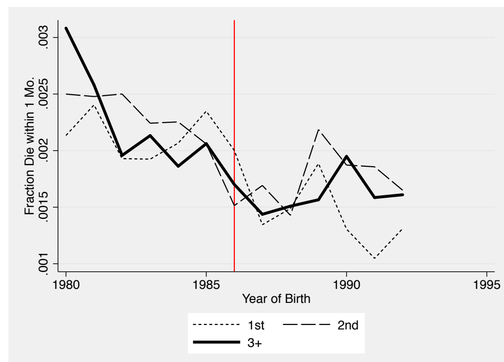


Figure 2: Female Mortality Rates within One Month of Life by Parity and Mother's Age

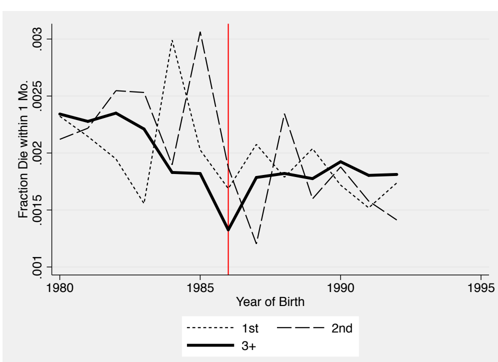
(a) Mother's Age 18-22



(b) Mother's Age = 23-28



(c) Mother's Age = 29-34



(d) Mother's Age = 35+

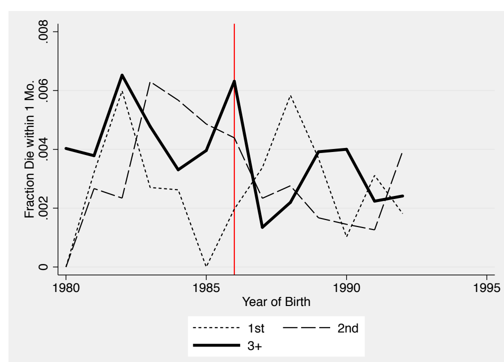
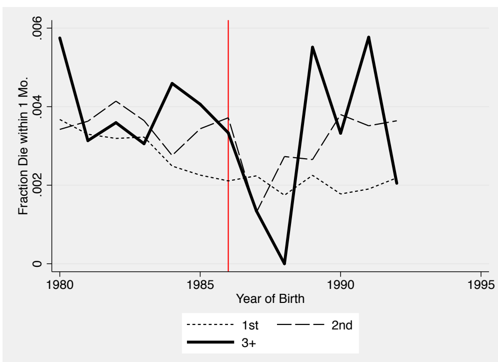


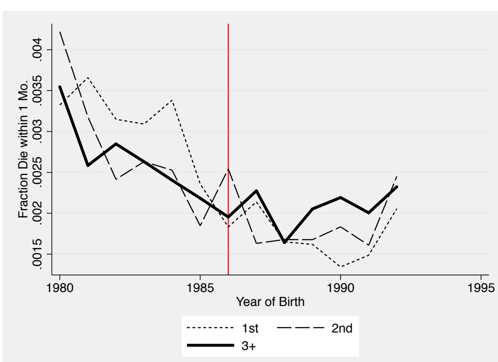
Figure 3: Male Mortality Rates within One Month of Life by Parity and Mother's Age

(a) Mother's Age = 18-22



(b) Mother's Age = 23-28

(c) Mother's Age = 29-34



(d) Mother's Age = 35+

Figure 4: Ratio of Female and Male Mortality Rates within One Month of Life by Parity for All Mothers – Demeaned by Mother's Age and Mother's Age × Post-Reform

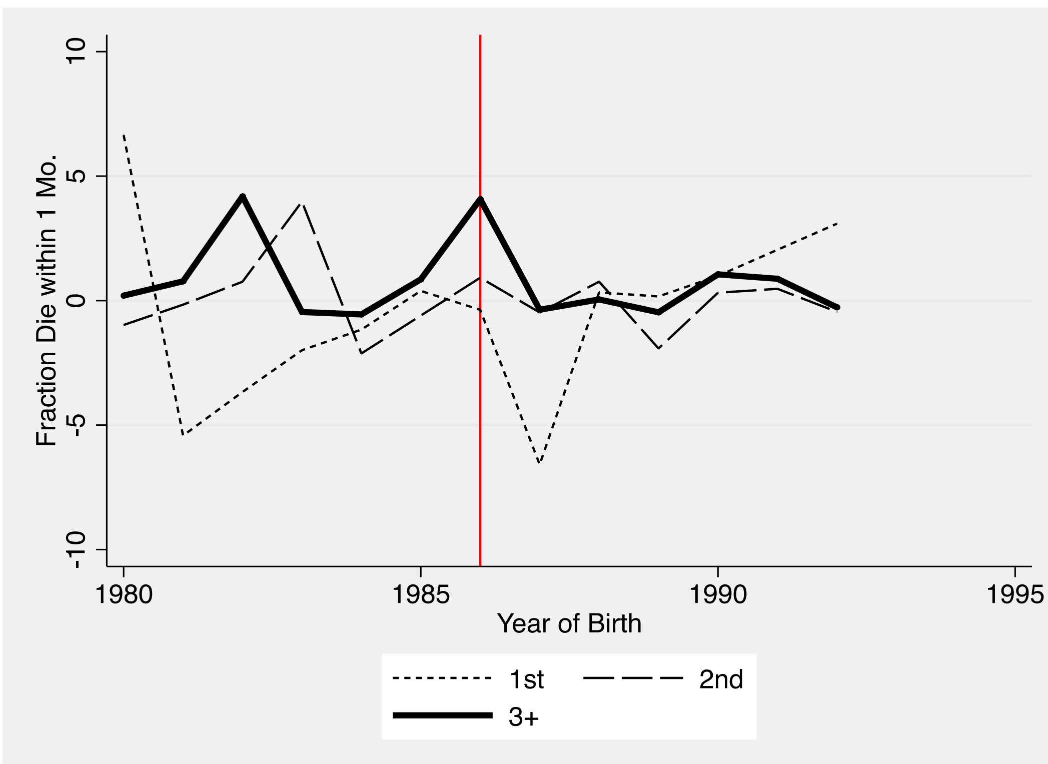
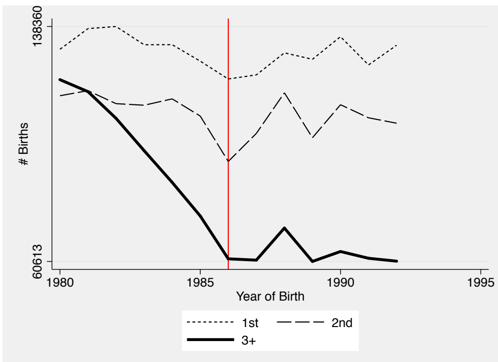
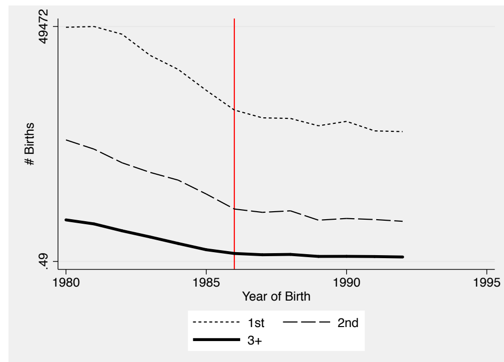


Figure 5: The Number of Births by Parity and Mother's Age

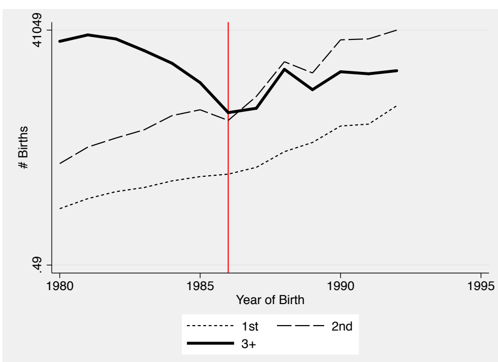
(a) All Births



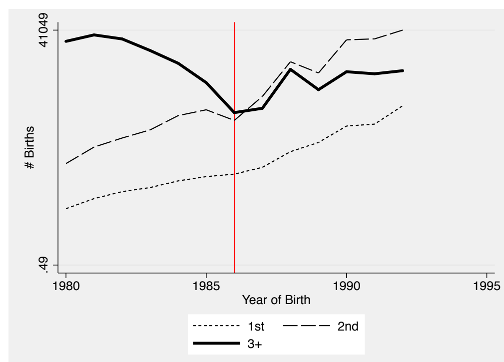
(b) Mother's Age = 18-22



(c) Mother's Age = 23-28



(d) Mother's Age = 29-34



(e) Mother's Age = 35+

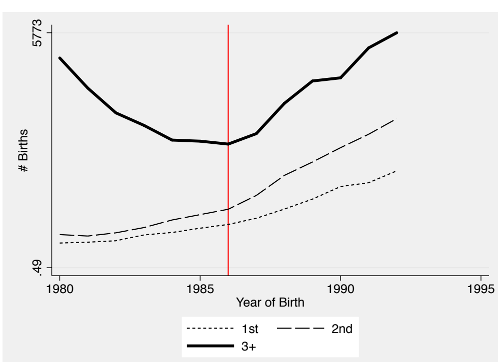


Figure 6: The Effect of the Reform for Each Year Since the Reform

(a) The Effect on Sex for 2nd and 3+ Parity Children

(b) The Effect on Relative Female Mortality Rates within One Month of Life

ONLINE APPENDIX

A Causes of Death

We identify the top ten causes of deaths and the share of deaths attributed to each cause for the first month of life for each birth parity in Online Appendix Table A.1. The most common causes of death include congenital anomalies of the heart; disorders relating to short gestation and low birthweight; superficial injuries of the head; pneumonia; multiple fractures on the torso and ribs; sudden death; ill-defined intestinal infections; septicemia (bacterial infection in the blood); and superficial injuries of the arms.

Interestingly, the importance of each cause varies across birth parities. Our focus on death that results from parental neglect makes it interesting to investigate whether the likelihood of death from neglect varies across parities. It is impossible to provide conclusive summary statistics on the contribution of parental neglect to mortality across parities with our data. However, we can perform a rough approximation by making a few simple assumptions. For example, we can we assume that death by congenital anomalies of the heart and disorders relating to short gestation and low birthweight are less likely to be due to parental neglect than death from injuries or sudden death. Therefore, we can approximate the percentage of deaths that may be due to parental neglect for each birth parity. Since this categorization is very speculative, we experiment with three definitions. The first includes deaths from any physical injury and sudden death. The second additionally includes deaths from pneumonia and septicemia. The logic for this is that a child's recovery from these diseases can be highly dependent on whether parents take the infant to the hospital when the first symptoms appear. The third group additionally includes deaths from unknown causes. Regardless of the definition we use, a comparison of the statistics at the bottom of each panel in Table A.1 suggests that parental neglect is likely to be more important for higher parity deaths.

B Simple Model of Fertility, Sex-Preference and Mortality

Our empirical analysis examines the the effect of the legalization of abortion, which we interpret as having reduced the cost of sex selection, on the sex and mortality of children born. To motivate and help interpret these estimates, we provide a simple framework to illustrate how the reform can affect the decision of a woman to give birth, the number of males born by the woman and the amount of effort the woman exerts towards providing care for the child, which we assume to increase a child's probability to survive.

Formally, consider a woman's decision to become pregnant. Suppose that the payoff from not having children is zero. Once pregnant, the woman learns the sex of the child, and she can abort the child at a cost c. If the child is born, the woman receives utility

$$(\theta - \lambda)\, e,$$

where $e = \{\underline{e}, \overline{e}\}$ is the amount of effort she puts into childcare with $\overline{e} > \underline{e} > 0$. The parameter λ captures the mother's cost of child-rearing. We interpret cost broadly to include physical, logistical, financial and psychological costs.

θ captures the mother's utility from the child. If a son is born, $\theta = \theta^*$. If a girl is born, $\theta = \theta^* - \mu$ for $\mu > 0$, which is a parameter capturing the preference for sons.

We can solve the woman's problem by backward induction. Once the child is born and θ is realized, the woman will exert high effort towards caring for the child only if $\theta \geq \lambda$. In this case, the mother's welfare is $(\theta - \lambda)\,\overline{e}$. Otherwise, she will exert low effort, and her welfare is $(\theta - \lambda)\,\underline{e}$.[34] Note that, if a girl is born, the level of effort is weakly decreasing in $\lambda + \mu$. In other words, the higher the cost of childcare and the higher the son preference, the lower the level of effort is exerted by the mother conditional on having a girl.

Now, consider the woman's decision to abort a girl, where it is clear in this case that $\theta = \theta^* - \mu$.[35] She will abort a girl if

$$I\{\theta^* - \mu \geq \lambda\}\,(\theta^* - \mu - \lambda)\,\overline{e} + (1 - I\{\theta^* - \mu \geq \lambda\})\,(\theta^* - \mu - \lambda)\,\underline{e} < -c, \tag{6}$$

where $I\{\theta^* \geq \lambda + \mu\}$ is an indicator which equals 1 if $\theta^* - \mu \geq \lambda$ so that high effort is chosen in rearing a girl. Note that it is clear that the left hand side of (6) is decreasing in λ and μ. In other words, the higher the cost of childrearing and the higher the preference for sons, the lower the payoff conditional on having a girl and the more likely a pregnant woman is to choose to abort.

Finally, consider the woman's decision to bear a child. Clearly, if λ is very high, she will never have a child, independent of the cost of abortion. Furthermore, if λ is very low, she will always have a child,

[34] We have assumed without any loss of generality that high effort is chosen if the mother is indifferent.

[35] In our simplified framework, a woman would only ever choose to abort a girl, since if she also chose to abort a boy she would have never chosen to be pregnant in the first place. Note that we have assumed that accidental pregnancies are not possible for illustrative purposes. The results discussed in this section are robust to allowing for this possibility. This extension is not presented for brevity and available upon request.

independent of the cost of abortion. The interesting case then is for intermediate levels of λ, where the cost of abortion matters for the pregnancy decision. In this situation, a woman will become pregnant and plan to abort her child if she learns that it is a girl. She will thus become pregnant if

$$.5\,(\theta^{*}-\lambda)\,\bar{e} > .5c, \tag{7}$$

where we have taken into account that ex-ante, the woman expects aa 50% probability of having a girl and that she would always choose high effort conditional on having a boy.[36]

Given these decisions, we can thus summarize the effect of a reduction in the cost of abortion c. For illustrative simplicity and without any bearing on the results, suppose the cost is reduced from ∞ to zero. At the abortion stage, this leads to an increase in the abortion of girls for women for whom the left hand side of equation (6) is below zero. Clearly, this effect is most pronounced for women for whom the cost of childrearing λ and the son preference μ are high. At the pregnancy stage, this leads to an increase in the pregnancies of women hoping to use the option value of abortion in order to only have boys. In other words, women for whom the left hand side of equation (7) exceeds zero. Again, this effect is also most pronounced for women for whom λ and μ are high.[37] These two channels thus imply that, in the aggregate, there is an increase in the number of boys born relative to girls.

A natural question regards the effect of the change in policy on the aggregate amount of effort exerted in the rearing of girls who are born. In order to assess this, one has to consider how the total effort exerted by the women who continue to bear girls after the change in policy compares to the total effort exerted by the women who stop bearing girls after the change in policy. Clearly, if women continue to bear girls despite the option to abort them, it implies that the left hand side of (6) exceeds zero, which can only be true if they exert high effort conditional on having girls. On the other hand, if they choose to abort them, it must be because that same term is below zero, which can only be true if they exert low effort conditional on having girls. This means that the legalization of abortion will lead to an increase in the level of effort exerted in raising girls in the aggregate because the women who would have otherwise exerted low effort towards girls prior to the reform choose to abort female fetuses after the reform.

[36]If she chose low effort, it would imply that $\theta^{*}-\lambda<0$, so that the payoff from having a boy is negative.

[37]Note that even though μ does not enter the left hand side of (7), it does influence the woman's original decision to have chosen to not become pregnant and therefore has an effect.

Table A.1: Most Prevalent Causes of Death by Birth Parity

Cause of Death (ICD Code)	% of Death
Parity = 1	
746 Other congenital anomalies of heart	12.6
765 Disorders relating to short gestation and low birthweight	6.78
910 Superficial injury of face, neck, and scalp except eye	6.78
486 Pneumonia, organism unspecified	5.82
819 Multiple fractures involving both upper limbs, and upper limb with rib(s) and sternum	5.66
Unknown	5.63
798 Sudden death, cause unknown	3.05
009 Ill-defined intestinal infections	2.51
038 Septicemia	2.4
913 Superficial injury of elbow, forearm, and wrist	2.31
Head Injury + Torso Injury + Sudden Death + Arm Injury	**17.8**
+ Pnuemonia + Septicemia	**26.02**
+ Unknown	**31.65**
Parity = 2	
746 Other congenital anomalies of heart	12.09
765 Disorders relating to short gestation and low birthweight	8.53
910 Superficial injury of face, neck, and scalp except eye	6.46
Unknown	5.64
486 Pneumonia, organism unspecified	5.57
819 Multiple fractures involving both upper limbs, and upper limb with rib(s) and sternum	5.38
798 Sudden death, cause unknown	3.28
009 Ill-defined intestinal infections	2.93
913 Superficial injury of elbow, forearm, and wrist	2.59
038 Septicemia	2.35
Head Injury + Torso Injury + Sudden Death + Arm Injury	**17.71**
+ Pnuemonia + Septicemia	**25.63**
+ Unknown	**31.27**
Parity = 3+	
746 Other congenital anomalies of heart	11.14
910 Superficial injury of face, neck, and scalp except eye	8.08
486 Pneumonia, organism unspecified	7.4
819 Multiple fractures involving both upper limbs, and upper limb with rib(s) and sternum	6.3
765 Disorders relating to short gestation and low birthweight	5.15
Unknown	4.94
009 Ill-defined intestinal infections	3.02
798 Sudden death, cause unknown	2.85
038 Septicemia	2.6
913 Superficial injury of elbow, forearm, and wrist	2.37
Head Injury + Torso Injury + Sudden Death + Arm Injury	**19.6**
+ Pnuemonia + Septicemia	**29.6**
+ Unknown	**34.54**

Source: Taiwan's National Death Registry, 1980-1996.

Table A.2: The Effects of the Reform on Sex at Birth for Each Year Since the Reform

		Dependent Variable: Fraction of Males					
		(1)	(2)			(3)	(4)
		coef	p-val			coef	p-val
Dep Var Mean				**0.52**			
Birth Parity = 2 x				Birth Parity = 3+ x			
	Year=1981	0.000196	(0.114)		Year=1981	0.00228	(0.017)
	Year=1982	0.00145	(0.019)		Year=1982	0.000789	(0.012)
	Year=1983	-0.00495	(0.117)		Year=1983	-0.00417	(0.023)
	Year=1984	-0.00275	(0.087)		Year=1984	-0.000218	(0.049)
	Year=1985	-0.00407	(0.108)		Year=1985	-0.00265	(0.023)
	Year=1986	-0.00564	(0.122)		Year=1986	-0.00331	(0.037)
	Year=1987	0.00145	(0.118)		Year=1987	0.00826	(0.029)
	Year=1988	-0.0025	(0.113)		Year=1988	0.00894	(0.049)
	Year=1989	-0.000953	(0.128)		Year=1989	0.0176	(0.023)
	Year=1990	0.00434	(0.120)		Year=1990	0.0296	(0.049)
	Year=1991	0.00159	(0.099)		Year=1991	0.0288	(0.023)
	Year=1992	-0.00178	(0.092)		Year=1992	0.0221	(0.049)
Observations				156			
R-squared				0.863			

Notes: Observations are at the cell-level, which are constructed by aggregating individual-level data by birth parity, birth year and mother's age. The regressions are weighted by cell size. p-values, clustered at the birth-parity and year level, are presented in the parentheses. The raw data are reported by the Taiwan's National Birth Registries, 1980-1992. The regression controls for birth parity fixed effects, birth year fixed effects, post x mother's age=22-28, post x mother's age =29-35, post x mother's age=35+ and mother's age fixed effects.

Table A.3: The Effects of the Reform on Relative Female Mortality (1 Month) for Each Year Since the Reform

	(1) coef	(2) p-val		(3) coef	(4) p-val
	Dependent Variable: # Deaths per 1,000 Births				
Dep Var Mean			**2.251**		
Girl x Birth Parity = 2 x			Girl x Birth Parity = 3+ x		
Year=1981	0.172	(0.000)	Year=1981	0.169	(0.000)
Year=1982	0.389	(0.000)	Year=1982	0.318	(0.048)
Year=1983	0.431	(0.000)	Year=1983	0.476	(0.001)
Year=1984	0.425	(0.000)	Year=1984	0.312	(0.007)
Year=1985	0.131	(0.024)	Year=1985	-0.674	(0.000)
Year=1986	-0.303	(0.000)	Year=1986	-0.388	(0.004)
Year=1987	0.299	(0.045)	Year=1987	-0.237	(0.304)
Year=1988	-0.025	(0.645)	Year=1988	-0.147	(0.152)
Year=1989	0.142	(0.046)	Year=1989	-0.701	(0.000)
Year=1990	-0.067	(0.367)	Year=1990	0.059	(0.659)
Year=1991	0.336	(0.000)	Year=1991	0.134	(0.241)
Year=1992	-0.300	(0.000)	Year=1992	-0.526	(0.000)
Observations			312		
R-Squared			0.827		

Notes: Observations are at the cell-level, which are constructed by aggregating individual-level data by birth parity, birth year, mother's age and the sex of the child. For birth parity, individuals are categorized into groups of 1st, 2nd and 3+ births. For mother's age, individuals are categorized into groups for mothers' ages of: 23-28, 29-34, 35+. The regressions are weighted by cell size. Bootstrapped p-values, clustered at the birth-parity and post-reform level, are presented in the parentheses. The raw data are reported by the Taiwan's National Birth Registries, 1980-1992. All regressions control for all underlying interaction and main effects. The controls are: Post x girl, Post x parity=2, Post x parity=3, girl x parity=2, girl x parity=3, a dummy variable for girl, two dummy variables for birth parity, birth year fixed effects, addition include: post x mother's age=23-28, post x mother's age=29-34, post x mother's age=35+, girl x mother's age=23-28, girl x mother's age=29-34, girl x mother's age =35+, and the four dummy variables for mother's age groups.