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MORE WOMEN MISSING, FEWER GIRLS DYING: THE IMPACT OF ABORTION ON SEX RATIOS AT BIRTH AND EXCESS FEMALE MORTALITY IN TAIWAN

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ABSTRACT

Many countries with "deficits" in their female population see banning sex-selective abortion as a way to curb the observed sex imbalance. However, they rarely discuss the potentially negative unintended consequences of this ban on female survival rates as parents may be forced to substitute post-natal for pre-natal sex-selection. This paper presents novel empirical evidence on the impact of access to abortion on sex ratios at birth and relative female infant mortality. We use the universe of birth and death registry data from Taiwan and exploit plausibly exogenous variation in the availability of sex-selective abortion caused legislative changes to identify the causal effects of sex-selective abortion on sex ratios at birth and excess female mortality. We find that sex-selective abortion increased the fraction of males at birth by approximately 0.7 percentage-points, accounting for approximately 100% of the observed increase in sex ratios at birth during the 1980s; and it decreased relative female neo-natal mortality by approximately 61%. We estimate that approximately 13 more female infants survived for every 100 aborted female fetuses.

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1 Introduction

Missing women, a term coined by Amartya Sen, refers to the observation that in countries such as China, India, Albania, Taiwan and South Korea, only 48.4% of the existing population is female, whereas in most of Western Europe and the U.S., the proportion is 50.1%. There is much concern that sex imbalance in the population could lead to increased crime rates or distort the marriage market.[1] And despite rapid economic growth and social "modernization", the phenomenon has only increased over time. A large part of the increase in the observed male biased sex imbalance is due to an increase in the fraction of males at birth, a trend particularly stark in Asia in the 1980s and 90s. Observers have speculated that this is due to improved access to sex-selective abortion in combination with preferences for smaller family sizes. This is certainly consistent with the data which show a tremendous increase in the percentage of males born during this period, particularly for higher birth parities, and a decrease in family size. Figures 1A-1C show fraction of males at birth for Taiwan, China and South Korea which increased from roughly around 0.50-0.52 (the same as for countries not known for boy-bias such as the U.S. or Western European countries) to much higher levels. Most of the increase is observed for higher parities. By 1990, the fraction of males at births for Taiwan, China and S. Korea had risen to approximately 0.54, 0.56 and 0.66. Given the large potential impact that sex-selective abortion has on sex ratios at birth, it is perhaps surprising that there are no studies to date which examine the causal impact of sex-selective abortion or quantifies its contribution to the observed imbalance.[2] It is all the more surprising when we note that all sides in the recent heated debate

[1] Angrist (2002) and Samuelson (1985) study the long-run impact of sex imbalance on the marriage market.

[2] Studies in demography such as Gu and Roy (1995) and Park and Cho (1995) have described the changes over time in sex ratios at birth by birth parity in several East Asian countries. And studies in economics such as Lin and Luoh (2007) and Abrevaya (2008) have also remarked on the differences by birth parity. While all of these studies remark on the role of sex-selective abortion, they do not link the changing trends by birth parity to a legislative reform that would allow causal

on the determinants of the observed sex-imbalance agree on one thing: the likely importance of access to sex-selective abortion.[3] Absent concrete evidence, policy makers in many countries (e.g. China, India and South Korea), have nevertheless attempted to curb sex imbalance by prohibiting pre-natal sex-selection. While this may lead to a decrease in sex ratios (heretofore defined as the fraction of males) at birth, it may also have unintended negative consequences for relative female survival rates by forcing parents with strong boy-preferences to substitute from pre-natal to post-natal selection.[4] This paper fills the gap in the literature and presents novel empirical evidence on the causal effect of access to sex-selective abortion on sex ratios at birth and EFM (which refers to female relative to male infant mortality in this paper) by exploiting variation in access to abortion caused by a legislative reform in Taiwan, when the technology for detecting sex prenatally was already available.

The principal methodological contribution of this study is to resolve identification issues that have typically hindered past studies of the effect of sex-selective abortion. A simple cross-sectional comparison of observed population sex imbalances between regions with access to this technology and regions without access faces the problem that adoption of the technology may be driven by a region's underlying demand for boys. If regions with stronger boy-preferences more readily adopt the technology, then the underlying preferences will

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identification of the impact of sex selective abortion.

[3] Burgess and Zhuang (2001), Edlund (1999) Grogan (mimeo), Gu and Roy (1995), Li (2002), presents mixed evidence on the relationship between income and sex imbalance. Ben Porath (1967, 1973, 1976), Burgess and Zhuang (2002), Clark (2000), Duflo (2002), Das Gupta (1987), Foster and Rosenzweig (2001), Qian (2007), Rholf et al. (2005), Rosenzweig and Schultz (1982), Thomas (1991) and Thomas et al. (1994) study the effect of relative female socio-economic status on outcomes for girls relative to boys. Ebenstein (2007), Li (2002) and Qian (2006) examine the effect of family planning policies on sex imbalance in China. Lin and Luoh (200), Norberg (2004) and Oster (2005) study the effect of biological causes on sex imbalance. Chu (2001) presents a descriptive analysis of the practice of prenatal sex selection in rural central China using detailed survey data.

[4] This rationale is similar to the one used in Donohue and Levitt's (1999) studies of the impact of legalizing abortion on crime rates. They argued that because access to abortion allowed parents to avoid having unwanted children, children born after abortion became legal were on average better treated and less likely to commit crime.

confound analysis of the causal effect of access *per se*. Hence, the correlation will overestimate the true effect of access to sex-selective abortion on the fraction of males at birth. The bias for estimating the effect on EFM is ambiguous because there are two possibilities. On the one hand, if regions that adopt sex-selective abortion also face lower costs in killing girls postnatally, then a negative cross-sectional correlation between access to sex-selective abortion and EFM will underestimate the magnitude of the true effect. On the other hand, if regions that adopt sex-selective abortion are regions with stronger preferences for substituting pre-natal sex-selection for post-natal sex selection, then the observed correlation will overestimate the magnitude of the true effect.

To address the issue of endogenous adoption, we exploit two sources of variation. First, we exploit the plausibly exogenous variation in access to sex-selective abortion caused by the legalization of abortion in Taiwan in 1985/86. Technology for pre-natal sex detection was already available in Taiwan when abortion was legalized. Hence, we interpret the legalization of abortion as a plausibly exogenous increase in access to sex-selective abortion. Second, we exploit variation in demand for boys associated with higher birth orders and older mothers. If parents wish to have a boy, then the preference should be more binding for parents who face more uncertainty (for financial or biological reasons) about their ability to have more children. In addition to examining sex ratios at birth and sex-differential infant mortality rates, we use the same empirical strategy to investigate the effects of access to abortion on the composition of children born and parental characteristics.[5]

[5] Previous studies such as Park and Cho (1995), Gu and Roy (1995) and Lin and Luoh (2007)\ have observed the increase in sex ratios at birth, and the faster increase in higher parity births in China, South Korea and/or Taiwan. However, these studies did not link the changes to any "exogenous"\ factors that would enable the identification of a causal effect. In China and South Korea, there was no legislation that legalized abortion. And in Lin and Luoh (2007), the Taiwanese data used were all collected after the abortion legalization.

Using an individual level dataset constructed from birth and death registries for all individuals born in Taiwan during 1982-89, we find that the legalization of abortion significantly increased the fraction of males born. The effect comes entirely from third and higher-parity births and children born to mothers over the age of 28. For those groups, abortion increased the fraction of males born by 0.7 percentage-points for post-reform cohorts on average (from 51.7 percentage-points in 1982-84 to 53.5 percentage-points by 1989), accounting for nearly 100% of the observed increase in sex imbalance during this period. The results on sex-differential mortality show that legalizing abortion decreased EFM by up to 61%. Our results suggest that approximately 13.3% of parents selecting postnatally before the reform would have substituted to abortion as a method of sex-selection. Taken literally, this suggests that for every 100 abortions of female fetuses, approximately 13 lives of girls born are saved.

Studying the effect of sex-selective abortion in Taiwan has both advantages and disadvantages. On the one hand, the data are much better than other countries with boy biased sex imbalances. The legislative reform legalizing abortion allows us to have plausibly exogenous variation on access to sex-selective abortion. And, unlike China and India, it was legal to reveal the sex of the fetus and there were no family planning policies which restricted the number of children. This makes interpreting the results beyond the direct context of the study relatively easier. On the other hand, Taiwan is wealthier than China and India and also has a much lower infant mortality rate on average. Our study may underestimate what the effects of a similar reform in those countries would be. Hence, caution should be used when applying our results outside of the Taiwanese context.

That said, our results should still show policy makers that that banning sex-selective abortion will have a large effect in decreasing the observed sex imbalance and that there is a tradeoff between decreasing the fraction of males at birth and increasing relative female survival.

Unless governments can also incentivize parents to care for girls born, a ban on sex-selective abortion may lead to an increase in EFM. Our results taken literally and applied to the mainland China and India contexts suggests that effectively banning sex-selective abortion could increase the number of girls born each year by approximately 1.6 million but cause female neo-natal mortality to increase by approximately 156,800 girls in the two countries combined.

The paper is organized as follows. Section two discusses the empirical strategy and background of the technology and policy reforms in Taiwan. Section three describes the data. Section four presents the empirical results. Section five interprets the results. Section six offers conclusions.

2 Empirical Strategy and Background

Sex-selective abortion requires two technologies: one that reveals the gender of the fetus and another that facilitates the miscarriage of the fetus. While there are several procedures prevalently used during pre-natal care in developed countries today that also can reveal the sex of the fetus (e.g. *amniocentesis*, *chorionic villus sampling*), the most inexpensive and easily available method in both developed and developing countries is Ultrasound B. This study estimates the effect of legalizing abortions given the existing technologies for sex-detection. It does not separately identify the effect for different methods of abortion. Hence, we will focus the background discussion on Ultrasound B. It is the technology that is most widely used in developing countries today, and was the most prevalent method in Taiwan for the period of our study. Ultrasound B was first introduced into Taiwan during the early 1980s. It can reveal the sex of the fetus beginning in the 16th week of gestation. Accuracy is greatly increased by the 20th week. Ultrasound B machines are inexpensive to manufacture and relatively easy to use. The procedure for revealing the sex is not invasive and the results can be easily interpreted by a trained technician. In Taiwan, Ultrasound B is used in standard pre-natal care and is available from registered medical doctors. Unlike China and India today, revealing the sex of the fetus has never been prohibited in Taiwan.

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*. During the mid-1980s, a growing demand for safe abortions as a method of family planning, and a growing feminist movement pushed Taiwanese legislators to make abortion legal. The law was initially relaxed in 1984 to allow couples with known a genetic disease to induce an abortion. At that time, if a physician performed an unauthorized abortion, he/she was fined approximately NT$20,000, roughly 15% of the contemporaneous per capita GDP. The law was further relaxed in 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 with physicians who performed abortions during the 1980s, the cost of an abortion was on average 1% of average household income at the time.

Our empirical strategy interprets 1985-1989 as the "text" reform period. This will likely 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. Officially, the relaxation of the *Eugenics Law* passed on Jan 1, 1985. It stated that women can get an abortion if (a) the women was raped (b) the fetus has some genetic disease or (c) the pregnancy would affect the mental health of the women and the family. (The added clause (c) effectively legalized non-medically motivated abortions). However, "practicing abortion" was not made explicitly legal. Hence, according to anecdotal evidence from interviews we conducted, doctors waited to see if they would be prosecuted if they performed abortions using reason (c). The government also made no attempts to publicize the reform or the details of implementation. For example, only "certified" hospitals were allowed to conduct the procedure (Liu, 1995). Hence, the policy was phased in over 1985 as practitioners gradually learned how and where to conduct legal abortions. Second, the law was applied to all contemporaneous pregnancies but abortion was allowed only up to the 24th week (6 months). Hence, the first cohort that was exposed to sex selective abortion are not born until 4 months into 1985. Therefore, the legislation was not in effect for a full calendar year, until 1986.[6]

To see if the legalization actually affected the number of abortions, we use survey data from the *Knowledge, Attitudes, and Practice of Contraception in Taiwan* (KAP) on the cumulative abortion histories of women between the ages of 18-44. This is the only source of data on abortion that we know of. The KAP data faces all of the problems of self reported data (e.g. women may not wish to report abortions undertaken for non-medical reasons.) And the survey does not happen with enough frequency for us to see if there is a trend break in the increase of abortions after the reform. That said, the data reported supports our argument that legalizing abortion effectively increased the number of abortions. The two years closest to the reform when the survey was conducted were 1985 and 1992. It asks women about whether they have ever had an abortion. It does not distinguish abortions for medical reasons from other abortions. The data show that the percentage of women who have ever had abortions increased from 23% in the 1985 to approximately 27% in 1992. Figure 2A shows the fraction of women who have ever had an abortion by age group. It shows that abortions increased between 1985 and 1992 for all age groups. Figure 2B shows the increase in the fraction of women who have ever had an abortion by

[6] Changing the cutoff from January 1985 to April 1985 does not affect the main DD estimates. However, changing the year-by-year specification to a month-by-month specification creates very noisy estimates due to the small number of births each month. Hence, for simplicity, we use January 1985 as a cutoff.

age group. It shows that women who were 29-35 experienced the largest increase (7.5 percentage-points). Figures 2C and 2D plot the fraction of women who have ever had abortions, and its increase over time for all education levels. There is no discernible pattern in the increase across different education levels.

At the time of the reform, Ultrasound B, amniocentesis and chronic virillus sampling were all available as different ways of detecting the sex of the fetus. The legalization of abortion combined with the use of these 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. 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 in Taiwan over time. The legalization of abortion in 1985 was followed by a large increase in the number of Ultrasound B machines.[7] Toshiba, who has had the largest market share in Ultrasound B machines in Taiwan, reports that doctors were quite open in their desire to use these machines to reveal the sex of the fetus. To obtain a machine for a private office, a physician must be a member of the *Society of Ultrasound in Medicine*. From 1984 to 1989, the number of doctors in this organization increased from 557 to 3024. While doctors' primary reasons for increasing the use of Ultrasound B machines was to meet the rising demand for sex-detection, using Ultrasound B during routine pre-natal care may have also increased the quality of pre-natal care more generally.

We exploit the legalization of abortion in 1985/86, when Ultrasound B was already available, to estimate the causal effect of sex-selection on the fraction of males born. In addition to birth year/cohort variation, we also exploit the variation in birth order and mother's age. For

[7] Hospitals are required to register "precious machines" which in the 1980s largely referred to ultrasound B. The number of ultrasound B machines registered by hospitals increased by many orders of magnitude.

parents who wish to have a boy, they are more likely to sex-select if their ability to have another child and try for a boy is lower. This decrease in ability may reflect either biological constraints due to the mother's age or financial constraints due to the existing household size.[8] Note that because Ultrasound B was already available when abortion was legalized, it would have been possible for parents to select the sex of the child using abortion illegally prior to the reform. Hence, our analysis examines the effect of increasing access to abortion by legalizing abortion rather than the effect of introducing abortion.

The identification for estimating the effect of sex-selective abortion on sex ratios at birth relies on the assumption that no other changes occurred at the time that abortion was legalized that would decrease the cost of sex-selection *and* decrease the cost more for higher birth parities and older mothers. For example, the increased use of Ultrasound B improved the quality of overall pre-natal care. If male fetuses are more vulnerable, then males may respond more positively to this improvement. In this case, the fraction of males at birth may increase even absent sex-selective abortion. However, this should be independent of birth order or the mother's age. In other words, the identification assumption is only violated if the improvement in pre-natal care affects males more positively than females *and* has larger effects at higher birth parities or with older mothers. There is no reason to believe this is true. To be cautious, we investigate this possibility by examining the effect of the reform on the composition of boys born relative to girls born.

Similarly, the identification for estimating the effect of sex-selective abortion on sex-differential mortality relies on the assumption that therewas no improvement in medical technology that would have affected infant mortality for higher-parity births more *and* affect

[8] The assumption that older mother's and higher birth-parities are more likely to be affected is consistent with the findings of Chu (2001).

girls and boys differentially.

We first estimate the effect of legalizing abortion by birth order and birth year. This has an advantage over a simple differences-in-differences specification in that it allows us to observe the timing of the effect of access to abortion. For example, if there was latent demand for sex-selective abortion, then we would expect the reform to affect sex ratios of individuals born close to the 1985/86. The simpler pre-post differences-in-differences specification has the pitfall that it would capture changes that occurred at any time after the reform.

$$\mathrm{Male}_{\mathrm{it}} = \sum_{\mathrm{i}=2}^{3} \sum_{\mathrm{t}=1983}^{1989} \beta_{\mathrm{it}} \left(\mathrm{Ord}_{\mathrm{i}} * \mathrm{Born}_{\mathrm{t}}\right) + \gamma_{\mathrm{i}} + \rho_{\mathrm{t}} + \varepsilon_{\mathrm{it}} \tag{1}$$

We regress the fraction of males of birth order *i* and birth year *t*, $Male_{it}$, on: the interactions of dummy variables for being the second birth and the third or higher-parity birth, Ord_i, and dummy variables for being born in year *t*, $Born_t$; birth order fixed effects, γ_{i}; and birth year fixed effects, ρ_{t}. The reference group is comprised of first-born children. It and all of its interactions are dropped. If access to abortion increased boy-biased sex selection, then the coefficients for $\beta_{2\mathrm{t}}$ and $\beta_{3\mathrm{t}}$ should be larger for individuals born after 1985. If parents are more likely to sex select at higher birth orders, then $\beta_{3\mathrm{t}} > \beta_{2\mathrm{t}}$. More specifically, the difference should be larger in magnitude for cohorts born after 1985 $\left(\beta_{3,\mathrm{t}=85} - \beta_{2,\mathrm{t}>=85}\right) > \left(\beta_{3,\mathrm{t}<85} - \beta_{2,\mathrm{t}<85}\right)$

Next, we estimate the effect of legalizing abortion on the sex imbalance by mother's age. We separate mothers into four age groups: 18-22, 23-28, 29-35 and 35 and above. We estimate the following equation

$$\mathrm{Male}_{\mathrm{mt}} = \sum_{\mathrm{t}=1983}^{1989} \sum_{\mathrm{m}=2}^{4} \beta_{\mathrm{mt}} \left(\mathrm{morn}_{\mathrm{m}} * \mathrm{Born}_{\mathrm{t}}\right) + \gamma_{\mathrm{m}} + \rho_{\mathrm{t}} + \varepsilon_{\mathrm{mt}} \tag{2}$$

We regress the fraction of males for individuals born to mothers age *m* and birth year *t*, $Male_{mt}$, on: the interaction s of dummy variables indicating the mother's age, *m*, is 23-28, 29-35 or greater than 35 mom_m and dummy variables for being born in year *t*, $Born_t$; mother's age fixed effects, γ_m; and birth year fixed effects, ρ_t. The reference group is comprised of children born to mothers who are 18 to 21 years of age. It and all of its interactions are dropped. If access to abortion increased boy-biased sex selection, then the coefficients for β_{2t}, β_{3t} and β_{4t} should be larger for individuals born after 1985. If older mothers are more likely to select boys, then $\beta_{4t} \geq \beta_{3t} \geq \beta_{2t}$. More specifically, the difference should be larger in magnitude for cohorts born after 1985.

After we check that the timing of the effect is consistent with our identification strategy, we estimate a simpler specification where we group individuals to those born before the reform and those born afterwards to better assess the magnitude and statistical significance of the effect,

$$\text{Male}_{it} = \sum_{i=2}^{3} \beta_{it}\left(\text{Ord}_i * \text{Post}_t\right) + \gamma_i + \rho_t + \varepsilon_{it} \tag{3}$$

We regress the fraction of males for individuals of birth order *i* and birth year *t* , $Male_{it}$, on: the interaction terms between dummy variables for being the second birth and third and higher-parity births, Ord_i ,and a dummy variable for being born in 1985 or afterwards, $Post_t$; birth order fixed effects, γ_i; and birth year fixed effects, ρ_t. The reference group is comprised of first-born children. It and all of its interactions are dropped. We estimate a similar equation to assess the magnitude and statistical significance of the effect of sex-selective abortion on outcomes by mother's age.

$$\text{Male}_{mt} = \sum_{m=2}^{4} \beta_{mt}\left(\text{morn}_m * \text{Post}_t\right) + \gamma_m + \rho_t + \varepsilon_{mt} \tag{4}$$

We regress the fraction of males for individuals of birth year *t* whose mother was *m* years old when giving birth, $Male_{mt}$, on: the interaction term between a dummy variable indicating that the mother is 22 to 28 years of age, 29 to 35, or 35 and over, mom_m, and a dummy variable for being born after 1985, Post_t; mother's age fixed effects, γ_{m}; and birth year fixed effects, ρ_{t}. The reference group is comprised of children born to mothers who are 18 to 21 years of age.

Finally, to examine the interaction effect of higher birth order and mother's age on sex ratios at birth, we further interact birth order fixed effects with a continuous variable for mother's age.

$$
\begin{aligned}
\mathrm{Male}_{\mathrm{itz}} = & \sum_{\mathrm{i}=2}^{3} \beta_{\mathrm{i}}\left(\mathrm{momage}_{\mathrm{z}} * \mathrm{Ord}_{\mathrm{i}} * \mathrm{Post}_{\mathrm{t}}\right) \\
& + \sum_{\mathrm{i}=2}^{3} \theta_{\mathrm{i}}\left(\mathrm{Ord}_{\mathrm{i}} * \mathrm{Post}_{\mathrm{t}}\right) + \sum_{\mathrm{t}=2}^{3} \pi_{\mathrm{i}}\left(\mathrm{Ord}_{\mathrm{i}} * \mathrm{momage}_{\mathrm{z}}\right) \\
& + \lambda\left(\mathrm{momage}_{\mathrm{z}} * \mathrm{Post}_{\mathrm{t}}\right) + \delta_{\mathrm{i}} + \gamma_{\mathrm{z}} + \varepsilon_{\mathrm{imt}}
\end{aligned} \tag{5}
$$

We regress the fraction of males for individuals of birth order *i* birth year *t* whose mother was *z* years old when giving birth, $Male_{itz}$, on: the triple interaction term between a birth order dummy variable, Ord_i, a continuous measure of the mother's age at birth, $momage_z$, and a dummy variable for being born after 1985, $Post_t$; the interaction terms between birth order dummy variables, Ord_i and $Post_t$; birth order dummy variables and mother's age, Ord_i and $momage_z$; mother's age at birth and being born in 1985 and after, $momage_z$ and $Post_t$; birth order fixed effects, δ_{i}; mother's age fixed effects, γ_{z}; and birth year fixed effects, ρ_{t}.

The positive effect of the legalization of abortion on fraction of males born at higher birth orders and for older mothers could be due to two possibilities: 1) parents of those groups are using abortion as a method of pre-natal sex-selection; and/or 2) the improvement in pre-natal care caused by the increased use of Ultrasound B benefited male fetuses more than female fetuses, and the benefit was larger at higher birth orders and for older mothers. While there is no anecdotal evidence or medical reason to believe the latter to have been the case, we can investigate it with the data. The latter hypothesis implies that the increase in number of boys born relative to girls born is due to an increase in the number of "marginal" births for boys relative to girls (e.g. children who would not be born absent the improvement in care). Hence, we can test this hypothesis by examining whether the reform caused the fraction of "marginal" births to increase for boys relative to girls. Our measures of marginal births are limited by our data. We will examine the fraction of singleton and LBW births. The fraction of singleton births is the fraction of multiple births subtracted from one. Multiple births (e.g. twins, triplets, etc.) tend to be more difficult. They are strongly correlated with premature delivery and low birth weight. The pregnancy is typically more difficult relative to singleton births. Hence, if the increase in boys in higher-parity births is caused by a sex-specific-parity-specific benefit from increased access to Ultrasound B, then we should observe that the reform also increased the fraction of LBW and multiple births (and decreased the fraction of singleton births). For the sake of brevity, we focus this analysis on using variation from birth order. We estimate the following triple difference equation.

$$\begin{aligned} \mathrm{Y}_{\mathrm{its}} = {} & \sum_{\mathrm{i}=2}^{3}\sum_{\mathrm{t}=1989}^{1989}\beta_{\mathrm{it}}\left(\mathrm{Ord}_{\mathrm{i}} * \mathrm{Born}_{\mathrm{t}} * \mathrm{Male}_{\mathrm{s}}\right) \\ & + \sum_{\mathrm{i}=2}^{3}\sum_{\mathrm{t}=1989}^{1989}\delta_{\mathrm{it}}\left(\mathrm{Ord}_{\mathrm{i}} * \mathrm{Born}_{\mathrm{t}}\right) + \sum_{\mathrm{t}=1989}^{1989}\pi_{\mathrm{t}}\left(\mathrm{Male}_{\mathrm{s}} * \mathrm{Born}_{\mathrm{t}}\right) \\ & + \sum_{\mathrm{t}=2}^{3}\lambda_{\mathrm{i}}\left(\mathrm{Male}_{\mathrm{s}} * \mathrm{Ord}_{\mathrm{i}}\right) + \mathrm{X}_{\mathrm{it}}\alpha + \gamma_{\mathrm{i}} + \rho_{\mathrm{t}} + \varepsilon_{\mathrm{it}} \end{aligned} \tag{6}$$

We regress outcome *Y* of individuals of birth order *I*, birth year *t* and sex *s*, Y_{its} on: the triple interaction terms between dummy variables for birth order, Ord_i, birth year, Born_t, and sex, Male_s; the full set of double interaction terms; a vector of controls such as mother's age, mother's education, and father's education, X_{it}; birth order fixed effects, γ_{i}; and birth year fixed effects, γ_{i}. If the increase in fraction of males born was due to sex-differential effects of the improvement in pre-natal care, then we may find that the reform decreased the fraction of singleton births for boys, $\left(\beta_{3,t>=85} - \beta_{2,t>=85}\right) \leq \left(\beta_{3,t<85} - \beta_{2,t<85}\right) \leq 0$.[9] For LBW births, the hypothesis predicts that the reform led to an increase in LBW births for boys at higher birth orders, $\left(\beta_{3,t>=85} - \beta_{2,t>=85}\right) \geq \left(\beta_{3,t<85} - \beta_{2,t<85}\right) \geq 0$.

To estimate the effect of sex-selective abortion on sex-differential mortality rates, we exploit variation by birth year, sex, birth order, and mother's age in one regression.

$$\begin{aligned} \text{Death}_{\text{itsm}} &= \sum_{i=2}^{3} \beta_i \left(\text{Ord}_i * \text{Post}_t * \text{Male}_s\right) + \text{T}_m \left(\text{mom}_m * \text{Post}_t * \text{Male}_s\right) \\ &+ \sum_{i=2}^{3} \delta_i \left(\text{Ord}_i * \text{Post}_t\right) + \sum_{t=2}^{3} \lambda_i \left(\text{Male}_i * \text{Ord}_i\right) + \lambda\left(\text{mom}_m * \text{Post}_t\right) \\ &+ \pi\left(\text{Male}_s * \text{Post}_t\right) + \text{X}_{\text{it}}\phi + \theta_s + \gamma_i + \rho_t + \varepsilon_{\text{its}} \end{aligned} \tag{7}$$

We regress the fraction of deaths occurring within a specified number of months for individuals of birth order *i*, birth year *t*, sex *s*, mother's age *m*, $Death_{itsm}$, on: the triple interaction terms of a dummy variable for birth order, Ord_i, a dummy variable for being born in 1985 or after, $Post_t$, and a dummy variable for being male, $male_s$; the triple interaction term of a linear measure for the mother's age mom_m, $Post_t$ and $male_s$; the full set of double interaction terms; controls for mother's age, mother's education, father's education, an indicator variable for low birth weight, and an indicator variable for whether it was a singleton birth, $\text{X}_{\text{it}}\phi$; the sex fixed effect θ_s; birth order fixed effects, γ_i; and birth year fixed effects, ρ_t. The reference group is comprised of first births. It and all of its interactions are dropped. If sex-selective abortion increased survival rates for girls relative to boys, then $\pi \leq 0$. If the effects are larger for higher birth orders, then $\beta_3 \geq \beta_2 \geq 0$.

[9] This dataset is compiled from individual level data, not birth-level data. Hence, multiple births are weighted more than singleton births. This means that we over-weigh multiple births; or that we under-estimate the effect of the reform on the fraction of multiple births. The estimates using birth-level data are similar and are not reported in the paper.

One concern for the identification strategy is that the trend break in the fraction of males at birth for higher parity births or older mothers reflects a general trend in Asian countries during this period rather than the legalization of abortion in Taiwan. To examine this, we collected data on sex at birth by parity from South Korea and China. This data shows that while there is a general trend towards more males being born over time, and that this increase is larger for higher parities, the trend break we observe in Taiwan is unique in both its timing and its distinct discontinuity. The divergence in fraction of males born in South Korea and China begin earlier in 1980 and 1982, respectively. And in both cases, the divergence is gradual (see Figures 1A-1C). Hence, we believe that our strategy is unlikely to be confounded by general regional trends.

3 Data

This study uses the universe of data from Taiwan's National Birth Registries from 1982-1989 and Death Registries from 1982-1991 which is comprised of approximately 2.8 million individuals. The data is linked at the individual level. It reports region and year of birth, sex, birth weight, birth order, whether the child was part of a multiple birth, whether the birth was premature, mothers' marital status, mothers' and fathers' age and level of education. The data from the death registry reports whether a child dies within one, two, three, four, five, six, nine, eighteen, twelve and twenty-four months after birth. Both for the sake of brevity and because EFM is more likely to occur soon after birth, we focus on death within one month but also present results for death within six and twelve months. We restrict our sample to individuals born to mothers who were 18-45 years of age at the time of birth. For examining the effect of sex-selective abortion on the fraction of males born by birth order, the data is aggregated to birth order (first, second, and third and higher), birth year and birth county cells. For examining the effect of sex-selective abortion on the fraction of males born by mother's age, the data is aggregated to mother's age (18-21, 22-28, 29-35 and over 35), birth year and birth county cells. For the analysis on survival, the data is aggregated to sex, birth order, birth year and birth county cells; and sex, mother's age, birth year, and birth county cells. Cell sizes are always retained so that all regressions are weighted. The weighted regression results are numerically identical to regressions using data at the individual level. See Chou et al. (2007) for a detailed discussion of the microdata.

Table 1 shows the total number of births before and after the reform. Panel A shows that the number of births have decreased over time in all birth parities; and the decrease is disproportionately large in third and higher parity births. Both facts are consistent with the observation of an increasing preference for smaller family sizes during this period. Panel B shows the number of births by mother's age. Interestingly, it shows that while the number of births has decreased for younger mothers, older mothers have *more* children after the reform. Table 2 shows the descriptive statistics by birth order for individuals born before the reform (Panel I) and after the reform (Panel II). Panel III is the difference in means. It shows that on average, there are more males born after abortion was legalized, especially for higher-order births. There is also an increase in the occurrence of low birth weight and multiple (non-singleton) births, which may reflect an improvement in pre-natal care that facilitated more difficult births during this period. Column III also shows that mothers of children born after the reform are older, more educated and less likely to be married at the time of birth. Figure 3A plots the fraction of males by birth order and birth year. It shows that the fraction of males is similar across parities before the reform at approximately 0.517. This is slightly higher than the 0.51-0.515 fraction of males at birth observed in countries not known to have male-biased sex preferences such as the U.S. (Mathews and Hamilton, 2005).[10] But it is still within the range that demographers typically accept as "natural". For first and second births, there is no change over time. However, for third and higher-parity births, there is a clear trend break: the fraction of males increases steadily for children born after abortion was legalized, up to approximately 0.535 in 1989. Figure 3B plots the fraction of males by mother's age and birth year. It shows that before the reform, the fraction of males born was similar for all age groups. After the reforms, there is no change for young mothers (under 28). However, for mothers who were 29 to 35, the fraction of males born increased for cohorts born after the reform.

Figure 4A plots the natural log of total births over time by birth parity. It shows that there is a steady decrease in the number of higher parity births during the early 1980s, which flattens out in 1986. To see if the stop in the decline of total births for higher parity births is due to the increase in the number of births that we observe for older mothers, we plot the natural log of total births for young and old mothers in Figures 4B and 4C, respectively. Figure 4B shows that the total number of births for mothers under 28 years of age is declining steadily through out the 1980s. Interestingly, Figure 3C shows that for mothers over 28, the total number of births decreases gradually during the early 1980s, but increases dramatically after abortion is legalized. Together with the fact that older mothers are more likely to have higher parity births, this suggests that the halt in decline of the number of higher parity births is produced by the offsetting effect of older mothers having more children after the reform.

Table 3 shows the fraction of deaths within one month and twelve months for children born before and after the reform by sex and birth order. Note that mortality within the first month account for approximately half of mortality within twelve months. This suggests that neo-natal

[10] Table A in Mathews and Hamilton (2005)\ show that fraction of males at birth in Western Europe during 1999-2002 typically ranges from 0.51 to 0.514.

mortality is an important contributor to total infant mortality rates. The means show that Taiwan had very low rates of infant mortality, approximately 3 deaths per 1,000 births. At the same time, Taiwan's higher income neighbors, South Korea and Japan, had infant mortality rates of approximately 6 per 1,000 births.[11] Columns (1)-(2) and (4)-(5) show that mortality rates were higher for boys across birth orders for all cohorts. This is consistent with the widely held belief that males are more vulnerable during infancy. Columns (3) and (6) show changes in mortality over time for girls and boys, respectively. For the post-reform cohort, mortality rates decreased for both boys and girls, which could reflect an improvement in medical technology and/or the post-natal benefit of not forcing parents to have unwanted children. Column (7) is the sex-differential changes in mortality after the reform (column (3) subtracted from column (6)). The differences show that while mortality rates decreased more for boys for all birth parities, the difference for death within one month was smaller in magnitude for higher-parity births.

4 Empirical Results

4.1 The Effect on Fraction of Males at Birth

We first estimate the effect of legalizing abortion on the fraction of males by birth order by estimating equation (1). The estimates for $\hat{\beta}_{2t}$ and $\hat{\beta}_{3t}$ and their robust standard errors are shown in Appendix Table A2 columns (1) and (2). They are statistically significant for post-reform cohorts at the 1% and 5% levels. The coefficients are plotted in Figure 5A. The figure shows that sex ratios were similar for second births and higher-parity births relative to first births before the reform. After the reform, the fraction of males increased dramatically for third and higher-parity births while staying the same for second births. The coefficients for the third

[11] Source: World Development Indicators.

and higher-parity births are plotted with their 95% confidence intervals in Figure 5B.Note that for third and higher-parity births, the reform had increased the fraction of males born by two percentage-points by 1989, which is the observed increase in fraction of males at birth in Figure 3A.

Next, we estimate the effect of legalizing abortion on the fraction of males at birth by mother's age by estimating equation (2). The coefficients and their robust standard errors are shown in Appendix Table A2, Columns (3)-(5). The estimates are statistically significant at the 5% and 1% levels for mothers aged 28-35 and 35 and over for children born in 1987-89. The coefficients are plotted in Figure 6. The figure shows that relative to mothers who were aged 18-22, sex ratios for children born were constant over birth years before the reform for all age groups. After the reform, the fraction of males increased in children born to older mothers.

To assess the statistical significance and the average effect of the reform, we estimate the simpler difference-in-difference equation (3). The estimates for $\hat{\beta}_{2t}$ and $\hat{\beta}_{3t}$ and their robust standard errors are shown in Table 4 column (1). It shows that for cohorts born after the reform, the introduction of sex-selective abortion increased the fraction of males born amongst third and higher-parity births by 0.7 percentage points. The estimate is statistically significant at the 1% level. There was no effect for second order births. The coefficients reported in columns (2)-(4) show that our estimate is robust to controlling for mother's age, mother's education and father's education. (The estimate is also robust to controlling for birth characteristics and mother's marital status. They are not reported in the paper for the sake of brevity). We estimate equation (4) to examine the effect of sex-selective abortion on the fraction of males born by mother's age. The estimates are shown in Table 4 column (5). It shows that sex-selective abortion increased the fraction of males born to mothers aged 28-35 by 0.8 percentage-points. The estimate is statistically significant at the 5% level. The effect is similar for children born to mothers aged 35 and older. However, the estimate is not statistically significant. This is most likely due to the small number of women who choose to have children after 35 years of age. There was no effect for mothers who were 22-28 years of age. To examine the interaction effect of mother's age and birth order, we estimate equation (5). The estimates for $\hat{\beta}_{2t}$ and $\hat{\beta}_{3t}$ and their robust standard errors are shown column (6) of Table 3. It shows that for third and higher-parity births, one additional year in mother's age increases the fraction of males by 0.6 percentage-points. The estimate is statistically significant at the 5% level.

4.2 The Effect on the Composition of Boys and Girls

To examine the effect of the reform on the composition of children, we estimate equation (6). The coefficients are reported in Table 5. Columns (1)-(4) show the estimates for the effect of abortion legalization the average mother's age, education, father's education, and mother's marital status for second and 3+ born boys. The coefficients for mother's age, education and father's education are positive but not statistically significant. The estimated effect of mother's marital status is near zero in magnitude and statistically insignificant. Next, we estimate the effect on child characteristics such as low birth weight or whether a child was a singleton birth. The estimates in Columns (6) and (7) show that there is no effect. To the extent that these crude measures capture the health status of an infant, these estimates show that the reform did not have any sex-differential effects on the health composition of children born.

4.3 The Effect on Sex-differential Infant Mortality

To estimate the effect on EFM, we first examine the effect of abortion by birth order. We estimate equation (7) with the fraction of deaths occurring within one, two, three, four, five, six, nine and twelve months as dependent variables. The coefficients and their robust standard errors are reported in Table 6. The triple difference estimates in Column (1) show that the reform increased male mortality within one month relative to female mortality by 0.06 percentable points for second births, and by 0.14 percentage points for third and higher parity births. When we look at longer time horizons in columns (2)-(8), we see the estimates are similar. This suggests that most of the differential neglect which leads to mortality occurs within the first month of life. This is consistent with Ebenstein (2007), which found that post-natal sex selection in Taiwan most often occurs during the first month of life.

Interestingly, the coefficients for the triple interaction term of mother's age, male and post reform shows that boys born to older mothers after the reform are *more* likely to survive relative to girls. This suggests that older mothers were not engaging in post-natal sex selection before the reform. Or at least older mother's who are engaging in sex-selective abortion are mostly not comprised of those who would have selected postnatally before the reform. The fact that older mothers are having more children after the reform suggests that those that are engaging in sex-selective abortion after the reform may not have had children at all if abortion was not legal. In other words, the reform caused older mothers to have more children because they can now engage in sex selective abortion. Being able to better select the sex increased the option value of giving birth for mothers with son preferences. The value goes up more for older mothers who bear a higher physical cost of being pregnant.

5 Interpretation

The results show that the legalization of abortion increased the fraction of males born in higher birth-orders and to older mothers. The estimates show that the access to abortion increased the fraction of males for higher-order births by 0.7 percentage-points on average, approximately 100% of the observed increase in the fraction of males during the late 1980s in Taiwan. The finding that the reform did not alter the average health characteristics of boys born relative to girls born supports the interpretation that the increase in number of boys born is due to parents using abortion to select for sex rather than the possibility that the increased use of Ultrasound B has larger benefits for male fetuses of higher birth parities or born to older mothers.

Mortality rates for third and higher-parity births for girls are on average 0.23 percentage-points (see Table 3 Panel A3). The results on neo-natal mortality show that the reform decreased female mortality relative to male mortality at higher birth orders by 0.14 percentage-points, approximately a 61% reduction in neo-natal mortality.

While it is beyond the scope of this paper to show the cause of death, we do not think that it is due to a reduction in intentional female infanticide. Studies such as Banister (2004) suggest that there is no evidence of this extreme method of selection in South Korea or Taiwan. A more likely cause of differential mortality in a country where overall infant mortality is so low is *marginal* differential neglect.[12] For example, when a child is unwell at night and the symptoms do not obviously suggest a serious illness, parents may decide to take a child to the hospital if he is a son, but wait until morning if she is a daughter. For infants, who can experience large biological fluctuations within a very short period (e.g. fluctuations of body temperature from 98.4 to 104 degrees Fahrenheit within a few hours), and more importantly, extreme outcomes within hours of displaying mild symptoms, this small level of differential neglect could have extreme results and be reflected in the mortality data.[13]

[12] There is evidence that parents in Taiwan do discriminate girls relative to boys. For example, Lin, Liu and Chou (2007) and evidence for parental neglect of LBW girls relative to LBW boys.

[13] The Nelson Textbook of Pediatrics (2008) gives a long list of conditions that can change within a short period of time

"Neonates who die tend to die quickly" (Lantos, Mokala and Meadow, 1997).

The increase in the fraction of boys born and the decrease in the relative female mortality implies that female mortality is decreasing because parents who really want a son are no longer forced to have a daughter. To quantify the extent of this effect, we calculate the fraction of parents who substitute from post-natal to pre-natal selection when abortion is legalized. The calculation is very straightforward. See Table 7. Since the effects we find are mainly for 3+ parity births, we use data for 3+ parity births for our calculations. We assume that the *natural* fraction of males at birth is 51% (the lowest observed fraction of males at birth in the U.S.).[14] Hence, the natural fraction of girls is 49%. Using the natural rate of 49%, the fraction of girls that are actually born in the pre and post period (48.3% and 47.6%), and the number of children born in each period, we can calculate the number of missing girls. Next, using the number of girls born and the mortality rate in each period, we can calculate the number of girls that die in each period. For mortality rate in the pre-reform period, we use the average for 3+ girls, 0.003, reported in Table 3. For the post-reform period mortality rate, we subtract the pre-reform rate by the estimate for the effect of legalizing abortion on 3+ mortality (the coefficient for $bors3*post*sex$ in column (1) of Table (5)), 0.0014. Hence, the post-reform mortality rate is 0.003-0.0014 = 0.0016. The fraction of parents who used to select postnatally and switch to abortion after it is legalized is the absolute value of the change in the number of girls dying divided by the number of girls missing in the pre-reform period, 101/757= 0.133. Conversely, if

[14] To be conservative, we use 0.51 males at birth as the benchmark for the natural fraction absent intervention rather higher fractions (0.51-0.52) that are often accepted as natural in the literature. Because the biology of sex determination of a fetus is not yet perfectly understood, benchmarks are made by observing the fraction of males at birth in countries that are assumed to have no son-preference. However, as many demographic studies have pointed out, this number varies widely across countries and over time (e.g. see Mathews and Hamilton, 2005); and recent studies such as Abrevaya have even asserted that there is evidence of boy-biased sex selective abortion by certain populations in the U.S. Hence, to be conservative, we used the lowest observed fraction of males at birth in the U.S. as the benchmark.

abortion was banned in the post period, then the fraction of parents who would switch from abortion to postnatal selection is the absolute value of the change in the number of girls dying divided by the number of girls missing in the post-reform period, 101/1029=0.098.

Our results can also shed a little light on the question of what type of parents substituted from postnatal to prenatal selection. Although older mothers were the group who were using abortion to select for sex, the findings that the reform increased relative female mortality rates for children born to older mothers suggest that the substitution across technologies must have been done by mothers who were younger at the time of birth. The results that abortion caused older mothers to have more boys and the observation that the total number of births for older mothers increased after abortion while it was decreasing for younger mothers have very interesting implications. They suggest that legalizing abortion actually increased the value of giving births for older mothers (with son preference) by making it possible for the latter to selectively abort females.

Hence, our study is able to identify the effects of legalizing abortion on the extensive and intensive margins. On the extensive margin, sex-selective abortion significantly increases the fraction of boys being born. Some older mothers who would not otherwise have had children, choose to do so because they are able to sex select with abortion. On the intensive margin, some parents who would select postnatally if abortion were not available, now choose to use abortion as a method of selection instead.

Caution should be used in interpreting these results beyond the context of this study. This study estimates the impact of sex-selective abortion for the first few years after abortion is legalized, when the technology is still being phased in. This is evident from Figure 1, which shows that legalized abortion caused the fraction of males at births at third and higher parties to begin a steady climb from 1986 until 1990, and then stabilize at the new higher level. It is clear then that if we compared the fraction of males at birth in the 1990s to the pre-reform fraction, the estimated impact of legalizing abortion would be much larger.

One should also carefully consider the Taiwanese context. Taken literally, our estimates give the effect of legalizing abortion conditional on easy access to pre-natal sex detection technology. We are also estimating the effect of legalizing sex-selective abortion in a society with son-preference that is also experiencing a secular decrease in the preferred number of children. Our results will obviously overestimate the effect of legalizing sex-selective abortion on sex ratios at birth in places where there is less son-bias or places where parents are less constrained on the number of children. On the other hand, there are contexts for which our estimates will underestimate the true effect. Taiwan has extremely low infant mortality rates relative to the rest of the world and we believe that the mortality results are being driven by very small changes in marginal differential neglect. It is difficult to predict the effect in a place where there is more extreme differential neglect or a place where infant mortality rates are higher overall (e.g. India). Very likely, the effect will be bigger there than in Taiwan. In places such as China, where strict family planning policies constrain parents in the number of children to only one or two, changes in access to sex-selective abortion will likely have a much larger effect than in Taiwan. That said, for the purpose of discussion, we will in our concluding remarks make the simplifying assumption that parents in China and India face similar constraints as parents in Taiwan to broadly understand the implications of our estimates for those other contexts.

6 Conclusion

This paper uses a straight forward empirical strategy to provide evidence for the impact of sex-selective abortion on sex ratios at birth and EFM. The results show that legalizing abortion had little effect on sex ratios for parents who can reasonably expect to have more children (low birth parities and young mothers). However, for parents who face relatively more uncertainty in their ability to have more children (high parity births and children born to older mothers), legalizing abortion dramatically increased the male-biased sex imbalance at birth. For third and higher-parity births, access to abortion increased the fraction of males from 51.7% to 53.5% in the late 1980s, which accounts for nearly 100% of the observed increase in sex imbalance at birth during this period. This leaves little doubt that access to sex-selective abortion has been by far the most important contributor to the recent increase in the observed population sex imbalance.

The stark results on relative female mortality show that access to abortion decreased female neo-natal mortality by 61% relative to males. They show that up to 13% of parents who were selecting postnatally before the reform would have substituted to pre-natal sex selection using abortion. In other words, for every 100 abortions, 13 lives of born girls are saved. If these results are interpreted literally for purely illustrative purposes, they suggest that in China and India, strictly enforcing the ban on sex-selective abortion would cause there to be 1.6 million more girls born but 156,800 more girls will die neutrally each year.[15]

For policy makers, this means that the welfare implications of banning sex-selective abortion depends on the weight placed on the welfare of unborn female fetuses relative to newly born girls and the additional disutility for parents to select the sex post- rather than pre-natally, as

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[15] China in 2000 had about 17.7 million births. At least 57% were boys. Hence, there is a seven percentage point de.cit of girls. If the effect of sex selective abortion in China is the same as Taiwan, then approximately 3 percentage-points is due to sex selective abortion. Hence, banning sex selective abortion will increase the number of girls born in China by 17.7 mil * 0.03= 531,000. And it will increase the number of girls dying by approximately 160,000.

India's statistics are similar to those of China. So, banning sex selective abortion in both countries will increase the number of girls born by almost 1.61 million, and the number of female neonatal mortality by approximately 160,000 annually.

well as the weights placed on the potential economic and social consequences from having unbalanced sex ratios. It also suggests that for increasing relative female welfare, policies which restrict access to sex-selective abortion compliment policies that subsidize the cost of raising daughters. In other words, policies that prohibit the use of sex-selective abortion should be coupled with policies that increase parents' incentives to invest in daughters after they are born.[16]

[16] 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).

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Figure 1A: Fraction of Males at Birth by Parity over Time in Taiwan (1980-1998)

Figure 1B: Fraction of Males by Birth Parity in China (1980-90)

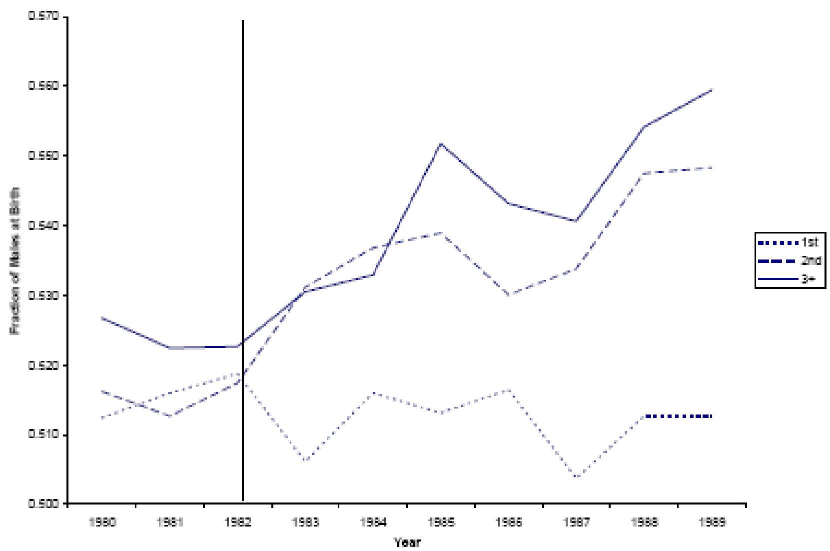

Figure 1C: Fraction of Males by Birth Parity in South Korea (1980-90)

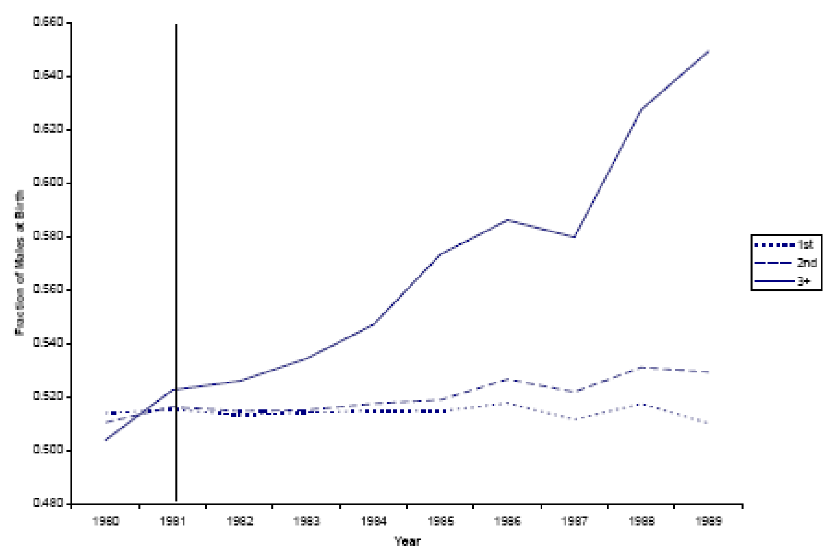

Figure 2A: Fraction of Women Reported to Have Ever Had an Abortion by Age

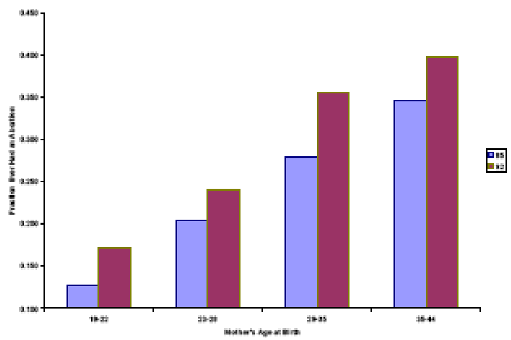

Figure 2C: Fraction of Women Reported to Have Ever Had an Abortion by Education

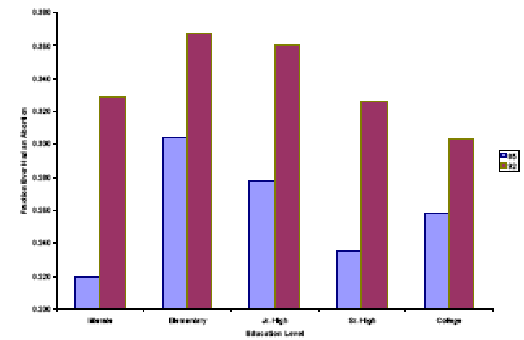

Figure 2B: Change in Fraction of Women who Ever Had an Abortion by Age

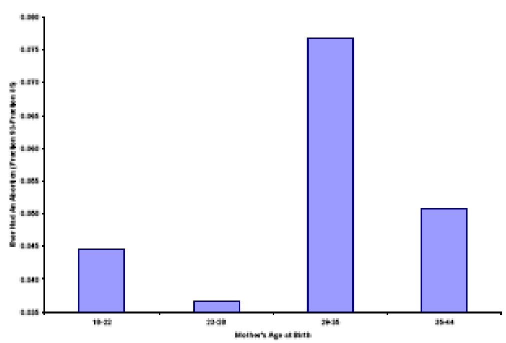

Figure 2D: Change in Fraction of Women who Ever Had an Abortion by Education Level

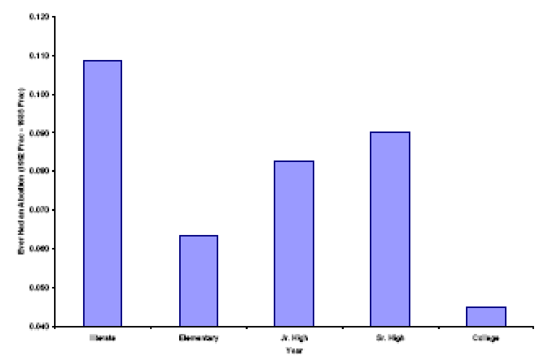

Figure 3A: The Fraction of Males by Birth Year and Birth Order

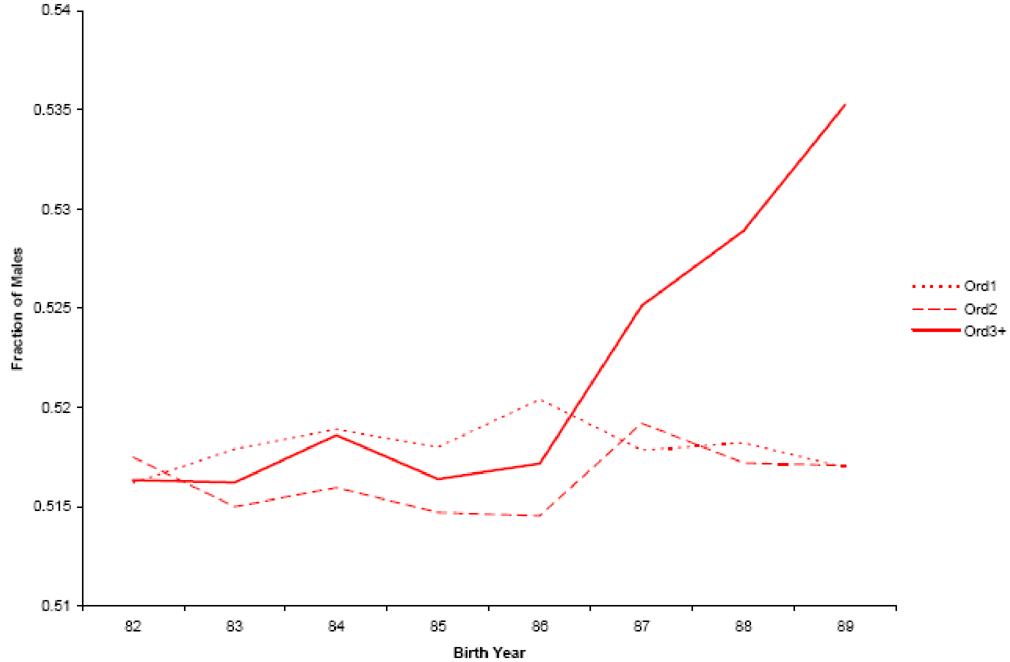

Figure 3B: The Fraction of Males by Birth Year and Mother's Age

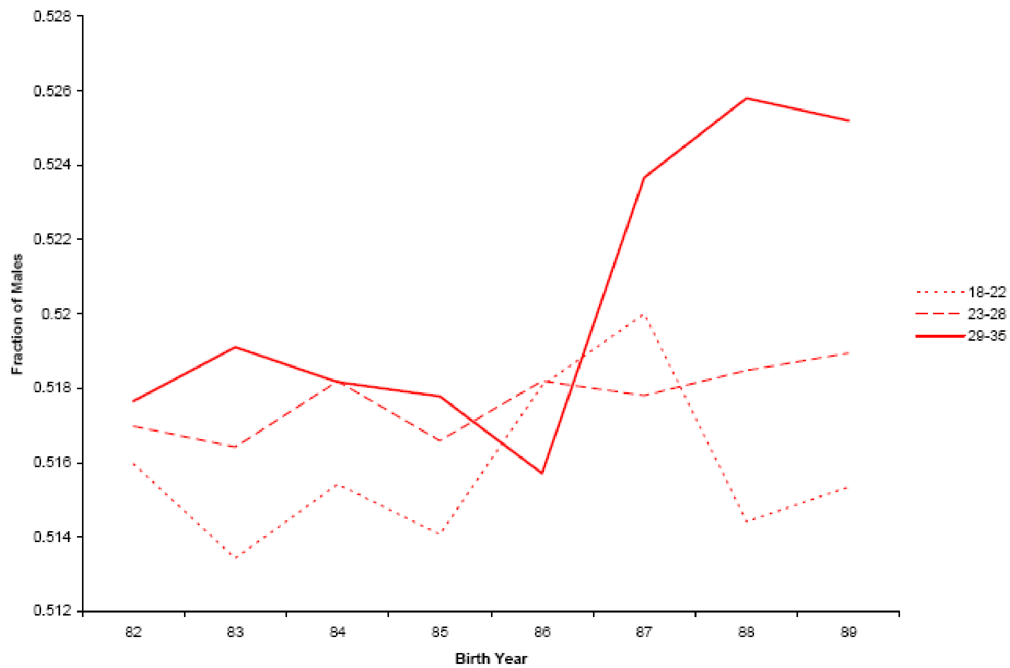

Figure 4A: Log (Total Births) by Birth Year and Birth Order

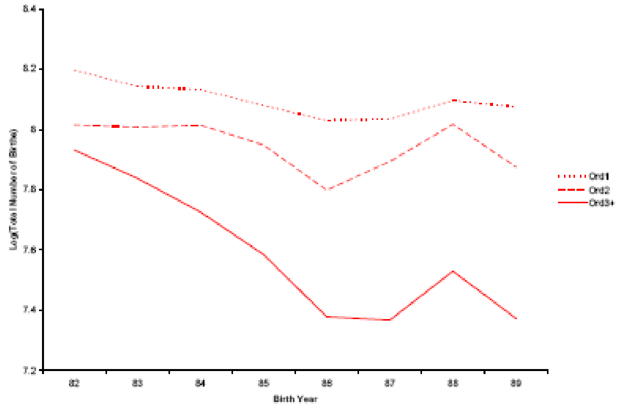

Figure 4B: Log (Total Births) by Birth Year and Mother's Age for Young Mothers (18-28)

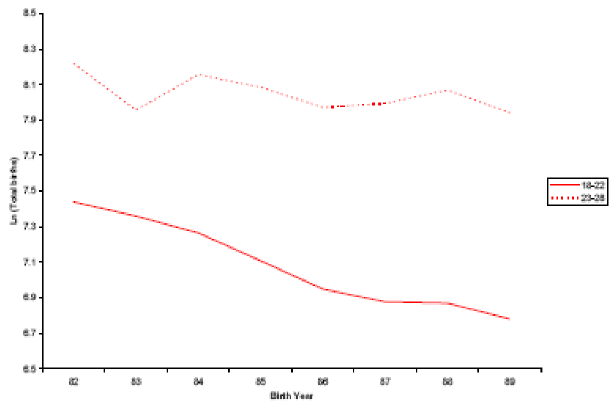

Figure 4C: Log (Total Births) by Birth Year and Mother's Age for Old Mothers (29+)

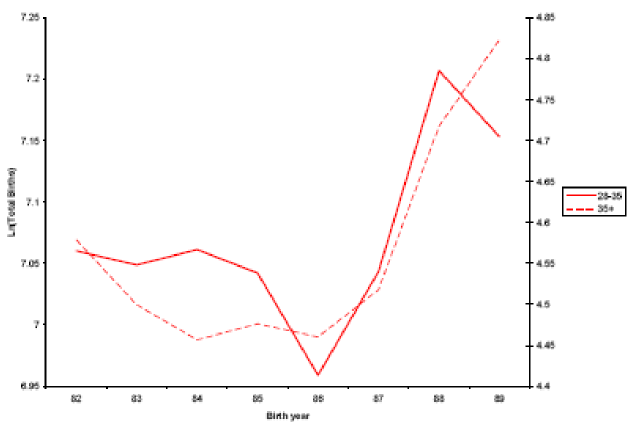

Figure 5A: The Effect of Abortion on Fraction of Males by Birth Order
Coefficients of the interaction terms of birth year dummy variables and birth order dummy variables

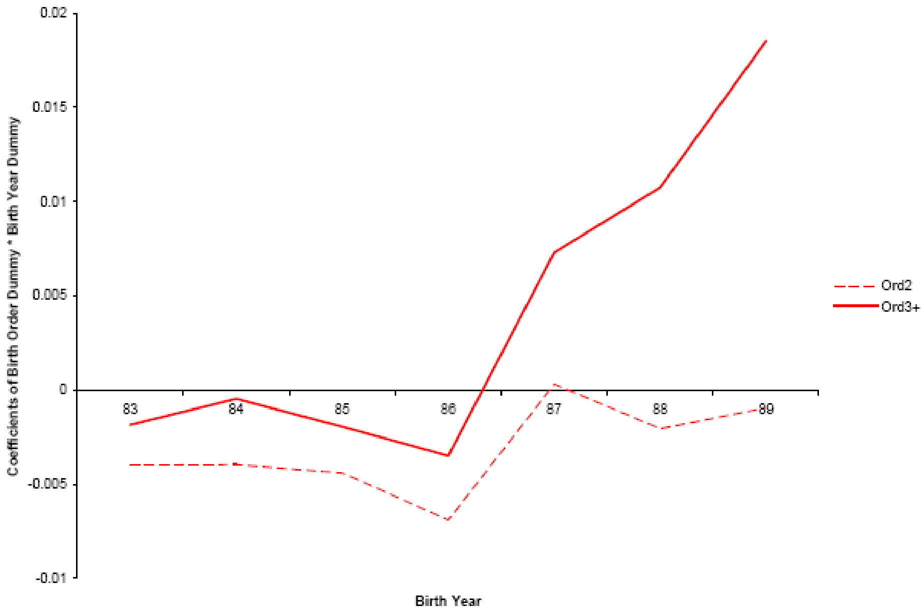

Figure 5B: The Effect of Abortion on Fraction of Males on 3+ Births
Coefficients of the interaction terms of birth year dummy variables and 3+ birth order dummy variable and their 95% Confidence Intervals

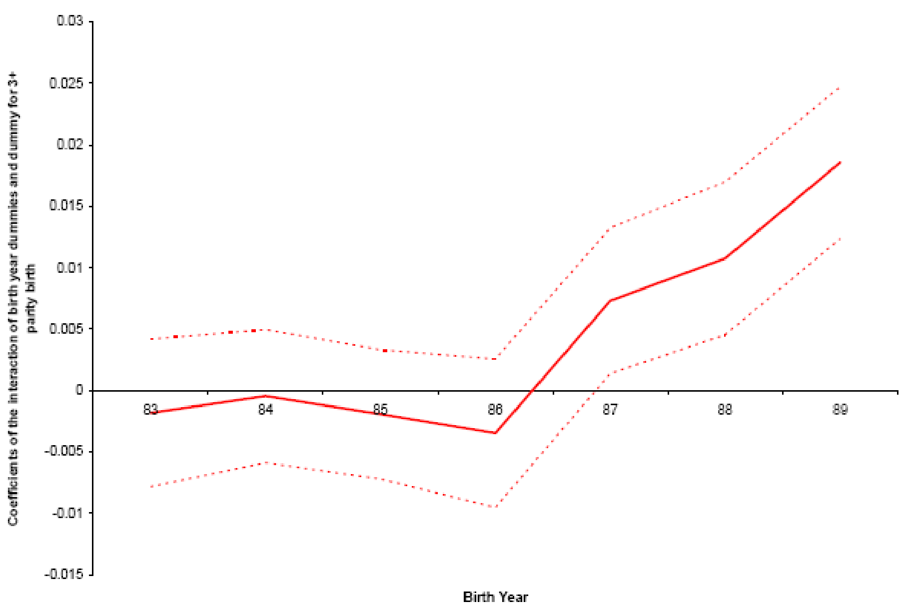

Figure 6: The Effect of Abortion on Fraction of Males by Mother's Age
Coefficients of the interaction terms of birth year dummy variables and mother's age dummy variables

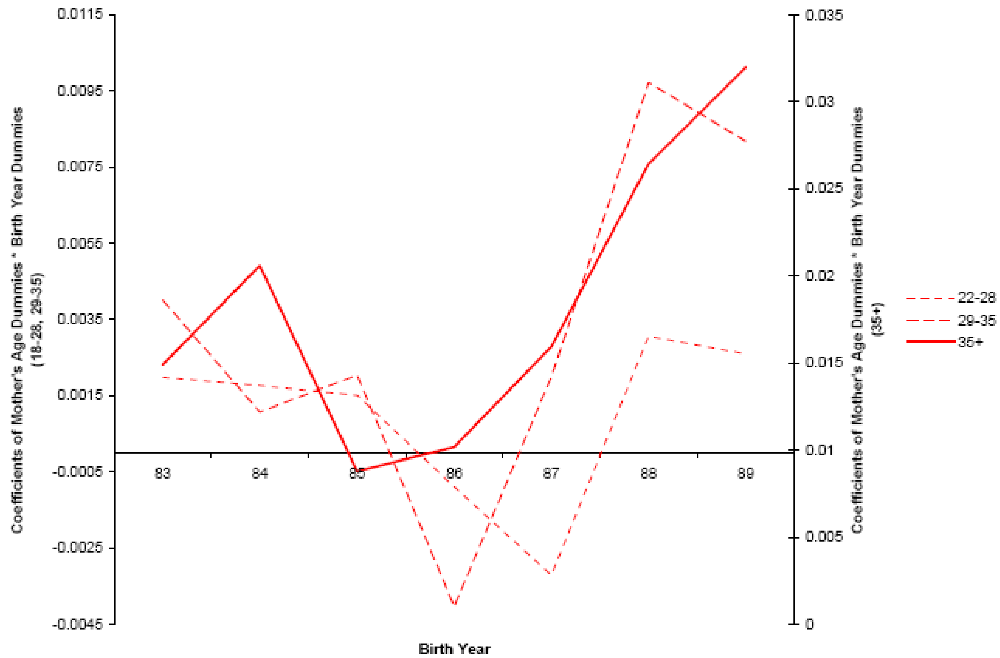

Table 1: The Change in Number of Births by Birth Order and Mother's Age

Total Number of Births Annually				
	1982-84	1985-89	Change	%Change
A. Birth Order				
1	146861	133755	-13106	-8.92%
2	125896	113699	-12198	-9.69%
3+	108162	73500	-34662	-32.05%
B. Mother's Age				
18-22	80799	52252	-28546	-35.33%
23-28	213459	178130	-35329	-16.55%
29-35	81042	83381	2339	2.89%
35+	6093	7537	1444	23.69%

Table 2: Descriptive Statistics on Birth and Parental Characteristics
by Birth Year and Birth Order

	I. Born 1982-84			II. Born 1985-89			III.
Variable	Obs	Mean	Std. Err.	Obs	Mean	Std. Err.	Diff
A. Birth Order = 1							
Male (Fraction)	99	0.518	0.001	165	0.518	0.001	0.001
Mother's Age	99	24.072	0.051	165	24.855	0.041	0.784
Mother's Education (Years)	99	9.378	0.044	165	9.954	0.028	0.576
Father's Education (Years)	99	10.172	0.043	165	10.542	0.027	0.370
LBW (Fraction)	99	0.065	0.000	165	0.066	0.000	0.001
Birth Weight (Grams)	99	3196.233	1.170	165	3185.754	0.921	-10
Singleton Birth (Fraction)	99	0.993	0.000	165	0.989	0.000	-0.004
Mother Married (Fraction)	99	0.976	0.000	165	0.969	0.000	-0.007
Death within 1 Month (Fraction)	198	0.003	0.001	330	0.002	0.001	-0.001
Death within 6 Months (Fraction)	198	0.005	0.002	330	0.004	0.002	-0.001
Death within 12 Months (Fraction)	198	0.006	0.002	330	0.005	0.002	-0.001
B. Birth Order = 2							
Male (Fraction)	100	0.516	0.001	165	0.517	0.001	0.000
Mother's Age	100	25.735	0.057	165	26.741	0.050	1.005
Mother's Education (Years)	100	8.884	0.048	165	9.653	0.035	0.769
Father's Education (Years)	100	9.778	0.047	165	10.373	0.032	0.595
LBW (Fraction)	100	0.054	0.001	165	0.055	0.000	0.001
Birth Weight (Grams)	100	3270.391	1.222	165	3257.464	1.089	-13
Singleton Birth (Fraction)	100	0.991	0.000	165	0.987	0.000	-0.004
Mother Married (Fraction)	100	0.991	0.000	165	0.989	0.000	-0.003
Death within 1 Month (Fraction)	199	0.003	0.002	330	0.003	0.001	-0.001
Death within 6 Months (Fraction)	199	0.006	0.002	330	0.005	0.002	-0.001
Death within 12 Months (Fraction)	199	0.008	0.002	330	0.006	0.002	-0.001
C. Birth Order = 3+							
Male (Fraction)	100	0.517	0.001	165	0.524	0.001	0.007
Mother's Age	100	27.984	0.043	165	28.773	0.046	0.789
Mother's Education (Years)	100	7.005	0.043	165	7.964	0.037	0.959
Father's Education (Years)	100	8.128	0.040	165	8.903	0.034	0.776
LBW (Fraction)	100	0.043	0.000	165	0.048	0.000	0.005
Birth Weight (Grams)	100	3352.349	1.638	165	3331.444	1.473	-21
Singleton Birth (Fraction)	100	0.990	0.000	165	0.985	0.000	-0.005
Mother Married (Fraction)	100	0.992	0.000	165	0.987	0.000	-0.005
Death within 1 Month (Fraction)	199	0.003	0.002	330	0.003	0.002	-0.001
Death within 6 Months (Fraction)	199	0.007	0.002	330	0.006	0.002	-0.001
Death within 12 Months (Fraction)	199	0.008	0.003	330	0.007	0.003	-0.001

Observations for all variables except mortality are birth year x birth order x birth county cell.
Observations for mortality variables are birth year x birth order x birth county x sex cells.

Table 3: Descriptive Statistics on Neo-natal Mortality by Birth year, Birth Order and Sex

	A. Girls					B. Boys					
	(1)		(2)		(3)	(4)		(5)		(6)	(7)
Death Within X Months	Born 1982-84 Obs	Mean	Born 1985-89 Obs	Mean	Diff (2)-(1)	Born 1982-84 Obs	Mean	Born 1985-89 Obs	Mean	Diff (5)-(4)	DD: (6)-(3)
	A1. Birth Order =1					B1. Birth Order=1					
1 Month	99	0.0028	165	0.0022	-0.0006	99	0.0037	165	0.0025	-0.0012	-0.0006
		(0.0001)		(0.0001)			(0.0001)		(0.0001)		
12 Months	99	0.0057	165	0.0047	-0.0010	99	0.0070	165	0.0052	-0.0017	-0.0007
		(0.0002)		(0.0001)			(0.0002)		(0.0001)		
	A2. Birth Order=2					B2. Birth Order=2					
1 Month	99	0.0032	165	0.0025	-0.0007	100	0.0038	165	0.0028	-0.0010	-0.0004
		(0.0001)		(0.0001)			(0.0002)		(0.0001)		
12 Months	99	0.0070	165	0.0057	-0.0012	100	0.0080	165	0.0066	-0.0015	-0.0002
		(0.0002)		(0.0002)			(0.0002)		(0.0002)		
	A3. Birth Order=3					B3. Birth Order=3					
1 Month	100	0.0030	165	0.0023	-0.0007	99	0.0035	165	0.0028	-0.0007	-0.0001
		(0.0001)		(0.0001)			(0.0002)		(0.0001)		
12 Months	100	0.0077	165	0.0069	-0.0008	99	0.0088	165	0.0072	-0.0016	-0.0008
		(0.0003)		(0.0002)			(0.0003)		(0.0002)		

Data are aggregated into cells by sex, birth year, birth county and birth order.

Table 4: The Effect of Abortion on Fraction of Males by Birth Order and/or by Mother's Age

	Dependent Variable: Fraction of Males at Birth					
	(1)	(2)	(3)	(4)	(5)	(6)
Ord2 * Born 1985-89	0.000	0.000	-0.001	0.000		0.031
	(0.001)	(0.001)	(0.001)	(0.001)		(0.064)
Ord3+ * Born 1985-89	0.007	0.007	0.006	0.006		-0.183
	(0.002)	(0.002)	(0.002)	(0.002)		(0.085)
Mother's Age * Ord2 * Born 1985-89						-0.001
						(0.003)
Mother's Age * Ord3 * Born 1985-89						0.006
						(0.003)
Mother 22-28 * Born 1985-89					-0.001	
					(0.003)	
Mother 29-35 * Born 1985-89					0.008	
					(0.004)	
Mother35+ * Born 1985-89					0.008	
					(0.010)	
Controls						
Mother's Age	N	Y	Y	Y	N	Y
Mother's Education	N	N	Y	Y	N	N
Father's Education	N	N	N	Y	N	N
Observations	794	794	794	794	1057	794

Columns (1)-(4), (6) uses data aggregated by birth order, birth year, and birth county.
Column (5) uses data aggregated by mother's age, birth year, and birth county.
All regressions control for birth year fixed effects.
Columns (1)- (4), (6) also controls birth order fixed effects.
Column (6) also controls for mother's age * birth order dummies, and mother's age * born 1985-89.

Table 5: The Effect of Abortion Legalization on the Composition of Children Born and Parental Characteristics

Coefficients of the interaction terms of birth order dummy variables, birth year dummy variables and a dummy variable for male

	(1)	(2)	(3)	(4)	(5)	(6)	(7)
Dependent Variables:	**Log(Mother's Age)**	**Log(Mother's Edu)**	**Log(Father's Edu)**	**Mother Married**	**Log(Births)**	**LBW**	**Singleton**
Ord2*Male*Born 1985-89	0.000	-0.001	-0.002	0.001	-0.003	0.002	0.000
	(0.006)	(0.015)	(0.014)	(0.001)	(0.298)	(0.002)	(0.001)
Ord3*Male*Born 1985-89	0.003	0.010	0.010	0.001	0.020	-0.001	0.000
	(0.005)	(0.016)	(0.014)	(0.001)	(0.285)	(0.002)	(0.001)
Observations	1586	1545	1578	1586	1586	1586	1586

Robust standard errors in parentheses.

All regressions control for the full set of double interactions terms, sex, birth order and birth year fixed effects.

Table 6: The Effect of Abortion on Sex-differential Neo-Natal Mortality by Birth Order
Coefficients of the interaction terms of birth order dummy variables and a dummy variable indicating if an individual was born after the reform.

	Dependent Variables: Death Within `x' Months									
	1 Mo.	2 Mo.	3 Mo.	4 Mo.	5 Mo.	6 Mo.	9 Mo.	12 Mo.	18 Mo.	24 Mo.
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)
Bord2*Sex*Post	0.00063	0.00054	0.00071	0.00053	0.00083	0.00092	0.00100	0.00125	0.00101	0.00096
	(0.00027)	(0.00017)	(0.00009)	(0.00011)	(0.00024)	(0.00036)	(0.00045)	(0.00045)	(0.00033)	(0.00029)
Bord3*Sex*Post	0.00142	0.00112	0.00150	0.00084	0.00130	0.00120	0.00116	0.00149	0.00124	0.00118
	(0.00054)	(0.00034)	(0.00019)	(0.00020)	(0.00048)	(0.00072)	(0.00094)	(0.00094)	(0.00070)	(0.00063)
Mage*Sex*Post	-0.00023	-0.00021	-0.00031	-0.00020	-0.00034	-0.00038	-0.00034	-0.00036	-0.00035	-0.00034
	(0.00014)	(0.00009)	(0.00005)	(0.00006)	(0.00013)	(0.00019)	(0.00024)	(0.00024)	(0.00018)	(0.00016)
Observations	1586	1586	1586	1586	1586	1586	1586	1586	1586	1586

Standard errors are clustered by birth order.
All regressions control for the full set of double interaction terms, and birth year, birth order, and sex main effects.

Table 7: Calculating the % of Parents who Substitute from Post- to Pre-natal Sex Selection

3rd+ Parity Births	1982-84 (1)	1985-89 (2)	Changes (3)
Average Number of Total Annual Births	108162	73500	-34662
Fraction of Boys	0.517	0.524	0.007
Fraction of Girls	0.483	0.476	-0.007
# of Girls Born	52242	34986	-17256
Fraction of Missing Girls= 0.49 - Fraction of Girls	0.007	0.014	0.007
# of Missing Girls = Fraction of Missing Girls * Average Number of Total Annual Births	*757*	*1029*	272
Girl's Mortality (Fraction) (0.003 - 0.0014)*	0.003	0.0016	-0.0014
# of Girls Dying = Girl's Mortality * # of Girls Born	157	56	*-101*
Fraction of Parents Substituting from post- to pre-natal selection =101/541			**0.133**
Fraction of Parents who would Select Post-natally if Abortion was Banned =101/1029			**0.098**

* 0.0014 is the estimate for the effect of abortion on relative mortality reported in column (1) of Table (5).

Table A1: Fraction of Males by Parity for Taiwan, China and South Korea

	Sex Ratios at Birth (# Boys/ # Girls)											
	S Korea				China				Taiwan			
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)	(11)	(12)
Year	1	2	3+	3+/1	1	2	3+	3+/1	1	2	3+	3+/1
1980	0.514	0.510	0.504	0.490	0.512	0.516	0.527	0.514	0.516	0.516	0.516	0.500
1981	0.515	0.516	0.522	0.507	0.516	0.513	0.522	0.506	0.516	0.517	0.518	0.501
1982	0.513	0.515	0.526	0.513	0.519	0.517	0.523	0.504	0.516	0.518	0.516	0.500
1983	0.514	0.515	0.534	0.520	0.506	0.531	0.531	0.524	0.518	0.515	0.516	0.498
1984	0.515	0.517	0.547	0.532	0.516	0.537	0.533	0.517	0.519	0.516	0.519	0.500
1985	0.515	0.519	0.573	0.559	0.513	0.539	0.552	0.539	0.518	0.515	0.516	0.498
1986	0.518	0.527	0.586	0.569	0.516	0.530	0.543	0.527	0.520	0.515	0.517	0.497
1987	0.511	0.522	0.580	0.568	0.504	0.534	0.541	0.537	0.517	0.520	0.526	0.508
1988	0.517	0.531	0.627	0.611	0.513	0.548	0.554	0.542	0.518	0.517	0.528	0.511
1989	0.510	0.529	0.649	0.640	0.513	0.548	0.559	0.547	0.517	0.517	0.534	0.518
1990	0.520	0.539	0.658	0.640					0.516	0.521	0.547	0.531
1991	0.514	0.529	0.646	0.633	0.516	0.557	0.559	0.543	0.518	0.520	0.546	0.528
1992	0.515	0.529	0.660	0.647					0.519	0.518	0.542	0.523
1993	0.516	0.534	0.674	0.660					0.517	0.516	0.530	0.513
1994	0.514	0.533	0.672	0.659					0.519	0.518	0.533	0.514
1995	0.514	0.528	0.643	0.630	0.516	0.585	0.607	0.592	0.517	0.513	0.534	0.517
1996	0.513	0.523	0.624	0.612					0.519	0.517	0.533	0.514
1997	0.512	0.515	0.575	0.563					0.519	0.517	0.536	0.518
1998	0.514	0.519	0.593	0.579					0.517	0.516	0.539	0.522
1999	0.514	0.518	0.589	0.575								
2000	0.515	0.518	0.590	0.575	0.517	0.603	0.614	0.598				
2001	0.513	0.516	0.586	0.573								
2002	0.516	0.518	0.585	0.570								
2003	0.512	0.517	0.577	0.566								
2004	0.513	0.515	0.570	0.558								

S. Korea Data Source: South Korea National Statistical Office..China Data Source: Data covering the period from 1960 to 1989 are taken from Gu and Xu (1994) and Gu and Roy (1995), with their calculations having been drawn from the Data Volumes of the National Reproduction and Birth control Sample Surveys, Chapter 3. The 1990 (census results), 1995 (1% population survey results) and 2000 (census results) data are taken from Yuan and Tu (2005). The 1992 data is the result of a 0.1% sample which was taken from the Chinese Population Statistical Yearbook. Taiwan Data Source: Taiwan National Birth Registries

Table A2: The Effect of Abortion on Fraction of Males by Birth Order or by Mother's Age
Coefficients of the interaction terms of birth year dummy variables and birth order dummy variables;
or the interaction terms of birth year dummy variables and mother's age dummy variables

	Dependent Variable: Fraction of Males				
	(1)	(2)	(3)	(4)	(5)
	Ord2	Ord3+	Mom 22-28	Mom 29-35	Mom35+
Born 1983	-0.004	-0.002	0.002	0.004	0.015
	(0.003)	(0.003)	(0.003)	(0.004)	(0.007)
Born 1984	-0.004	0.000	0.002	0.001	0.021
	(0.003)	(0.003)	(0.003)	(0.004)	(0.011)
Born 1985	-0.004	-0.002	0.002	0.002	0.009
	(0.003)	(0.003)	(0.003)	(0.004)	(0.007)
Born 1986	-0.007	-0.003	-0.001	-0.004	0.010
	(0.003)	(0.003)	(0.003)	(0.004)	(0.008)
Born 1987	0.000	0.007	-0.003	0.002	0.016
	(0.003)	(0.003)	(0.003)	(0.004)	(0.007)
Born 1988	-0.002	0.011	0.003	0.010	0.026
	(0.002)	(0.003)	(0.003)	(0.004)	(0.007)
Born 1989	-0.001	0.019	0.003	0.008	0.032
	(0.002)	(0.003)	(0.003)	(0.004)	(0.007)
Observations	794			1057	
R-squared	0.17			0.11	

Robust standard errors in parentheses.
Columns (1) and (2) are estimated from one regression. Columns (3)-(5) are estimated from one regression.
All regressions control for birth year fixed effects.
Data for Columns (1) and (2) are aggregated to sex x birth order x birth year x birth county cells.
Data for Columns (3)-(5) are aggregated to sex x mother's age x birth year x birth county cells.