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 without discussing 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 by the legalization of abortions 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 25%. We estimate that approximately 15 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%. The phenomenon has increased over time even when there is rapid economic growth and social "modernization". Figure 1 shows that in Taiwan, the fraction of males at birth increased from approximately 51.5% in 1970 to 52.5% by 2003. Many believe this severe sex imbalance may lead to increased crime rates, or affect the marriage market.[1] There is also the concern that to select the sex of the child, parents must selectively abort female fetuses, or resort to means that result in excess female mortality (EFM) such as differential neglect, or in extreme cases, female infanticide. Economists and demographers, wishing to understand the determinants of the missing women phenomenon, have heatedly argued over the amount of the observed sex imbalance that can be attributed to economic, cultural and biological causes.[2] All sides of the debate agree on one thing: that a significant portion of the increase in sex imbalance in recent decades is driven by the introduction of technology that enables sex-selective abortion. However, there is a surprising lack of empirical evidence on the impact of this technology.[3]

Absent concrete evidence, policy makers in many countries, including China and India, 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 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 presents novel empirical evidence on the causal effect of access to sex-

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

[2] 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) examines 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.

[3] Chu (2001) presents a descriptive analysis of the practice of prenatal sex selection in rural central China using etailed 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

selective abortion on sex ratios at birth and EFM (which refers to female relative to male infant mortality in this paper). For identification, we exploit variation in access to abortion caused by a legislative reform in Taiwan, when the technology for detecting sex pre-natally was already available.

The principal 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 imbalance 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 adopt the technology, then a positive correlation between access to the technology and sex imbalance will confound the causal effect of access to sex-selective abortion on sex imbalance with the effect of the underlying preferences on both sex-selection and technology adoption. 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

crime.

of children born and parental characteristics.

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 51.5 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 25%. Our results suggest that approximately 15% of parents selecting post-natally before the reform would have substituted to abortion as a method of sex-selection. Taken literally, this suggests that for every 100 abortion of female fetuses, 15 lives of girls born are saved.

For policy makers, our results show that banning sex-selective abortion should have a large effect in decreasing the observed sex imbalance. However, unless governments can also force parents to care for girls born, a ban on sex-selective abortion may lead to an increase in EFM. For example, our results applied to the mainland China and India contexts suggest that effectively banning sex-selective abortion could increase the number of girls born each year by 1.6 million but cause female neo-natal mortality to increase by 160,000 girls in the two countries combined.

Studying the effect of sex-selective abortion in Taiwan has several advantages. First, the data are much better than other countries with boy biased sex imbalances. Second, the legislative reform legalizing abortion allows us to have plausibly exogenous variation on access to sex-selective abortion. Finally, 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.

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

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. It is the technology that is most widely used in developing countries today, and was the only method available 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 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 1985/86, when it became legal for women to induce an abortion for social as well as medical reason up to the 24th week of pregnancy. The service was inexpensive and safely conducted for 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. This paper interprets the relaxation in 1985/86 as the effective legalization of induced abortion.

Ultrasound B was already a well-established technology when abortion was legalized. The legalization of abortion combined with the use of Ultrasound B enabled parents to use abortion as a method of sex selection. This is reflected in both anecdotal evidence and the data on 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.[5] 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 reason for increasing the use of Ultrasound B machines was to meet the rising demand for sex-selection, 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 parents who wish to have a boy, they are more likely to sex-select if the their ability to have another child and try for a boy again 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.[6] 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 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 occured at the time that abortion was legalized that would also 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

[5] 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.

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

in pre-natal care affect males more positively than females, *and* have 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 there was no improvement in medical technology that would have affected infant mortality for higher-parity births more, *and* affect 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 specification has the pitfall that it would capture changes that occurred at any time after the reform.

$$Male_{it} = \sum_{i=2}^{3} \sum_{t=1983}^{1989} \beta_{it}(Ord_i * Born_t) + \gamma_i + \rho_t + \varepsilon_{it} \tag{1}$$

We regress the fraction of males of birth order i and birth year t, $Male_{it}$, on: the interaction terms between 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 β_{2t} and β_{3t} should be larger for individuals born after 1985. If parents are more likely to sex select at higher birth orders, then $\beta_{3t} > \beta_{2t}$. More specifically, the difference should be larger in magnitude for cohorts born after 1985 $\left(\beta_{3,t>=85} - \beta_{2,t>=85}\right) > \left(\beta_{3,t<85} - \beta_{2,t<85}\right)$.

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

$$Male_{mt} = \sum_{t=1983}^{1989} \sum_{m=2}^{4} \beta_{mt}(mom_m * Born_t) + \gamma_m + \rho_t + \varepsilon_{mt} \tag{2}$$

We regress the fraction of males for individuals born to mothers age m and birth year t, $Male_{mt}$, on: the interaction terms between 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,

$$Male_{it} = \sum_{i=2}^{3} \beta_{it}(Ord_i * Post_t) + \gamma_i + \rho_t + \varepsilon_{it} \tag{3}$$

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.

$$Male_{mt} = \sum_{m=2}^{4} \beta_{mt}(mom_m * Post_t) + \gamma_m + \rho_t + \varepsilon_{mt} \tag{4}$$

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.

$$\begin{aligned} Male_{itm} &= \sum_{i=2}^{3} \beta_i (momage_m * Ord_i * Post_t) \\ &+ \sum_{i=2}^{3} \theta_i (Ord_i * Post_t) + \sum_{i=2}^{3} \pi_i (Ord_i * momage_m) \\ &+ \lambda (momage_m * Post_t) + \delta_i + \gamma_m + \rho_t + \varepsilon_{imt} \end{aligned} \tag{5}$$

We regress the fraction of males for individuals of birth order i, birth year t whose mother was m years old when giving birth, $Male_{itm}$, on: the triple interaction term between a birth order dummy variable, Ord_i, the mother's age at birth, $momage_m$, 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 $momage_m$; mother's age at birth and being born in 1985 and after, $momage_m$ and $post_t$; birth order fixed effects, δ_i; mother's age fixed effects, γ_m; 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.

$$
\begin{aligned}
Y_{its} = & \sum_{i=2}^{3}\sum_{t=1983}^{1989} \beta_{it}(Ord_i * Born_t * Male_s) \\
& + \sum_{i=2}^{3}\sum_{t=1983}^{1989} \delta_{it}(Ord_i * Born_t) + \sum_{t=1983}^{1989} \pi_t(Male_s * Born_t) \\
& + \sum_{i=2}^{3} \lambda_i(Male_s * Ord_i) + \gamma_i + \rho_t + \varepsilon_{it}
\end{aligned} \tag{6}
$$

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, 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, $(\beta_{3,t>=85} - \beta_{2,t>=85}) \leq (\beta_{3,t<85} - \beta_{2,t<85}) \leq 0$.[7] For LBW births, the hypothesis predicts that the reform led to an increase in LBW births for boys at higher birth orders, $(\beta_{3,t>=85} - \beta_{2,t>=85}) \geq (\beta_{3,t<85} - \beta_{2,t<85}) \geq 0$.

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

[7] 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 underestimate 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.

$$
\begin{aligned}
Death_{its} \quad = \quad & \sum_{i=2}^{3} \beta_i (Ord_i * Post_t * Male_s) \\
& + \sum_{i=2}^{3} \delta_i (Ord_i * Post_t) + \sum_{i=2}^{3} \lambda_i (Male_s * Ord_i) \\
& + \pi (Male_s * Post_t) + \theta_s + \gamma_i + \rho_t + \varepsilon_{its}
\end{aligned} \tag{7}
$$

We regress the fraction of deaths occurring within a specified number of months for individuals of birth order i, birth year t, and sex s, Y_{its}, 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 full set of double interaction terms; 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 $\beta_3 \geq \beta_2 \geq 0$.

We estimate a similar equation to examine the effect of sex-selective abortion on sex-differential mortality rates by mother's age.

$$
\begin{aligned}
Death_{mts} \quad = \quad & \sum_{m=2}^{4} \beta_m (mom_m * Post_t * Male_s) \\
& + \sum_{m=2}^{4} \delta_m (mom_m * Post_t) + \sum_{m=2}^{4} \lambda_m (Male_s * mom_m) \\
& + \pi (Male_s * Post_t) + \theta_s + \gamma_m + \rho_t + \varepsilon_{mts}
\end{aligned} \tag{8}
$$

We regress the fraction of deaths occurring within a specified number of months for individuals born to mothers who are m years old, birth year t, and sex s, $Death_{mts}$, on: the triple interaction terms of dummy variables indicating that the mother is 22-28, 29-35 or above 35 years of age, mom_m, a dummy variable for being born in 1985 or afterwards, $Post_t$, and a dummy variable for being male, $male_{mts}$; the full set of double interaction terms; sex fixed effects, θ_s; mother's age fixed effects, γ_i; and birth year fixed effects, ρ_t. The reference group is comprised of children born to mothers 18-21 years of age. It and all of its interactions are dropped. If sex-selective abortion increased survival rates or test scores for all girls relative to boys, then $\pi \leq 0$. If the effects are larger for older mothers, then $\beta_4 \geq \beta_3 \geq \beta_2 \geq 0$.

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 comprises 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 at very young ages, 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 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 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 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 are also more low birth weight babies born, and the occurrence of multiple (non-singleton) births increased. This most likely reflects an improvement in pre-natal care 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 2A 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 51.5%. This is significantly higher than the 50.1% fraction of males at birth observed in countries not known to have male-biased sex preferences (e.g. U.S.). The difference most likely reflects illegal access to abortions before the legislative reform. (See section on Interpretation for further discussion). 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 53.5% in 1989. Figure 2B 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.

Figures 3A-3F plot the logarithm of total births, mother's age, educational attainment and marital status at the time of birth, and the fraction of singleton and LBW births by birth year and birth order. They show that the total number of births decreased with the introduction of abortion. Most of the decrease is observed for third and higher-parity births. This is consistent with the anecdotal evidence that the legalization of abortion was in part driven by parents' desire to use abortion as a method of family planning. Mother's age and education increases over time. The probability that a mother is married at the time of birth decreases over time. The fraction of singleton births decreases while the fraction of LBW births increases over time. Changes for all outcomes are similar across birth orders. Figures 4A-4F plot the same outcomes by mother's age at the time of birth and birth year. They show that while the total number of births decreased steadily over time for younger mothers, the number of births for older mothers (age 29-35) remained constant over time. Mother's educational attainment increased while the fraction of mothers who were married at the time of giving birth decreased for all age groups. The fraction of singleton births decreased while the fraction of LBW births increased. These patterns are similar across age groups.

Table 2 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 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.[8] Columns (1)-(2) and (3)-(4) 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. In other words, survival rates for girls of third and higher-parity births born after the reform were increasing relative to boys.

Figures 5A and 5B plot neo-natal mortality rates (within one month) by birth order for girls and boys, respectively. They show that infant mortality was decreasing at the same rate across birth orders for both sexes.

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 A1 columns (1) and (2). They are statistically significant for post-reform cohorts at the 1% and 5% levels. The coefficients are plotted in Figure 6A. 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. 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

[8] Source: World Development Indicators.

Figure 1A.

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 A1 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 6B. The figure shows that relative to mothers who were 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 $\hat{\beta}_2$ and $\hat{\beta}_3$ and their robust standard errors are shown in Table 3 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. We estimate equation (4) to examine the effect of sex-selective abortion on fraction of males born by mother's age. The estimates are shown in Table 3 column (2). 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 $\hat{\beta}_2$ and $\hat{\beta}_3$ and their robust standard errors are shown column (3) 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). First, we examine mothers' age, educational attainment and marital status at the time of birth. The coefficients are plotted in Figures 7A-7C. Figures 7A and 7B show that the reforms caused boys of higher birth parities to be more likely born to older and more educated mothers. The first finding is not surprising since the earlier results show that the reform caused older mothers to have more boys. The second finding is likely to partially come from the fact that mother's age at birth and educational attainment are highly positively correlated. It also may reflect that more educated mothers have better access to new technology or may be better able to afford sex-selective abortion. Figure 7C shows that the reform did not have any sex-differential effect on the marital status of mothers. Next, we estimate the same equation with the fraction of singleton and LBW births as dependent variables. The coefficients are plotted in Figures 7D and 7E. They show that the reform did not have any sex-differential effects on the health composition of children born. None of the estimated coefficients in this specification are statistically significant. For brevity, they are not reported in the paper.[9] We also examined the effects on father's age and educational attainment. The effects are similar as on mother's age and educational attainment because these outcomes are very correlated for mothers and fathers. The estimates are smaller in magnitude relative to those for mothers and are not statistically significant. The estimates are not reported in the paper.

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, six and twelve months as dependent variables. The coefficients and their robust standard errors are reported in Table 4 Panel columns (1)-(3). The triple difference estimates in Column (1) show that the reform had no effect on relative mortality within one month for second order births. But for third and higher-parity births, it increased mortality for males relative to females by 0.06 percentage-points. The effect is statistically significant at the 5% level. Columns (2) and (3) show that there is no effect for survival into the sixth and twelfth months. For these measures, the fraction of deaths is positively associated with low birth weight, and negatively correlated with being a singleton birth, mother's age and the mother being married.[10]

[9] Appendix Table A2 Panel A reports the coefficients for a simpler specification where we interact birth order dummy variabes with sex and a dummy variable for being born after the reform (rather than each birth year dummy variable). The estimates are not statistically significant. Panel B reports the estimates of the effect of the reform by mother's age at birth instead of birth order. They are also not statistically significant.

[10] These estimates are not reported in the paper.

Next, we examine the effect of the reform by mother's age by estimating equation (8). The coefficients and their robust standard errors are shown in Table 4 Panel B columns (1)-(3). Like the analysis on fraction of males at birth, the estimated coefficients are larger in magnitude for older mothers. However, the coefficients show that following the reform, boys born to older mothers to had lower mortality relative to girls. Column (1) shows that following the reform, the fraction of death within one month for boys born to mothers who were 28-35 years of age to be 0.05 percentage-points lower than girls; and those born to mothers over 35 years of age to be 0.13 percentage-points lower than girls. The estimates are statistically significant at the 1% level and similar for all three outcomes.

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 2 Panel A3). The results on neo-natal mortality (within one month) show that the reform decreased female mortality relative to male mortality at higher birth orders by 0.06 percentage-points, approximately a 25% reduction in neo-natal mortality. The fact that there is no effect on sex-differential mortality within six and twelve months probably reflects two factors. First, differential neglect resulting in EFM is most likely to occur to very young infants. This can either be because parents are less attached to very young infants and/or that infants are most vulnerable in the neo-natal period. This is consistent with the fact that approximately 50% of mortality within 12 months is attributable to mortality within the first month.[11] The finding that there is no effect on mortality within 6 months and 12 months may reflect the possibility that absent differential treatment from parents, male infants are more vulnerable relative to female infants. While the mechanism is not well-understood, it is a common finding in the health, demography and epidemiological literature that male infant mortality is higher relative to females. This is a good avenue for future research.

Aside from the results presented in this paper, we also collected data on the total fertility history of mothers and matched them to the birth and death registry data to investigate whether the impact of abortion on sex ratios at birth or differential mortality varied by the sex of the existing children. We found that parents were equally likely to use abortion to select for boys whether the first two children were both girls, both boys, or if there were one of each sex. Similarly, we found no differential effect for sex-differential mortality. This suggests that parental sex preferences are not such that they wish to have at least one boy. If this were the case, then parents who only have girls would be more likely to select for boys than parents with only boys. Rather, the results suggest that parents desire to have as many boys as possible.

If the improvement in relative female survival is entirely due to the fact that parents who used to select post-natally are now able to abort pre-natally, then the results show that approximately 9% of parents who were using abortion to select for sex were those who would have selected post-natally absent access to abortion. (See Appendix Table A3 for calculation). If we apply this to the pre-reform level of sex-selection, it means that approximately 15% of parents who were sex-selecting post-natally before the reform would have substituted to sex-selective abortion.[12] Our results can also shed a little light on the question of what type of parents substituted from post-natal to pre-natal 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

[11] Our findings are consistent with Ebenstein (2007) which finds that sex-selection in Taiwan today mostly occurs pre-natally vias abortion or post-natally during the first month.

[12] Assume that the fraction of males at birth with no intervention is 50.1% and that the pre and post- reform sex ratios at birth were 51.5% and 52.4% males. Hence, the proportion of parents who had selected before the reform and would have substituted to abortion is

$$[(53.5 - 50.1)0.09] \times (51.5 - 50.1) = 0.15$$

technologies must have been done by mothers who were younger at the time of birth.

While this study explains the observed increase in the fraction of males at birth, it cannot explain the pre-existing sex imbalance. Before abortion was legalized, the fraction of boys at birth was approximately 51.5%. The most likely explanation according to anectdotal evidence was that parents were illegally using abortion to select the sex prior to the reform. We attempted to investigate this in two ways. First, we attempted to examine sex ratios at birth during the 1950s and 60s when Ultrasound B was not available. If the male-biased sex ratios at birth prior to the legalizaion of abortion was caused by illegal sex-selective abortions, then sex ratios at birth before Ultrasound B was available should be closer to 50%. Unfortunately, Taiwan only began collecting data on sex ratios at birth starting in 1970, when Ultrasound B was already available. Second, we collected data on maternal mortality and examined maternal mortality before and after abortion was legalized. If illegal abortions are unsafe relative to legal abortions, then the legalization of abortions should decrease maternal mortality rates. (This assumes that death caused by a badly performed illegal abortion would still be reported as maternal mortality). We observe a dramatic decrease in maternal mortality during this era. However, because of the extremely low maternal mortality rates (approximately 40 per year in all of Taiwan), we are not able to divide the data by birth order to statistically analyze whether the decrease was relatively higher for mothers of third and higher parity births or older mothers.

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 51.5% 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 sex imbalance.

The stark results on relative female mortality show that access to abortion decreased female neo-natal mortality by 25% relative to males. They show that up to 15% of parents who were selecting post-natally before the reform would have subsituted to pre-natal sex selection using abortion. In other words, for every 100 abortions, 15 lives of born girls are saved. If we take these results 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 160,000 more girls will die neotally each year. [13]

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. 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.[14]

[13] China in 2000 had about 17.7 million births. At least 57% were boys. Hence, there is a seven percentage point deficit 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 80,000.

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.

[14] 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 1: Fraction of Males at Birth Over Time

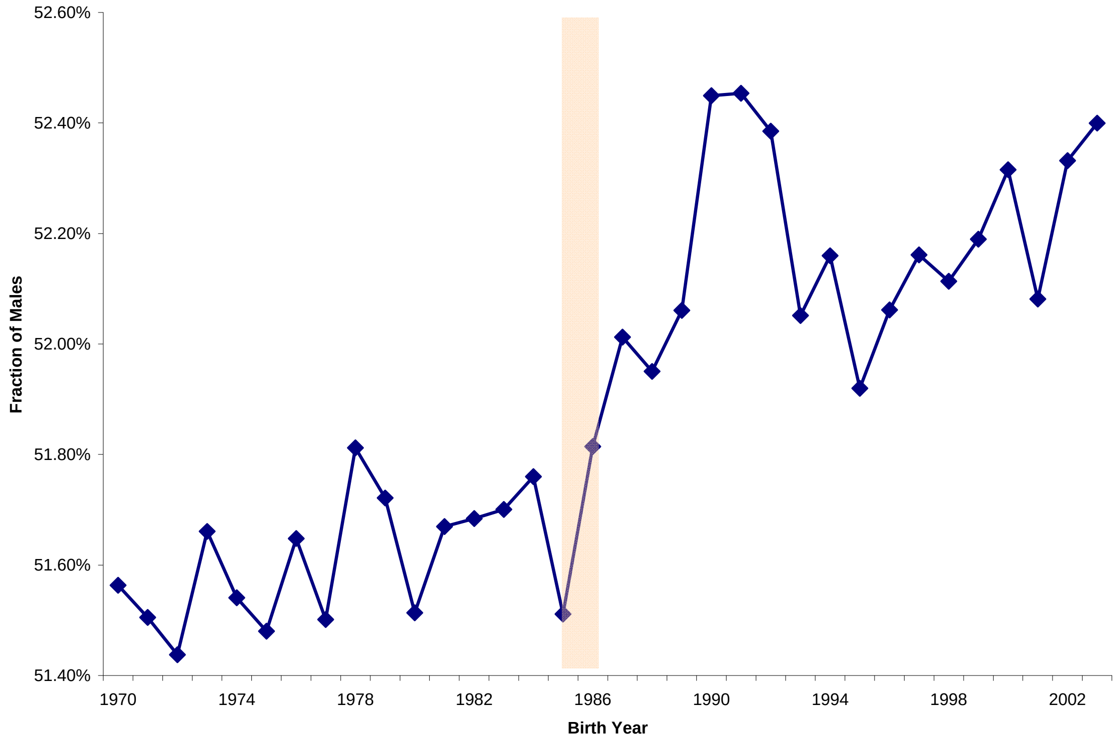

Figure 2A: The Fraction of Males by Birth Year and Birth Order

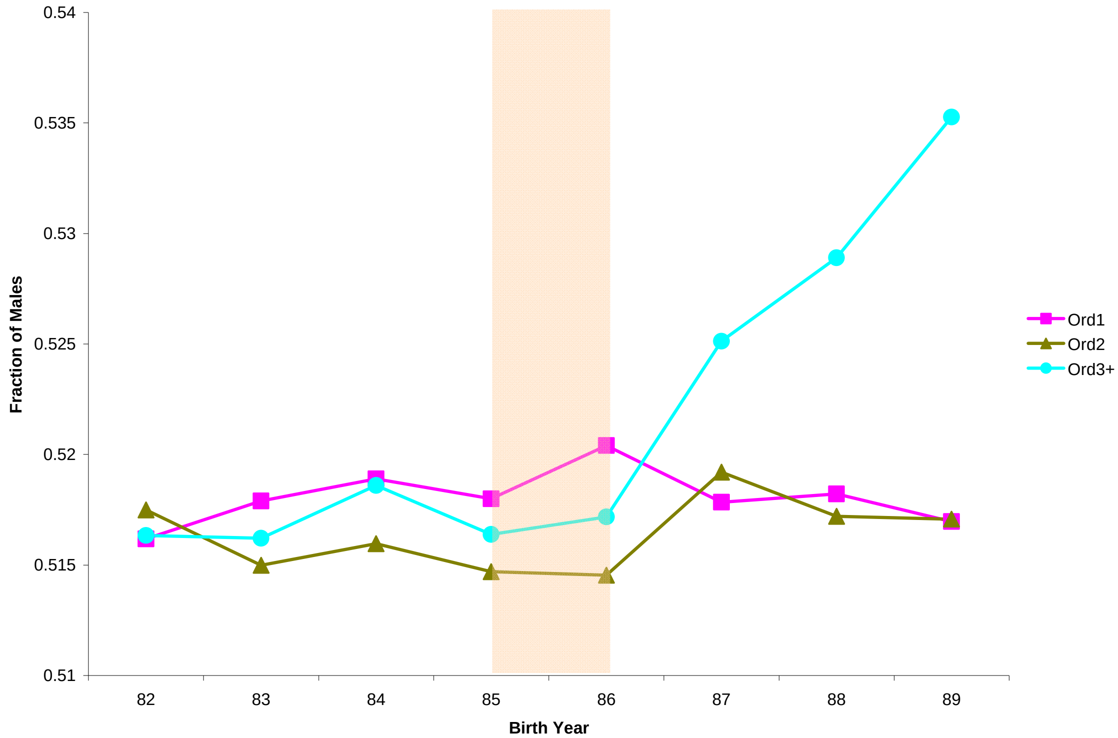

Figure 2B: The Fraction of Males by Birth Year and Mother's Age

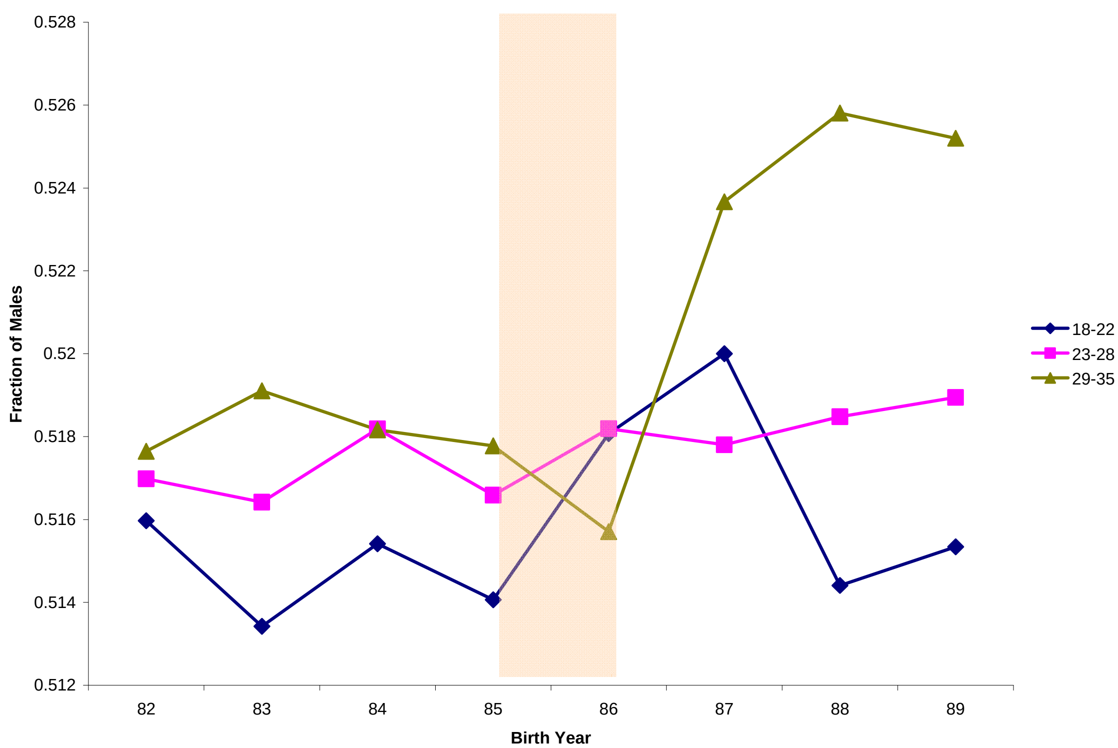

Figure 3A: Log (Total Births) by Birth Year and Birth Order

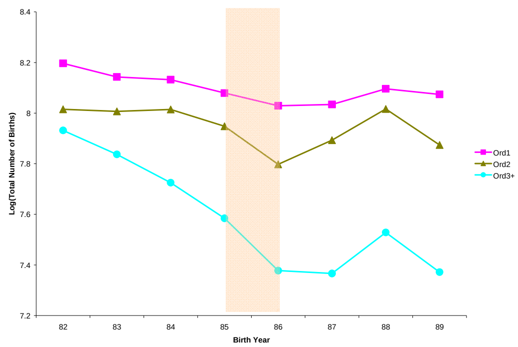

Figure 3B: Mother's Age by Birth Year and Birth Order

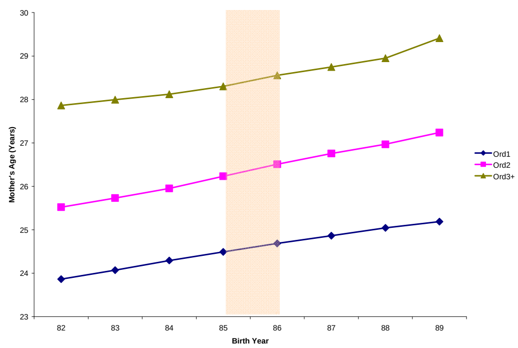

Figure 3C: Mother's Educational Attainment by Birth Year and Birth Order

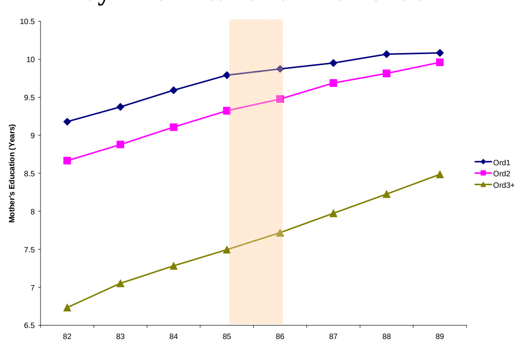

Figure 3D: Fraction of Mother's Married at Birth by Birth Year and Birth Order

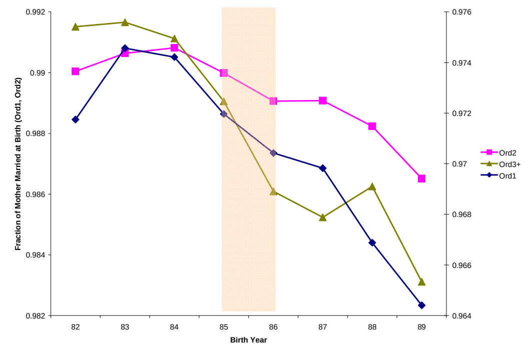

Figure 3E: Fraction of Singleton Births by Birth Year and Birth Order

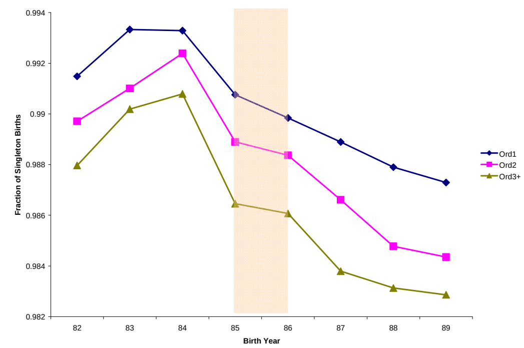

Figure 3F: Fraction of LBW Births by Birth Year and Birth Order

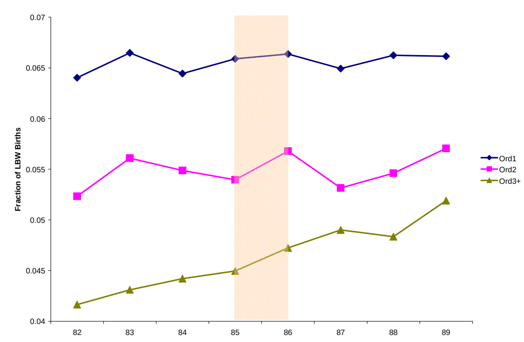

Figure 4A: Log (Total Births) by Birth Year and Mother's Age

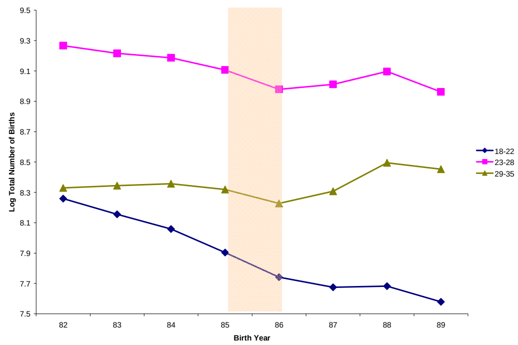

Figure 4D: Fraction of Singleton Births by Birth Year and Mother's Age

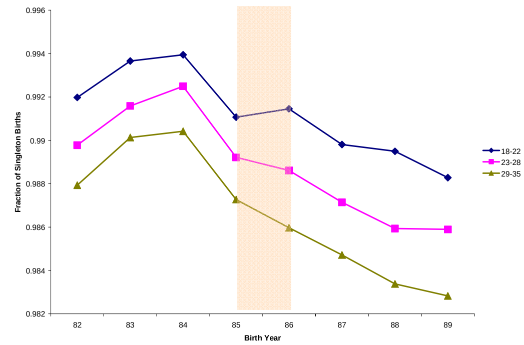

Figure 4B: Mother's Education by Birth Year and Mother's Age

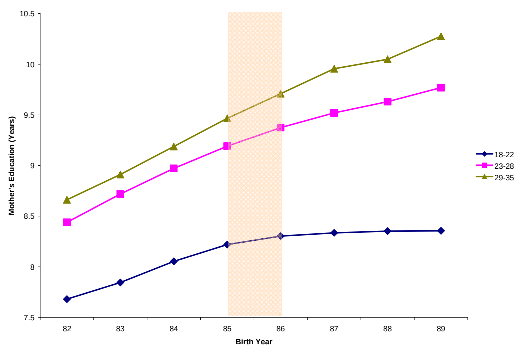

Figure 4E: Fraction of LBW Births by Birth Year and Mother's Age

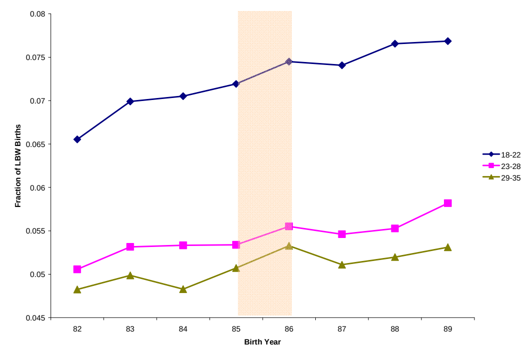

Figure 4C: Fraction of Mother's Marital Status at Birth by Birth Year and Mother's Age

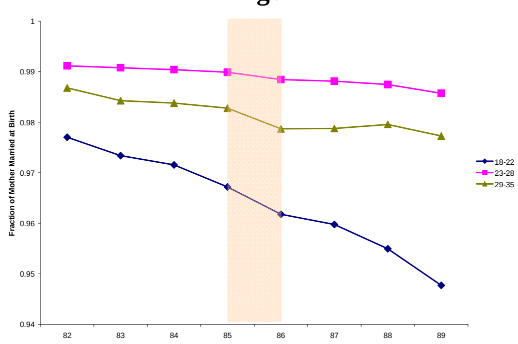

Figure 5A: Female Infant Mortality by Birth Order and Birth year

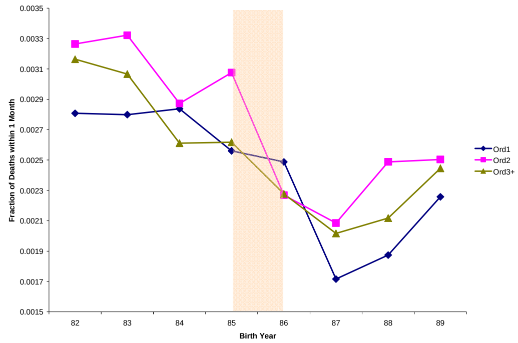

Figure 5C: Female Infant Mortality by Mother's Age and Birth Year

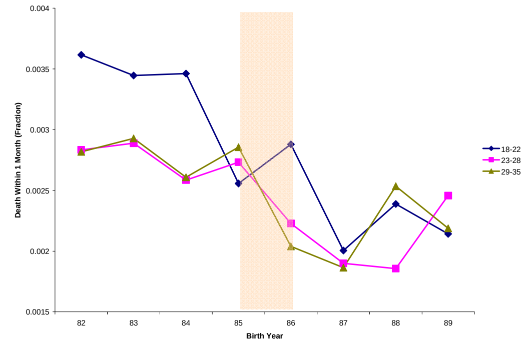

Figure 5B: Male Infant Mortality by Birth Order and Birth year

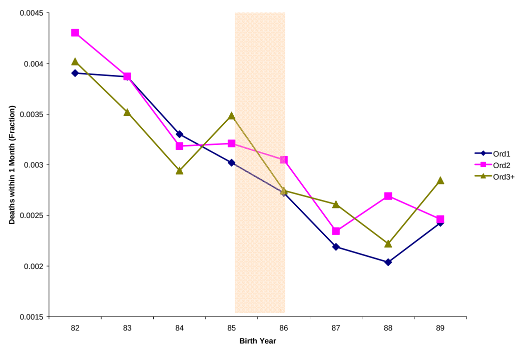

Figure 5D: Male Infant Mortality by Mother's Age and Birth Year

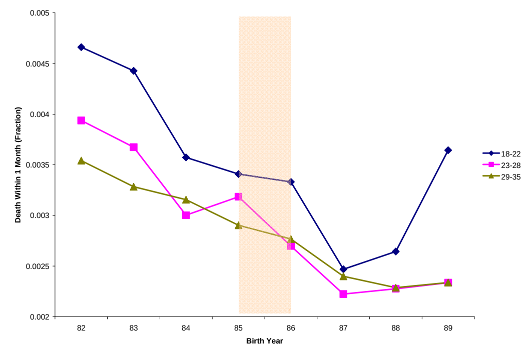

Figure 6A: The Effect of Abortion on Fraction of Males by Birth Order

Coefficients of the interaction terms of birth year dummy variables and birth order dummy variables

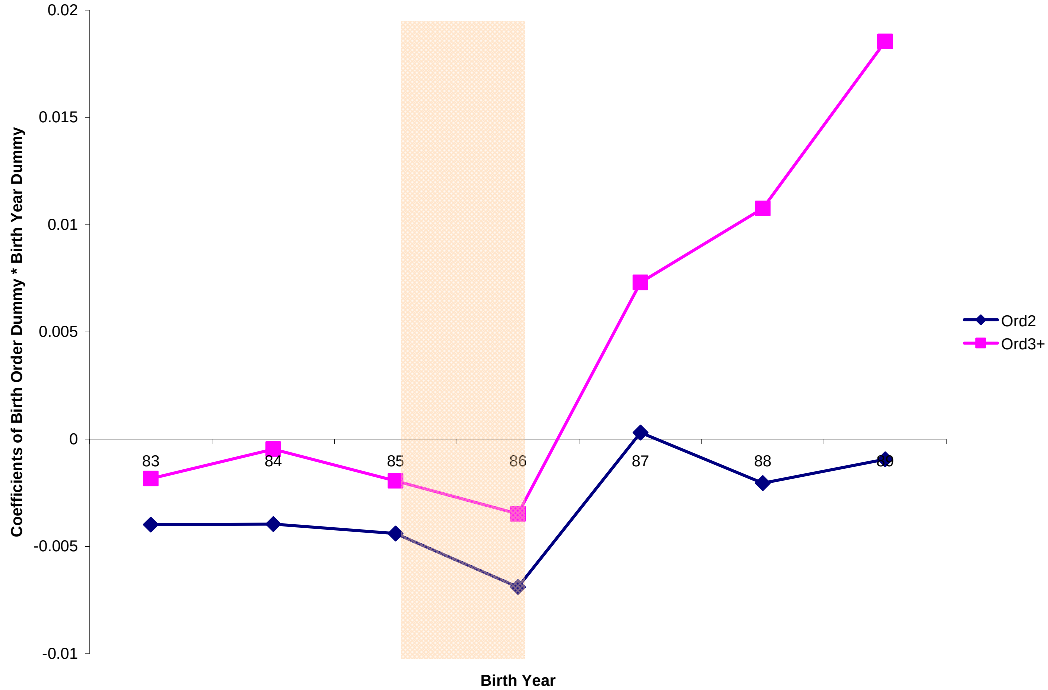

Figure 6B: The Effect of Abortion on Fraction of Males by Mother's Age

Coefficients of the interaction terms of birth year dummy variables and mother's age dummy variables

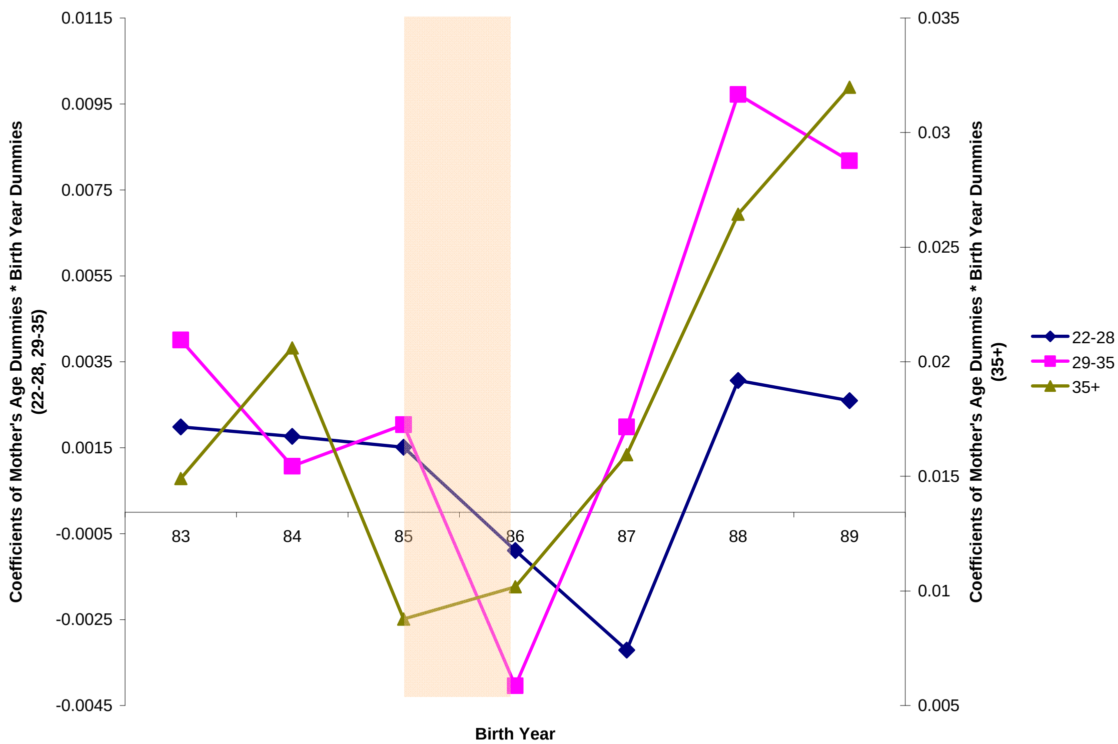

Figure 7A: The Effect of Abortion on Mother's Age by Birth Order and Sex

Coefficients of the interaction terms between male, birth order dummy variables, and birth year dummy variables

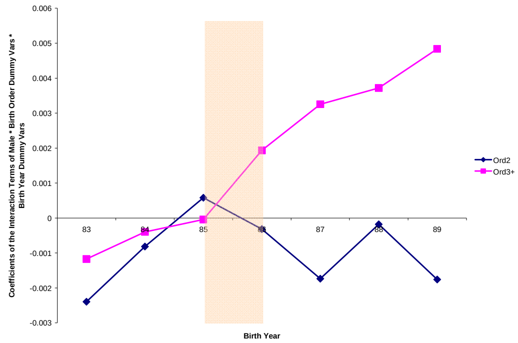

Figure 7B: The Effect of Abortion on Mother's Education by Birth Order and Sex

Coefficients of the interaction terms between male, birth order dummy variables, and birth year dummy variables

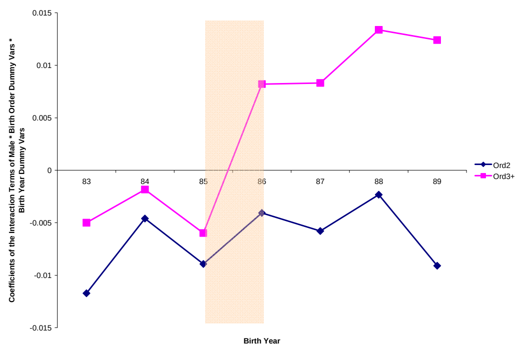

Figure 7C: The Effect of Abortion on Mother's Marital Status by Birth Order and Sex

Coefficients of the interaction terms between male, birth order dummy variables, and birth year dummy variables

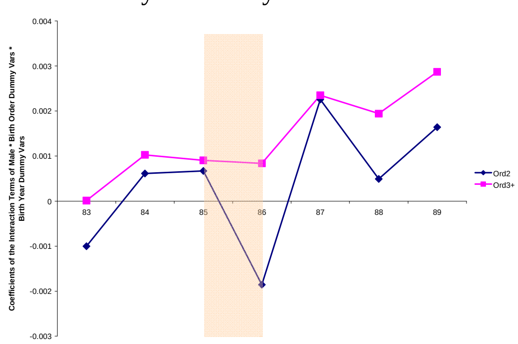

Figure 7D: The Effect of Abortion on the Fraction of Singleton Births by Birth Order and Sex

Coefficients of the interaction terms between male, birth order dummy variables, and birth year dummy variables

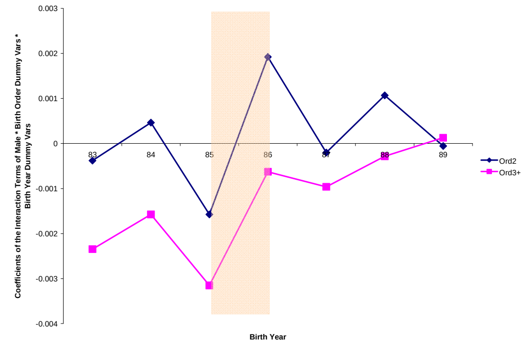

Figure 7E: The Effect of Abortion on the Fraction of Low Birth Weight Births by Birth Order and Sex

Coefficients of the interaction terms between male, birth order dummy variables, and birth year dummy variables

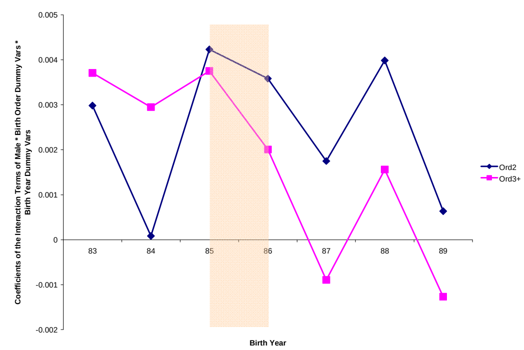

Table 1: Descriptive Statistics on Birth and Parental Characteristics by Birth Year and Birth Order

Variable	I. Born 1982-84 Obs	Mean	Std. Err.	II. Born 1985-89 Obs	Mean	Std. Err.	III. 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 2: 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	Born 1982-84		Born 1985-89		Diff	Born 1982-84		Born 1985-89		Diff	DD: (6)-
Months	Obs	Mean	Obs	Mean	(2)-(1)	Obs	Mean	Obs	Mean	(5)-(4)	(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 3: The Effect of Abortion on Fraction of Males by Birth Order and/or by Mother's Age

Dependent Variable: Fraction of Males			
	(1)	(2)	(3)
Ord2 * Born 1985-89	0.000		0.031
	(0.001)		(0.064)
Ord3+ * Born 1985-89	0.007		-0.183
	(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)	
Observations	794	1057	794
R-squared	0.09	0.01	0.14

Robust standard errors in parentheses.

Columns (1) and (3) use data aggregated by birth order, birth year, and birth county.

Column (2) uses data aggregated by mother's age, birth year, and birth county.

All regressions control for birth year fixed effects.

Column (1) also controls for birth order fixed effects.

Column (2) also controls for mother's age main effects.

Column (3) also controls for mother's age, mother's age * birth order dummies, and mother's age * born 1985-89.

Table 4: 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 Variable: Death Within X Months		
	1 Month	6 Months	12 Months
	(1)	(2)	(3)
A. By Birth Order			
Ord2 * Male * Born 1985-89	0.0002	0.0002	0.0005
	(0.0003)	(0.0005)	(0.0005)
Ord3 * Male * Born 1985-89	0.0006	-0.0002	0.0001
	(0.0003)	(0.0005)	(0.0006)
Observations	1586	1586	1586
B. By Mother's Age			
Mom22-28 * Male * Born 1985-89	-0.0004	-0.0006	-0.0007
	(0.0002)	(0.0003)	(0.0004)
Mom29-35 * Male * Born 1985-89	-0.0005	-0.0007	-0.0009
	(0.0003)	(0.0004)	(0.0004)
Mom35+ * Male * Born 1985-89	-0.0013	-0.0015	-0.0017
	(0.0007)	(0.0009)	(0.0009)
Observations	2111	2111	2111

Robust standard errors in parentheses

Data for Panel A are aggregated to sex x birth order x birth year x birth county cells.

Data for Panel B are aggregated to sex x mother's age x birth year x birth county cells.

All regressions control for the full set of double interaction effects, birth order or mother's age fixed effects, and birth year fixed effects.

Table A1: 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.

Table A2: The Effect of Abortion 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; or the interaction terms of mother's age dummy variables, birth year dummy variables, and a dummy variable for male.

	(1)	(2)	(3)	(4)	(5)	(6)
Dependent Variables:	**Log(Mother's Age)**	**Log(Mother's Edu)**	**Log(Father's Edu)**	**Mother Married**	**LBW**	**Singleton**
A. By Birth Order						
Ord2*Male*Born 1985-89	0.000	-0.001	-0.002	0.001	0.002	0.000
	(0.006)	(0.015)	(0.014)	(0.001)	(0.002)	(0.001)
Ord3*Male*Born 1985-89	0.003	0.010	0.010	0.001	-0.001	0.000
	(0.005)	(0.016)	(0.014)	(0.001)	(0.002)	(0.001)
Observations	1586	1545	1578	1586	1586	1586
A. By Mother's Age						
Mom22-28 * Male * Born 1985-89		-0.001	0.002	-0.001	0.002	-0.001
		(0.007)	(0.006)	(0.001)	(0.001)	(0.001)
Mom29-35 * Male * Born 1985-89		-0.002	0.000	-0.001	0.003	0.000
		(0.009)	(0.008)	(0.001)	(0.001)	(0.001)
Mom35+ * Male * Born 1985-89		0.011	0.013	0.004	-0.004	0.001
		(0.015)	(0.014)	(0.003)	(0.003)	(0.002)

Robust standard errors in parentheses.

Data for Panel A are aggregated to sex x birth order x birth year x birth county cells.

Data for Panel B are aggregated to sex x mother's age x birth year x birth county cells.

All regressions control for the full set of double interaction effects, birth order or mother's age fixed effects, and birth year fixed effects.

Table A3: The Proportion of Change in Sex Ratios Cause by a Change in Sex-Differential Mortality

	(1)	(2)	(3)	(4)
	Number Born	Fraction Death Within 1 Month	Number Left after 1 Month = (1) x (2)	Fraction of Change in Girls Accounted by Change in Mortality = [1-(3)/(1)]
Pre	100	0.0030	99.70	
Post =(Pre-Prex0.007)	99.3	0.0024	99.06	
Difference	-0.7		-0.64	**0.09**