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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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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

This paper presents novel empirical evidence on the impact of access to abortion on sex ratios at birth (SRB), excess female mortality (EFM) and fertility in Taiwan. For identification, we exploit plausibly exogenous variation in the availability of sex-selective abortion caused by the legalization of abortion. Our results show that the legalization of abortion accounts for almost all of the observed increase in SRB during the 1980s and decreased EFM by approximately 20%. Approximately ten more female infants survived for every one hundred that were aborted. Interestingly, we find that while abortion reduced overall fertility, it increased fertility for older mothers.

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1 Introduction

Missing women, a term coined by Amartya Sen, refers to the observation that a male-*biased* sex imbalance exists in many Asian countries. For example, in China, India, 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%.[1] There is much concern that sex imbalance in the population could lead to increased crime rates or distort the marriage market (Angrist, 2002; and Samuelson, 1985).[2] In countries such as China, Taiwan and South Korea, the phenomenon has only increased over time, despite rapid economic growth and social "modernization". 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.[3] For example, Figure 1 shows that the fraction of males at birth for Taiwan increased from approximately 0.515 in the early 1980s (e.g. the same as for countries not known for boy-bias such as the U.S. or Western European countries) to much higher levels by the 1990s. Almost all of the increase is observed for higher parity births. By 1990, the fraction of males at births for Taiwan had risen to approximately 0.54. Given the large potential impact that sex-selective abortion could have had 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 quantify its contribution to the observed imbalance. It is all the more surprising when one notes that all sides in the recent heated debate on the determinants of the observed sex-imbalance agree on one thing: the potentially large contribution of access to sex-selective abortion.[4]

[1] It has been observed that the same male-biased sex ratios exists in many Muslim countries.

[2] For example, Angrist (2002) and Samuelson (1985) study the long-run impact of sex imbalance on the marriage market. More recently, Edlund et al. (2007) finds that increased male-biased sex imbalance increases crime.

[3] Das Gupta and Bhat (1997) refers to the observation of the simultaneous decrease in fertility and increase in EFM as the "intensification" effect and the fact that boy-biased sex imbalance is higher at higher birth parities as the "parity" effect.

[4] 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 (2008),

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.[5] They are supported by international policy makers. According to the *United Nations Program of Action*, adopted by the *International Conference on Population and Development* in Cairo, the increasing use of technologies to determine fetal sex and pre-natal sex selection are "harmful and unethical". The objective of the *Program* was to stop such activities (UN 1994).[6] These policy makers do not seem to have considered the potential for unintended negative consequences of banning sex selective abortion for relative female survival rates (after birth). Forcing parents with strong boy-preferences to give birth to a girl may force some of those parents to select post-natally.[7] Until now, empirical testing of this potential substitution effect has been hindered by the lack of data and appropriate research design (Goodkind, 1999).[8]

This paper addresses these important questions by offering novel empirical evidence on the causal effects of access to sex-selective abortion on sex ratios at birth (heretofore defined as the fraction of males at birth) and excess female mortality (EFM or female infant mortality relative to male infant mortality) after birth. To establish causality, 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 methodological contribution of this study is to resolve identification issues

Rholf et. al. (2005), Rosenzweig and Schultz (1982), Thomas (1991) and Thomas et al. (1994) study the effect of relative female socioeconomic 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 (2008), 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. Anderson and Ray (2008) describes the age-specific patterns of sex-imbalance across different regions of the world.

[5]For policies in China, see Greenlaugh and Li (1995); for those in India see Balakrishnan (1994), Bumiller (1990), Nandan (1993), Parikh (1993) and Roggencamp (1985).

[6]The *Program* stated its objectives as "to eliminate all forms of discrimination against a girl child.. which results in harmful and unethical practices regarding female infanticide and prenatal sex selection" (UN1994, Article 4.15).

[7]The rationale is similar to the one used in Donohue and Levitt's (2001) study 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.

[8]Goodkind (1996) argues that the ability to sex-select post-natally and the high levels of sex imbalance before abortion was available in some countries makes the subsitiution between prenatal and post-natal sex selection a troubly concern.

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 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. 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.

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. For that group, 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 20%. These two findings together suggest that approximately 10% 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 hundred abortions of female fetuses, approximately ten lives of girls born are saved.

Interestingly, we find that the introduction of abortion had heterogeneous effects on fertility. While it reduced fertility overall, it *increased* fertility for older mothers. Our evidence also suggests that older mothers were more likely to have sons and that relative survival rates for female infants born to older mothers did not improve after abortion was legalized. Together, these results are consistent with the hypothesis that by increasing the option value of pregnancy for mothers with son-preference, legalizing abortion induced older mothers who are commonly believed to bear a higher biological cost of child rearing, to have children. See Sections 2.1 and 5 for more detailed discussions.

Studying the effect of sex-selective abortion in Taiwan has both advantages and disadvantages. On the one hand, the data are much better than many 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.[9] Moreover, 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 it easier to relate the reduced form results to the reduction in the cost of sex selection. On the other hand, Taiwan is very small geographically. The majority of the population lives in half of an island that is approximately 275 miles long and 87 miles wide. Expecting parents effectively have access to health care facilities in the entire country. Hence it does not make sense to exploit cross-sectional variation for identification purposes. More importantly, Taiwan is wealthier than countries such as mainland China and India. And relative to those countries, it has a much lower rate of infant mortality and a significantly lower rate of post-natal sex selection. 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.

Our paper makes several important contributions. First, it is the first to provide rigorous empirical evidence for the enormous role that increased access to sex-selective abortion has

[9] Appendix Figures A1 and A2 show that similar trends over time existed for South Korea and China. (The figures are based on the data shown in Appendix Table A1). This is consistent with the belief that the increase in sex selective abortion tehcnologies in these countries were increasing the sex imbalance of births. That the technology diffused slowly and organically (without obvious sources of exogenous variation such as legislative reforms) is consistent with the observation that the changes in SRB in South Korea and China are not discrete over time as it is in Taiwan. This also means that it is difficult to establish causality in these other contexts.

contributed to the Missing Women phenomenon. In showing that the boy-biased sex imbalance almost exclusively exists for higher birth parities, our study is most closely related to the recent work by Abrevaya (2009).[10] Second, we provide evidence for two novel insights: that there is a tradeoff between allowing sex-selective abortion and increasing relative female survival because parents are no longer forced to have unwanted girls; and that reducing the cost of abortion can induce women who bear a high cost of child rearing but who have son preference to have children. To the best of our knowledge, no previous studies have discussed or studied these two mechanisms. The only other study that has considered the potential for heterogeneous treatment effects under a rigorous framework in the context of sex-selection is the recent work by Ebenstein (2009).[11]

These results have important policy implications. First, they suggest that banning sex-selective abortion could increase mortality for female infants. 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 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 161,000 girls in the two countries combined. Second, the results show that reducing the costs of abortion in a society with son preference and where mothers face heterogeneous costs for child rearing has an ambiguous *a priori* effect on total fertility. It would depend on the reduction in the number of births from unwanted pregnancies relative to the increase in the number of births from women who have children because abortion lowers the cost of sex selection.

The paper is organized as follows. Section 2 discusses background, conceptual framework and empirical strategy. Section 3 describes the data. Section 4 presents the empirical results.

[10] Abrevaya (2009) shows that there is stark inequality for higher parities amongst populations of Asian decent in the 1980 U.S. Population Census and discusses the role of sex-selective abortion as a contributing factor to this phenomenon. Studies in demography such as Das Gupta (1987), Das Gupta (1997), Gu and Roy (1995), Muhri and Preston (2001), Park and Cho (1995) and Pebley and Amin (1991) have described the changes over time in sex ratios at birth by birth parity in several East Asian countries. Arnold, Kishor and Roy (2002) have associated access to abortion with increasing sex ratios. 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 identification of the impact of sex selective abortion.

[11] Ebenstein (2009) estimates a dynamic model of sex selection where parents have access to sex-selective abortion. Using data on fines for violating the One Child Policy in China, he finds that raising the costs of fertility and sex selection have heterogenous effects across households of different wealth and education.

Section 5 interprets the results. Section 6 offers conclusions.

2 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 can also reveal the sex of the fetus (e.g. *amniocentesis*, *chorionic villus sampling*), the least expensive and most 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 effects for different methods of aborting a fetus. 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 "post" 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 January 1, 1985. It stated that women can get an abortion if (a) the woman was raped (b) the fetus has some genetic disease or (c) the pregnancy would affect the mental health of the woman and the family. The added clause (c) effectively legalized non-medically motivated abortions. However, "practicing abortion" was not made explicitly legal. According to the 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). Therefore, the first cohort that was exposed to sex selective abortion were not born until 4 months into 1985. 1986, then, was the first year that the legislation was in effect for the full twelve months.[12]

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 Series* (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

[12]Changing the cutoff from January 1985 to April 1985 does not affect the main 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. Note that our stragey divides individuals according to the calendar year of their birth. We could alternatively define each yearly cohort to be those born from April 1st - March 31st. Doing so results in very similar estimates as those reported in this paper. Therefore, for simplicity, we use calendar years and only report these estimates. The other estimates are available upon request.

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 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 age group. It shows that women who were 29-35 experienced the largest increase (7.5 percentage-points).[13] 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 virilus 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.[14] 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-detection, using Ultrasound B during routine pre-natal care may have also increased the quality of pre-natal care more generally.

[13]Survey data from KAP on abortion should be interpreted very cautiously because the sample size is very small. For example, in 1985, a total of only 57 women report as ever having had an abortion. Therefore, it is difficult to further split the data by birth parity and mother's age. For example, Freedman, Chang and Sun (1994) show that in the KAP survey, older mothers (age 30-39) on average have more abortions at fourth and fifth and higher parity births in 1992 ersus 1985. However, this is not true for 3rd parity births.(See Table 16 of Freedman, Chang and Sun, 1994).

[14]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.

2.1 Conceptual Framework

To motivate the empirical estimates, we provide a simple framework to illustrate how the effect of a decrease in the cost of abortion caused by the legalization can affect the type of mothers who choose to give birth, the number of children born, and the fraction of males born. We consider two extreme cases where only the cost of abortion differs. In the first case, the cost is infinitely high (e.g. no abortion). In the second case, the cost is zero (e.g. legalization and free access of abortion).

We assume that in the population, every woman has a cost c of child rearing which include all costs associated with bearing and raising a child. There are three types of women $i = \{1, 2, 3\}$ with costs $\{0, c, \infty\}$ and they represent populations of size p_i and we assume that $\sum_{i=1}^{3} p_i = 1$ so that the population is normalized to 1. We assume that there is son preference such that the utility of having a boy is $B > 0$ and the utility of having a girl is $G < 0$. The ex-ante probability of having a boy from nature is 1/2. We make the following assumption:

$$B > c > .5B + .5G > 0 > G \tag{1}$$

This means that Type 1 women always want to have children. Type 2 women have children if they have a reasonable probability of having a boy. This probability increases as the cost of sex-selective abortion decreases. Type 3 women never want to have children. The order of events are as follows.

1. Some women decide to get pregnant. A fraction of women π have accidental and unwanted pregnancies.

2. Nature chooses a signal z which indicates the probability of boy. We will say that this signal is uniformly distributed between zero and one , $z \sim U[0, 1]$.

3. Women decide whether or not to have an abortion.

Case #1: No Abortion, $c = \infty$

Imagine if abortion is not possible. Then only Type 1 women choose to have children. Type 2 and 3 women will have children if they accidentally become pregnant. The total number of kids is then

$$p_1 + (p_2 + p_3)\,\pi \tag{2}$$

and the total number of boys is

$$\frac{p_1 + (p_2 + p_3)\,\pi}{2} \tag{3}$$

which implies that the fraction of boys is 1/2.

Case #2: Free Abortion $c = 0$

In this situation women can freely abort the child if the benefit of having the child exceeds the cost of rearing them. More specifically, let us define cutoffs z_1^* and z_2^* :

$$z_1^* B + (1 - z_1^*)\,G = 0 \tag{4}$$

$$z_2^* B + (1 - z_2^*)\,G = \eta \tag{5}$$

Our assumptions ensure that these cutoffs are between 0 and 1 and that $z_2^* > z_1^*$. Therefore, all Type 1 and Type 2 women will become pregnant. Type 3 women will become pregnant with probability π (same as if abortion was not available).

This illustrates the first insight: that a sufficient decrease in the cost of abortion could induce some women to become pregnant by increasing the option value of pregnancy. In particular, Type 2 mothers who bear a high cost of child rearing are now willing to become pregnant because they will be able to abort if it is not a son. (Without abortion, the only way to reveal whether a child will be a son is to give birth after nine months of pregnancy. Therefore, access to abortion saves mothers the costs associated with the last few months of pregnancy and child birth).

Next, we investigate the effect on the number of children born. Mothers of Types 1 and 2 will have an abortion if $z_i < z_i^*$. Type 3 women will always abort if pregnant. Given the assumption that z is distributed uniformly, the probability of keeping a child for women of Type i in the population is $1 - z_i^*$. It follows that the number of children being born is now:

$$(1 - z_1^*)\,p_1 + (1 - z_2^*)\,p_2 \tag{6}$$

To see the effect of a decrease in the cost of abortion on the total number of children born, we can compare equation (6) above to equation (2) from the first case. It follows that a decrease in the cost of abortion will decrease the number of children born if the following is true.

$$(p_2 + p_3)\,\pi + z_1^* p_1 > (1 - z_2^*)\,p_2 \tag{7}$$

If the sum of the number of women who were having children due to unwanted pregnancies and the number of Type 1 mothers who now abort because they do not wish to have a daughter exceeds the number of Type 2 mothers who are now induced into giving birth, then decreasing the cost of abortion will decrease the total number of children born.

The second insight follows from the above: because abortion serves both as an instrument for reducing unwanted births and for improving the technology for sex selection, its effect on the number of births is ambiguous ex-ante. There are two forces at work here. On the one hand, abortion will decrease the number of births by allowing unwanted pregnancies to be aborted. On the other hand, by increasing the expected value of giving birth (e.g. the probability of having a son), it has increased the number of births from mothers who bore a high cost of child bearing and who otherwise would not have had children. Therefore, the net effect of a decrease in the cost of abortion on the number of children born will depend on parameter values in equation (7): the probability of an unwanted pregnancy, the thresholds z_1^* and z_2^*, and the relative size of the population of each type of women. For example, if the probability of an unwanted pregnancy is very high relative to the number of women that are induced to giving birth, then the reduction in unwanted births from decreased costs of abortion will more likely dominate the increase in the number of births from the latter mechanism. In this case, the net effect of a decrease in abortion costs will likely be to reduce the number of births.

Third, we can investigate the effect on the fraction of boys born. The number of boys born will be revealed by the conditional distributions for each group of mothers. Because Type 3 mothers always abort if pregnant, we only need to consider Type 1 and Type 2 mothers here. Amongst these mothers, those who received signals above z_i^* such that $z_i > z_i^*$ will have children, where the probability of each mother getting a boy is z_i and the probability for getting a girl is $1 - z_i^*$. This is true for all $z_i > z_i^*$. The probability of having a boy for this population on average will be $E[z \mid z_i > z_i^*]$. Since the distribution of z is uniform, this will equal $\frac{1+z^*}{2}$. Therefore, the total number of boys will be:

$$\frac{(1 + z_1^*)(1 - z_1^*)}{2} p_1 + \frac{(1 + z_2^*)(1 - z_2^*)}{2} p_2 \tag{8}$$

which simplifies to

$$\frac{1 - {z_1^*}^2}{2} p_1 + \frac{1 - {z_2^*}^2}{2} p_2 \tag{9}$$

so that the fraction of boys born now equals:

$$\frac{\frac{1 - {z_1^*}^2}{2} p_1 + \frac{1 - {z_2^*}^2}{2} p_2}{(1 - z_1^*) p_1 + (1 - z_2^*) p_2} \tag{10}$$

Therefore decreasing the cost of abortion to zero will increase the fraction of boys born if the expression above in equation (10) is greater than the fraction of boys born in Case 1, which was 1/2. In other words, the following needs to be true for the legalization of abortion to cause an increase in the fraction of boys born.

$$\frac{(1 - {z_1^*}^2) p_1 + (1 - {z_2^*}^2) p_2}{(1 - z_1^*) p_1 + (1 - z_2^*) p_2} > 1 \tag{11}$$

Note that this will always be true because we have assumed that $z_2^* > z_1^*$ and $z_i^* \in [0, 1]$ for all i. Therefore, the third insight is that the decrease in the cost of abortion will lead to an increase in the fraction of boys born. In particular, equation (11) shows that the effect of decreasing the cost of abortion on the fraction of males born is increasing in threshold values for having children z_i^*'s (which can also be interpreted as the probability of having a son, or very loosely as son preference). It is also increasing in the proportion of Type 2 mothers relative to the proportion of Type 1 mothers.

In summary, the stylized framework in this section illustrates several insights on the effect of decreasing the cost of abortion and sex-selective abortion in an environment where there is son preference and heterogeneity in the cost of bearing children. First, it will increase the option value of pregnancy and therefore induce mothers who bear a high cost of child bearing to become pregnant and try to have a son. Our empirical results will test for this by examining whether abortion increased the number of children born to older mothers, who bear a much larger biological cost in child bearing according to existing medical studies.[15] Second, the framework illustrates how the reduction in unwanted births and the increase in births from mothers that we described in the first mechanism have opposing effects on the total number of births. The empirical results will estimate the net of these two effects by estimating the effect of the decrease in the cost of abortion on the total number of births. Third, the framework illustrates how the decrease in the cost of abortion, by improving the technology for sex-selection, will increase the fraction of males born. The empirical analysis will estimate the magnitude of this effect by examining the impact of the legalization of abortion on the fraction of males at birth.

Finally, there is an important effect that we have not formally described: decreasing the cost of abortion can increase the quality of children more generally. Thus far, the only dimension of quality in our framework is the sex of the child. Another potential dimension is the health of the child. The mechanism is straight-forward. If children that result from unwanted births receive lower investment, then the decrease in unwanted births will increase in the average quality of children (Donahue and Levitt, 2001). More specifically, because there will be a disproportionately larger decrease in the births of girls, we would expect to see quality increase disproportionately for girls born after abortion is legalized. The section on the empirical analysis will test for this by examining the effect of the legalization of abortion on relative female infant mortality.[16]

2.2 Empirical Strategy

The empirical strategy exploits 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

[15] For example, studies such as Augensen and Bergsje (1984), Dundaram, Liu and Laraque (2005) and Marmol et al. (1967) find that women over the age of 35 are significantly more likely to experience pregnancy-related deaths. Studies have also found that women over the age of 35 are more likely to experience pregnancy-related complications such as gestational diabetes, high blood pressure, placental problems, premature birth and stillbirths (Cleary-Goldman et al., 2005; Joseph, 2005; Usta and Nassar, 2008; Bahtiyar, 2008).

[16] Note that the empirical analysis with look at outcomes of interest by birth parity, a dimension which has been omitted from the conceptual framework. This omission for was done for the sake of simplicity. In the empirical analysis, we will find that all of the effects of abortion occurs for third and higher parity births. Therefore, we have made the simplifying assumption that births of lower parity children are exogenous and mothers only make decisions regarding the third and higher parity births.

born. In addition to birth year/cohort variation, we also exploit the variation in birth order. For parents who wish to have a boy, they are more likely to sex-select if their desire 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.[17] 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 occured at the time that abortion was legalized that would decrease the cost of sex-selection *and* decrease the cost more for higher birth parities. 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. 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. There is no reason to believe this is true. To be cautious, we will investigate this possibility by examining the effect of the reform on the composition of boys born relative to girls born across birth parities.

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 1985/86. The simpler pre-post differences-in-differences specification has the pitfall

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

that it would capture changes that occurred at any time after the reform.

$$Male_{it} = \sum_{i=2}^{3} \sum_{t=1983}^{1989} \beta_{it}(Ord_i \times Born_t) + \gamma_i + \rho_t + \varepsilon_{it} \tag{12}$$

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. Robust standard errors are reported for all estimations in this paper. The reference group is comprised of first-born children. It and all of its interactions are excluded from the regressions. 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 $(\beta_{3,t>=85} - \beta_{2,t>=85}) > (\beta_{3,t<85} - \beta_{2,t<85})$.

To examine whether the effect differs by mothers' age, we separate mothers into four age groups: 18-22, 23-28, 29-35 and 35 and above, and re-estimate equation (12) for each subsample of mothers. If older mothers who are more uncertain about the possibility of having more children are more likely to take up abortion as a method of selection, then the effect will be increasing in mothers' age. In the main specification, we address this by controlling for mothers' age and the interaction terms between mothers' age and post reform dummy variables. These interaction terms allow the effect of mother's age to vary over time.

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.

$$\begin{aligned} Male_{imt} &= \sum_{i=2}^{3} \beta_i(Ord_i \times Post_t) \\ &\quad + \sum_{m=2}^{4} \alpha_m(Mage \times Post_t) \\ &\quad + \mathbf{X}_{it}\psi + \phi_m + \gamma_i + \rho_t + \varepsilon_{it} \end{aligned} \tag{13}$$

We regress the fraction of males for individuals of birth order i, born to mothers that are age

m, in birth year t, $Male_{imt}$, on: the interaction terms between birth parity dummy variables, Ord_i, and a dummy variable for being born in 1985 or afterwards, $Post_t$; the interaction term between a dummy variable indicating that the mother is 22 to 28, 29 to 35, or 35 years of age and over, $Mage$, and $Post_t$; a vector of controls such as mother's education, $\mathbf{X}_{it}$; mother's age fixed effects, γ_m; 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 excluded from the regressions.

To estimate the effect of abortion on fertility, we repeat the previous estimates and replace the dependent variable with the natural logarithm of the number of births. The birth registries are at the birth level and we are not able to link mothers and calculate the total fertility rate of each woman. The total number of births is a reasonable proxy in the context of our study as the population of women of child bearing age does not significantly change in size during the seven years of our study.

To estimate the effect of sex-selective abortion on sex-differential mortality rates, estimate the following equation with the fraction of deaths within x months after birth as the dependent variable..

$$\begin{aligned} Death_{imts} = & \sum_{i=2}^{3} \beta_i (Ord_i \times Post_t * Male_s) + \sum_{m=2}^{3} \alpha_m (Mage_m \times Post_t \times Male_s) \\ & + \sum_{i=2}^{3} \delta_{it} (Ord_i \times Post_t) + \varphi\,(Mage_m * Post_t) + \pi_t (Male_s \times Post_t) \\ & + \sum_{i=2}^{3} \lambda_i (Male_s \times Ord_i) + \varpi\,(Mage_m \times Male_s) \\ & + \mathbf{X}_{it}\psi + \theta_s + \gamma_i + \rho_t + \phi_m + \varepsilon_{imts} \end{aligned} \quad (14)$$

We regress the fraction of deaths within x months of life for individuals of birth order i, born to mothers of age m in birth year t, and who are sex s, $Death_{imts}$ on: the triple interaction terms between dummy variables for birth order, Ord_i, a dummy for post, $Post_t$, and sex, $Male_s$; the triple interaction terms between dummy variables for mothers' age, $Mage_m$, $Post_t$, and $Male_s$; the full set of double interaction terms; a vector of controls such as mother's education, X_{it}; birth order fixed effects, γ_i; mothers' age fixed effects,ϕ_m;

a dummy for being male, θ_s; and birth year fixed effects, γ_i. If legalizing abortion increased relative female survival, then $\hat{\beta}_3 > \hat{\beta}_2$.

There are two main concerns for the identification strategy. First is the worry 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. More importantly, in both cases, the divergence is gradual. This is consistent with the fact that the introduction of sex selective abortion into Taiwan was largely brought by a legislative reform, whereas in the other two countries, technological adoption was organic and thereby arguably more gradual. See Figure 1 and Appendix Figures A1 and A2. The y-axes of the latter are scaled for comparison. Hence, we believe that our strategy is unlikely to be confounded by general regional trends.

Second is the concern that the reform induced parents with certain characteristics to give birth, or that the increased use of ultrasound during prenatal care during this time changed the characteristics of higher parity children born such that an improvement in relative female survival will reflect these factors rather than the decrease in the cost of sex selection. We will investigate whether this was the case by re-estimating equation (14) with characteristics of parents and births as dependent variables.

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.

In this study, we take the birth and death registries at face value. According to these data, over 97% of children are born in hospitals. And to the best of our knowledge, there have been no controversies surrounding misreporting in these data, systematic misreporting due to sex selection or any changes in reporting patterns due to the increase in sex-selective abortion.

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 have decreased for younger mothers, older mothers have had *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. (Matthews and Hamilton, 2005).[18] 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 throughout the 1980s. Interestingly, Figure 4C 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. Note that in examining the unconditional sample means, we cannot distinguish the effect of mothers' age from the effect of birth parity as the two are highly correlated. In the regression analysis, we will control for mother's age.

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 accounts 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

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

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.[19] 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 Fertility

To investigate whether abortion was used as a method of birth control as suggested by the descriptive statistics, we estimate the effect of legalizing abortion on the number of births. We do this by placing the natural logarithm of the number of births as the dependent variable in equation (13). The estimates with and without controls are shown in Table 4. Column (5) shows that legalizing abortion decreased the number of second and third parity births. The decrease is larger for the latter group. This supports the claim that abortion was used as a form of birth control. The estimates for the interaction of mother's age and post show that the legalization caused older mothers to have more children relative to younger mothers. In fact, the effect is monotonically increasing with mother's age. All the estimates are statistically significant at the 1% level.

[19] Source: World Development Indicators.

4.2 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 (12). The estimates for $\hat{\beta}_{2t}$ and $\hat{\beta}_{3t}$ and their robust standard errors are shown in Appendix Table A2 column (1). 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 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 same equation for mothers of different ages. The coefficients and their standard errors are reported in Appendix Table A2 columns (2)-(4). We plot the coefficients in Figures 6A-6C. Figures 6A and 6B show that abortion had the same effect, that is to say – no effect, on second and third parity births for young mothers. However, Figure 6C shows that for older mothers, abortion increases the fraction of boys for third and higher parity births. The y-axes of the figures are scaled for comparison. These figure shows that it is older mothers having their third or higher parity children who are using abortion to sex select after abortion is legalized.

To assess the statistical significance and the average effect of the reform, we estimate the simpler difference-in-difference equation (13). The estimates for $\hat{\beta}_2$, $\hat{\beta}_3$, $\hat{\alpha}_2$, $\hat{\alpha}_3$ and $\hat{\alpha}_4$ and their robust standard errors are shown in Table 5. We focus our discussion on the estimates in column (5). They show that legalizing abortion has no effect on the fraction of males born at lower parities. For third and higher parities, the reform increased the fraction of boys born by 0.7 percentage points. The estimate is statistically significant at the 1% significance level. There was no effect on younger mothers. But for older mothers, abortion seemed to have increased the fraction of males born; the magnitude of the coefficient shows as 0.4 percentage point increase, though the estimate is not statistically significant.

In these estimates, we use first born children as the reference group and exclude them from the regressions. Alternatively, we could exclude the birth year dummy variables and examine the effect of the fraction of males amongst first born children. The results from both the yearly estimate and the pre and post estimate show that the reform had no effect on the sex ratios at birth for first born children. For brevity, these results are not reported from the paper.[20] But they are evident from the plot of sex ratios at birth by parity in Figure 1.

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 (14) with the fraction of deaths occurring within one, two, three, four, five, six, nine and twelve months as dependent variables. For brevity, we only report estimates for deaths within one, three, six, and twelve months. Table 6 Panels A and B present the estimates not controlling and controlling for mother's age. We will focus our discussion on the latter.

Panel B shows that the coefficients are negative and often near zero in magnitude for second parity groups. In contrast, the estimated coefficients for third and higher parity births are consistently positive. The estimates for deaths within one and six months of life, shown in columns (1) and (3), are statistically significant at the 1% and 10% levels. Those in columns (2) and (4) are not significant. But they are very similar in magnitude as those in columns (1) and (3). These estimates show that legalizing abortion increased relative female survival rates by approximately 0.062-0.074 percentage points. The coefficients shown in columns (1)-(5) are relatively stable in magnitude, meaning that most of the deaths occur within the first month of life. The results show that overall, legalizing abortion, and presumably the increase in usage of ultrasound increased relative survival rates of male infants, who have often been observed to be more fragile. However, third and higher parity births, the group we know to for which abortion is being used to sex select, relative survival rates increased for female infants.

Note that the estimated coefficients for the interaction of mother's age and post are negative, often statistically significant, and typically increasing in magnitude with mother's age. And that the estimates of the interaction terms of birth parity and post are only stable across the number of months of life if we control for the interaction terms with mother's age (as we do in Panel B). This reflects the fact that having higher parity births is strongly correlated

[20]They are available upon request.

with mother's age, and that older mothers behave differently with respect to investment in the health of their sons relative to their daughters. We will explore this further in the following sections of the paper.

4.4 The Effect on Birth Composition

To examine whether the reform affected the composition of children born for higher parity boys differently than for higher parity girls, we estimate equation (14). The coefficients and their standard errors are reported in Table 7. First, we focus on the characteristics of parents. Columns (1)-(2) show that the reform induced more unmarried mothers and educated mothers to have more higher parity births sons. The estimates are statistically significant at the 1% level. The former may reflect the fact that unmarried mothers, like others, prefer sons to daughters. And the latter may reflect that more educated mothers are more willing to take up new technologies. The estimates show that there is little difference in the marital status across mothers' age groups after the reform. Interestingly, column (2) shows that older mothers who had sons after the reform were less educated. This is consistent with the estimates in column (3) which show that sons born to older mothers after the reform typically had less educated fathers.

Next, we investigate whether the reform had differential effects on the composition of male and female infants along biological characteristics. The outcomes we examine are the fraction of LBW and multiple births. LBW and multiple births are more fragile during pregnancy and infancy. Hence, if the increased use of ultrasound during prenatal care increased the number of LBW and multiple births for higher parity males more than for females, then the often observed fact that males are more fragile during infancy could cause us to find that the reform improved relative female survival. The estimates in columns (4) and (5) show that this does not appear to be the case. Abortion has no effect on the fraction of LBW or multiple births for higher parity births. The estimates are very small in magnitude and statistically insignificant. Hence, the increase in relative female survival for third and higher parity births cannot be caused by the changes in composition induced by the reform (to the extent that those changes are captured by LBW and multiple births).

Interestingly, the estimates in column (4) suggests that sons born to mothers over 35 after the reform are more likely to have LBW. The estimate is only statistically significant at the 15% level, but the coefficient is rather large in magnitude. We will discuss this in the section on interpretation.

4.5 Quantifying the Results

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 8. 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.).[21] 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 from the pre-reform rate the estimate for the effect of legalizing abortion on 3+ mortality (the coefficient for $bord3 \times post \times sex$ in column (1) of Table (6)), 0.0006. Hence, the post-reform mortality rate is 0.003-0.0006 = 0.0024. The fraction of parents who used to select post-natally 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, 73/757= 0.10. Conversely, if abortion was banned in the post period, then the fraction of parents who would switch from abortion to

[21] 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 (2008) 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.

post-natal 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, 73/1029=0.07.

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 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 higher parity boys relative to girls 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.

Mortality rates for third and higher-parity births for girls were on average 0.3 percentage-points before abortion was legalized (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 0.06 percentage-points, approximately a 20% reduction in neo-natal mortality. Like Ebenstein (2007), we find that almost all of the mortality occurs in the first month of life. There is very little additional mortality afterwards.

While it is beyond the scope of this paper to show the cause of death, we have strong priors that intentional female infanticide does not play an important role in this context. 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.[22] 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

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

level of differential neglect could have extreme results and be reflected in the mortality data.[23]

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

One of the interesting findings of this study is that older mothers respond very differently to the legalization of abortion relative to younger mothers. They are more likely to have children after abortion is legalized, and at the same time, there is less improvement in relative female mortality for children born to older mothers and they are more likely to use sex-selective abortion once it is legalized. (The estimates for the last effect are not statistically significant at conventional levels and therefore should be interpreted cautiously as suggestive evidence only. See Table 5 columns (4) and (5)). The results are consistent with the predictions we outlined in the conceptual framework in Section 2.1 and provides evidence for the insight that older mothers base their decision to have children on the availability of abortion as a method of sex-selection. This is consistent with the widely-held belief that the physical costs of pregnancy and child delivery increases with age. Access to sex-selective abortion reduces the number of months necessary for learning the sex of the child from nine to six months of pregnancy, effectively cutting out the most physically taxing and risky three months of pregnancy.

We can investigate this argument further. If abortion induces mothers who face higher health risks for themselves and for the child to be more willing to give births, then we may find that the children born to theses mothers are "weaker". Medical studies have consistently shown that women over the age of 35 are more likely to experience pregnancy-related complications such as gestational diabetes, high blood pressure, placental problems, premature birth and stillbirths (Cleary-Goldman, 2005; Joseph, 2005; Usta and Nassar, 2008; Bahtiyar, 2008). The crudeness of our data limits our health measure to birth weight, which has been linked to many of the conditions listed above and has been shown to adversely affect outcomes later in life.[24] Indeed, we find suggestive evidence that legalizing abortion causes older mothers to be more likely to have LBW babies (see Table 7 Column (4)). (The estimate is large in

[23]The *Nelson Textbook of Pediatrics* (2008) give a long list of conditions that can change within a short period of time.

[24]Birth weight is very commonly used as the outcome variable of interest in studies of the effects of policy interventions such as welfare reform, health insurance, and food stamps on infant welfare (see, for example, Currie and Gruber, 1996), and in analyses of the impact of maternal behavior on infant health (see, for example, Currie and Moretti, 2003). See Black, Devereux and Salvanes (2007) for a discussion on the literature on LBW.

magnitude but only significant at the 15% significance level). To the best of our knowledge, this is the first study that we know of to have ever noticed that the availability of abortions, by allowing parents to select the "quality" of children, may potentially induce certain parents to have *more* children. The heterogeneous effects of abortion would be an extremely interesting avenue of future research.

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. Taiwan during the 1980s was a society with both strong son-preference and a secular decline 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 greater 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 straightforward empirical strategy to provide evidence for the impact of sex-selective abortion on sex ratios at birth, EFM and fertility. The results show that legalizing abortion had little effect on sex ratios for parents who can reasonably expect to have more children (e.g. those with low birth parities). However, for parents who faced relatively more uncertainty in their ability to have more children, 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 20% relative to males. They show that up to 10% of parents who were selecting post-natally before the reform substituted to pre-natal sex selection using abortion. In other words, for every one hundred abortions, ten 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 161,000 more girls will die after birth each year.[25] The relatively low amount of substitution is not surprising considering how traumatic a death of a child can be to parents.

Interestingly, our results also show that the legalization of abortion does not decrease fertility for all women. Older mothers, who bear a higher cost of child rearing, actually have more children after the legislative reforms. The most obvious explanation for this is that sex selective abortion increased the option value of pregnancy, which induced older mothers with son preference to give birth.

[25] 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 53,100.

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 161,000 annually.

Our findings provide three important facts for policy makers. First, the results suggest that the enormous psychological cost of selecting post-natally will probably prevent the vast majority of parents from substituting from pre-natal to post-natal sex selection if the former is banned. (It is beyond the scope of this paper to estimate the welfare implications of banning sex-selective abortion more precisely. That will depend 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 well as the weights placed on the potential economic and social consequences from having unbalanced sex ratios). Second, for increasing relative female welfare, policies which restrict access to sex-selective abortion complement 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.[26] Finally, the introduction of abortion will have heterogeneous effects on fertility in contexts where there is son-preference and pre-natal sex detection technology is readily available. On the one hand, it will reduce the number of children born from unwanted pregnancies. On the other hand, a decrease in the costs of sex-selection from sex-selective abortion could induce women who bear a high cost of child rearing to have children.

[26] 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 by Parity over Time in Taiwan (1980-1998)

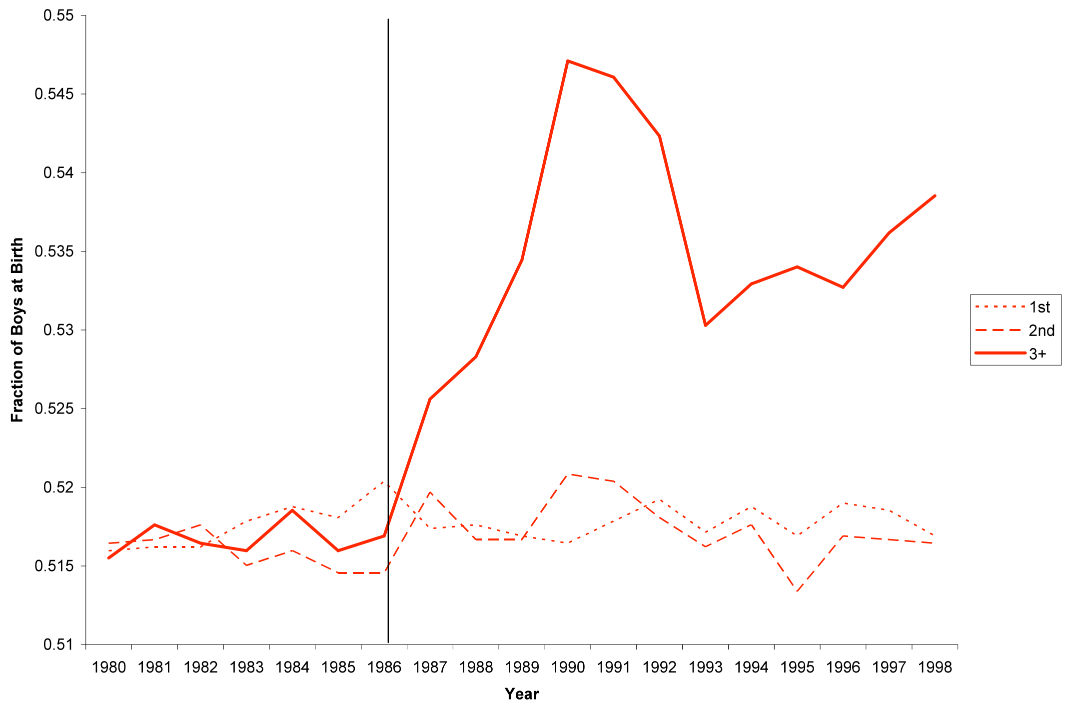

Figure 2A: Fraction of Women Reported to Have Ever Had an Abortion by Age

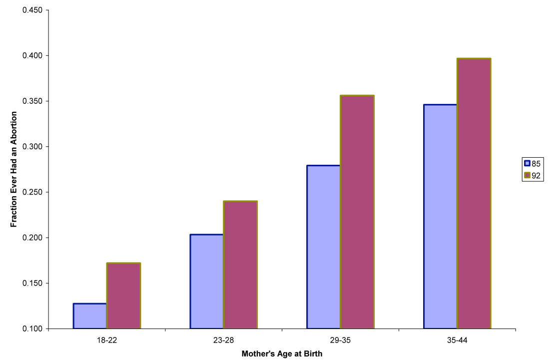

Figure 2C: Fraction of Women Reported to Have Ever Had an Abortion by Education

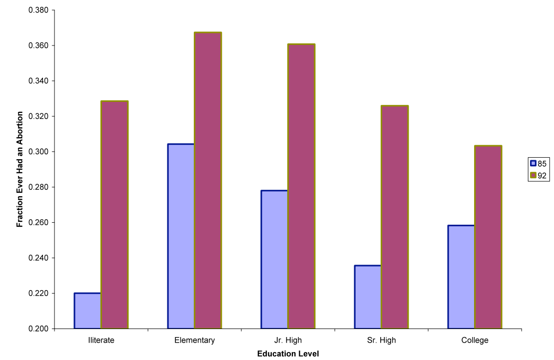

Figure 2B: Change in Fraction of Women who Ever Had an Abortion by Age

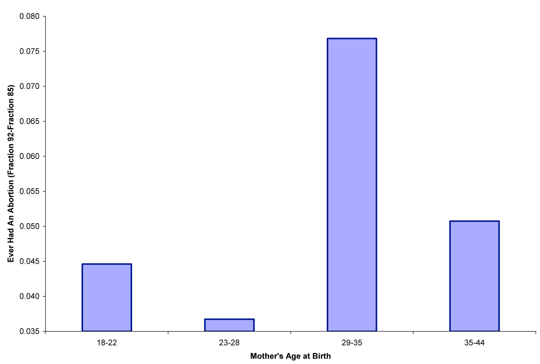

Figure 2D: Change in Fraction of Women who Ever Had an Abortion by Education Level

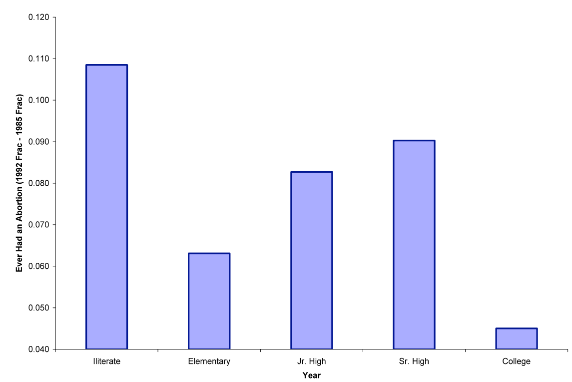

Figure 3A: The Fraction of Males by Birth Year and Birth Order

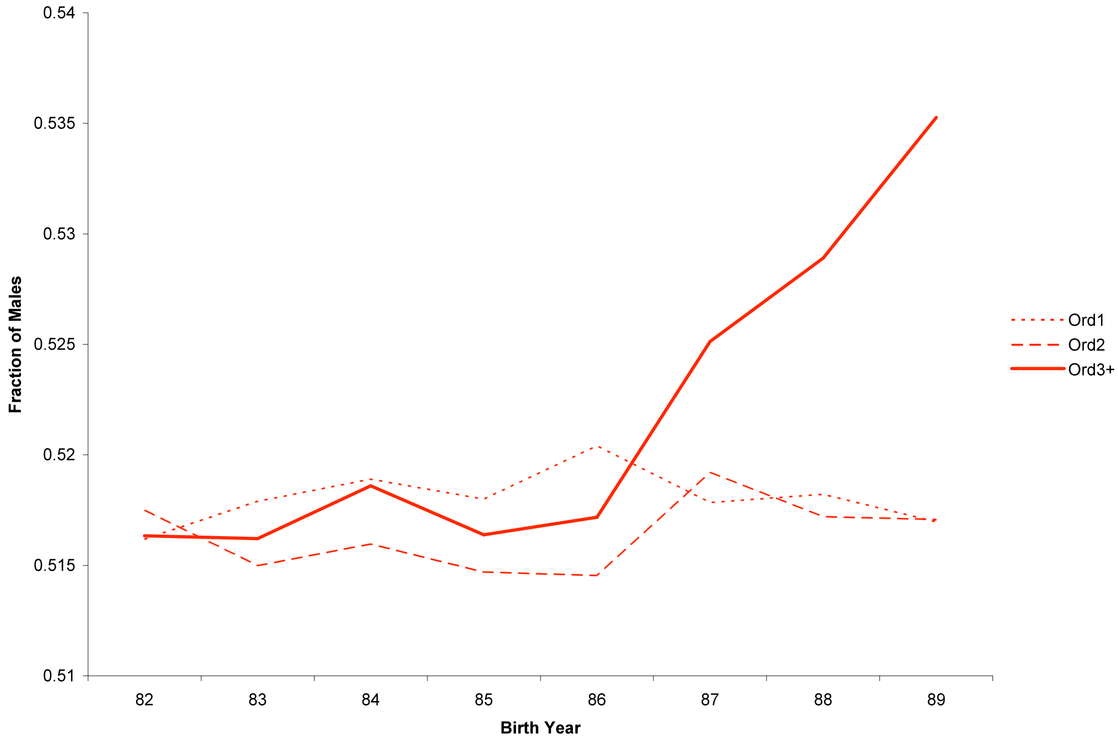

Figure 3B: The Fraction of Males by Birth Year and Mother's Age

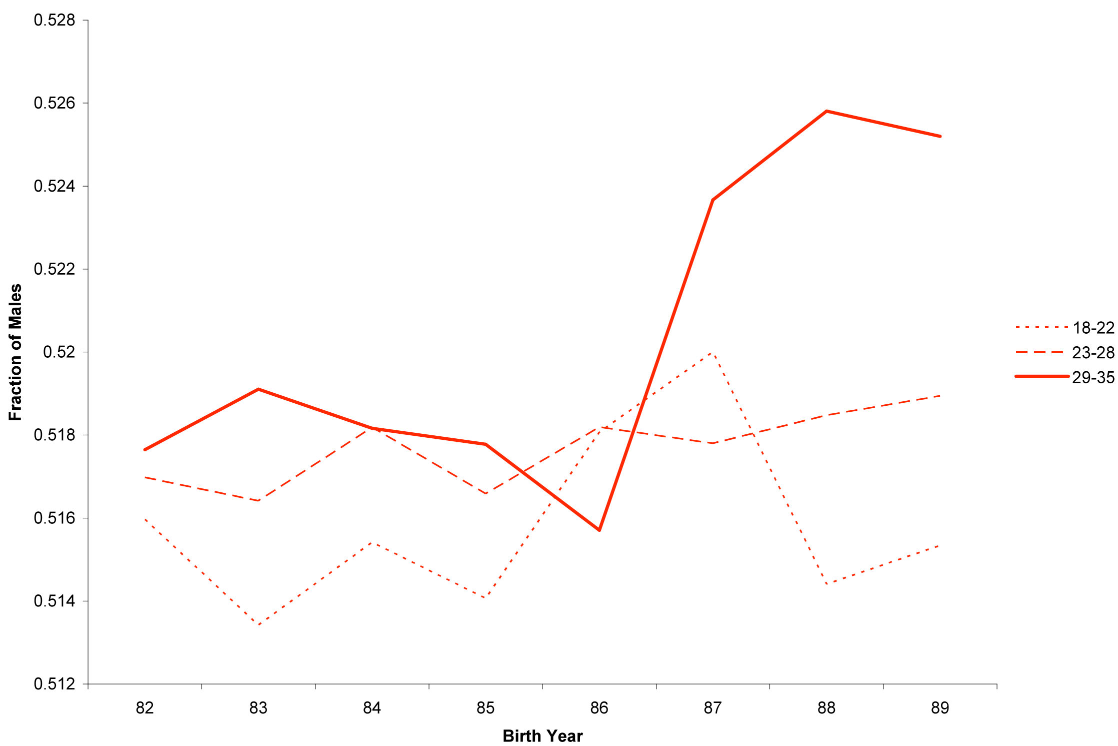

Figure 4A: Log (Total Births) by Birth Year and Birth Order

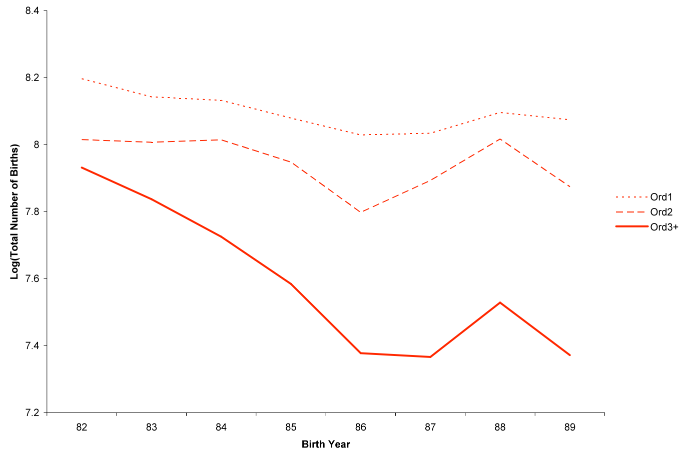

Figure 4B: Log (Total Births) by Birth Year and Mother's Age for Young Mothers (18-28)

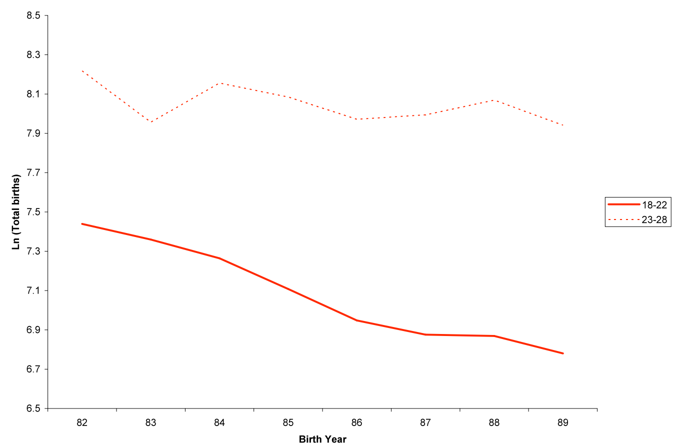

Figure 4C: Log (Total Births) by Birth Year and Mother's Age for Old Mothers (29+)

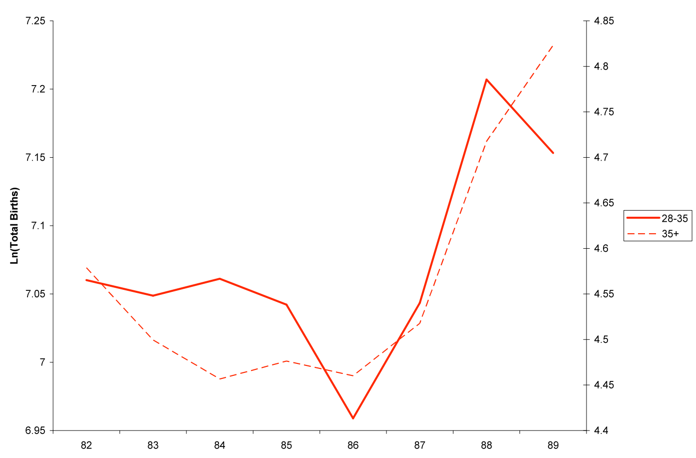

Figure 5A: The Effect of Abortion on Fraction of Males by Birth Order

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

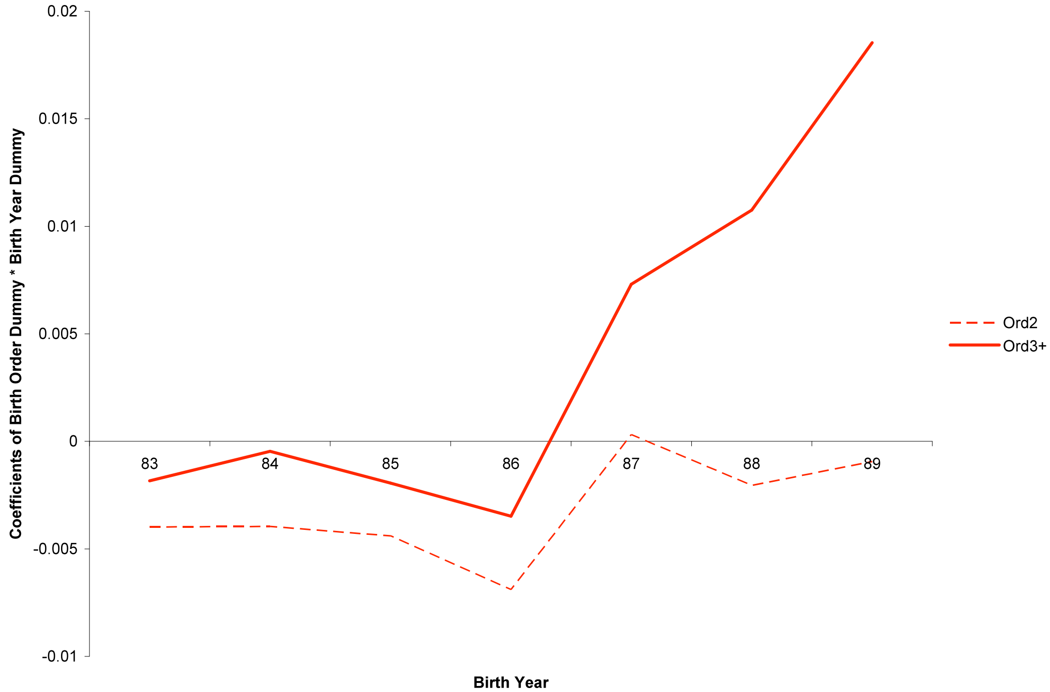

Figure 5B: The Effect of Abortion on Fraction of Males on 3+ Births

Coefficients of the interaction terms of birth year dummy variables and 3+ birth order dummy variable and their 95% Confidence Intervals

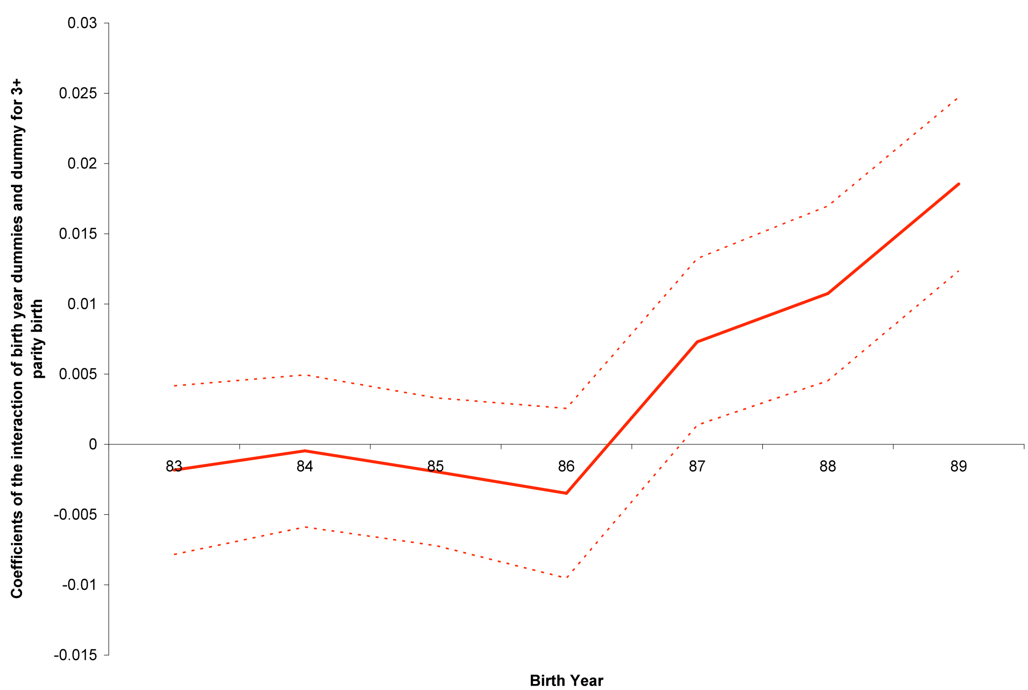

Figure 6A: The Effect of Abortion on Fraction of Males by Birth Order (Mothers 18-22)
Coefficients of the interaction terms of birth year dummy variables and birth order dummy variables

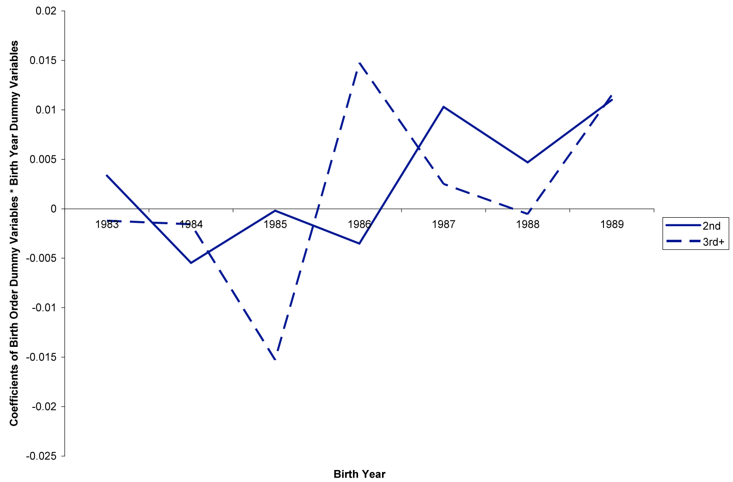

Figure 6B: The Effect of Abortion on Fraction of Males by Birth Order (Mothers 23-28)
Coefficients of the interaction terms of birth year dummy variables and birth order dummy variables

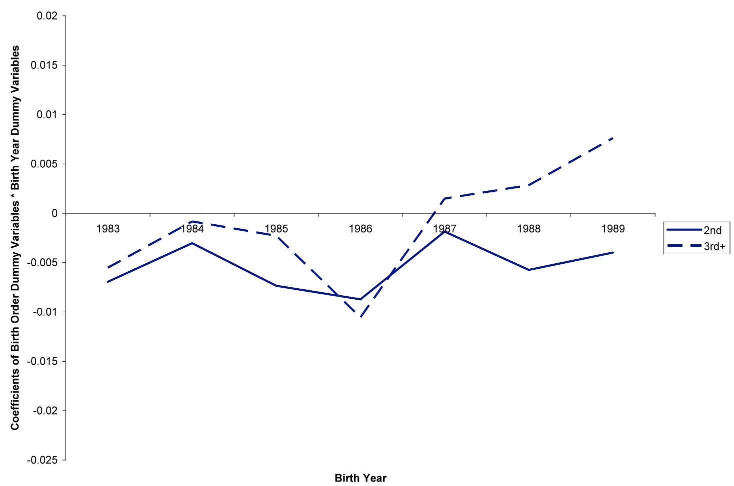

Figure 6C: The Effect of Abortion on Fraction of Males by Birth Order (Mothers 28+)
Coefficients of the interaction terms of birth year dummy variables and birth order dummy variables

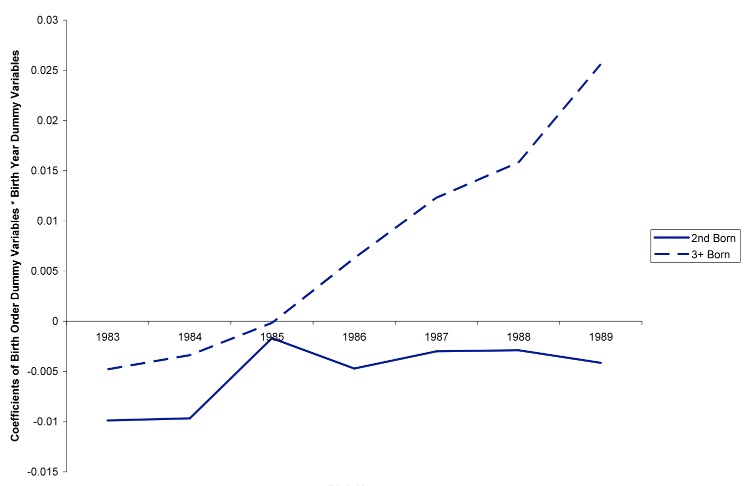

Table 1: The Change in Number of Births by Birth Order and Mother's Age

	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	Born 1982-84		Born 1985-89		Diff	Born 1982-84		Born 1985-89		Diff	DD: (6)-
X 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 4: The Effect of Abortion on Number of Births by Birth Order and/or by Mother's Age

	Dependent Variable: Ln(Number of Births)				
	(1)	(2)	(3)	(4)	(5)
2nd Born * Post	-0.00365	0.05272	-0.30231		-0.47547
	(0.18472)	(0.14688)	(0.11922)		(0.11823)
3rd+ Born * Post	-0.27445	-0.14960	-1.07289		-1.49228
	(0.16827)	(0.14252)	(0.10843)		(0.11682)
Mage 23-28 * Post				0.20874	0.50636
				(0.00181)	(0.10252)
Mage 29-35 * Post				0.31626	1.01467
				(0.00417)	(0.11297)
Mage 35+ * Post				0.50773	1.91630
				(0.00635)	(0.13727)
Birth Order Dummies	Y	Y	Y	N	Y
Mother's Age Dummies	N	Y	Y	Y	Y
Mother's Education	N	N	Y	N	Y
Observations	3159	3159	3159	3159	3159

All regressions control for birth year and birth county fixed effects.

Robust standard errors in parentheses

Table 5: 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)	(4)	(5)
2nd Born * Post	-0.00002	-0.00013	-0.00080		-0.00010
	(0.00142)	(0.00141)	(0.00145)		(0.00143)
3rd+ Born * Post	0.00692	0.00685	0.00511		0.00658
	(0.00166)	(0.00165)	(0.00171)		(0.00170)
Mage 23-28 * Post				-0.00084	-0.00173
				(0.00176)	(0.00177)
Mage 29-35 * Post				0.00188	-0.00025
				(0.00210)	(0.00212)
Mage 35+ * Post				0.00695	0.00389
				(0.00435)	(0.00438)
Birth Order Dummies	Y	Y	Y	N	Y
Mother's Age Dummies	N	Y	Y	Y	Y
Mother's Education	N	N	Y	N	Y
Observations	3159	3159	3159	3159	3159

All regressions control for birth year and birth county fixed effects.
Robust standard errors in parentheses.

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 Variable: Death within `X' Months			
	1 Month	3 Months	6 Months	12 Months
	(1)	(2)	(3)	(5)
A. Not Controlling for Mother's Age * Male * Post				
Born 2nd * Male * Post	0.00020	0.00014	0.00021	0.00051
	(0.00023)	(0.00029)	(0.00037)	(0.00041)
Born 3rd+ * Male * Post	0.00051	0.00024	-0.00032	-0.00005
	(0.00018)	(0.00031)	(0.00043)	(0.00053)
Observations	6258	6258	6258	6258
B. Controlling for Mother's Age * Male * Post				
Born 2nd * Male * Post	-0.00004	-0.00092	-0.00057	-0.00018
	(0.00018)	(0.00013)	(0.00024)	(0.00035)
Born 3rd+ * Male * Post	0.00062	0.00074	0.00060	0.00074
	(0.00028)	(0.00053)	(0.00037)	(0.00054)
Mage 23-28 * Male * Post	-0.00052	-0.00060	-0.00077	-0.00101
	(0.00025)	(0.00019)	(0.00024)	(0.00030)
Mage 29-35 * Male * Post	-0.00060	-0.00034	-0.00013	-0.00047
	(0.00031)	(0.00029)	(0.00032)	(0.00032)
Mage 35+ * Male * Post	-0.00215	-0.00229	-0.00284	-0.00255
	(0.00047)	(0.00072)	(0.00078)	(0.00086)
Observations	6258	6258	6258	6258

All regressions control for birth order, mother's age and birth year fixed effects.
Regressions in Panel B also control for the full set of interaction terms.
Robust standard errors are reported in the parenthesis.

Table 7: 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

	Dependent Variables				
	(1)	(2)	(3)	(4)	(5)
	Single Mother	Mother's Edu	Father's Edu	LBW	Multiple Birth
2nd Born * Post * Sex	0.0041	0.0054	-0.0067	0.0028	-0.0004
	(0.0016)	(0.0089)	(0.0066)	(0.0023)	(0.0012)
3rd+ Born * Post * Sex	0.0082	0.0456	0.0089	0.0000	-0.0021
	(0.0020)	(0.0124)	(0.0098)	(0.0038)	(0.0021)
Mage 23-28 * Post * Sex	0.0028	0.0042	-0.0031	0.0009	-0.0007
	(0.0017)	(0.0082)	(0.0073)	(0.0019)	(0.0009)
Mage 29-35 * Post *Sex	0.0008	0.0044	-0.0182	0.0001	-0.0002
	(0.0022)	(0.0099)	(0.0082)	(0.0026)	(0.0013)
Mage 35+ * Post *Sex	-0.0041	-0.0278	-0.0245	0.0108	-0.0015
	(0.0065)	(0.0153)	(0.0134)	(0.0064)	(0.0032)
Observations	6258	6258	6258	6255	6258

All regressions control for birth order, mother's age and birth year fixed effects.

Robust standard errors in parentheses.

Table 8: Calculating the % of Parents who Substitute from Post- to Pre-natal Sex Selection

	3rd+ Parity Births	1982-84 (1)	1985-89 (2)	Changes (2) - (1)
A.	Average Number of Total Annual Births	108162	73500	-34662
B.	Fraction of Boys	0.517	0.524	0.007
C.	Fraction of Girls = 1-B	0.483	0.476	-0.007
D.	# of Girls Born = A x C	52242	34986	-17256
E.	Fraction of Missing Girls= 0.49 - C	0.007	0.014	0.007
F.	# of Missing Girls = E x A	*757*	*1029*	272
G.	Girl's Mortality (Fraction)	0.003	0.0024	-0.0006
H.	# of Girls Dying = D x G	157	84	*-73*
I.	**Fraction of Parents Substituting from post- to pre-natal selection = 73/757**			**0.096**
J.	**Fraction of Parents who would Select Post-natally if Abortion was Banned =73/1029**			**0.071**

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 and 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 at Birth			
	(1)	(2)	(3)	(4)
	Full	Mage 18-22	Mage 23-28	Mage 29+
Second Born * Born 1983	-0.004	0.003	-0.007	-0.010
	(0.003)	(0.006)	(0.004)	(0.007)
Second Born * Born 1984	-0.004	-0.005	-0.003	-0.010
	(0.003)	(0.006)	(0.003)	(0.007)
Second Born * Born 1985	-0.004	0.000	-0.007	-0.002
	(0.003)	(0.006)	(0.004)	(0.007)
Second Born * Born 1986	-0.007	-0.004	-0.009	-0.005
	(0.003)	(0.007)	(0.004)	(0.007)
Second Born * Born 1987	0.000	0.010	-0.002	-0.003
	(0.003)	(0.007)	(0.004)	(0.007)
Second Born * Born 1988	-0.002	0.005	-0.006	-0.003
	(0.003)	(0.006)	(0.003)	(0.006)
Second Born * Born 1989	-0.001	0.011	-0.004	-0.004
	(0.003)	(0.006)	(0.004)	(0.006)
Third+ Born * Born 1983	-0.002	-0.001	-0.006	-0.005
	(0.003)	(0.008)	(0.004)	(0.007)
Third+ Born * Born 1984	-0.001	-0.002	-0.001	-0.003
	(0.003)	(0.010)	(0.003)	(0.007)
Third+ Born * Born 1985	-0.002	-0.015	-0.002	0.000
	(0.003)	(0.010)	(0.004)	(0.007)
Third+ Born * Born 1986	-0.004	0.015	-0.011	0.006
	(0.003)	(0.010)	(0.004)	(0.007)
Third+ Born * Born 1987	0.007	0.003	0.001	0.012
	(0.003)	(0.014)	(0.004)	(0.007)
Third+ Born * Born 1988	0.011	-0.001	0.003	0.016
	(0.003)	(0.011)	(0.004)	(0.007)
Third+ Born * Born 1989	0.018	0.012	0.008	0.026
	(0.003)	(0.012)	(0.004)	(0.007)
Observations	3159	792	794	1573

All regressions control for birth order fixed effects and birth year fixed effects.
Robust standard errors are reported in parentheses.

APPENDIX Figure A1: Fraction of Males by Birth Parity in China (1980-90)

Figure A2: Fraction of Males by Birth Parity in South Korea (1980-90)