Divorce, Abortion and Sex Ratio at Birth: The Effect of the Amended Divorce Law in China*

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Abstract

This paper explores whether and to what extent the relative circumstances of men and women following marital dissolution affect sex-selection behavior within marriages. China's new divorce law, which was enacted in 2001, reduced divorce costs, especially for women, by granting the right to divorce and claim damages in the cases of domestic violence and extra-marital relationships and by securing women's property rights upon divorce. We model the legal change as a decrease in women's divorce costs in a household in which the spouses have non- transferable utility, and all the marital surplus accrues to the husband. We show that the new law results in fewer sex-selective abortions for the second pregnancy if the first pregnancy produced a daughter, and that the sex ratio should decline the most in historically low-divorce-rate regions. Both predictions are consistent with the empirical evidence, and the spatial variations in the decline of the sex ratio helps rule out concomitant changes in household income and relative returns to male and female children. With a Difference-In-Difference approach, we also find that women with a later first pregnancy, which increases the health-related costs of performing sex-selective abortions, are more responsive to the changes in the divorce law. In addition, we perform an exercise of "timing regression discontinuity" to separate the effect of the divorce law from that of other policy changes; examine the outcomes of induced abortions and birth spacing to address the concern about underreporting of female births; and explore the changes in household consumption, which show a pattern consistent with the increase in women's position within households.

Keywords: Intra-household Allocation; Divorce Law; Missing Women.

JEL: D1, D13, J12, J13.

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1 Introduction

In addition to a substantial interest in the large number of "missing girls" in China, in the past decade, economists have focused on the fact that the primary source of this gap is the abnormally high sex ratio at birth (the ratio of male births to female births) among couples with one previous daughter in areas where China's family planning policy allows a second birth only if the first-born was a girl (Park and Cho, 1995; Gao, 1996; Poston et al., 1997; Zhang, 1998; Banister, 2004). The skewed sex ratio at birth among women with a firstborn daughter is higher than two boys to every girl. What is perhaps less well known is that there is also a higher level of divorce among women whose first child was a daughter, as shown in Table 1. Moreover, we find that the increase in divorce among these women occurs along with the decline in the sex ratio evident in Figure 1. (See Appendix A for a detailed explanation of the sex ratio.)

[Figure 1 is about here]

The changes in both the sex ratio and the divorce followed the enactment of China's new divorce law on April 30th, 2001. This law has two key components. First, it makes unilateral divorce possible in the cases of domestic violence or extra-marital relationships, and it allows the innocent party to claim damages from the guilty party. Both of these changes favor women more than men because, according to a 2002 Survey by the All-China Women's Federation, husbands commit 90 percent of domestic violence. Second, four new clauses protecting property rights upon divorce also favor women because women forfeit their property in most divorce cases in rural China (Platte 1988).

This paper investigates the hypothesis that the 2001 divorce law in China has improved women's utility upon divorce and their well being within marriage, and that these changes have led to fewer health-damaging sex-selective abortions (SSA) if the firstborn is a daughter. By exploring this hypothesis, this paper will contribute to the literature in at least two areas. One is how intra-household resource allocation, especially public-good provision, will be affected by the relative circumstances of men and women following marital dissolution. The other is whether and to what extent sex selection, which is the most salient manifestation of gender differentials in China and other Asian countries, can be decreased by policies that aim to improve women's well being in the short term.

From a household-theory standpoint, the relative circumstances of men and women upon divorce are often used as a reference point in models of household bargaining. However, there is substantial skepticism about the ability of changes in divorce laws or other legal protections to fundamentally alter the distribution of household resources in ways that will improve the well being of women and their children.[1]. Much of the research finds evidence of the divorce law's impact on household outcomes, such as labor supply (Gray, 1998; Stevenson, 2007), the well being of children (Gruber, 2004), household specialization (Stevenson, 2007) and domestic violence (Stevenson and Wolfers, 2007). However, some skepticism is still warranted in a traditional society with low levels of divorce, where women have little day-to-day interaction with the legal institutions. Under such circumstances, one might expect the new divorce law to have little or no impact on household behavior.

Moreover, even when changes in such laws shift a greater fraction of private resources towards women,

[1] While divorce may be the ultimate threat to both spouses and is a possible destination for marriages in which bargaining has failed, it is not the only possible "threat point" from which bargaining could proceed. Instead, other non-cooperative equilibria can be reasonable candidates for the "threat point," especially in areas with historically low divorce rates. For example, Tauchen, Witte, and Long (1991) suggest that family violence can be a non-cooperative equilibrium within a marriage. Lundberg and Pollack (1993, 1996), following the research suggestion of Wooley (1988), introduce a non-cooperative marriage in which the spouses receive some benefits due to joint consumption of public goods.

it is not clear whether the public-good provision in a household will be altered, especially the provision on specific aspects that may be of policy interest, such as the allocation of resources towards children. Mansor and Brown (1979, 1980) and Horney and McElroy (1981) provide a theoretical foundation of how divorce options affect intra-household allocation through "bargaining in the shadow of divorce". They suggest a Nash bargaining decision-making procedure. The uniqueness of the bargaining solution requires a convex feasible set. Therefore, in most of the standard Nash bargaining models, to make the models tractable, transferable utility is assumed, and, therefore, these models automatically predict that a public good should not be impacted by bargaining positions or divorce options.

However, recent empirical "reduced form" research has found that divorce laws alter the allocation of "public goods" such as the schooling of children (Gruber, 2004) and domestic violence (Stevenson and Wolfers, 2007); and that improving the relative economic standing of women raises the well being of female children (Duflo and Udry, 2005 and Qian, 2008) without explaining how women's bargaining power should impact household public-good provisions.

This paper provides a model that is a junction of non-transferable utility (NTU) and husbands' "dictatorship" decision-making to bridge this gap. Specifically, in this paper, by assuming non-transferable utility, we allow for asymmetric change in the marginal utility of private consumption between the husband and wife when efforts to produce the household public good—the sex composition of children—have been exerted. In particular, we notice the fact that since women bear the lion's share of sex-selective abortion—which hurts their physical and psychological health—it is plausible to think that, all else equal, the health-related cost of sex selective abortion lower their marginal utility of private consumption, and at the same time, the husband's remains intact. We show that even under husband dictatorship decision-making, increasing divorce options for women will decrease sex-selective abortions because it is more expensive for the "dictator" husband to compensate his wife to have an abortion.

Noticing that the prediction following transferable utility[2] is often rejected by empirical evidence, Chiappori, Iyigun, and Weiss (2007) explore a similar question from a different standpoint. They focus on showing the complexity of the divorce outcome when transferable utility assumption does not hold within marriage or upon divorce. Specifically, they show that the divorce outcome depends on the initial sharing rule, match quality and property division upon divorce. We follow Chiappori, Iyigun, and Weiss (2007) by assuming that TU is unlikely to hold in both marriage and divorce. Different from their model focusing on that the marginal rate of substitution (MRS) between public goods and private goods is different upon divorce from that within marriage, our model focuses on the fact that within marriage, certain public-good production behavior—sex-selective abortion, for example—will change the marginal utility of private consumption. In addition, we provide a complete model—which includes the utility-functional form of both the spouses' and the decision-making procedure—to solve the mechanism through which the changed "marriage exit option" leads to a redistribution of a public good. Finally, in this paper, we find a large decrease in sex-selective abortion, but a very small increase in divorce, this should not be regarded as evidence supporting the "Coase-Becker Theorem." Rather, assuming non-transferable utility function, that the wife' efforts of producing public goods depreciate marginal utility of private consumption and vary with marital status, the households on the margin of changing their efforts on public good provision within marriage are different from those on the margin upon divorce.

The second area of literature that this paper contributes to is that on gender disparity, especially

[2] Holding total family income fixed, transferring income from the husband to the wife should have no impact on household consumption except on those goods that are used to transfer utility.

regarding the skewed sex ratio at birth. The traditional literature in this field has focused on whether the sex ratio would be affected by policy changes in the short term and has tried to understand the mechanism of the effect. Much of the research on factors affecting the economic value to the household of women relative to men find that gender-specific infant mortality responds to economic change (Rosenzweig and Schultz, 1982; Qian, 2008). However, there is also reasonable skepticism about the existing empirical evidence: first, is the economic change sufficient to affect the outcome of infant mortality (Drèze and Sen, 1998; Foster and Rosenzweig, 2001), especially when the son preference is rooted deeply in the culture (Almond, Edlund and Milligan, 2009); and second, when the relative economic value of women changes, it is not clear which mechanism causes the mortality rate to change. It is hard to empirically distinguish among an increase in consumption on daughters, an increase in investment in female infants, and a shift in bargaining power toward women. For example, using shocks to the economic value to the household of women relative to men, it is difficult to establish the household bargaining interpretation and rule out the others, although there is some suggestive evidence from the former research (Qian 2008).

This paper exploring the effect of the pro-women divorce law on sex-selective abortions in rural China has several advantages over previous empirical studies. First, it links the female's position in a household to the children's sex composition by examining a much "cleaner" shock to women's welfare within the household: the divorce law. Previous empirical studies, such as Qian (2008) and Duflo (2009), use reforms that increased women's bargaining power and the relative returns of having daughters, as well as the total household income. This paper can rule out the increase in total income, while it cannot completely rule out the relative returns of having a daughter.[3] However, the change in relative returns itself cannot explain the finding that the divorce rate increased only among women whose only child is a daughter.[4] Second, this study improves upon past studies by considering the impact of fertility (through divorce) when examining the sex ratio. Finally, this paper is different from past research on sex-differential investment or "postnatal selection" in the sense that we explore the outcome of sex-screening and induced abortions. Most postnatal selection is more complicated—a consequence of differences in investment that aren't meant to produce differential mortality, yet do so as a result of the stochastic nature of mortality. Postnatal selection would also be very hard to contract for, especially if women are the primary caregivers. Therefore, investigating the change of behavior in sex-selective abortions—a discrete, predictable, and contractible decision—can help to rule out many competing explanations.

The main empirical difficulty of this paper is that the divorce law is universal, and there is no obvious research design that can rule out its effect through channels other than improving women's position in the household.[5] The two-fold goal of the model in this paper is to better understand the decline in the sex ratio caused by women's improved divorce options, and to look for possible comparative groups that would rule out concomitant changes in the relative returns to male and female children. In this model, the legal change is formalized as a decrease in women's divorce costs. Then, by imposing a very general assumption of single-peaked symmetry distribution of women's intrinsic divorce options, we demonstrate that the sex ratio at birth should decline the most in historically more-traditional regions with a low

[3]The relative return of having a daughter could increase after the change in the divorce law because if the daughter's property rights are more secure in her marriage, she might be more capable of helping her parents.

[4]If a daughter is more valuable to a household or, put differently, if spouses are able to derive more marriage surplus when they have a daughter, then the divorce rate should not increase.

[5]A related paper using similar data, Sun (2012), exploits the multiple dimensional regression discontinuity to examine the household outcomes along the boundary of eligible marriage age of men and women. The marriage is protected by law only if both spouses attain the eligible marriage age. If either spouse was below the eligible age when marrying, the government will void the marriage instead of granting a divorce. But Sun (2012) examines only abortions around the age cutoff.

divorce rate. Therefore, the comparison of the sex ratio between historically high and low divorce-rate regions allows us to drop the time trend and other concomitant changes. However, a caveat is that, different from a Difference-in-Difference design, the historical divorce rate should not be regarded as the intensity of the new divorce law.

Using the 25 percent sample of China's 2005 One-Percent Population Survey, we find that the fraction of male children declined from .69 to .64 among children conceived nine months after the implementation of the divorce law in provinces where a second birth is allowed when the first is a daughter, and the fraction decreased further afterwards. Furthermore, the sex ratio at birth declined the most in more-traditional regions with an historically low divorce rate. We also find that, after the implementation of the divorce law, the chance of having a son decreased by 19.7 percent when the mother's age at first pregnancy increased by one year among women getting married at a later age. This is consistent with our hypothesis that women with higher health-related costs of abortions are more responsive to the change in the divorce law if women's risk of pregnancy and SSA increases with age at pregnancy. Using the same data set, we find that the birth spacing between the first and second child in households in which both children are female decreased by 1.9 months after the divorce law, suggesting fewer sex-selective abortions within marriages. Using China Health and Nutrition Survey (CHNS) data, we find that the likelihood of induced abortion after having a firstborn daughter decreased by 6-11.8 percent after the divorce law. We also find evidence of change in the household consumption pattern: The husband's cigarette consumption decreased by -2.505 per day; the husband's frequency of drinking alcohol also decreased at ten-percent significance level. The wife's average daily protein intake increased by 2.61 grams and calorie intake increased by 88.694, while the husband's average protein and calorie intake showed no significant change. These patterns in consumption provide suggestive evidence that the divorce law improved women's well being in the household.

It's important to keep in mind that the divorce law might also affect the matching function of the marriage market. In this paper, we confine the sample to those who married before April 2001. Thus, the effect of the new divorce law on intra-household allocation will not be contaminated by the change in the marriage market (Chiappori 2007). However, the change in the matching pattern in the remarriage market might still have an effect on divorce and, therefore, change the sample composition. Although, as shown in the empirical section, given the low levels of divorce, it seems unlikely that divorces drive the pattern of sex ratio decrease, it is important to keep this caveat in mind.

The paper is organized as follows: In Section 2, we describe the policy's background. Next, we present the theoretical model and derive three predictions in Section 3. In Section 4, we describe the empirical strategy showing the effect of the divorce law under the model's predictions. The empirical results are presented and interpreted in Section 5. In Section 6, we discuss the results of the robustness checks. Finally, we offer concluding remarks.

2 Sex Selection and Divorce in China

2.1 The Family Planning Scheme, Son Preference and Sex-Selective Abortions

In China, son preference is rooted in both institutional (sons are primarily responsible for financially supporting their elderly parents) and cultural factors (only a son can carry the family name). The family- planning policy has allowed only one birth per urban resident since 1980. The majority (Han ethnicity) of rural residents are subject to the same "One-Child Policy" in six provinces.[6] In 19 other provinces, the "Girl Exception," or the so-called "1-Son-2-Child" policy, is applied, meaning that if the firstborn is a daughter, a couple can have another child, but two is the maximum number of births allowed. In the remaining five provinces, two births are generally allowed. To prevent people from using divorce to avoid the family-planning policy, remarried couples are qualified to have a child only when one party does not already have children and the other party does not have more than one.[7] Given the restrictive family-planning policy, to increase the chance of having a son, it is more feasible to perform certain sex selection than to increase fertility.

Prior research has found that sex selection occurs mainly through prenatal ultrasound B screening and induced abortions (Zeng, 1993; Li, 2005, 2007; Das Gupta et al., 2009). The sex-screening ultrasound B technology was first introduced in the early 1980s. After that, the sex ratio started to increase, and after the 1987 census, the central government realized the problem and issued an "emergency notice" to forbid ultrasound sex screening nationwide in 1989. The "Law on Maternal and Infant Health Care" restated the ban of SSA in 1994. However, because ultrasound machines are very mobile and inexpensive, the ban has been proved to be less restrictive by the increasing sex ratio at birth (SRB) in the following years. In addition, the performance of local government officials is evaluated only by whether they successfully curb the population, regardless of whether they keep a balanced sex ratio. Therefore, local governments have no incentive to ban sex screening in practice (Chu,2001).

SSA is seldom used for the first birth, but is very common for the second birth after a firstborn daughter. The sex of a fetus cannot be determined until the second trimester or later, and an abortion at that period of pregnancy risks women's reproductive capacity. Given the health-related cost of SSA and that in most provinces, a second child is allowed if the first is female, most couples prefer to have the first child regardless of the sex, knowing that they will have another chance if it is a girl (Li, 2007). Li, Wei and Feldman (2005) find that temporal trends and patterns of variation in induced abortion and SRB are fundamentally consistent. They also find that among women with only two children, the risk of having an induced abortion to end the next pregnancy is significantly higher for those whose first child is a daughter than for those whose first child is a son.

Behind the sex selection performance, there is no concrete evidence showing that women, in general, have a weaker son preference than their husbands. Very few surveys ask questions about both the husband's and the wife's "ideal sex composition" of children. The survey of Population and Reproductive Health in China includes this question, but only for women, and women may not want to reveal their true thoughts if they are not consistent with the values espoused by the government. Chu (2001) surveyed 820 rural women whose families were embedded in the same extended family networks of the author's, and the answers to the questions about attitude could be more reliable. Chu finds that "[W]hen asked if your first child is a girl and your second is a girl too, will you try a third birth for a son?', almost every woman said 'Yes, definitely.'" Chu also indicates in this research that women would risk their lives to have SSAs to win the respect of their families and communities. Li and Wu (2011) find that a firstborn son increases women's bargaining power. Keeping all this evidence in mind, it is reasonable to cast skepticism on the assumption that women have a daughter preference or at least a weaker son preference than their

[6] Beijing, Tianjin, Shanghai, Jiangsu, Chongqing and Sichuan.

[7] The restriction has been relaxed a little in a few provinces, such as Zhejiang. However, these changes happened after 2008.

husbands.

However, in the same survey, Chu finds that" [E]ven for women who had had SSA, a very large proportion believed it was not right or not fair to girls. 'But I have no choice. . . .' Some women resorted to SSA under pressure of the husband or parents-in-law'." This reveals the potential costs of SSA: Specifically, the cost of the SSA is the financial cost of performing the abortion and the health-related cost to the wife. The former is a minor concern because ultrasound B sex screening and induced-abortion technology have became cheaper and more efficient in past decades.[8] However, the cost to women's physical and, especially, psychological health is substantial. An abortion after the twelfth week of pregnancy is more dangerous and harmful for women than a first-trimester abortion. Second-trimester abortions require higher doses of misoprostol, a drug used to induce labor. Time to complete the abortion is longer and side effects are more common. In the 1997 survey of Population and Reproductive Health in China, 75.46 percent of women stated that abortion has a negative impact on their physical and psychological health, with 18.58 percent rating the impact as major. (11.13 percent of the women responded "no idea" and 13.33 of responded "no impact.") In this paper, we argue that the asymmetric health-related costs of abortion is the key determinant of the SSA decision.

2.2 Divorce Costs for Women and the New Divorce Law

China's divorce rate has been low historically: Only about three divorces per hundred marriages occur in rural China.[9] There are two main reasons behind the low divorce rate. First, the Chinese government has long considered divorce "a factor of social instabilities," and the People's courts, village committees and even the work units are required to mediate, even when the divorce is mutually agreed upon by the spouses. This condition changed greatly after the 1980s, especially after the Supreme People's Court, in December 1989, issued fourteen conditions under which a unilateral divorce could be granted.[10]

The second reason for the low divorce rate is that women generally have had very few divorce options. Marriage in old China was conceived as the transfer of a woman from the power of her family to that of her husband; therefore, women deserve no property when the relationship ends. In a case study by Liu and Chan (1999), the researchers find that women who would like to divorce, but who still endure their unhappy marriages, assumed that, were they to leave their marriages, they would face problems with housing and finances. Liu and Chan conclude that the reason these women do not consider divorce is their fear of having insufficient resources for independent living. Before the 2001 law, two marriage laws had existed since the establishment of the People's Republic of China (PRC). Although both laws upheld the principle of protection of women and children, they lacked specific regulations. This, together with the patriarchal norms, created a substantial cost of divorce for women. A divorced woman would not only suffer a drop-off of family income, but would also risk losing her housing and land, which the state generally supplied through her husband's family.

In the new divorce law passed on April 18, 2001 , there were two main revisions. First, for the first time, the law clearly established four grounds[11] for unilateral divorce, including domestic violence and extra-

[8] In December 2009, a doctor was caught providing ultrasound B sex determination to pregnant women in Danzhou city, Hainan province. He charged only 150 RMB (around 20 US dollars) for each ultrasound B sex screening.

[9] The age-specific divorce rate is calculated using 2000 Census and 2005 One-Percent Population Survey samples.

[10] However, the costs of divorce were still very high. For example, a reference letter from one's work unit or village committee is required on divorce. This requirement was abolished in October 2003 to protect privacy. But this policy change is more effective for urban residents. For rural residents, households in the same village are generally embedded in extended family networks and know about personal lives of other people anyway.

[11] The four new grounds are: domestic violence (Article 32 (3) (2)); extra-marital affairs (Article 32 (3) (1)); habitual

marital relationships. The innocent party could now also seek damages from the guilty party.[12] Second, the amended divorce law establishes stipulations regarding the right of visitation and circumstances regarding joint custody of children after divorce. In the system of property-ownership division, items regarding the management of property held in joint possession, the distinction between a common debt and individual debt, and the use of common or separate property to meet debts are all clearly stated. In addition, newly-added Article 47 improved enforcement.[13] The amended law also restates rural women's rights to land and housing upon divorce and emphasizes that one party is responsible to support the other in the case of difficulty of self-support. To summarize, the new divorce law can be plausibly considered pro-women in two main senses: 1) it emphasizes the conditions for unilateral divorce—relied upon mostly by women—so that women will have a greater chance of obtaining a divorce once they initiate one; and 2) it enables women to have a larger share of property upon divorce by allowing them to claim damages and by specifying property division.

It is important to clarify that the 2001 divorce law does not lower the general divorce costs for both spouses. This addresses the concern that if the new divorce law makes divorce easier for men as well, then husbands can simply walk away and start a new family when the first birth is a daughter or when the spouses find that the fetus is female during the second pregnancy. The combination of a few policies makes it difficult for men to simply abandon their wives. In addition to the new pro-women divorce law, the family-planning scheme imposes restrictions to prevent the abandonment of family members when the firstborn is a girl. First, if a remarried man already has a daughter from his previous marriage, his new wife must not have had any children before if the couple is to qualify to have their own child. In addition, the family-planning policy in some provinces (Hubei, etc.) states that a remarried man may not have more children if he divorced his ex-wife because she gave birth to a female child. Second, the 1980 Marriage Law states that "the husband should not initiate a divorce during the wife's pregnancy, the first year after delivery or the first six months after abortion, unless the wife also wants a divorce." Therefore, upon divorce, although it is possible for a man to obtain a great share of the family income, it is very difficult, if not impossible, for him to be exempted from the obligation of taking care of his daughter(s) and starting a new family.

Table 1 shows the fertility history of women in their first marriage by marital status in 2000 and in 2005, respectively. Although the Population Census does not question the timing of divorce, we can still use the codebook in Appendix C to infer the fertility history of women upon divorce in 97.5 percent of the cases. We confine the sample to rural women in their fertile years, between ages 20 and 35. In our sample, most rural women have at least one child when they are 35, and more than 75 percent have had two children before age 35. The statistics in Table 1 show that most divorces happen when the women have no children or only one child. Another notable point is that for those who are divorced, including currently divorced and remarried, the percentage having one daughter is slightly larger than that of having one son in the previous marriage. This difference is also statistically significant at the

gambling and drug abuse and other vices that remain incorrigible after frequent attempts at rehabilitation (Article 32 (3) (3)); and separation for two years as a result of failure to maintain a loving relationship (Article 32 (3) (4)). In addition, not only should divorce be granted in the above cases, but the innocent party can also further seek remedies in damages for domestic violence (Article 46 (3) and (4)) or from the party involved in bigamous and extra-marital relationships (Article 46 (1)).

[12] For the stipulation on damage compensation for domestic violence, see Article 46 (3) and (4), and for the damage compensation for bigamous and extra-marital relationships, see Article 46 (1).

[13] Article 47: "During the divorce proceedings, if one party attempts to conceal, transfer, sell or destroy community property, or falsify liability with the intension of misappropriating the other party's property, when property is partitioned in court, the court may award to the party at fault either a smaller amount of property or none at all."

one-percent level.[14]

3 The Model

The model incorporates the features of non-transferable utility and the husband's dictatorship. The key assumption is that the wife's effort to provide public goods lowers her marginal utility of private consumption, and that this drives the prediction that the wife will exert less effort when her reservation utility upon divorce increases. We present a specific example to illustrate the intuition and then discuss the potential generalizations.

3.1 Setup

Denote the husband's utility within marriage by u and the wife's utility within marriage by v;

$$\begin{aligned} u &= x_H + \eta \\ v &= \frac{x_W}{e} + \eta \end{aligned}$$

x_H and x_W are the private consumption of the husband and wife, respectively. The gender composition of the children is the only public good.[15] A representative couple places their values $\eta \in \{\sigma, \gamma\}$ on a son and a daughter, respectively. We focus only on the sex of the second birth, given that the first birth is a girl (SRB of the second birth after a firstborn daughter is abbreviated as SRB2 hereafter). Assume that a firstborn daughter brings utility σ. For simplicity, assume that a second daughter does not improve the spouses' utility. γ is the utility level of having a son. $\gamma > \sigma$,indicates the preference for sons. A second born son increases utility by $\gamma - \sigma$. e is the effort that the wife expends to increase the chance of having a son, and e is greater if the wife performs SSA. Namely, $e \in \{1, e_{NA}, e_A | e_A > e_{NA} > 1\}$. e_{NA} and e_A represent the efforts of fertility without and with SSA, respecitvely. Since $\frac{1}{e}$ enters the wife's utility function as a discount factor, $e = 1$ means not exerting any effort or not having a second child.[16]

We assume that the husband and wife have the same preference/value of sex composition of their children.[17] Another realistic simplifying assumption is that the spouses can detect the sex of the fetus

[14] Theoretically, it is not clear whether households with a firstborn daughter or a firstborn son are at greater risk for divorce. On the one hand, if both spouses have a strong son preference, a son means greater gain from the marriage. On the other hand, having a son could also change the reservation utility when the marriage ends. For example, when a woman has a son, she may have a better bargaining position and feel less subject to her husband's will (Li and Wu, 2011). One possible reason is that a woman can depend on her son to support her in her old age, which makes divorce less unaffordable. In rural China, given women's low divorce options in general, the improvement in women's position by having a son is unlikely to cause a divorce. This is consistent with differece of divorce between women with a firstborn son and those with a firstborn daughter, and with the finding of Li and Wu (2011) that women's well being increases within marriage (rather than higher propensity of marriage disruption) after a firstborn son.

[15] The main predictions hold in the case of multiple public goods.

[16] The feature of e is not necessary for the main predictions of the model. Put differently, we can reach the same predictions by simply assuming that rural Chinese women prefer to have daughters and men prefer to have sons. But we would keep the feature of asymmetric health-related costs for at least two reasons. First, as suggested in the literature reviewed in this paper and the statistical descriptions about the desired sex composition of children, the evidence shows that women have a son preference, as well. In addition, the statistical descriptions using the 1997 sample survey of Population and Reproductive Health in China suggest that women's health-related costs upon abortion are a likely driver of the potential controversy over SSA. In addition, a salient nature of prenatal selection is the asymmetric cost between the husband and wife. This feature of health-related costs has testable empirical implications, which will be discussed in this paper.

[17] Again, it does not matter much if the wife's son preference is weaker/stronger than that of the husband. In this model, the asymmetric health-related costs are the primary mechanism determining the husband's decision on SSA.

and perform SSA for free, which does not contradict the low cost of ultrasound B sex screening and induced abortions.

The budget constraint of a representative household is:

$$x_H + x_W \leq M$$

where M is the household income and is assumed to be exogenous once the match is formed. To exclude the case that husband initiates a divorce to marry a wealthier wife, we assume that the husband brings M to the marriage.

The couple chooses to split up if the utility of staying married is less than the utility of splitting up minus the divorce costs for either of the spouses:

$$\begin{aligned} u_d - c_H &> x_H + \eta \\ v_d - c_W &> \frac{x_W}{e} + \eta \end{aligned}$$

where c_i is the cost of divorce for spouse i, $i = \{H, W\}$. The direct effect of the divorce law is formalized as a decrease in c_W in this model. We impose an assumption on the parameters that $c_W > \sigma$.[18]

u_d and v_d describe the best options of the spouses upon divorce. Given the focus of this paper, we assume the husband as dictator, and do not specify the determination of v_d.We assume that (1) v_d is the combination of alimony, child support, property division, etc.,which is determined by the wife's intrinsic characteristics– i.e., fertility, remarriage options, etc.; and that (2) c_W reflects mainly the costs imposed by legal institutions, such as the costs of going through legal procedures, etc.. Assume that the husband does not recognize the "type" of the wife v_d until after they marry. To get a closed form of SRB2, assume that v_d follows a normal distribution. This assumption will be relaxed in the discussion of comparative statics.

$$f(v_d) = \frac{1}{\sqrt{2\pi}\sigma_v} e^{-\frac{(v_d - \mu_v)^2}{2}}$$

Despite the divorce costs c_H, a husband can still initiate a divorce if it is cheaper to improve his children's sex composition in a subsequent marriage. To focus on the decision-making within the current marriage, we do not formalize the expected utility derived from a second potential marriage. Instead, assume that this expected utility to be $Eu_0 = g(f(v_{d2}), \gamma, \sigma, e_A, e_{NA}\}$, which enters the model as a constant.[19] It is noteworthy that $g > 0$ only if the husband has one or no birth from the previous marriage. Otherwise, $g = 0$. Remember that the family-planning policy restricts the couple from having any child in their current marriage if either of the spouses has two or more children from a previous marriage.

[18] This assumption is not critical to derive the decline in SRB2. It is only to assure that c_W, which is the amount that the husband has to pay to the wife only upon divorce, is sizable enough compare to other parameters in the model, so that we can exclude some unlikely outcomes in the household optimization procedure. should be sizable enough compared to the gains from marriage. So that we can exclude some unlikely outcomes in household optimization procedure.

[19] Eu_0 is determined by the equilibrium of the remarriage market: the distribution of the wife's type in the remarriage market and the parameters γ, σ, e_A and e_{NA}. It is plausible to believe that the distribution of v_{d2},the type of the second wife, upon remarriage is different from that of v_d in the marriage market. The ceveate is that the divorce law might affect Eu_0 through affecting the remarriage market. But this simplification should not have a major impact on SRB in the population given the low divorce rate.

3.2 The Time Line

[Figure 2 is about here]

The timeline of the model is depicted in Figure 2, where N represents "nature" and H "husband", respectively.

The game begins when the spouses having a firstborn daughter. The dictator husband chooses to divorce or stay married.[20] If he chooses to stay married, they will have a second child,[21] and then the sex of the fetus is determined by nature. If the fetus turns out to be male, the couple will have a son and stay married. If it is female, the husband will again decide whether to divorce or stay married. Joint with this decision, he will also determine whether to have an SSA or to have a female child. If he chooses to have an SSA and stay married, then nature determines the sex of the fetus in the next pregnancy.

In Figure 2, we omit the player wife and the sub-game in the stage of consumption division, where the wife can choose to either accept or decline the husband's offer of consumption. If she declines, the marriage ends, the wife receives her reservation utility upon divorce, and the husband has to bear the divorce cost in addition to the compensation he must give her. Note that if she declines by initiating a divorce, there is a probability that the petition may not go through. We can regard v_d as an expected utility of the wife upon divorce, which is determined by the chance of being granted a divorce and the property division, and v_d follows a Normal distribution.

We assume that there is no unwanted pregnancy; the couple has complete information about the cost of abortion and fertility, and a woman can have only one abortion due to biological constraints.[22]

3.3 The Decision-Making

3.3.1 Husband Dictatorship

We first calculate the payoff for the husband and wife in each contingency. The dictator-husband model predicts $v^1 = v^2 = \cdots = v^7 = v_d - c_W$. We assume that the natural probabilities of bearing a male or a female fetus are equal.[23] The payoffs for the husband in all contingencies are calculated as follows:

- The spouses divorce after having the firstborn daughter; $u^1 = M - v_d + g - c_H$.
- The spouses stay married; the fetus is revealed as male; the wife has a male birth; u^2=$M - (v_d - c_W - \gamma)e_{NA} + \gamma$.
- The spouses stay married; the fetus is revealed as female; the spouses stay married and the wife has a female birth; u^5=$M - (v_d - c_W - \sigma)e_{NA} + \sigma$.
- The spouses stay married; the fetus is revealed as female; the wife has an SSA; the spouses stay married and the second child is a son; u^7=$M - (v_d - c_W - \gamma)e_A + \gamma$.

[20] This is a simplification of the scenario in which the divorce petition will be denied by the court with a probability θ.

[21] We do not consider the case in which a couple voluntarily ends fertility when the spouses are permitted to have another birth. According to the sample from the census and population survey, very few families have only one child by the time the firstborn girl is ten years old. The ratio of voluntarily choosing not to have another child might be even lower if some of these couples cannot have another child because of low fecundity. In addition, we do not find any evidence of a change in fertility after the implementation of the new divorce law.

[22] This assumption can be extended to the case of multiple abortions.

[23] There is research showing that the probability of bearing a male fetus is a little higher than 0.5 (around 0.53) (Oster 2005; Pongou, 2010;), but this will not affect the propositions derived from this reduced-form model.

- The spouses stay married; the fetus is revealed as female; the wife has an SSA; the spouses stay married and the second child is a daughter; $u^6 = M - (v_d - c_W - \sigma)e_A + \sigma$.
- The spouses stay married; the fetus is revealed as female; the spouses divorce and the wife has a female birth; $u^4 = M - (v_d - \sigma)e_{NA} + \sigma - c_H$.
- The spouses stay married; the fetus is revealed as female; the spouses divorce and the wife has an SSA; $u^3 = M - (v_d - \sigma)e_A + \sigma + g - c_H$.

We first rule out two outcomes. Divorce after the wife's second female birth is strictly dominated. It is easy to see that $u^5 > u^4$. This is consistent with the low proportion of divorced women having two children, as shown in Table 1. [24] We can also rule out the outcome of divorce after the wife has an SSA.[25] Then we use backward induction to solve the husband's strategy based on the wife's type v_d and the wife's divorce cost c_W.

We break the game into two stages, in stage one (after having a firstborn daughter but before the second pregnancy), the dictator husband determines whether to divorce the wife. If he stays married, in stage two (after the second pregnancy), he determines whether to initiate an SSA when the fetus in the second pregnancy is female. Solve the decision in stage two first. If the husband decides against the SSA, he gets $u^5 = M - (v_d - c_W - \sigma)e_{NA} + \sigma$; If he chooses the SSA, he gets $\frac{u^6+u^7}{2} = M - (v_d - c_W - \frac{\gamma+\sigma}{2})e_A + \frac{\gamma+\sigma}{2}$.Thus, if he stays married, his strategy is to initiate SSA if $\frac{u^6+u^7}{2} > u^5$, which is $v_d < \check{v} \equiv \frac{\frac{\gamma+\sigma}{2}e_A - \sigma e_{NA} + \frac{\gamma-\sigma}{2}}{e_A - e_{NA}} + c_W$, and distribute the wife x_W so that she will not initiate a divorce.

Now, consider stage one. If $v_d < \check{v}$, the husband's expected utility if he stays married is

$$Eu^A = M + \frac{3\gamma+\sigma}{4} + \frac{\gamma e_{NA}}{2} + \frac{(\gamma+\sigma)e_A}{4} + \frac{(e_{NA}+e_A)c_W}{2} - \frac{(e_{NA}+e_A)}{2}v_d \tag{1}$$

If $v_d > \check{v}$, the husband's expected utility if he stays married is

$$Eu^{NA} = M + \frac{\gamma+\sigma}{2} + \frac{(\gamma+\sigma)e_{NA}}{2} + e_{NA}c_W - e_{NA}v_d \tag{2}$$

If he divorces, he gets u^1, which is

$$Eu^d - c_H = M + g - c_H - v_d \tag{3}$$

It is clear that the lines $Eu^A(v_d)$ and $Eu^{NA}(v_d)$ will cross. $Eu^A(v_d)$ has a higher slope because it is more expensive to compensate the wife to $v_d - c_W$ after an SSA. The intercept of line $Eu^d - c_H$ is the smallest and the slope is the flattest of the three. Therefore, line $Eu^d - c_H$ crosses lines $Eu^A(v_d)$ and $Eu^{NA}(v_d)$.

Plot expressions (1)-(3) in the space of (v_d, Eu) are in Figure 3.

[Figure 3 is about here]

[24] In the 2000 sample, among women currently divorced, 9.18 percent have two children from the previous (first) marriage, compared with 59.86 percent having one child from the previous (first) marriage and 30.95 percent having no children from the previous marriage. A similar pattern appears in the sample of remarried women and in the sample from the 2005 Population Survey.

[25] If $\frac{u^6+u^7}{2} > u^3$,then the husband stays married. Otherwise, it is easy to show that $u^1 > \frac{u^2+u^3}{2}$ if $\frac{u^6+u^7}{2} < u^3$.*Proof*: $\frac{u^6+u^7}{2} < u^3 \Leftrightarrow \sigma + g - c_H \geq e_A(c_W + \frac{\gamma-\sigma}{2}) + \frac{\gamma+\sigma}{2}$.At the same time, $u^1 > \frac{u^2+u^3}{2} \Leftrightarrow \sigma + g - c_H \geq e_{NA}(\frac{c_W+\gamma}{2}) + \frac{\sigma(e_A+1)}{2} + \frac{\gamma+\sigma}{2}$. It is easy to see that $\frac{u^6+u^7}{2} < u^3$ follows $u^1 > \frac{u^2+u^3}{2}$ when $c_W > \sigma$.The intuition is that if divorce is a good enough option so that it is not dominated in stage 2, then divorce is likely already applied in stage 1.

The intersection of lines Eu^A and Eu^{NA}, is:

$$\check{v} = \frac{\frac{\gamma+\sigma}{2}e_A - \sigma e_{NA} + \frac{\gamma-\sigma}{2}}{e_A - e_{NA}} + c_W \tag{4}$$

Solving the intersection of lines Eu^{NA} and $Eu^d - c_H$, the cutoff between divorce and staying married is:

$$\hat{v} = \frac{\frac{\gamma+\sigma}{2}(e_{NA}+1) - g + c_H}{e_{NA}-1} + \frac{e_{NA}}{e_{NA}-1}c_W \tag{5}$$

In Figure 3, to enable two kinks along the UPF, we need to impose an extra necessary assumption on parameters so that the three lines satisfy the condition $\check{v} < \hat{v}$. Otherwise, there will be only one kink along the UPF and *Divorce* (D) will dominate *Non-Abortion* (NA). In Appendix A, using the 1997 Survey of Population and Reproductive Health in China, we show that 59.54 percent of the married women in the sample would never use SSA. Compared with the small number of divorces in Table 1, this supports the assumption $\check{v} < \hat{v}$.[26]

Figure 3 depicts the utility Pareto frontier (UPF) for each realization of v_d. The horizontal axis is the wife's reservation utility upon divorce; for each v_d, the feasible set of utility is composed of three points. The points on the top compose the utility Pareto frontier. Figure 3 clearly illustrates the critical role of the non-transferable utility function. If utility is transferable, then the utility frontier should be a straight line. By contrast, under the setup of this model, the line pivots under different choices corresponding to the efforts exerted by the wife. Now, consider three representative households (three different types of wives) as described at the bottom of Figure 3. The wife's reservation utility v_d is realized as v_d1, v_d2, v_d3,respectively. The husbands of household 1, 2 and 3 choose the action on the UPF.

Thus, the strategy of the dictator husband in equilibrium is described as follows:

$$s(v_d; c_W) = \begin{cases} A & \text{if } v_d < \check{v} \\ NA & \text{if } \check{v} < v_d < \hat{v} \\ D & \text{if } v_d > \hat{v} \end{cases}$$

,where $\check{v}$ and $\hat{v}$ are functions of v_d, and c_W is a parameter capturing the women's divorce costs imposed by the leagal environment. Namely, the husband chooses to *divorce* (D) the wife after having a firstborn daughter if her divorce option is very high ($v_d > \hat{v}$); otherwise, he stays married and produces another child with his wife. If the fetus is revealed as female, he initiates SSA only when the wife's divorce option is very low ($v_d < \check{v}$), and compensates the wife for the abortion and next birth (A). He does not initiate SSA (NA) if the wife's divorce option is above $\check{v}$ (but below $\hat{v}$).

We would like to emphasize that the decisions about SSA and divorce after having a firstborn daughter are determined by the wife's type v_d. Therefore, SSA when the fetus is female is determined once the household is formed.

[26]To focus on the discussion around the margin of $\check{v}$ and $\hat{v}$, we do not impose restrictions that $Eu >= 0$ or $v_d > 0$. Therefore, we avoid the complexity in the corner and we can also apply the property of continuous distribution when calculating SRB2.

3.4 Sex Ratio at Birth for the Second Birth (SRB2)

Under the assumption of v_d's density, the share of population taking each action from $\Theta = \{A,\ NA,\ D\}$ can be depicted as the areas below the density function in Figure 4.

[Figure 4 is about here]

We derive the propensity to have a second-born son after a firstborn daughter in the entire population with a firstborn daughter as:

$p = \frac{0.75\int_{-\infty}^{\check{v}} f(v_d)dv_d + 0.5\int_{\check{v}}^{\hat{v}} f(v_d)dv_d + 0.5\pi\int_{\hat{v}}^{+\infty} f(v_d)dv_d}{\int_{-\infty}^{\hat{v}} f(v_d)dv_d + \pi\int_{\hat{v}}^{+\infty} f(v_d)dv_d}$[27] ,where the number 0.75 represents for the probability of having a son for abortion couples, the probability for *non-abortion* (NA) couples is 0.5, and the probability that a divorced woman will have her second child in the new marriages is π.[28]

3.5 Comparative Statics

Denote the cutoffs in Figure 4, $\check{v} \equiv a + c_W$ and $\hat{v} \equiv b + \beta c_W$. It is clear that $\frac{\partial \check{v}}{\partial c_W} > 0$, and the decrease in c_W shifts the margin $\check{v}$ leftward, which means fewer SSAs after the new divorce law. We also have $\frac{\partial \hat{v}}{\partial c_W} > 0$, and the margin $\hat{v}$ also shifts leftward, which indicates that the propensity to divorce is higher after the implementation of the new divorce law.

Proposition 1 *The ratio of the divorced population* $1 - \int_{-\infty}^{\hat{v}} f(v_d|M)dv_d$ *should increase after the new pro-women divorce law.*

The propensity of divorce increases because the wife's efforts differ by marital status and the improved divorce options for women increase her utility within marriage.

Consider the effect of decreasing c_W on p. Rewrite p as

$$p = 0.5 + \frac{0.25\Phi(\frac{\check{v}-\mu_v}{\sigma_v})}{(1-\pi)\cdot\Phi(\frac{\hat{v}-\mu_v}{\sigma_v}) + \pi} \tag{6}$$

Holding $\hat{v}$ constant, on the one hand, shifting $\check{v}$ leftward will decrease p because while the fertility (denominator of p) is the same, the abortions in the whole population (the numerator) decrease. On the other hand, holding $\check{v}$ constant, shifting $\hat{v}$ leftward will increase p because the original NA women now switch to D, and are less likely to have a child (the probability of divorced women having a child in the new marriage is $\pi < 1$), so that the denominator gets smaller but the numerator remains unchanged.[29]

Case 1 *Very low divorces:* $\hat{v} \to \infty$

In this case, $p \to 0.5 + 0.25\Phi(\check{v} - \mu_v)$. $\frac{\partial \check{v}}{\partial c_W} > 0$ clearly predicts the decline in p.

Case 2 $\hat{v} < \infty$

[27] Sex ratio at birth is an increasing function of p, $SRB = \frac{p}{1-p}$. To derive the prediction directly testable in the empirical work, we use p instead of SRB in the following analysis. It is easy to prove that this monotonic transformation does not change the properties in the following.

[28] π is the probability that a divorced woman with a daughter marries a man with no children. π is determined by the remarriage market equilibrium. g and π are exogenous in this model, because, given the low divorce levels, on average, the decline in SRB is less likely driven by divorce and remarriage. We incorporate π only to indicate that the chance of having another child after divorce is smaller than 1 under the current family-planning policy.

[29] $\pi = .37$ in 0.25% sample of 2005 Population Census.

Rewrite $p = 0.5 + 0.25\frac{\Phi(\frac{a+c_W}{\sigma_v})}{(1-\pi)\cdot\Phi(\frac{b+c_W}{\sigma_v})+\pi}$. Taking the partial derivative of p with respect to c_W, we have:

$$\frac{\partial p}{\partial c_W} > 0 \tag{7}$$

if

$$\frac{\phi(\frac{a+c_W-\mu_v}{\sigma_v})}{\Phi(\frac{a+c_W-\mu_v}{\sigma_v})} > (1-\pi)\beta\frac{\phi(\frac{b+\beta c_W-\mu_v}{\sigma_v})}{(1-\pi)\cdot\Phi(\frac{b+\beta c_W-\mu_v}{\sigma_v})+\pi} \tag{8}$$

Though the sign of $\frac{\partial p}{\partial c_W}$ is ambiguous, inequality (8) clearly clarifies the condition under which p will decrease: If the initial divorce rate is low enough (or the density of divorce margin $\phi(\frac{b+\beta c_W-\mu_v}{\sigma_v})$ is small enough), compared to the abortion rate (or the density of abortion), the effect of fewer abortions will dominate and the sex ratio should decrease. Note that this intuition remains when the assumption of Normal distribution is relaxed to any single-peaked density function. This leads to the second proposition:

Proposition 2 *Under the assumption of single-peaked density of v_d, the sex ratio at birth should decrease toward the natural level after the new divorce law when the initial divorce rate is low enough, compared to the abortion rate.*

Before continuing, we should clarify that in this paper, the discussion on divorce rate is not necessary for the purpose of deriving the decline of SSA behavior per se, but there are two reasons why we discuss divorce in the model. First, since we aim to explain the decline in SRB2, which is determined not only by SSA, but also by fertility, we should carefully examine women who switch from *NA* to *D* because a second birth is less likely for them under the current family-planning policy. Second, although the very low divorce rate in rural China suggests that the shifting of the margin between *NA* and *D* should not have a major impact on SRB2 of the whole population, which is indicated by case 1, we still would like to explore the p in case 2 because there are significant regional variations in divorces (Zeng 1989). This [30] reveals information about the distribution of v_d across regions, and we want to take advantage of these regional variations to look for comparative groups. Imagine two places, for example—a traditional and a more modern society, with different initial divorce rates before the implementation of this new divorce law. Assume that the shape of the distributions is the same across regions. Therefore, the differing divorce rate captures the mean of women's intrinsic divorce options in that place. In Figure 5, the shadowed areas in the right tail of the two density functions depict the divorce rate in the two regions, respectively. Figure 5 illustrates the importance of the regional variations in the divorce rate: Although there are few women on the right tails in different places, the difference still reflects the differing means of the intrinsic divorce-option distributions. If the margin of SSA has a greater density, as shown in Figure 5, the differing means imply very different magnitude of the change in SRB2.

[Figure 5 is about here]

Under the assumption of single-peaked density function, the portions of *A* calculated in Appendix A, and *D* shown in Table 1, indicate that women who opt for an abortion and women who divorce are likely distributed around the left tail and the right tail, respectively, for both high- and low-divorce regions. Then, in Figure 5, if the new divorce law shifts the margins $\check{v}$ and $\hat{v}$ leftward and if the initial divorce rate for both places is low enough, we can see that the law's effect on decreasing SRB2 should be greater in places with initially low divorce rates because in these places, the density of $\check{v}$ (which determines the

[30] Also see Table A1.

effect of shifting $\check{v}$ on p) is greater, and the density of $\hat{v}$ (which determines the opposite effect on p) is small enough for both regions to drive a major difference. Formally, take the cross derivative of p with respect to c_W and μ_v. To save notations without losing the essence of the model, let $\sigma_v = 1$ and $\pi = 0$. Denote the ratio $\frac{\Phi(\check{v}-\mu_v)}{\Phi(\hat{v}-\mu_v)}$ as λ.We have,

$$\begin{aligned}\frac{\partial}{\partial \mu_v}\left(\frac{\partial \lambda}{\partial c_W}\right) &= \frac{1}{[\Phi(b+\beta c_W-\mu_v)]^2}\{-\phi'(a+c_W-\mu_v)\Phi(b+\beta c_W-\mu_v)+\beta\Phi(a+c_W-\mu_v)\phi'(b+\beta c_W-\mu_v) \\ &\quad +\frac{2\phi(b+c_W-\mu_v)}{\Phi^2(b+\beta c_W-\mu_v)\Phi(a+c_W-\mu_v)}[\frac{\phi(a+c_W-\mu_v)}{\Phi(a+c_W-\mu_v)}-\beta\frac{\phi(b+c_W-\mu_v)}{\Phi(b+c_W-\mu_v)}]\}. \end{aligned} \tag{9}$$

Since 59.54 percent of married women would never have an SSA according to the data, we will have $\check{v} < \mu_v$ and $\hat{v} > \mu_v$; then $\phi'(a + c_W - \mu_v) > 0$ and $\phi'(b + \beta c_W - \mu_v) < 0$. The first two terms are negative. Inequality (8) indicates that the last term is positive. But when the density of the divorce margin $\phi(\frac{b+\beta c_W-\mu_v}{\sigma_v})$ is very small, such that the last term is negligible compared to the first term, the cross-derivative is negative. Note that when $\phi(\frac{b+\beta c_W-\mu_v}{\sigma_v})$ is small, and since the continuous density is single-peaked, $\phi'(b+\beta c_W-\mu_v)$ should also be very small. The first term $-\phi'(a+c_W-\mu_v)\Phi(b+\beta c_W-\mu_v)$ is likely to dominate. The intuition can be extended to other single-peaked density functions; We should also impose the assumption of symmetry density to assure $\check{v} < \mu_v$. It follows in Proposition 3:

Proposition 3 *Under the assumption of single-peaked symmetry density of v_d and a very low divorce rate in general, places with comparatively low initial divorces (big μ_v) should experience a greater decline in the sex ratio of children, while places with relatively high initial divorces (small μ_v) should experience a smaller decline in the sex ratio of children.*

Proposition 3 directly applies to the empirical settings of the pro-women divorce law and SRB2. Note that Proposition 3 seems heavily dependent on the assumption of distribution, which could cast doubt on the empirical implication it conveys. Taking a closer look at the assumptions we made, only two are critical to deliver proposition 3: (1) $\check{v}$ should have greater density in low-divorce regions, and (2) the divorced population should be very small compared to the population that would have SSAs. The first component is to ensure that the shift on margin $\check{v}$ has a larger impact in reducing the occurrence of SSAs in low-divorce regions, and the second component ensures that the offsetting force through lower fertility of the original NA couples is very small, so that the shift on margin $\hat{v}$ should not revert the sign of whole cross derivative to positive.

These two requirements are satisfied by any single-peaked symmetry density. This assumption is also consistent with patterns derived with data from other sources. For example, Figure A1 shows a clear negative relationship between SRB and divorce across provinces, using the data from *China's Year Book*. This pattern is well explained by Figure 5.

Before continuing, we briefly discuss the generalization of this model. The two key assumptions are non-transferable utility function and the dictator-husband decision-making. The utility function in the specific case analyzed above can be generalized to any utility that satisfies $\frac{\partial}{\partial e}(\frac{\partial v}{\partial x_W}) < 0$. The second assumption seems extreme, but the underlying intuition is just that the increase in v_d should increase the marriage surpluses distributed to the wife. Nash bargaining with divorce as a reference point cannot be applied because the non-transferable utility results in a non-convex feasible set. However, the intuition of the model applies to other ways of surplus division that set up a monotone positive relationship between v_d and the wife's utility within marriage.

4 Data and Empirical Strategies

4.1 Data

The estimation of the new divorce law's effects on divorce propensity and SRB2 uses the 0.1-percent sample of the 2000 Population Census and the 25-percent sample of the 2005 One-Percent Population Survey. The 2000 Census and 2005 Survey contain questions on sex, date of birth, marital status, date of the first marriage, educational attainment, migration and relationship to the head of household. For women over age 15, the data contain information about fertility history—i.e., the number and sex of births and the number of surviving male and female children, respectively.

We also use China's Health and Nutrition Survey (CHNS) as an independent data source to estimate the effect of the amended divorce law on induced abortion and intra-household allocation of consumption. CHNS contains information on women's induced abortions, especially the timing of an induced abortion during the pregnancy, so that we can infer whether an abortion is triggered by the incentive of sex selection. CHNS also provides data on the consumption of cigarettes and liquor and the three-day average of different nutrients (calories, fat, protein, carbohydrates) for each household member.

4.2 Empirical Strategies

We first estimate the divorce law's effects on divorce propensity and SRB2 using the .1-percent samples of the 2000 Population Census and the 25-percent sample of the 2005 One-Percent Population Survey. Regarding SRB2, the variations in cohorts (by the timing of birth), calendar year, and region will all be examined. First, we show the increase in divorce and decrease in SRB2 after the implementation of the new divorce law across provinces. Second, with the guidance of the model, we show the spatial variations in the decline of SRB2 to rule out concomitant changes that affect all provinces in the same way. We further examine the outcome of birth spacing to find the similar pattern of regional variations. Third, using the approach of Difference-In-Difference, we show, using the reduced-form model, that the divorce law's effect is greater among women who were older at the time of their first pregnancy, which is used as a proxy for health-related costs of SSA. Fourth, we use a "timing Regression Discontinuity" design and an "event study" method to rule out the effect of the Law of Population and Family Planning. Finally, we provide consumption evidence that is consistent with the hypothesis of improving women's well being within households.

4.2.1 Divorce

Since we do not observe the exact timing of a divorce, to approximately estimate the change in the divorce propensity around 2001, we combine the samples from the 2000 Population Census and the 2005 Population Survey as two cross sections to compare the share of currently-divorced women in 2000 and 2005. Using the information in the Population Census and Survey, we can investigate if there is any change in the number of divorces among women who are in the same fertile-age window and have the same fertility history.[31] We focus on women who are between 25 and 35 and who married before April 2001 in "1-Son-2-Daughter" provinces. First, since the eligible marriage age is 20 for women, those married before age 20 may not be comparable to women married in their 20s; therefore, we restrict the sample to

[31] *China Statistical Year Book* provides the information of the number of new divorces on a yearly basis. But these data are aggregated on a provincial level.

women over age 25 (who were over 20 before 2001) in the 2005 Population Census sample. Second, our sample is confined to women below age 35 because in the 2005 sample, 90 percent of the mothers had had their first child before age 29 and their second child before age 35, and the divorces of interest should have happened before the second pregnancy. The descriptive statistics on women between ages 25 and 35 and married before the amended divorce law are listed in Table 2. Finally, we eliminate from our sample women who married after the new divorce law because the change in the law can also affect marriage patterns. But if we omit women married after April 2001, then we will leave out individuals only from the 2005 sample because all recorded marriages in the 2000 sample took place before April 2001. Thus, the women in the after-policy sample are older than those in the before-policy sample. Given the large sample size, we also drop women married after April 1996 in the 2000 census to balance the age structure between the 2000 and 2005 samples.

The empirical model to estimate the effect on divorce propensity is:

$$d_{ipt}= \alpha + post{\cdot}\beta + divorce_p \times post_t{\cdot}\delta + X_{ipt}\eta + Z_{pt}\rho + \lambda_p+\pi_l+\varepsilon_{ipt} \tag{10}$$

d_{ipt} is an indicator of whether woman i is currently divorced at the survey point t in province p. The dummy variable *post* indicates whether the observation is from the 2005 sample (after the divorce law). cdr_{p_2000} is the provincial crude divorce rate in 2000 (prior to the divorce law). X_{ipt} is a series of control variables, including women's education years, age at the time of survey, quadratic term of age, and first-marriage age; Z_{pt} is a vector of the provincial-level control variables, including GDP per capita, population, women's labor participation rate, and the sex ratio of adults (ages 20-40 at the survey point)[32]. λ_p is the province fixed effect. π_l is the cohort (birth year) fixed effect of women.

The coefficient of the dummy variable *post* captures the effect of the divorce law amendment and time trend. Under the assumption of single-peaked symmetry density of women's intrinsic divorce options, high-divorce regions should experience a greater increase in divorce. This spatial variation allows us to drop the time trend and other simultaneous changes if these factors are the same across provinces. Another two potential reference groups are households with firstborn sons in "1-Son-2-Child" provinces and households with one child, no matter if it is a daughter or a son, in "One-Child Policy" provinces. For these households, under the restrictions of the family planning policy, the wives are done with fertility and no further efforts will be exerted. Put another way, when facing a wife with a strong divorce option, the husband's trade-off between paying the costs of divorce to improve the children's sex composition through a new marriage and compensating his wife in the current marriage to have another child does not exist in these households. Therefore, we use women whose firstborn was a son in "1-Son-2-Child" provinces, and women with one child in "One-Child Policy" provinces as placebo reference groups. The story of household decision-making is more convincing if we do not see a significant change in the divorce rate within the placebo groups. In addition, if the time trend is same across groups with different firstborn genders, and if other shocks affect the divorce rate of both groups to the same extent, then these effects can be dropped out.

Finally, under the assumption of the model—i.e. the distribution is symmetric and single-peaked – δ

[32]The provincial sex ratio of adults is to measure the "tightness" of the marriage market and the likelihood of remarriage. The adult sex ratio is driven by many factors, such as the function of the marriage market, migration and so forth. Then, I use 1990 Population Census data to calculate the provincial sex ratio of those aged from ten to 30 as a proxy for the adult sex ratio for adults between 20 and 40 in year 2000, and I calculate the sex ratio of those aged from five to 25 in the sample of the 1990 Census to proxy the sex ratio of adults between ages 20 and 40 in the 2005 data.

should be positive.

We use the status "currently divorced" as the outcome variable. The alternative is "ever divorced." We choose the status of "currently divorced" in the benchmark model because the newly increased divorces might be diluted by pre-existing divorces. β and δ in the regressions using the two indicators essentially should capture the same thing unless some shocks between 2000 and 2005 precluded/facilitated remarriage or extended/shortened the time for divorced women to remarry.[33]

4.2.2 SRB2: the trend break and spatial variations

The empirical analysis of the SRB2 uses mainly the 25-percent sample of the 2005 One-Percent Population Survey. We first examine the change in SRB2 by cohorts. The basic strategy is to categorize children into different cohorts by the timing of conception, and we compare the SRB2 of cohorts conceived before and after the implementation of the new divorce law. Since we use only one cross- section, the sex ratio in the analysis is that of surviving children who completed residency registration in China's *Hukou* system. We shall discuss this approximation after presenting the benchmark empirical strategies.

We define the cohorts as all babies conceived in each six-month period following February 1997. The trend of the approximated average SRB2 is depicted in Figure 1, which shows that the decline of SRB2 occurred among cohorts conceived after January 2002. The sample is limited to the second birth, given that the first was a daughter, for the majority ethnic group (Han) in rural areas. Table 3 summarizes the descriptive statistics on parents' age at the time of the birth, the schooling years, and the marital status at the time of the survey.

Applying the same exercise as that depicted in Figure 1, we do the regression as follows:

$$male_{ipc} = \alpha + \Sigma_{l=1}^{15} 1\{c = l\} \cdot \beta_l + X_{ipc}\gamma + \lambda_p + \varepsilon_{ipc} \tag{11}$$

The dependent variable is an indicator of whether child i is male. X_{ipc} is the vector of control variables, including the parents' age when the mother gave birth to the child, age squared, the parents' education level, and whether the mother is an immigrant. λ_p is the province fixed effect. The sample is confined to intact families. The cohort conceived between February and July 1997 is the reference group.

In this section, we apply two methods to establish the linkage between the new divorce law and this decline in SRB2. First, we explore the spatial difference in the divorce law's effect. We find that the provinces with historically low divorce rates are most affected. The model in this paper not only makes sense of the salient spatial difference in the decline of SRB2, but also guides us to take advantage of this spatial difference in the empirical exercise to exclude the effect of time trend—if it is the same for all provinces— especially those affecting total income and the relative returns to having male and female children. Second, following the research of Chay and Greenstone (2005), we explore the timing "discontinuity" of the trend in SRB2 nine months after the implementation of the new divorce law. The very first cohort affected by the implementation can be detected by a series of regressions using an indicator of "cohort l or younger" as a proxy of the indicator of "exposed to implementation," where l represents one cohort among the 15 cohorts in our sample. When certain l is the very first cohort exposed to the new law, the correlation between the sex ratio and the proxy variable should be the strongest and,

[33] However, it is natural to imagine that ε_{ipt} follows $AR(1)$ with a negative time series correlation. For example, if there was a shock before 2000 causing most couples on the margin of divorce to split up, then there would be fewer divorces in the period between 2000 and 2005. In this case, δ could be negative. So as a robustness check, we replace the 2000 crude divorce rate with that of year 1996 to 1999. The coefficients are of similar magnitude and significance levels.

therefore, β_l has the largest magnitude of t-ratio. If the timing of conception within the cohort detected matches that of implementation of the new law, this should be a piece of strong evidence about the linkage between the divorce law and SRB2 decline. That is because this exercise does not need the assumption that places with different historical divorce rates have parallel trends of SRB2. In addition, this exercise can help us rule out the effect of other policies if conception of the first cohort exposed to the new law is earlier than the implementation of other policies.

We first compare the SRB2 in historically high- and low-divorce-cost regions for cohorts conceived before and after the enactment of the new divorce law. The regression function is:

$$male_{ipc}= \alpha + \Sigma_{l=1}^{15} 1\{c = l\} \times \operatorname{div}orce_p \cdot \omega_l + \Sigma_{l=1}^{15} 1\{c = l\} \cdot \beta_l + X_{ipc}\gamma + \lambda_p + \varepsilon_{ipc} \tag{12}$$

All the features in equation (12) are similar to those in equation (11), except that we include the interactions of the cohort dummy variables and the historical divorce rate in province p, which is the crude divorce rate in 2000. Because the effect of the divorce law should not be linear according to the model, as an alternative method, we categorize the sample into two groups by the median of the historical divorce rate, and then into four groups by the quantiles of the historical divorce rate. In this case, $divorce_p$ is a dummy variable indicating that a province is below the median divorce rate (or a vector of three variables indicating the quantiles of the historical divorce level).

Then, to detect the timing of the SRB2 "trend break," we do the regression as follows:

$$male_{ipc}= \alpha + 1\{c \geq l\} \cdot \beta_l + X_{ipc}\gamma + \lambda_p + \varepsilon_{ipc} \tag{13}$$

with all other features similar to those in equation (11); instead of including all cohort dummy variables, we only include one dummy variable: "cohort l or younger." We run the regression using a different cutoff l from the 15 cohorts defined in Figure 1.[34] If certain l is the first cohort exposed to the implementation, the proxy variable should have the strongest correlation with the outcome variable, and the magnitude of t-ratio of β_l in regression (13) using that l as a cutoff should peak.

4.2.3 Measurement error on SRB2 and direct evidence on abortions

In the empirical analysis on SRB2 using the 25-percent sample from the 2005 Population Survey, we have used the sex ratio of surviving children with residency registration as the approximation of SRB2. The underlying assumption is that there is no significant gender difference in the mortality rates for those 0-4 years old, or in the propensity to misreport births. Put another way, if there are other driving forces behind the skewed sex ratio, the decline of SRB2 in Figure 1 can be attributed to fewer SSAs only if the divorce law does not affect those forces, and other policies or the time trend in the same time window should not affect those forces.

Hull (1990) describes the trend of rising reported SRB in China and presents three possible explanations: female infanticide, prenatal sex identification followed by gender-specific induced abortion, and underreporting of births. Zeng (1993) rules out the likelihood of infanticide and uses the data from the 1990 census, the 1988 Two-per-Thousand Fertility and Contraception Survey, the 1987 One-Percent Population Survey, and the hospital records of a surveillance for birth defects in 29 provinces to indicate that

[34] We can define more cutoffs. For example, we can check whether beta l reaches the peak of magnitude every month instead of every six-month.

the skewed sex ratio at birth reflected in the census data is likely to be caused by prenatal sex-selective abortions and sex-differential underreporting.[35] Goodkind (2004) analyzes the 2000 Census data and finds that despite a total of nearly 37 million children missing from the census, sex differences in child underreporting were fairly minor.

Regarding infant mortality, the model in this paper requires only asymmetric (physical or psychological health-related) costs between the husband and wife. So, as long as we believe that women bear higher costs than men do in the case of either prenatal or postnatal sex selection, the model should predict a decline in the sex ratio of children. However, in general, there is a conceptual difference between prenatal and postnatal selection: Postnatal selection can be considered the outcome of insufficient investment in baby girls, while prenatal selection is a choice rather than an outcome or consequence of other intra-household allocations, especially the investment of time and nutrition in children of a specific sex. Therefore, the interpretation of the coefficients estimated in (11)-(13) would be different in the case of postnatal selection. In addition, in the case of postnatal selection by infanticide, neglect, or abandonment, it is hard to determine the effect of the change in the divorce law because the selection can happen at any time during infanthood, or even later.

As a "clearer" way to distinguish mortality and focus only on SRB, we use fertility data provided by the 2000 Census and 2005 Population Survey. Each census reports the fertility information of women between ages 15 and 50 if they gave birth in the 12 months prior to the survey. The reason why we prefer using the sex ratio of surviving children as the benchmark result is that the fertility information in the 2000 Census and 2005 Survey does not include the SRB2 for babies born between November 2000 and October 2004. Therefore, it is impossible to detect the timing of the SRB2 trend break. We use the fertility data mainly for the purpose of comparison with the results in the benchmark analysis. If the magnitude of the estimands is similar, the story of prenatal selection is more convincing. Thus, we use the fertility record from the repeated cross-sections by combining the 2000 and 2005 samples, and the regression function is:

$$male_{ipt} = \alpha + post \times divorce_p \cdot \beta + post_t \cdot \delta + X_{ipt}\gamma + \lambda_p + \varepsilon_{ipt} \tag{14}$$

The dependent variable is the indicator of whether the birth i is male in province p reported in survey year t; *post* is the indicator of whether the observation is from the 2005 (post implementation) sample. λ_p is the provincial fixed effect, and X_{ipt} is the vector of control variables, including the parents' age when the mother gave birth to the child, age squared, parents' education level, and whether the mother is an immigrant. We also use the births between November 1988 and October 1990 from the ten-percent sample from the 1990 Census as a placebo group to look for any "pre-trend" of SRB2.

The other concern is systematic underreporting of female births. Using fertility data cannot address this concern, especially when the fertility history and demographic information are collected in the same survey. Although it sounds unlikely that a potential policy or policies could cause nationwide, systematically less underreporting of female births in 2005, we will still carefully address this concern by directly examining SSA behavior. Specifically, we provide two pieces of empirical evidence. First, using a smaller sample from the *China Health and Nutrition Survey* (*CHNS*) data, which was administered by the

[35] Specifically, couples who have a high-order female birth will try to hide the birth from the authorities in one of the following ways: (1) giving the girl to someone for adoption or sending her to friends and relatives living elsewhere; (2) not reporting the birth of the girl, but reporting her as an immigrant at a later time; or (3) simply not reporting the birth of a girl at all, whether she lives with her own parents or other relatives.

Population Center at the University of North Carolina in 1991, 1993, 1997, 2000, 2004, 2006 and 2009, we look at the incidences of abortion, especially induced abortion. CHNS provides the pregnancy history (including abortions) of 1109 women in eight provinces, seven of which are girl-exception provinces.

$$abortion_{ipt}= \alpha + \Sigma_{\tau=1997}^{2009} 1\{t = \tau\} \cdot \beta_{\tau} + \tau \cdot \theta + X_{ipt}\gamma + \lambda_p + \varepsilon_{ipt} \quad (15)$$

The dependent variable is the indicator of the occurrence of an abortion for individual i in province p in year t. $1\{t = \tau\}$ is the indicator of whether the sample is drawn from wave τ. θ captures the linear trend of the calendar year[36]. X_{ipt} is the vector of control variables similar to equation (14), and λ_p is the provincial fixed effect. We omit the 1991 wave because of the common occurrence of forced abortions to achieve the goal of curbing population size in the early 90s in China (Hemminki et al.,2005). The occurrence of abortions in 1991 is also abnormally high in this dataset. We use the 1993 wave as the reference group. Note that the decline in abortions detected in the "simple difference regression" as in equation (15) may not be safely attributed to the divorce law, although the linear calendar-year trend is controlled for.[37] However, it is important to keep in mind that the goal of equation (15) is to show the decline in SSAs along with the decline in SRB2. Put another way, we show that the decline in SRB2 cannot be simply driven by any change in underreporting female births.

As another piece of parallel evidence, we examine whether the birth spacing gets closer for households that have two births, the first being a daughter. Using the sample of China's 2000 population census, Ebenstein (2010) finds that the birth spacing between two children is significantly longer if the second birth is a son. Porter (2010) uses a similar strategy to detect SSAs in India. We explore the same outcome of birth spacing using the sample from the One-Percent 2005 Survey. The regression functions are as follows:

$$Birth\_Spacing_{ipc} = \alpha + post_{ipc} \cdot \omega + \tau \cdot \theta + X_{ipc}\gamma + \lambda_p + \varepsilon_{ipc} \quad (16)$$

$$Birth\_Spacing_{ipc} = \alpha + post_{ipc} \times divorce_p \cdot \beta + post_{ipc} \cdot \omega + \tau \cdot \theta + X_{ipc}\gamma + \lambda_p + \varepsilon_{ipc} \quad (17)$$

In equation (16), the dependent variable is the birth spacing (months) between the first and second births for household i in province p in cohort c. θ captures the linear calendar-year trend. All other features are the same as those in equation (12). We expect to see ω negative, meaning that a lower occurrence of SSAs shortens the birth spacing between a firstborn girl and the second birth. If more women are on the margin of SSA in historically low-divorce-rate regions, we should see a stronger pattern in these regions. Therefore, in equation (17), β should be positive.

4.2.4 Health-related costs: A Difference-In-Difference design

We are motivated to find another comparison group in a reduced-form analysis for the following reasons. First, the regional variation in divorce rates allows for a comparison across provinces in order to drop out simultaneous changes in the relative returns of male and female children if these changes impact the high- and low-divorce-rate provinces in the same way. However, the skepticism about this assumption is well placed, in that these provinces could be heterogeneous in many other ways besides the distribution

[36] In equation (15) and all following equations that control the linear trend, higher order of polinomials are added as robustness checks.

[37] CHNS data contain only seven provinces under the "1-Son-2-Child" policy. The variation in the divorce rate is small and because of the small sample size, the exercise of timing discontinuity is not feasible.

of intrinsic divorce options for women. Second, Proposition 3, which leads to the spatial comparison, depends on the assumption that the density of women's intrinsic divorce options is symmetric and single-peaked. Although we argue that this assumption is very general and that the predictions of the model are consistent with many other patterns of sex ratio and divorce, the empirical evidence will be more convincing if the assumption on symmetric, single-peaked density can be further relaxed or removed. Finally, in this paper, we hypothesize that women's position in the household affects SSA decisions because women bear the lion's share of health-related costs of an abortion, and these costs depreciate women's marginal utility of private consumption. In line with our hypothesis, we provide direct evidence that women with higher health-related costs are more responsive to the divorce-law amendment.

The idea can be illustrated using Figure 4. When the health-related costs of abortion are minor, the UPF of abortion will be flatter and might dominate that of non-abortion if the psychological cost is not a major concern or if the desire to have a son outweighs the psychological cost. In this case, the couple will not respond to the change in the divorce law. However, if the health-related costs of abortion are sizable enough and hence the UPF of abortions has a much higher slope for pregnancies in older women and should intersect with the UPF of non-abortion, which is the case depicted in Figure 4. Thus, when women's relative divorce options increase, some households will change their decision from abortion to non-abortion.

This hypothesis implies a Difference-In-Difference design, which is described by the regression function as follows:

$$male_{ipc} = \alpha + Duration_{ipc} \times post_{ipc} \cdot \omega + post_{ipc} \cdot \beta + Duration_{ipc} \cdot \theta + X_{ipc}\gamma + \lambda_p + \varepsilon_{ipc} \tag{18}$$

In equation (18), we use the duration between marriage and first pregnancy as the proxy variable for health-related cost instead of using that between marriage and the second pregnancy because the latter is a consequence of SSA. Women's risk of pregnancy and SSA increases with age at pregnancy.[38] Therefore, it is plausible to expect that women who become pregnant at a later age will have a high cost of SSA. Controlling for the marriage age, the longer is the duration, the later age the woman becomes pregnant. ω has the interpretation of the "treatment effect on the treated" under the assumption that the trend of sex ratio of women with different health-related costs are parallel and that the span between marriage and first pregnancy as a proxy variable does not capture other factors related with the magnitude of the law's effect. The first assumption ensures the validity of the Difference-In-Difference design, and the second assumption ensures the validity of using the duration as the proxy variable for intensity. The two assumptions are plausible if this duration is "random" after controlling "enough" characteristics, such as the age at first marriage, etc..

The "randomness" is unlikely in that the pregnancy timing is the outcome of household optimization in most cases. However, taking a closer look at the third row of Table 2, panel A and B shows that the median woman's first pregnancy occurs only six or seven months after marriage. We infer that the timing of first-pregnancy is more likely determined by fecundity rather than by women's power to optimize their own life-long trajectory of fertility and labor-force participation, especially in rural China, where women's career development is not a major concern and couples start trying to have children immediately after getting married.

[38] Complications of pregnancy that increase with age include elevated blood pressure, gestational diabetes, premature labor and bleeding disorders such as placental abruption.

Notice that in Table 2, the density of the duration between marriage and first pregnancy has a long right tail. The determination on the "delayed" pregnancy for women in this long tail could be different from the remaining of the sample in many ways. As a robustness check, we drop out women in the quantile with the longest duration. $Duration_{ipc}$ in equation (18) can be an arguably random assignment only if θ does not have any significant impact on sex ratio.

Another important point is that in equation (18), the underlying assumption is that ω is the same across women with different ages at first marriage. However, the duration before getting pregnant is more likely to increase health-related costs for women who get married at a later age. Therefore, we catgorize the sample into two groups by whether the age at first marriage is above the age of 25, when women will be at their most fertile, and most likely to get pregnant. We expect that ω is statistically significant only among women who get married at a later age.

5 Empirical Results and Interpretations

5.1 Results on Divorce Propensity and SRB2

The estimates of equation (10) are shown in columns (1) and (2) of Table 4. They show that the ratio of currently-divorced women increased by 0.005 times of the 2000 crude divorce rate between 2000 and 2005 for women between ages 25 and 35 whose firstborn was a daughter. The crude divorce rate in 2000 was between 0.5 (Hainan) and 2.86 (Xinjiang) per hundred married women. Therefore, the ratio of currently-divorced women increased by 0.003 to 0.014. This result is comparable to the difference in the number currently divorced, using raw data. The nationwide proportion of currently-divorced women with one daughter increased by 0.009 between 2000 and 2005.

There is no significant change in the proportion of currently-divorced women in households with a firstborn son in both "1-Son-2-Child" places and "One-Child Policy" places. The results are shown in columns (3)-(4) and (7)-(8) of Table 4.

In column (5), the ratio of being currently divorced increased by .010 times of the 2000 crude divorce rate, but it is only significant on the margin, and the effect goes away put controlling for individual and provincial characteristics. The possible reason why this estimate in column (5) is significant is that in these provinces, around 30 percent of the households with firstborn daughters still have a second child. A more-detailed analysis of SRB in "One-Child Policy" areas is presented in Appendix B.

Appendix Table A2 shows the results from using the number of "ever being divorced" as the dependent variable. The sign of both the indicator of post-implementation and the interaction are less stable, suggesting that the measure of "ever being divorced" may include much "older" divorces, which dilutes the effect we want to detect.

The empirical results of equation (11) are shown in Table 5, column (1), which demonstrates that the trend in SRB2 in Figure 1 is statistically significant. Columns (1) and (2) show the results using the observations from the 24 "Girl Exception" or "Two-Children" provinces. In column (1), the decline of SRB2 starts to be statistically significant with the cohort conceived in February 2002; the possibility of being a second-born son after a firstborn daughter in a household decreased by 6-8 percent for cohorts conceived after February 2002 but before August 2003. For younger cohorts, this possibility dropped by 12-13 percent. The reason for the greater coefficients for younger cohorts may be that the divorce law was fully implemented for younger cohorts. However, the estimates could be contaminated by other shocks during the period—such as the "Law of Population and Family Planning Policy," which was implemented from late 2002 through mid-2003 (Shanghai implemented its Act in December 2003) across China's provinces. We will discuss the "Law of Population and Family Planning Policy" in detail in the robustness checks. For now, as a simple way to address the concern of concomitant changes, we omit the cohorts conceived after 2003, and the results are shown in column (2). The coefficients of cohort indicators are of similar magnitude to those in column (1). The change in the SRB2 can be calculated using the formula $\Delta SRB = \frac{1}{(1-p)^2}\Delta p$, where p is the probability of the second child being a son. There has been at least a six-percent decline in the probability of having a son since February 2002. The SRB2 before the implementation was 2.3-2.4 boys to one girl, and the ratio has declined to 1.56-1.6 boys to one girl since the implementation, holding other factors constant.

Then, we categorize the provinces into historically high- and low-divorce-rate regions using the median of the 2000 crude divorce rate. We show the coefficients and robust standard errors of cohort indicators using the sub-samples of low- and high-divorce-rate regions in Table 5, columns (3) and (4), respectively. The coefficients are consistent with the pattern in Figure 6 by averaging the raw data without controlling for any household characteristics or provincial fixed effect. In low-divorce-rate regions, the first drop-off of SRB2 happened in the cohort conceived in February 2002, and SRB2 further declined for younger cohorts. In high-divorce-rate regions, the decline is not very obvious from Figure 6.[39] However, in Table 5, column (4), the Wald test shows that the coefficients of dummy variables indicating cohorts conceived after the implementation are jointly significantly different from zero at the one-percent level.

[Figure 6 and Figure 7 are about here]

We then investigate the spatial difference using the interaction term(s) as shown in equation (12). Before showing the results, we must be very cautious in explaining the ω_l in equation (12). In the usual "Difference-In-Difference" design, in which the interactions are generally those of the program intensity and the indicator of implementation, and, thus, the coefficients are interpreted as the "treatment effect on the treated." However, in this paper, we assume universal implementation, or no difference in intensity. The prediction about the difference in the magnitude of the effect on SRB2 is driven by the initial distribution of women's intrinsic divorce options rather than by intensity of the implementation. Keeping this in mind, when reading Table 6, we would put more emphasis on the pattern that the SRB2 declines more in historically low-divorce rate regions and less on pinning down the meaning of the exact magnitude of the differences.

We first use the crude divorce rate in 2000 as a measure of historical divorce. Note that the effect of divorce costs on SRB2 could be non-linear. Therefore, we also use discrete measures such as the indicator of whether a province is "below the median crude divorce rate" and three dummy variables to indicate whether a province belongs to the first, second or third quantile, according to the 2000 crude divorce rate, respectively. Figures 6 and Figure 7 show that the provinces with traditionally low divorce rates have historically high SRB2, which is consistent with the pattern in Appendix Figure A1, using China's Year Book data. If SRB2 in low-divorce-rate provinces declines most, as Proposition 3 predicts, we should see SRB2 converge across provinces with different historical divorce rates. This convergence is clearly shown in Figures 6 and 7. Moreover, in Figure 6, there seems to be no "pre-trend" of SRB2 for both high- and low-divorce-rate provinces, which gives us more confidence that the change in the spatial difference in

[39] There are two spikes in the trend (the cohort conceived between August 2001 and January 2002 and the cohort conceived between February 2003 and July 2003). The spikes are driven mainly by the three northeastern provinces: Heilongjiang, Jilin and Liaoning.

SRB2 is driven by the divorce law.

The estimates of equation (12) are shown in Table 6. Column (1) shows that starting with the cohort conceived in August 2001, the probability of having a son between the households in traditionally high- and the households in traditionally low- divorce-rate regions shrinks significantly since the cohort conceived in August 2001. Column (2) compares the first quantile with the third and fourth quantiles; column (3) compares the second with the third and fourth quantiles, and both regressions find that the gap shrinks starting with the cohort conceived in August 2001. Column (4) shows a similar pattern with the linear measure. The results in columns (1)-(3) show that the decline in the sex ratio is driven mainly by provinces with a below-median divorce rate. Note that the SRB2 gap shrinks beginning with the cohort conceived in August 2001 according to the results in Table 6. However, from Figure 6, the decrease in the SRB2 gap for that cohort is more likely to be driven by the spike in high-divorce regions.

To detect the first cohort exposed to the divorce-law amendment, we present the results of the exercise described by equation (13) in Figure 8.

[Figure 8 is about here]

From Figure 8, we see that the magnitude of the t-ratio peaks at the cohorts conceived after January 2002 for low-divorce-rate provinces. Moreover, the "concave" shape of the t-ratio using different "potential" cutoffs indicates that the cohort conceived in February 2002 is the first one exposed to the effect of the new divorce law. The t-ratios using different potential cutoffs in high-divorce-rate regions do not have a clear pattern. However, the magnitude of the t-ratio starts to be consistently greater than 1.65 with the cohort conceived in February 2002.

Then, using repeated cross-sectional fertility data from the samples of the 2000 Population Census and 2005 Population Survey, we show the results of spatial difference in the decline of SRB2 in Table 7. Although the fertility data covers only the babies conceived from February 1999 to January 2000 and those conceived from [February 2004 to January 2005, it is still a useful check to see if the estimands are of similar magnitude to those using the sex ratio of surviving children with residency registration in the benchmark empirical work. If so, we will be more confident that the sex ratio among surviving children with residency registration is a good approximation for SRB.

Regarding the quantile with the lowest divorce rate, the SRB of the babies born from November 2004 to October 2005 (conceived from February 2004 to January 2005) declined by 14.7 percent, compared to those born between November 1999 and October 2000 (conceived from February 1999 to January 2000). The magnitudes of the coefficients are comparable to those in Table 6. The coefficient of the interaction of the crude divorce rate in 2000 and the indicator of implementation is -.026, which is smaller than those in Tables 5 and 6. The reason is that the second quantile does not experience a statistically significant decline in the SRB2 in contrast to the results of the sex ratio of surviving children. This is consistent with the pattern shown in Figure 7, where the cohort conceived from February 2004 to January 2005 in the second-lowest-divorce-rate quartile has a higher sex ratio than previous cohorts conceived after the new divorce law, although the reason behind this spike is unclear. We also use the fertility data in the 1990 Population Census to look at those born between November 1988 and October 1990 as a placebo group; the results are shown in columns (3) and (4). There is no significant divorce-related gradient in the change in SRB2 in either of the columns, which indicates again that the pattern of low-divorce-rate places experiencing a greater decline in SRB2 is not caused by any "pre-trend" before the implementation of the divorce law.

Finally, we present the evidence on abortions. The results of equation (15) are shown in Table 8. For each pregnancy, CHNS collects the information of how the pregnancy ended—by a spontaneous abortion, an induced abortion, a stillborn fetus, a stillborn full-term baby, or a live birth. Because the respondents might have an incentive to report an induced abortion as a spontaneous abortion since SSA is illegal, we look at all abortions and induced abortions, respectively. The abortions that we are interested in are those performed after having a firstborn daughter. Therefore, we should confine the abortion sample using the fertility history. However, only around one half of the women in the pregnancy-history data set can be matched to the data set of fertility history. Thus, for each dependent variable, we show the results using the whole sample, to increase the power, and the sample confined to households with firstborn daughters using fertility history information.

In columns (1) (3) (5) and (7), we include only the indicators of waves to show the time variation of abortions. In (2) (4) (6) (8), we define a dummy variable post=1 if the wave is post- implementation; post=0 otherwise, and we control the linear trend in these columns. In doing so, we hope to detect the "mean shift" or "trend break" after the divorce-law amendment. The coefficients of indicators for 2004 and later have negative signs in columns (1) and (3), though not statistically significant mainly due to the measurement error from including spontaneous abortions. The coefficients of indicators after 2000 are negatively significant at the five-percent level in column (5), but not significant due to a smaller sample size in column (7). To control the linear trend of abortion using a small sample, we aggregate the wave indicators into one dummy variable "post." The coefficients of "post" are on the margin of the ten-percent significance level in columns (4) and (6). After controlling for the linear trend of abortions, the magnitude of the dummy variable coefficients is greater for the sample of those with firstborn daughters. Using the estimand in column (6) as a lower bound, the propensity toward SSA in households with firstborn daughters decreases by 6-11 percent. All the evidence in Table 8 is consistent with the hypothetical link between the new divorce law and fewer induced abortions, and it rules out the possibility that the pattern we find using the Population Census/Survey sample is driven only by a change in the underreporting of female births.

Another piece of parallel evidence is shown in Table 9. First, in columns (1) and (2), we confine the 2005 Population Survey sample to households in which the spacing between the first- and second-born child is less than 72 months. The birth spacing of 24 and 48 months has the highest density in the 2005 Population Survey sample, and more than 75 percent of households have birth spacing no longer than 72 months. We exclude the quantile with longest birth spacing because the decrease will be very hard to detect, considering the big "noise" during the long time window. The density peaks at 48 months because in most places, the family-planning policy also stipulates a minimum of 48 months' birth spacing if the mother is under age 28 when giving birth to the second child. Therefore, in Table 9, columns (3) and (4), we confine the sample to households in which the birth spacing is between 48 and 72 months, or in which the birth spacing is less than 48 months but the mother is older than 28 when giving birth to the second child. Within this smaller sample, birth spacing is more likely to be the outcome of household optimization rather than policy restrictions. After controlling for linear time trend, the birth spacing between first- and second-born children in households with firstborn daughters decreases by 1.6-2 months. The coefficients are significant at the one-percent level in both columns (2) and (4).

Finally, we show the result of the reduced form as equation (18) in Table 10. In Table 10, we show the result using whole sample in column (1). Then, we confine the sample to those who get pregnant within the two-year period after marriage, and the result is shown in column (2). θ turns to insignificant in column (2) after dropping out the quantile with the longest durations. We split this confined sample into two by whether a women is above 25. The results for women below and above 25 are shown in column (3) and (4), respectively. The interaction term of the post implementation indicator and duration is significant only for women get married after the age 25. In column (4), we find that for these women the chance of having a son decreased by 19.7 percent when the duration before getting pregnant increased by 1 year. ω is significant at the five-percent level. In addition, θ remains insignificant. Under the assumption that health-related costs increase with women's pregnancy age, this result is consistent with the hypothesis that households with higher costs of SSA are more likely to respond to the divorce-law amendment.

6 Robustness

6.1 Other policies

The estimators of equation (12) could capture the effect of other policies during this period if the policies affect high- and low-divorce-rate regions differently. Policies that can contaminate the estimands include the change in family-planning policy, the change in the prohibition of ultrasound B sex screening, the change in the price and quality of health services for pregnant women, and the change in the strength of son preference by changing the institutional factors that attribute more value to sons than to daughters.

We reviewed the policies in this time window and found one policy change that may contaminate the estimation: the "Law of Population and Family Planning"(LPFP hereafter) announced in December 2001.[40] Based on the national law, from September 2002 to December 2003, provinces issued their own "Act of Population and Family Planning." The central government later announced two more documents to strengthen enforcement of the law: "The Prohibition of Ultra-sound B Sex Screening and Induced Abortions" and "The Comprehensive Views on Curbing the Rising Sex Ratio," which were announced in January and June 2003, respectively. The ban on ultra-sound B sex-screening and SSA are stated in both the LPFP and the subsequent documents.

However, LPFP was not the first to prohibit SSA.[41] As reviewed in the part of backgrounds, the ultrasound machine is very available and inexpensive and is affordable to a lot of private clinics or even to the pregnant women themselves (Chu 2001), which allows sex screening to be done anywhere, making it even harder to detect. In addition, officials' performance is evaluated only by their ability to curb the population and not by keeping a balanced sex ratio.

LPFP essentially only restates everything already stipulated in the original "family-planning policy."

[40] Another relevant program is "Care for Girls Program" (*Guan Ai Nv Hai Xing Dong*), which was initiated in 2003, aiming to mitigate the discrimination against women and girls and eventually curb the rising SRB. However, until February 2004, the program covers only 11 counties. Therefore, the nationwide decline of SRB2 is not likely to be driven by this program.

[41] The sex-screening ultrasound B test was first introduced in the early 1980s. The SRB started to rise then, and, after the census in 1987, the central government realized the problem and sent an "emergency notice" to forbid ultrasound sex screening nationwide in 1989. On October 27, 1994, the National People's Congress of China passed the "Law on Maternal and Infant Health Care," which not only stated that ultrasound sex screening is illegal, but also established the details of the punishment: "Where personnel engaged in the work of maternal and infant health care, in violation of the stipulations of this Law, issue fake medical certificates, or undertake sex identification of the fetus, medical and health institutions or administrative departments of public health shall in light of the circumstances give them administrative sanctions; if the circumstances are serious, they shall be disqualified for practice of their profession according to law(Article 37)". The previous articles proved to be less restrictive, judging by the abnormally high SRB in the following years. Even the new LPFP does not clarify the details of the potential punishments for hospitals, the doctors/practitioners and pregnant women for engaging in ultrasound B sex screening. For example, it is not clear under which condition a fine will be charged; nor does it specify the amount of the fine.

Many official promotional materials claim that "family planning is for the first time set as a law in China, but the current policy was only restated in the law. There is no tightening or relaxation of the articles, which shows the family planning policy will be stable in the long-run." But two concerns should be noted. First, we may suspect that the enforcement of a law is stronger than in the former policy, especially because of the two accompanying documents issued afterwards. Second, the "Act of Population and Family Planning" has been slightly changed in many provinces.[42] Although these changes are minor and are estimated to affect fewer than one in ten thousand, we may still suspect that the cohorts conceived after late 2002 may not be perfectly comparable to their older peers.

A simple method to address this issue is to omit the cohort conceived after 2003, and the results are shown in Table 5, column (2). A second way to rule out the effect of LPFP is to perform an "event study." Taking advantage of the variations in the timing of implementation, which are shown in Appendix Table A1, we define the children conceived in the first six-month period following the implementation of LPFP in the locality as "normalized cohort 0." The children conceived in the second sex-month period following implementation are defined as "normalized cohort 1," and those conceived in the six-month period prior to implementation are defined as "normalized cohort -1," and so on.

We confine the sample in the same way as in estimate equations (11) and (12), so the sample includes only the second birth for a household with a firstborn daughter among the majority ethnic group (Han) in rural areas. The reference group is composed of the children conceived five years before the implementation of LPFP in a locality, which accounts for 3.73 percent of the sample. Then, we redo the exercise in Figure 1 and Figure 6 and show the results in Figure 9 and Figure 10. We also redo the regressions in(11) and (12) and show the results in Appendix Tables A4 and A5.

[Figure 9 and Figure 10 are about here]

We plot the average sex ratios of each cohort (within each category) in Figure 9 and 10. It is clear from Figure 9 that the drop-off of SRB2 happened at some point (around -3) before LPFP. Figure 10 shows that the pattern of decrease in the gap of SRB2 between high- and low-divorce-rate regions shown in Figure 6 is not driven by LPFP. In Figure 10, the decrease of the gap of SRB2 also happened earlier than cohort 0. As described in the note of Table A3, before the provincial people's congress issue their own Act, it is unlikely for households to respond to the general law issued by the central government.

6.2 Effect of the new divorce law though other channels

It is possible that the link between fewer SSAs and the pro-women divorce law was caused by the increased returns to having a daughter. If the divorce law made married daughters more financially secure and capable of helping their elderly parents, there would be less incentive to avoid a female birth. However, we use two pieces of evidence to show that this channel is unlikely.

The first argument is that the divorce propensity in households in which the firstborns are girls increased significantly, while there was no significant increase in divorce for those with firstborn sons. If the increase in returns to having daughters is the main channel, we should have seen a decline in the divorce rate for households with firstborn daughters, compared with households with firstborn sons.

[42] For example, in Hubei province, the 1987 version of the "Act of Population and Family Planning" states that "the couple is qualified to have a second child if one party of the spouses is disabled." This article was reversed in the December 2002 Act. In the 1987 Act, a remarried couple is qualified to have one child of their own only if one party has had no child and the other party has had no more than one child. The 2002 Act relaxed these rules a little: If one party is widowed before the current marriage, and the other party has had no child before, the spouses are qualified to have their own child even if the widowed party had two children in the previous marriage.

Because the public good (a daughter) of the marriage would be more valuable after the new divorce law, more marriage surpluses could be derived from having a daughter, and divorce should be less likely than before. Our observations contradict this hypothesis.

Second, in line with other empirical research on women's position in a household, we resort to evidence on consumption. In the dictator-husband decision-making procedure model, women's consumption could go either way. If the husband switches from abortion to non-abortion, the wife's consumption could be even lower when her reservation utility improves because, without SSA, it is now comparatively inexpensive to compensate the wife. However, according to this model, women's consumption will unambiguously increase for households already finished having children. Put another way, for these households, the new divorce law cannot affect either their fertility or SSA decision, and the improvement in the wife's utility within marriage can only be reflected by the increase in consumption. However, the competing hypothesis of higher returns to having a daughter does not predict any increase in the wife's consumption relative to the husband's.

Although it could be arbitrary to distinguish between "women's consumption" and "men's consumption" in an empirical analysis (Chiappori and Browning, 1998), a great deal of empirical research finds that when women have more income resources, the change in the household's consumption pattern is characterized by an increase in nutritional intake and a decrease in the consumption of liquor and cigarettes (Thomas, 1989; Udry and Duflo, 2003). It is a useful check to see if a similar pattern can be observed within households who finished having children before the enactment of the new divorce law.

We use CHNS data for this analysis. The results are shown in Table 11. Controlling the linear-year trend, as shown in Table 10, column (2), the husband's cigarettes consumption decreased by -2.505 per day, which is significant at the one-percent level. In column (4), the husband's frequency of drinking alcohol also decreased at the ten-percent significance level. In columns (6) and (8), the wife's daily protein intake increased by 2.61 grams, and her calorie intake increased by 88.694 at ten-percent significance level. The husband's protein and calorie intake did not have change significantly, as shown in columns (10) and (12), respectively. The results in odd-numbered columns have larger magnitude and less significance compared with the results in corresponding even-numbered columns because of the smaller sample size for each wave. These results are in line with the hypothesis that women's welfare in the household increased.

7 Conclusion

This paper builds on the literature of household theory by addressing the question of whether a pro-women change in the divorce law can change the public-good allocation, such as the investment in the sex composition of children within a household. To shed light on the abnormally high SRB in Asian countries such as China and India, it also clarifies the mechanism through which the decision on sex-selective abortion is made and the related policy implications. The paper provides detailed empirical evidence to try to establish a causal link between divorce law and the drop-off of SRB2.

Taking a closer look at the sample of 2005 One-Percent Population Survey, we find that in the population of having a firstborn daughter, three main actions are observed when the firstborn is a daughter: 1) the spouses can divorce; 2) they can have a second child without performing an SSA; or 3) they can have a second child, using SSA to avoid a female birth. To frame the three choices in one simple model, we consider the possibility that the different choices are conditional on women's intrinsic divorce options, or on their reservation utility upon divorce. All else equal, women with the best divorce options can afford a divorce, and women with the worst divorce options stay married and are more likely to use SSA, despite its health-related costs, to produce a higher expected sex composition of children. Women who are not on either extreme stay married, but provide the service of fertility without SSA. This viewpoint is consistent with a husband-dictatorship model combining with the utility that the wife's efforts to produce public goods depreciates her marginal utility of private consumption. In this model, the improvement in women's divorce options should increase the wife's well being within marriage and the most costly service (SSA) becomes unaffordable when women have better divorce options. Therefore, the improvement in women's divorce options will affect public-good provision by women.

We find that the average SRB2 before the implementation of the new law is 2.3-2.4 boys to one girl, while the ratio declined to 1.98-2.18 boys to one girl (1.56-1.6 if controlling for other factors) after implementation. We find that between the high- and low-divorce-rate regions, the difference in the probability of having a second-born son given that the firstborn is a daughter decreased by 0.1 (the difference in SRB2 between the regions decreased by 0.3). This means that most of the SRB2 decline was driven by the decline in low-divorce-rate provinces.

The policy recommendation from these results is clear. One way to reduce the abnormally high SRB is to reduce the costs of divorces, and especially the costs to women. Most importantly, this does not necessarily mean that the divorce rate will increase significantly. On the contrary, this paper shows that a substantial decrease in the SRB can be achieved by increasing women's well-being within marriages. At the same time, the divorce rate remains at a comparatively low level. This result is not generated in a contest of Coase-Becker Theorem, which requires "generalized transferable utility" (GTU). By contrast, we show that by assuming non-transferable utility function, that the wife's efforts of producing public goods depreciate marginal utility of private consumption and vary with marital status, the margin of switching SSA behavior can be different from the margin of divorce.

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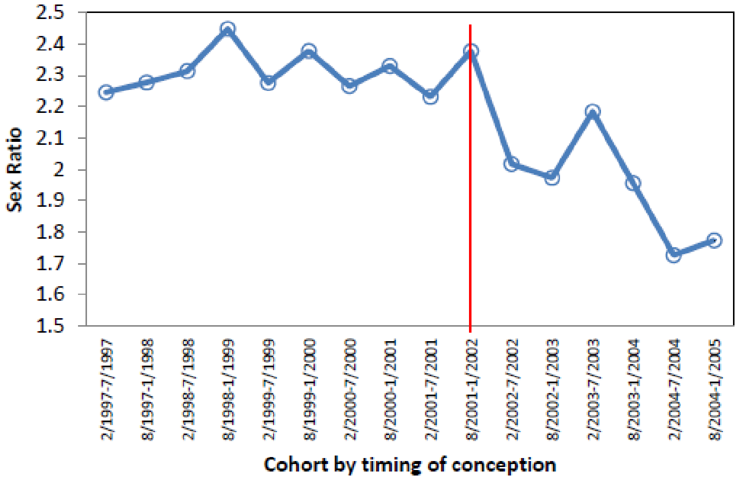

Figure 1. Sex ratio of the second-born children after having firstborn daughters

Source: The 25-percent sample of the 2005 One-Percent Population Survey.

Note: The children are categorized by the timing of their conception. Each cohort is defined as all babies conceived in each six-month period following February 1997. The sample includes only the second birth for a household with a firstborn daughter, among the majority ethnic group (Han) in rural areas.

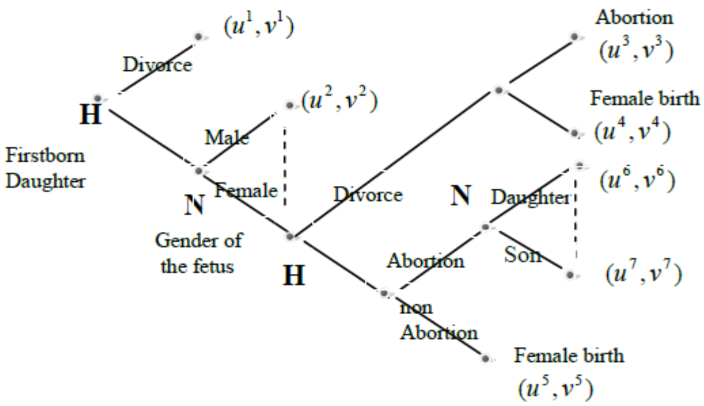

Figure 2. Time line

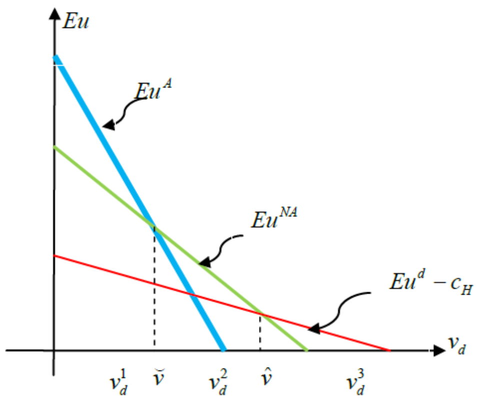

Figure 3. The wife's type and husband's optimization

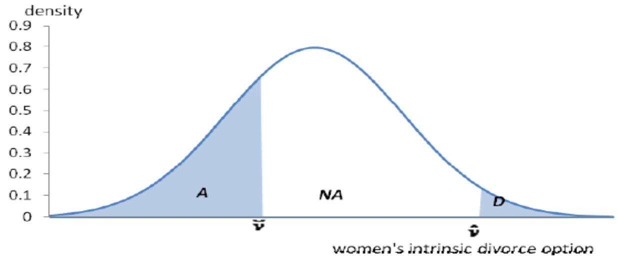

Figure 4. Sex ratio in the population

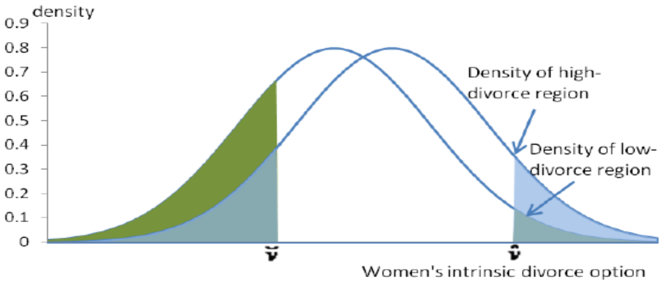

Figure 5. Regional variations in the policy effect on SRB2

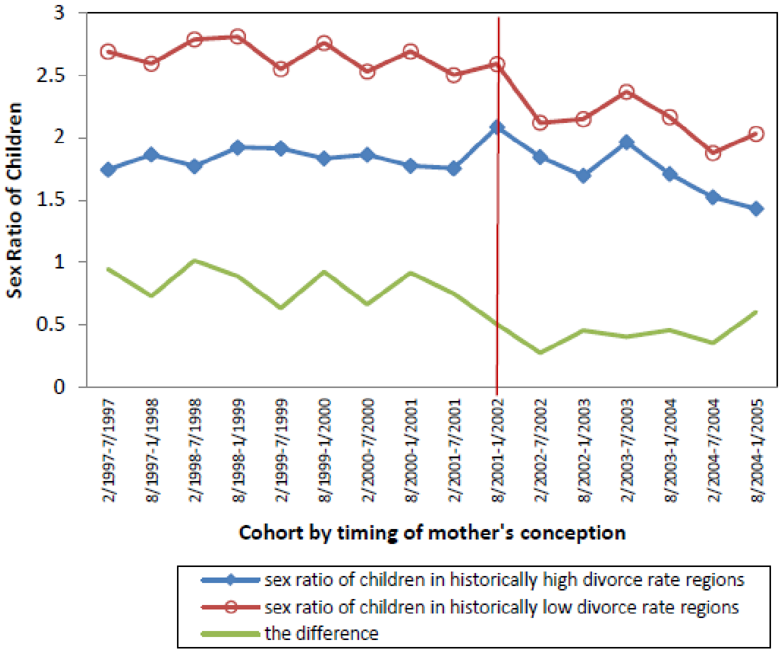

Figure 6. Trend of sex ratio of surviving children registered in the residency system (*Hukou*) by median of historical divorce rate

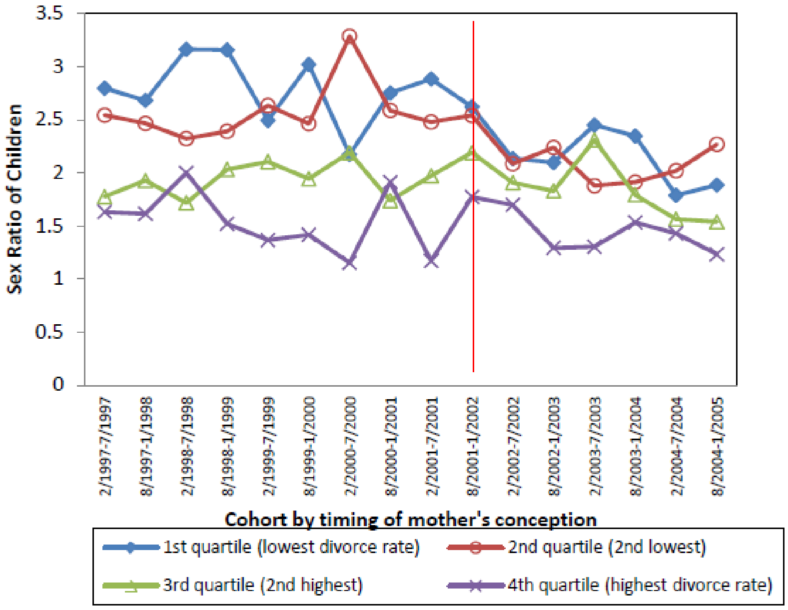

Figure 7. Trend of the sex ratio of survival children registered in the residency system (*Hukou*) by quantiles of historical divorce rate

Source: Figure 6 and Figure 7 use the 25-percent sample of the 2005 One-Percent Population Survey

Note: The sample includes only the second birth for a household with a firstborn daughter, among the majority ethnic group (Han) in rural areas. Figure 6 and Figure 7 plot the average sex ratio over cohorts defined by mother's conception time.

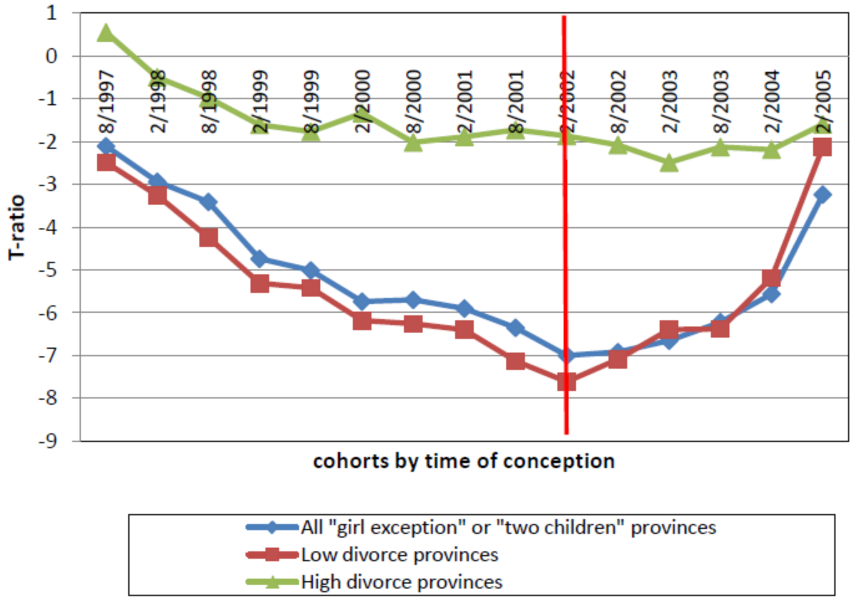

Figure 8. Coefficient T-ratio of the potential indicators for being first exposed to the implementation

Source: The 25-percent sample of the 2005 One-Percent Population Survey.

Note: The sample includes only the second birth for a household with a firstborn daughter, among the majority ethnic group (Han) in rural areas. Figure 8 plots the t-ratio of β_l in a series of regressions

$male_{ipc} = \alpha + 1\{c \geq l\} \cdot \beta_l + X_{ipc}\gamma + \lambda_p + \varepsilon_{ipc}$, where l is the timing of mother's conception marked in the horizonal axis.

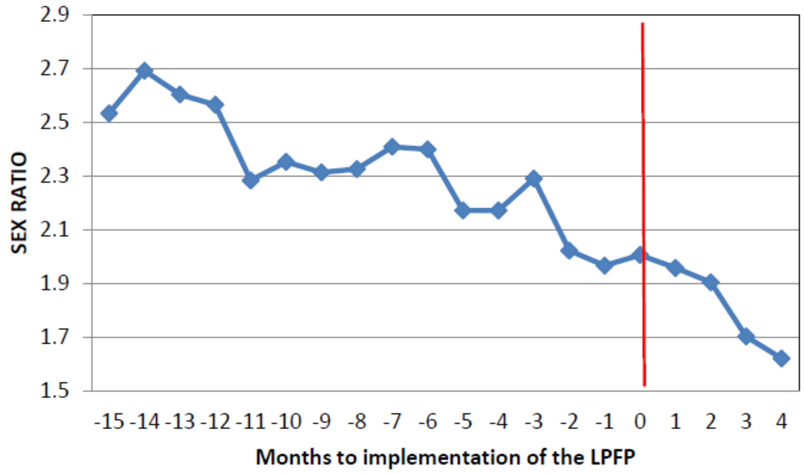

Figure 9. Sex ratio by length to implementation

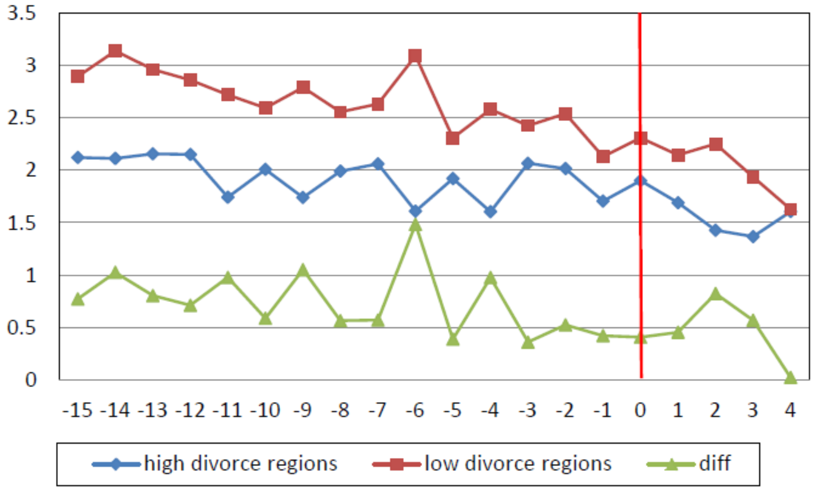

Figure 10. Sex ratio by length to implementation in high and low divorce regions

Source: The 25-percent sample of the 2005 One-Percent Population Survey.

Note: Define the children conceived in the first six-month period following the implementation of LPFP in the locality as "normalized cohort 0." The children conceived in the second sex-month period following implementation are defined as "normalized cohort 1," and those conceived in the six-month period prior to implementation are defined as "normalized cohort -1," and so on. Figure 9 and Figure 10 plot the average sex ratio over cohorts defined defined as above.

Table 1-Fertility History in the First Marriage of Rural Women Below Age 35 by Marital Status at the Time of the Survey

	Number of Observations	Pct.	Percentage/distribution by fertility history in the first marriage					
			No children	1 Son	1 Daughter	1 son and 1 daughter	2 Daughters	2 Sons
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
Panel A	*2000 Sample*							
In the first marriage	82,540	98.42	12.16	37.15	30.02	13.14	2.99	4.53
Currently divorced	441	.53	30.95	26.87	32.99	4.08	3.40	1.70
Remarried	881	1.05	22.17	29.71	34.15	7.54	2.88	3.55
Panel B	*2005 Sample*							
In the first marriage	147,149	94.71	9.84	34.59	24.68	19.59	4.87	6.43
Currently divorced	1,092	.70	27.29	29.30	30.22	7.05	2.75	3.39
Remarried	7,118	4.58	30.90	25.13	26.54	7.69	4.74	5.00

Source: The .1-percent sample of 2000 Population Census and the 25-percent sample of the 2005 One-Percent Population Survey

Notes: The 2000 Census and 2005 Population Survey do not include a question about the timing of divorce or the timing for each birth. We use the codebook provided by Lin and Zhao (2009) to infer the ferility history in women's first marriages. Table 1 can, to some extent, show the hazard of divorce by fertility history.

Table 2-Descriptive Statistics on Married Rural Woman below Age 35

	Obs	Mean	Std	Min	25 pc	50pc	75pc	Max
Panel A	*2000 Sample*							
First marriage age	94,512	21.477	2.780	16	21	22	24	31
Span between First pregnancy and first marriage(mos)	57,815	10.502	13.442	1	3	6	13	165
Age when gave 1st birth	57,815	22.949	2.445	15	21	23	24	35
Age when gave 2nd birth	34,666	24.733	3.179	19	22	24	27	35
Education years	94,512	7.376	2.454	0	6	9	9	16
Migrant	94,512	0.036	0.187	0	0	0	0	1
Panel B	*2005 Sample*							
First marriage age	202,212	21.590	2.631	16	20	21	23	35
Span between First pregnancy and first marriage(mos)	151,225	13.934	18.964	0	3	7	17	125
Age when gave 1st birth	151,225	24.175	3.294	12	22	24	26	35
Age when gave 2nd birth	57,666	28.737	4302	16	26	29	31	35
Education years	202,212	7.666	2.638	0	6	9	9	19
Migrant	202,212	.065	.247	0	0	0	0	1

Source: The .1-percent sample of 2000 Population Census and the 25-percent sample of the 2005 One-Percent Population Survey.

Note: The sample is confined to rural women below age 35 because the decision on SSA or divorce comes before the second birth. In the samples from both sources, most women are done having babies before age 35.

Table 3-Statistical Descriptions of Cohorts Conceived in Each Certain Six-month Period Following February 1997

	#Female Birth	*#Male Birth*	*SRB*	*Father's age at this birth*	*Mother's age at this birth*	*Father's school years*	*Mother's school years*	*Share of father ever divorced*	*Share of mother ever divorced*
Cohort	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)
2/1997	289	649	2.245	30.001	28.429	8.393	7.362	.013	.019
				(3.753)	(3.427)	(1.912)	(2.369)	(.113)	(.138)
8/1997	310	706	2.277	29.642	28.231	8.291	7.473	.008	.019
				(3.845)	(3.48)	(2.097)	(2.527)	(.09)	(.136)
2/1998	265	613	2.313	30.406	29.009	8.412	7.452	.008	.019
				(4.141)	(3.639)	(1.969)	(2.521)	(.087)	(.138)
8/1998	277	677	2.444	30.306	28.584	8.447	7.491	.014	.018
				(4.062)	(3.495)	(1.984)	(2.389)	(.119)	(.133)
2/1999	289	660	2.283	30.793	29.39	8.449	7.495	.01	.026
				(3.931)	(3.601)	(1.97)	(2.279)	(.098)	(.16)
8/1999	294	699	2.377	30.286	28.949	8.348	7.65	.011	.02
				(3.9)	(3.453)	(1.976)	(2.313)	(.105)	(.141)
2/2000	312	710	2.275	31.144	29.727	8.411	7.509	.009	.014
				(4.082)	(3.648)	(1.955)	(2.314)	(.096)	(.117)
8/2000	294	685	2.329	30.696	29.186	8.362	7.302	.014	.011
				(3.632)	(3.447)	(2.01)	(2.488)	(.117)	(.106)
2/2001	272	606	2.227	31.478	29.902	8.427	7.691	.006	.01
				(3.598)	(3.397)	(2.002)	(2.31)	(.078)	(.101)
8/2001	305	724	2.373	31.108	29.648	8.478	7.473	.012	.013
				(4.09)	(3.718)	(1.858)	(2.509)	(.11)	(.112)
2/2002	336	678	2.017	31.943	30.184	8.389	7.627	.007	.02
				(4.458)	(3.768)	(1.912)	(2.277)	(.083)	(.139)
8/2002	341	674	1.976	31.379	29.799	8.554	7.692	.021	.014
				(4.243)	(3.983)	(1.948)	(2.356)	(.144)	(.117)
2/2003	343	748	2.18	31.877	30.423	8.54	7.614	.008	.017
				(4.155)	(3.903)	(1.901)	(2.297)	(.087)	(.128)
8/2003	364	712	1.956	31.317	29.927	8.516	7.743	.011	.014
				(3.797)	(3.605)	(1.939)	(2.264)	(.106)	(.117)
2/2004	420	727	1.78	32.28	30.675	8.464	7.745	.007	.019
				(4.339)	(3.869)	(1.812)	(2.194)	(.084)	(.137)
8/2004	292	517	1.77	32.055	30.184	8.583	7.88	.015	.014
				(4.173)	(3.956)	(1.762)	(2.14)	(.121)	(.116)

Source: The 25-percent sample of the 2005 One-Percent Population Survey.

Note: The children are categorized by the timing of their conception. Each cohort is defined as all babies conceived in each six-month period following February 1997. The sample described in this table includes only the second birth for a household with a firstborn daughter, among the majority ethnic group (Han) in rural areas.

Table 4-The Change in Women's (Ages 20 to 35) Propensity to be 'Currrently Divorced' After the Implementation of the Amended Divorce Law

Dependent variable: The indicator of whether the woman was "currently divorced"								
	"Girl Exception"or"two Children"				*"One-Child Policy"*			
	Firstborn daughter		Firstborn son		Firstborn daughter		Firstborn son	
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
Post	-.003	-.001	.001	.004	-.010	-.024	.004	.025
implementation	(.002)	(.003)	(.001)	(.002)*	(.007)	(.025)	(.004)	(.017)
Historical Crude	.005	.005	.002	-.001	.010	-.011	-.003	.045
Divorce×post	(.003)**	(.002)**	(.003)	(.002)	(.006)*	(.051)	(.004)	(.030)
Controls	No	Yes	No	Yes	No	Yes	No	Yes
Woman Cohort F.E.	No	Yes	No	Yes	No	Yes	No	Yes
Provincial F.E.	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Adjusted R^2	.0028	.0146	.0124	.0138	.0019	.0050	.0032	.0034
Observations	59,251	58,671	84,362	82,404	12,205	12,205	15,534	15,534

Source: The repeated cross-sections from the .1-percent sample of 2000 Population Census and the 25-percent sample of the 2005 One-Percent Population Survey

Note: The divorce rate in the interaction terms is the crude divorce rate in 2000. Control variables include women's years of education, age at the time of survey, quadratic term of age, and age at first marriage; and provincial GDP per capita, population, women's labor-participation rate, and the sex ratio of adults (aging from 20-40 at the survey point). We use 1990 Population Census data to calculate the provincial sex ratio of those aged ten to 30 as a proxy for the adult sex ratio for adults between 20 and 40 in year 2000, and we calculate the sex ratio of those aged five to 25 to proxy the sex ratio of adults between 20 and 40 in 2005 data.

Columns (1) (2) (5) and (6) use the sample composed of women with firstborn daughters in their first marriages. Columns (3) (4) (7) and (8) use the sample composed of women with firstborn sons in their first marriages. The imbalanced sample size is partly due to the strategy to infer the firstborn gender with the limited information from the 2000 Population Census and the 2005 Survey. (See the codebook provided by Lin and Zhao (2009).)

In columns (1)-(8), the numbers in parentheses are the clustered standard errors at provincial level.

Table 5- Sex Ratio by Cohort: Each Cohort is Defined as All Babies Conceived in a Certain Six-month Period as following the Date Listed in the Table (cohort conceived between 2/1997 and 7/1997 is the reference group)

	Dependent Variable: The indicator of whether the second child is male							
	All "Girl Exception" or "Two Children" Provinces				*Low divorce Provinces*		*High divorce Provinces*	
	(1)		(2)		(3)		(4)	
8/1997	-0.007	(0.025)	-0.004	(0.025)	-.019	(.029)	.039	(.045)
2/1998	-0.015	(0.026)	0.012	(0.026)	-.011	(.031)	.017	(.047)
8/1998	0.005	(0.025)	0.006	(0.025)	-.009	(.029)	.046	(.045)
2/1999	-0.026	(0.026)	-0.024	(0.025)	-.024	(.029)	.019	(.046)
8/1999	-0.015	(0.025)	-0.014	(0.025)	-.026	(.029)	.023	(.047)
2/2000	-0.04	(0.025)	-0.042	(0.026)	-.028	(.025)	-.006	(.046)
8/2000	-0.04	(0.025)	-0.038	(0.025)	-.023	(.025)	-.003	(.046)
2/2001	-0.036	(0.026)	-0.035	(0.026)	-.026	(.026)	-.001	(.048)
8/2001	-0.035	(0.025)	-0.035	(0.026)	-.031	(.026)	.037	(.045)
2/2002	-0.064	(.026)**	-0.068	(.026)**	-.086	(.026)***	-.003	(.040)
8/2002	-0.074	(.026)***	-0.075	(.026)***	-.093	(.026)***	-.012	(.040)
2/2003	-0.06	(.025)**			-.073	(.026)***	.016	(.040)
8/2003	-0.123	(.026)***			-.090	(.026)***	-.045	(.041)
2/2004	-0.131	(.026)***			-.134	(.025)***	-.071	(.041)*
8/2004	-0.13	(.028)***			-.124	(.028)***	-.097	(.047)**
F-value(wald test)							2.97	
Prob > F							.0068	
Controls	Yes		Yes		Yes		Yes	
Provincial F.E.	Yes		Yes		Yes		Yes	
Adjusted R^2	.028		.0379		.0388		.0327	
Observations	10721		7909		7283		3438	

Source: The 25-percent sample of the 2005 One-Percent Population Survey.

Note: The sample includes only the second birth for a household with a firstborn daughter, among the majority ethnic group (Han) in rural areas. The variables of interest are the dummy indicators of the time when the second child was conceived. The time listed in the first column is the starting point of each six-month period. The control variables include mother's age when giving birth, mother's age squared, mother's education level, father's education level, whether the mother is an immigrant and the baby's age at the survey point. The numbers in parentheses are the clustered standard errors by province using Moulton (1986) factors.

TABLE 6-THE DIFFERENCE OF SEX RATIO BETWEEN HISTORICALLY HIGH- AND LOW-DIVORCE-RATE REGIONS ACROSS COHORTS (cohort conceived between 2/1997 and 7/1997 as the reference group)

Dependent Variable: The indicator of whether the second child is male								
	High vs Low (low div=1)		*1st quartile v.s. 3rd and 4th quartiles*(1stqt=1)		*2nd quartile v.s. 3rd and 4th quartiles*(2ndqt=1)		*Crude divorce rate In 2000*	
	(1)		(2)		(3)		(4)	
Interact8/1997_divorce	-0.068	(.054)	-0.077	(0.058)	-0.059	(0.066)	0.154	(.083)*
Interact2/1998_divorce	-0.048	(.056)	-0.027	(0.06)	-0.085	(0.069)	0.089	(0.079)
Interact8/1998_divorce	-0.064	(.053)	-0.041	(0.058)	-0.098	(0.066)	0.075	(0.08)
Interact2/1999_divorce	-0.07	(.055)	-0.073	(0.059)	-0.066	(0.067)	0.125	(0.077)
Interact8/1999_divorce	-0.065	(.055)	-0.058	(0.059)	-0.073	(0.066)	0.039	(0.089)
Interact2/2000_divorce	-0.061	(0.054)	-0.085	(0.059)	-0.019	(0.066)	0.09	(0.076)
Interact8/2000_divorce	-0.062	(0.055)	-0.058	(0.059)	-0.067	(0.068)	0.095	(0.087)
Interact2/2001_divorce	-0.056	(0.056)	-0.047	(0.06)	-0.068	(0.068)	0.081	(0.081)
Interact8/2001_divorce	-0.112	(.054)**	-0.113	(.058)*	-0.11	(.066)*	0.159	(.079)**
Interact2/2002_divorce	-0.114	(.054)**	-0.102	(.058)*	-0.132	(.068)*	0.194	(.069)***
Interact8/2002_divorce	-0.144	(.054)***	-0.143	(.058)**	-0.141	(.068)**	0.171	(.081)**
Interact2/2003_divorce	-0.152	(.053)***	-0.136	(.057)**	-0.175	(.066)***	0.127	(.075)*
Interact8/2003_divorce	-0.078	(.054)	-0.077	(.059)	-0.079	(.067)	0.117	(.073)
Interact2/2004_divorce	-0.121	(.053)**	-0.12	(.057)**	-0.12	(.066)*	0.133	(.071)*
Interact8/2004_divorce	-0.054	(.058)	-0.064	(.063)	-0.036	(.071)	0.111	(.076)
Controls	Yes		Yes		Yes		Yes	
Provincial F.E.	Yes		Yes		Yes		Yes	
Adjusted R^2	0.0459		0.0533		0.0277		0.0434	
Observations	10,721		7,951		6,208		10,721	

Source: The 25-percent sample of the 2005 One-Percent Population Survey.

Note: The sample includes only the second birth for a household with a firstborn daughter, among the majority ethnic group (Han) in rural areas. In column (1), we compare the gap in the SRB between historically high- and low-divorce-rate regions, categorized by the median of the 2000 divorce rate. We then categorize provinces into four groups by quantiles of divorce rate in 2000. In column (2), we compare the SRB of the first quantile with that of the third and fourth quantiles. In column (3), we compare the SRB of the second quantile with that of the third and fourth quantiles. The control variables include parents' age at time of this birth, parents' schooling years, whether the mother is an immigrant and the baby's age at the survey point. Only cohorts conceived after February 1997 are included in this sample. The numbers in parentheses are the clustered standard errors by province using Moulton (1986) factors.

Table 7-The change in Sex Ratio for children under age one between 2000 and 2005 (Using only the fertility data reported by women between ages 15 and 49 years old regarding having given birth in the 12 months prior to the survey)

Dependent Variable: The indicator of whether the second child is male				
	Comparison between 2000 and 2005		*Comparison between 1990 and 2000*	
	(1)	(2)	(3)	(4)
Lowest divorce Quartile×post	-.147 (.051)***		.001 (.008)	
2nd Lowest divorce Quartile×post	-.002 (.046)		.006 (.009)	
2nd highest divorce Quartile×post	-.013 (.045)		.009 (.01)	
Divorce rate×post		.026 (.016)*		-.026 (.021)
Post	.05 (.074)	-.06 (.02)***	.087 (.008)***	.101 (.008)***
Province F.E.	Yes	Yes	Yes	Yes
R-square	.0368	.0268	.1412	.1414
Observations	3228	3228	177,204	177,204

Source: The repeated cross-sections of the .1-percent sample of the 2000 Census and the 25-percent sample of the One-Percent Population 2005 Survey.

Note: The control variables include mother's age when giving birth, mother's age squared, father's age when the baby was born, mother's education level, father's education level, whether the mother is an immigrant in columns (1) and (2), and the baby's age at the survey point. The GDP per capita, population, women's labor-participation rate, the sex ratio of adults (aged from 20-40 at the survey point) are also controlled. We use the 2000 crude divorce rate in column (2) and to categorize quantiles in column (1). We use the 1990 crude divorce rate (The data are from Zeng(1993)) for the same purposes. The numbers in parentheses are the clustered standard errors by province using Moulton (1986) factors.

Table 8- The Propensity for Selective Abortions Pre- and Post-enactment of the Divorce Law

Dependent Variable: The indicator of having abortion(s) 10 months prior to the survey								
	Any abortion(s)				*Induced abortion(s)*			
	Whole Sample		*Firstborn Daughter*		*Whole Sample*		*Firstborn Daughter*	
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
Wave1997	.009		-.005		-.041		-.031	
	(.021)		(.072)		(.036)		(.092)	
Wave 2000	.003		.041		-.030		.066	
	(.043)		(.075)		(.041)		(.093)	
Wave 2004	-.017		-.049		-.081		-.002	
	(.039)		(.068)		(.035)**		(.087)	
Wave 2006	-.064		-.027		-.096		-.081	
	(.050)		(.082)		(.045)**		(.145)	
Wave 2009	-.030		-.053		-.080		-.027	
	(.046)		(.074)		(.041)**		(.125)	
Post		-.058		-.199		-.062		-.118
		(.049)		(.139)		(.038)*		(.133)
Linear trend		.001		.015		.000		.008
		(.005)		(.013)		(.003)		(.013)
Controls	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Provincial F.E.	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Adjusted R^2	.3698	.3476	.1036	.2787	.226	.1101	.2409	.2384
Observations	891	857	244	244	900	872	245	245

Source: The 1993, 1997, 2000, 2004, 2006 and 2009 waves of the longitudinal data of the China Health and Nutrition Survey (CHNS). We do not include wave 1991 because of the abnormally high occurrence of induced abortions in that wave. The same pattern was found by much of the research, such as Hemminki et al. (2005). These induced abortions were mostly forced abortions of the fetuses without the birth quota.

Note: The 1993 wave is the reference group. The control variables include mother's age when giving birth, mother's age squared, and mother's education level. The results are similar when higher-order terms of cohort conception time are controlled. The numbers in parentheses are the clustered standard errors by province using Moulton (1986) factors.

Table 9-The Difference of Birth Spacing Pre- and Post-enactment of the Divorce Law in Households with Firstborn Daughters

Dependent Variable: the Birth Spacing (months) between the first and the second child				
	Birth Spacing<72 months		Birth spacing between 48 and 72 months Or birth spacing <48months but mother above 28	
	(1)	(2)	(3)	(4)
Divorce rate×post		3.168		2.827
		(1.021)***		(1.058)***
Post implementation	-1.291	-3.843	-1.309	-3.508
	(.649)**	(1.056)***	(.694)*	(1.103)***
Linear trend	1.509	1.535	1.287	1.314
	(.072)***	(.072)***	(.076)***	(.072)***
Controls	Yes	Yes	Yes	Yes
Province F.E.	Yes	Yes	Yes	Yes
Adjusted R^2	.2665	.2639	.1586	.1662
Observations	7,395	7,395	6,404	6,404

Source: The 25-percent sample of the 2005 One-Percent Population Survey.

Note: The control variables include mother's age when giving birth, mother's age squared, father's age when the baby was born, mother's education level, father's education level, whether the mother is an immigrant and the baby's age at the survey point. The sample includes only households in which both spouses are in their first marriage and cohorts are conceived after February 1997. The results are similar when higher-order terms of cohort conception time are controlled. The numbers in parentheses are the clustered standard errors by province using Moulton (1986) factors.

Table 10-The Difference of Sex Ratio between women with different first pregnancy age across Cohorts (cohort conceived between 2/1997 and 7/1997 as the reference group)

Dependent Variable: The indicator of whether the second child is male				
	(1)	(2)	(3)	(4)
	Whole sample	Duration<=2yrs	Duration<=2yrs	
			marriage age<25	>25
Duration bw marriage	-.003	.006	.005	-.197
and first pregnancy	(.006)	(.031)	(.033)	(.112)**
Post	-.071	-.071	-.066	-.037
	(.019)***	(.043)	(.045)	(.031)
1st-pregnancy age	.010	.009	-.005	.109
	(.005)**	(.016)	(.018)	(.067)
Controls	Yes	Yes	Yes	Yes
Provincial F.E.	Yes	Yes	Yes	Yes
Adjusted R^2	.0384	.0392	.0381	.0513
Observations	6,592	5,232	4,709	523

Source: The 25-percent sample of the 2005 One-Percent Population Survey.

Note: Duration is defined as that between the marriage and women's first pregnancy. The control variables include mother's first-marriage age, father's age when the baby was born, mother's education level, father's education level, whether the mother is an immigrant and the baby's age at the survey point. The sample includes only households in which the wives gave birth to the first child 1 year after marriage, and cohorts conceived after February 1997. The numbers in parentheses are the clustered standard errors by province using Moulton (1986) factors.

Table 11-Consumption in Households That Were Done Having Children Before 2001

Dependent Variable	*Liquor and cigarettes consumption of the husband*				*Wife's nutrition intake 3-day average*				*Husband's nutrition intake 3-day average*			
	#cigarettes per day		Frequency alcohol		Protein (g)		Calories		Protein(g)		Calories	
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)	(11)	(12)
Wave1997	1.099		-.145		.849		19.060		-4.124		-40.414	
	(1.385)		(.262)		(2.345)		(77.741)		(2.257)		(75.544)	
Wave 2000	.672		-.436		5.867		200.726		-1.654		138.671	
	(2.123)		(.405)		(3.625)		(124.257)		(3.429)		(117.254)	
Wave 2004	-.977		-.732		9.673		332.774		-3.323		191.743	
	(3.148)		(.597)		(5.243)*		(181.481)*		(4.985)		(180.538)	
Wave 2006	-1.034		-.862		13.299		379.244		-1.467		259.323	
	(3.672)		(.695)		(6.158)**		(217.256)*		(6.028)		(190.323)	
Wave 2009	-.183		-1.082		17.284		493.375		-.624		315.804	
	(4.462)		(.841)		(7.374)**		(256.512)*		(7.137)		(256.673)	
Post		-2.505		-.205		2.610		88.694		.671		57.448
		(.639)***		(.116)*		(1.377)*		(51.487)*		1.362		(54.152)
Linear trend	.159	.278	.058	.005	-4.250	-.708	-48.771	-30.864	-.268	-.346	-32.179	-30.864
	(.248)	(.064)***	(.049)	(.009)	(1.737)**	(.114)***	(14.925)***	(6.690)***	(.427)	(.106)***	(15.234)**	(6.690)***
Controls	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Community F.E.	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Adjusted R^2	.1653	.1645	.1011	.1010	.1867	.1794	.1909	.1577	.1071	.0864	.3146	.1247
Observations	3479	3479	5197	5197	4069	4069	4586	4586	3267	3267	3830	3830

Source: The 1991,1993, 1997, 2000, 2004, 2006 and 2009 waves of the longitudinal data of the China Health and Nutrition Survey (CHNS).

Note: The control variables include age, education years, spouse's age, and education years. The numbers in parentheses are the clustered standard errors by province using Moulton (1986) factors. The sample is confined to rural households that were done having children before 2001—specifically, those who had one son, one son and one daughter, or two daughters by 2001. The reference waves are wave 1991 and wave 1993. The 1989 wave does not encompass the variables of liquor and cigarettes (in the PE data set). The sample size for diet data is only a half of the sample in other waves. So, we use the sample from 1991 for all regressions. In columns (3) and (4), the frequency of having alcohol =0 if never having any alcohol in last year; =1 if less than once a month; =2 if 1-3 times per month; =3 if 1-2 times per week; =4 if 3-4 times per week; =5 if drinking every day.

A Appendix: The distribution of willingness to have SSAs

Figure 1 shows that the sex ratio for the second birth after a firstborn girl is around 2.2-2.3 for those conceived in early 1997. (The probability of having a son is around 68-69 percent.) The 1997 survey of Population and Reproductive Health in China asks women questions to recall the history of each pregnancy. The national sample allows us to do an exercise to check how much of the skewed sex ratio can be explained by sex-selective abortion behavior, or put another way, if we know the frequency/occurrences of sex-selective abortions, we are able to check whether the abortions in the 1997 survey of Population and Reproductive Health sample can result in the sex ratio in Figure 1.

The 1997 survey of Population and Reproductive Health in China suggests among women in their first marriages and whose first child was a girl, 79.77 percent reported that they never had a abortion; 15.01 percent had one abortion prior to the survey; 3.43 percent had two abortions; and the remaining 1.79 percent had more than two abortions prior to the survey. However, the abortions observed are not the same as would be seen among the real number of women who would have abortions if they were carrying a female fetus because there is around a 0.5 percent chance of having a male and, thus, no need to have an abortion. Or put another way, an SSA could possibly happen only when a woman carries a female fetus, but we are concerned with women who would have SSA if the fetus is female. Therefore, we adjust these numbers accordingly. Denote the portion of "would have 0 SSA" as p_0, the portion "would have 1 SSA" as p_1, and the portion "would have 2 SSA" as p_2. We regard "having more than two abortions" as three abortions, and denote the portion as p_3. Assume that the probability of having a son is 0.5 and that this probability in each pregnancy is independent.

$$\begin{aligned} p_0 + 0.5(p_1 + p_2 + p_3) &= 79.77 \\ 0.5p_1 + 0.25(p_2 + p_3) &= 15.01 \\ 0.25p_2 + 0.125p_3 &= 3.43 \\ 0.125p_3 &= 1.79 \end{aligned}$$

59.54 percent of the women would never have had abortions to avoid a female fetus; 19.58 percent would have had one abortion; 6.56 percent would have had two abortions; and 14.32 percent would have had more than two abortions to improve the probability of having a son.

Thus, the probability of having a son in this population is $0.5p_0 + 0.75p_1 + 0.875p_2 + 0.9375p_3 =$ 63.62%.Then, we see that using the national sample of the 1997 survey of Population and Reproductive Health in China, the probability of having a male birth after a firstborn daughter is .63, which is 93.6 percent of that calculated using the 2005 One-Percent Population Survey. Appendix A also illustrates how the skewed SRB can be decomposed to populations with different willingness to have SSAs.

B Appendix: The Divorce Law in "One-Child Policy" Provinces

The bench mark model in this paper can be extended to predict the change in sex ratio and divorce rate in "One-Child Policy" provinces, where a second child is not allowed.[43] In this appendix, we focus on empirical evidence about children's sex ratio in "One-Child Policy" provinces.

Column (1) in Table A6 shows that there is no significant decrease in sex ratio of the firstborn in "One-Child Policy" provinces. One possible explanation is that SSA was rarely used in these provinces even before the new divorce law, which is indicated by the sex ratio of 1.08. Given the limited sample size, the change in SRB, if any, could be very hard to detect. In Figure A2, we plot the sex ratio of the firstborn in both "One-Child-Policy" provinces and provinces in which a second child is allowed after a firstborn daughter. We observe that SRB is only slightly above the natural level of 1.05 for most of the cohorts. [44]

[43] We develop a model with the functional form $u = x_H \cdot \eta$ and $v = \frac{x_W}{e} \cdot \eta$ in an earlier version of this paper so that it can incorporate four scenarios after the fetus of the first pregnancy is revealed as female: (1) stay married after a female birth; (2) divorce after a female birth; (3) stay married after having an SSA and have a second birth (without SSA): and (4) divorce after an SSA.

[44] TThe square-line bumps up and down because of the limited sample size. Each cohort includes only 150-250 children,

Much research discusses the reason for fewer SSAs in "One-Child Policy" provinces. Das Gupta et al. (2009) attribute the less-skewed SRB in these provinces to the better pension system and higher social status of women. Qian (2008) finds suggestive evidence that remote areas are more likely to receive loose restrictions on fertility because of the high cost of enforcing family-planning policy in these regions. By contrast, the six "One-Child Policy" provinces are the most urbanized areas in China, and the enforcement the "One-Child Policy" and the ban of SSAs should be less expensive and more effective, compared with other provinces.

Restricted by the "One-Child Policy," there are still couples in these regions having a second child. These exceptions are either because of the variation in local family-planning policy,[45] or due to paying a fine to the local family-planning committee. We investigate the SRB of the second birth after a firstborn daughter in "One-Child Policy" provinces; the result is shown in column (2) of Table A6. The sex ratio for the second birth shows a similar but less significant pattern of the decline after the new divorce law, compared to that in "Girl Exception" or "Two-Child" provinces, as shown in Table 5. This result suggests that since both abortion and having another birth without birth quota are illegal, the spouses are more likely to have a second birth instead of having an SSA for the first pregnancy. Thus, the margin affected by the divorce law is unlikely to be the first birth, even in "One-Child Policy" provinces.

C Appendix figures and tables

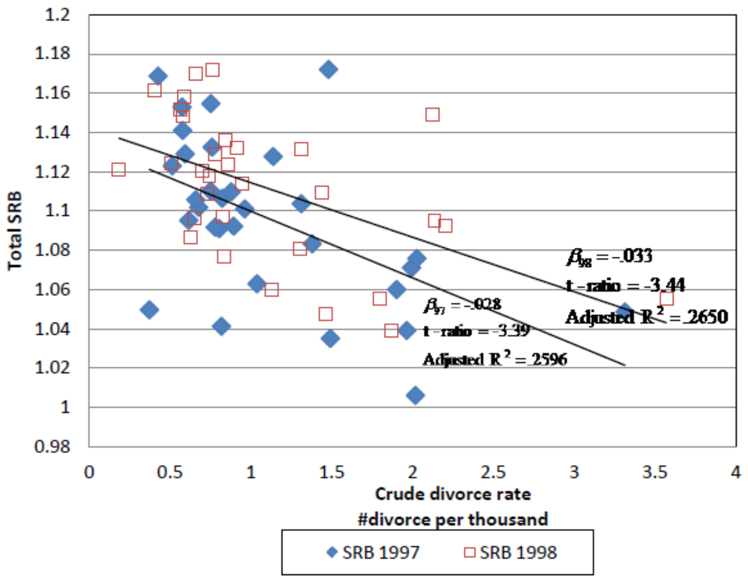

Figure A1: Negative relationship between SRB and divorce

Note: Total sex ratio at brith (SRB) is the number of male births to female births across parties in year 1997 and 1998, respectively. The crude divorce rate is the ratio of divorces in every one thousand of population. Regressing total SRB on the crude divorce rate, the negative correlation is statistically significant. Data source: China's Year Book. The total SRB data are available only for 1996-1998, but the crude divorce rate in 1996 is measured differently. Therefore, we show only 1997 and 1998 using this independent data source.

while along the diamond-line, which represents "Girl exception" or "2-Child" provinces, each cohort includes around 600 children.

[45] CHNS data suggest that there are variations in family-planning policy at the county level and below, while the principle is stipulated by the provincial "Family Planning Act."

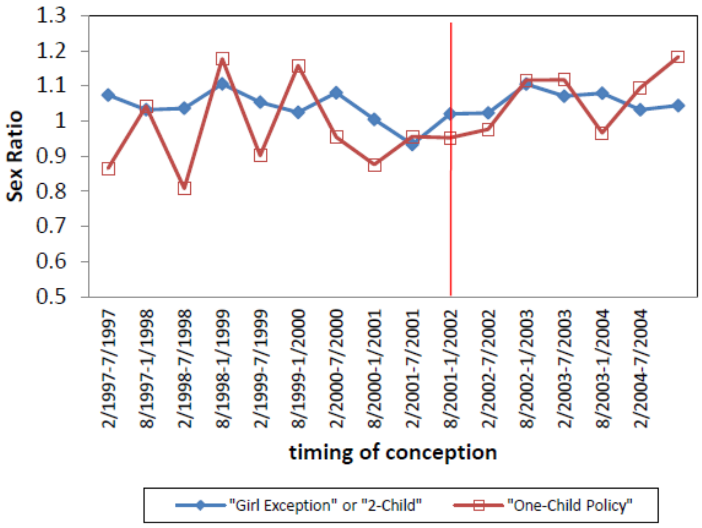

Figure A2: Sex ratio of the firstborn in "One-Child Policy" provinces

Table A1. Regional Differentials of Divorce Rates

Provinces, by Group		Crude divorce rate (%)[a]				Age specific divorce rate[b]	
		1982	1990	2000	2005	2000	2005
1	Xinjiang	4.16	3.79	2.86	3.88	147.15	147.57
2	Heilongjiang	0.67	1.56	2.03	2.51	40.69	45.27
	Jilin	0.59	1.59	1.87	2.69	44.85	46.85
	Qinghai	0.80	1.24	1.29	1.47	48.32	52.69
	Liaoning	0.57	1.45	1.93	2.63	26.76	30.27
3	Beijing	0.59	1.35	1.92	2.21	35.43	32.12
	Shanghai	0.46	1.25	1.90	2.19	36.94	40.66
4	In. Mongolia	0.51	0.85	1.37	1.03	26.48	24.06
	Ningxia	0.50	0.76	1.10	1.51	47.81	52.92
	Shanxi	0.77	0.77	0.72	0.86	38.14	36.77
	Guizhou	0.48	0.67	0.74	1.02	61.23	55.56
	Shaanxi	0.44	0.75	0.89	1.08	44.24	45.59
	Chongqing			1.43	2.61	34.84	36.52
	Sichuan	0.30	0.75	1.21	1.72	37.55	43.56
	Yunnan	0.39	0.65	0.85	1.52	48.97	49.83
5	Tianjin	0.40	0.68	1.33	1.82	38.77	30.46
	Gansu	0.38	0.63	0.74	0.89	48.92	48.42
	Hunan	0.39	0.61	0.85	1.05	31.68	36.48
	Hubei	0.35	0.61	0.79	1.33	35.94	35.31
	Guangxi	0.37	0.52	0.64	1.03	46.62	51.35
	Hebei	0.46	0.61	0.79	1.21	25.91	27.50
	Henan	0.34	0.56	0.70	0.97	38.04	38.87
6	Jiangxi	0.35	0.47	0.59	0.90	23.87	29.48
	Zhejiang	0.27	0.51	0.92	1.49	29.73	31.05
	Hainan		0.41	0.50	0.72	27.58	28.26
	Guangdong	0.35	0.41	0.55	0.88	27.71	38.11
	Fujian	0.14	0.32	0.83	1.10	29.13	35.26
	Anhui	0.23	0.42	0.61	0.95	26.21	38.31
	Shangdong	1.26	1.47	0.67	1.07	40.32	37.56
	Jiangu	0.24	0.39	0.75	1.30	28.58	31.37

Note: [a]The crude divorce rate is the number of divorces per 1,000 population. The crude divorce rates in 1982 and 1990 are from Zeng (2000). The rates in 2000 and 2005 are calculated using the data from China's Year Book.

[b]The age specific divorce rate is calculated using the formula $\frac{\#\text{divorces among persons of a given age group}}{\text{population of persons in given age group}} \times 100,000$. The age specific divorce rates are calculated the .1-percent sample of the 2000 Census and 2005 One-Percent Population Survey.

Table A2-The Change in Ever Being Divorced of women between ages 20 and 35 after the Implementation of the amended divorce law

Dependent variable: the indicator for whether the woman was "ever divorced"								
	"Girl Exception"or"two Children"				*"One-Child Policy"*			
	Firstborn daughter		Firstborn son		Firstborn daughter		Firstborn son	
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
Post	-.004	-.004	.000	.001	-.004	-.034	.007	-.032
implementation	(.003)*	(.005)	(.001)	(.002)	(.009)	(.032)	(.006)	(.028)
Historical Crude	.005	.005	-.004	-.001	.005	-.054	-.012	-.102
Divorce×post	(.002)*	(.002)**	(.002)*	(.002)	(.008)	(.068)	(.006)**	(.065)
Controls	No	Yes	No	Yes	No	Yes	No	Yes
Woman Cohort F.E.	No	Yes	No	Yes	No	Yes	No	Yes
Provincial F.E.	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Adjusted R^2	.0129	.0148	.0058	.0138	.0016	.0045	.0032	.0034
Observations	59,251	58,671	84,362	82,404	122,05	122,05	15,534	15,534

Source: The repeated cross-sections of the .1-percent sample of the 2000 Census and 2005 One-Percent Population Survey.

Note: The divorce rate in the interaction terms is the crude divorce rate in 2000. Control variables include women's education years, age at the time of survey, quadratic term of age, and age at first marriage; and provincial GDP per capita, population, women's labor-participation rate, and the sex ratio of adults (ages 20-40 at the survey point). We use 1990 Population Census data to calculate the provincial sex ratio of those aged ten to 30 as a proxy for the adult sex ratio for adults between 20 and 40 in year 2000, and we calculate the sex ratio of those aged five to 25 to proxy the sex ratio of adults between 20 and 40 in 2005 data.

Columns (1) (2) (5) and (6) use the sample composed of women with firstborn daughters in their first marriages. Columns (3) (4) (7) and (8) use the sample composed of women with firstborn sons in their first marriages. The imbalanced sample size is partly because of the strategy to infer the firstborn's gender with the limited information from the 2000 Population Census and the 2005 Survey. (See Appendix C for the codebook.)

In columns (1)-(8), the numbers in parentheses are the clustered standard errors by province using Moulton (1986) factors.

Table A3. The variations in LPFP implementation across provinces

Province	Provincial Act passed	Implementation	Province	Provincial Act passed	Implementation
Beijing	18-Jul-03	1-Sep-03	Henan	30-Nov-02	1-Jan-03
Tianjin	11-Jul-03	1-Sep-03	Hubei	1-Dec-02	1-Jan-03
Hebei	25-Mar-03	1-Oct-03	Hunan	29-Nov-02	1-Jan-03
Shanxi	28-Sep-02	1-Nov-02	Guangdong	25-Jul-02	1-Sep-02
Inner Mongolia	25-Sep-02	1-Nov-02	Guangxi	27-Jul-02	1-Sep-02
Liaoning	16-Jan-03	1-Apr-03	Hainan	24-Oct-03	1-Dec-03
Jilin	27-Sep-02	1-Nov-02	Chongqing	25-Sep-02	1-Nov-02
Heilongjiang	18-Oct-02	1-Jan-03	Sichuan	26-Sep-02	1-Oct-02
Shanghai	31-Dec-03	15-Apr-04	Guizhou	29-Sep-02	29-Sep-02
Jiangsu	23-Oct-02	1-Dec-02	Yunnan	25-Jul-02	1-Sep-02
Zhejiang	3-Sep-02	3-Sep-02	Shaanxi	29-Sep-02	29-Sep-02
Anhui	28-Jul-02	1-Sep-02	Gansu	7-Sep-02	7-Sep-02
Fujian	26-Jul-02	1-Sep-02	Qinghai	20-Sep-02	1-Jan-03
Jiangxi	29-Jul-02	1-Sep-02	Ningxia	7-Nov-02	1-Jan-03
Shandong	28-Sep-02	28-Sep-02	Xinjiang	28-Nov-02	5-Dec-02

Note: Though the law was passed at the end 2001, it was only enacted on September 1, 2002. Moreover, in China, each province has its own "act of population and family planning" based on the "law of population and family planning" issued by the central government. The central government has clearly stated since the beginning of its family-planning policy that there is no unified rule about family planning, and all provinces should be allowed to have their own policy based on their specific conditions.

The procedure, in practice, is that the central government first issues a law in the National People's Congress, and then each Provincial population and family-planning commission discusses the law and drafts the provincial "act of population and family planning." The draft is sent to the national population and family-planning commission for review, which may be followed by lobbying and negotiation between the central and provincial governments. Thus, it can take two to three years for the provincial act to be approved and finally passed in the provincial people's congress and enacted. It's important to point out that the "act of population and family planning" varies a lot in many detailed articles across provinces. Therefore, in December 2001, there was no way for households to respond to the new "law of population and family planning" because the specific law in the province had not been drafted at that time, and the decline in the sex ratio before September 2002 should not be attributed to the effect of the "law of population and family planning."

Table A4-Sex Ratio by Cohort (Each Cohort is Defined as All Babies Conceived in a Certain Six-month Period Relative to the Implementation of the Law of Population and Family Planning in the Locality)

Dependent Variable: the dummy variable indicator for whether the second child is male						
	All "Girl Exception" or "Two Children" Provinces		*Low divorce Provinces*		*High divorce Provinces*	
	(1)		(2)		(3)	
c-10(10th half year before)	-.015	(.021)	-.002	(.033)	-.028	(.053)
c-9 (9th half year before)	-.015	(.028)	-.053	(.033)	.065	(.053)
c-8 (8th half year before)	-.027	(.028)	-.041	(.033)	.002	(.055)
c-7 (7th half year before)	-.031	(.021)	-.072	(.034)**	.018	(.053)
c-6 (6th half year before)	-.009	(.021)	-.028	(.033)	-.007	(.055)
c-5 (5th half year before)	-.023	(.020)	-.041	(.032)	-.026	(.055)
c-4 (4th half year before)	-.041	(.021)**	-.083	(.033)**	.016	(.055)
c-3 (3rd half year before)	-.039	(.021)*	-.057	(.034)*	-.037	(.056)
c-2 (2nd half year before)	-.054	(.022)**	-.104	(.034)***	.013	(.054)
c-1 (1st half year before)	-.023	(.021)	-.059	(.033)*	.017	(.054)
c 0 (1st half year after)	-.074	(.017)***	-.132	(.031)***	.014	(.050)
c1 (2nd half year after)	-.084	(.021)***	-.122	(.034)***	-.039	(.054)
c2 (3rd half year after)	-.121	(.022)***	-.153	(.034)***	-.089	(.050)*
c3 (4th half year after)	-.135	(.023)***	-.180	(.036)***	-.080	(.057)
c4 (5th half year after)	-.147	(.030)***	-.178	(.043)***	-.120	(.066)*
F-value(wald test)					3.10	
Prob > F					.0146	
Controls	Yes		Yes		Yes	
Provincial F.E.	Yes		Yes		Yes	
Adjusted R^2	.0381		.0414		.0250	
Observations	10721		7283		3438	

Source: The 25-percent sample of the 2005 One-Percent Population Survey.

Note: The sample is confined to the cohort conceived beginning in November 1997. The reference group is composed of the children conceived five years before the implementation of LPFP in a locality, which accounts for 3.73 percent of the sample. The sample includes only the second birth for a household with a firstborn daughter, among the majority ethnic group (Han) in rural areas. The variables of interest are the dummy indicators of the time when the second parity was conceived. The time listed in the first column is the starting point of each six-month period. The control variables include mother's age when giving birth, mother's age squared, mother's education level, father's education level, whether the mother is an immigrant and the baby's age at the survey point. The numbers in parentheses are the clustered standard errors by province using Moulton (1986) factors.

Table A5- The Difference of Sex Ratio between Historically High- and Low-Divorce-Rate Regions across Cohorts (Each Cohort is Defined as All Babies Conceived in a Certain Six-month Period Relative to the Implementation of the Law of Population and Family Planning in the Locality)

Dependent Variable: the indicator for whether the second child is male								
	High vs Low (low div=1)		*1st quartile v.s. 3rd and 4th quartiles*(1stqt=1)		*2nd quartile v.s. 3rd and 4th quartiles*(2ndqt=1)		*Crude divorce rate In 2000*	
	(1)		(2)		(3)		(4)	
Interact_10th half year before	-.136	(.048)***	-.112	(.052)**	-.171	(.059)***	.063	(.070)
Interact_9th half year before	-.059	(.050)	-.022	(.054)	-.112	(.061)*	-.018	(.082)
Interact_8th half year before	-.109	(.048)**	-.090	(.053)*	-.134	(.059)**	.092	(.067)
Interact_7th half year before	-.040	(.049)	-.045	(.054)	-.029	(.057)	-.028	(.078)
Interact_6th half year before	-.035	(.049)	-.013	(.053)	-.065	(.059)	-.008	(.069)
Interact_5th half year before	-.120	(.049)**	-.129	(.053)**	-.099	(.059)*	.094	(.069)
Interact_4th half year before	-.041	(.050)	-.005	(.054)	-.098	(.062)	-.036	(.079)
Interact_3rd half year before	-.139	(.049)***	-.136	(.054)**	-.136	(.060)**	.107	(.066)
Interact_2nd half year before	-.097	(.048)**	-.104	(.052)**	-.074	(.059)	.138	(.057)**
Interact_1st half year before	-.159	(.051)***	-.145	(.056)**	-.173	(.061)***	.080	(.079)
Interact_1st half year after	-.175	(.047)***	-.160	(.051)***	-.193	(.058)***	.152	(.068)**
Interact_2nd half year after	-.107	(.048)**	-.095	(.052)*	-.118	(.058)**	.012	(.062)
Interact_3rd half year after	-.086	(.047)*	-.064	(.051)	-.116	(.060)*	.034	(.059)
Interact_4th half year after	-.122	(.051)**	-.115	(.055)**	-.118	(.067)*	.120	(.067)
Interact_5th half year after	-.077	(.065)	-.064	(.069)	-.089	(.088)	-.040	(.104)
Controls	Yes		Yes		Yes		Yes	
Provincial F.E.	Yes		Yes		Yes		Yes	
Adjusted R^2	.0392		.0479		.0217		.0387	
Observations	10,721		7,951		6,208		10,721	

Source: The 25-percent sample of the 2005 One-Percent Population Survey.

Note: The sample is confined to the cohort conceived beginning in November 1997. The reference group is composed of the children conceived five years before the implementation of LPFP in a locality, which accounts for 3.73 percent of the sample. The sample includes only the second birth for a household with a firstborn daughter, among the majority ethnic group (Han) in rural areas. The variables of interest are the dummy indicators of the time when the second parity was conceived. The time listed in the first column is the starting point of each six-month period. The control variables include mother's age when giving birth, mother's age squared, mother's education level, father's education level, whether the mother is an immigrant and the baby's age at the survey point. The numbers in parentheses are the clustered standard errors by province using Moulton (1986) factors.

Table A6- Sex Ratio by Cohorts in "One-Child-Policy" provinces: Each Cohort is Defined as All Babies Conceived in a Certain Six-month Period as following the Date Listed in the Table (cohort conceived between 2/1997 and 7/1997 is the reference group)

Dependent Variable: the indicator for whether the second child is male				
	Firstborn		Second-born after a firstborn daughter	
	(1)		(2)	
8/1997	.043	(0.032)	-.049	(0.068)
2/1998	-.017	(0.034)	-0.019	(0.067)
8/1998	.064	(0.034)*	-0.184	(0.081)**
2/1999	-.007	(0.034)	.052	(0.063)
8/1999	.050	(0.035)	-0.023	(0.066)
2/2000	-0.009	(0.036)	0.002	(0.065)
8/2000	-0.009	(0.036)	-0.115	(0.074)
2/2001	0.008	(0.041)	-0.1	(0.073)
8/2001	0.025	(0.048)	-0.088	(0.067)
2/2002	0.019	(0.056)	-0.080	(0.073)
8/2002	0.015	(0.065)	-0.139	(0.071)**
2/2003	0.045	(0.075)	-0.091	(0.068)
8/2003	0.050	(0.073)	-0.15	(0.071)**
2/2004	0.010	(0.049)	-0.101	(0.017)
8/2004	0.018	(0.052)	-0.113	(0.077)
Control var	Yes		Yes	
Provincial F.E.	Yes		Yes	
Adjusted R^2	0.0081		0.035	
Observations	4090		922	

Source: The 25-percent sample of the 2005 One-Percent Population Survey.

Note: The sample includes only the second birth for a household with a firstborn daughter, among the majority ethnic group (Han) in rural areas. The variables of interest are the dummy indicators of the time when the second child was conceived. The time listed in the first column is the starting point of each six-month period. The control variables include mother's age when giving birth, mother's age squared, mother's education level, father's education level, whether the mother is an immigrant and the baby's age at the survey point. The numbers in parentheses are the clustered standard errors by province using Moulton (1986) factors.