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Research Article

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Contents

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Abstract

OBJECTIVE
Most research on population sex imbalance in China has focused on the One-Child Policy era. However, because much of China's fertility decline occurred during the 1970s, we investigate the possibility that sex ratios began rising during this period (as predicted by theory) before the One-Child Policy.

RESULTS
Analyzing sex ratios between 1960 and 1987 by birth order and sibship sex composition, we find that among the subset of couples expected to have the greatest demand for sons (those at higher parities without previous sons), sex ratios at birth reached 115–121 boys per 100 girls during the 1970s – implying approximately 840,000 to 1,100,000 girls missing from Chinese birth cohorts during the 1970s. Importantly, these results do not appear to be driven by differential under-reporting of living girls, or instances of adoption. Given the absence of ultrasound technologies prior to 1979, they imply the presence of postnatal sex selection in China during the 1970s.

CONTRIBUTION
Our work makes several important contributions to existing literature. First, we focus on the subset of couples among whom the demand for sons is predicted to be the strongest: higher parity couples not yet having a boy. Second, we estimate sex ratios by single year of age (from birth to age 4), distinguishing differential rates of infant death from more gradual neglect of girls as they age throughout childhood. Third, we

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combine graphical and multivariate statistical analyses to test for meaningfully imbalanced sex ratios. Finally, we measure potential irregularities in the reporting of living girls, including the adoption of girls, and we generate new estimates of unreported females.

1. Introduction

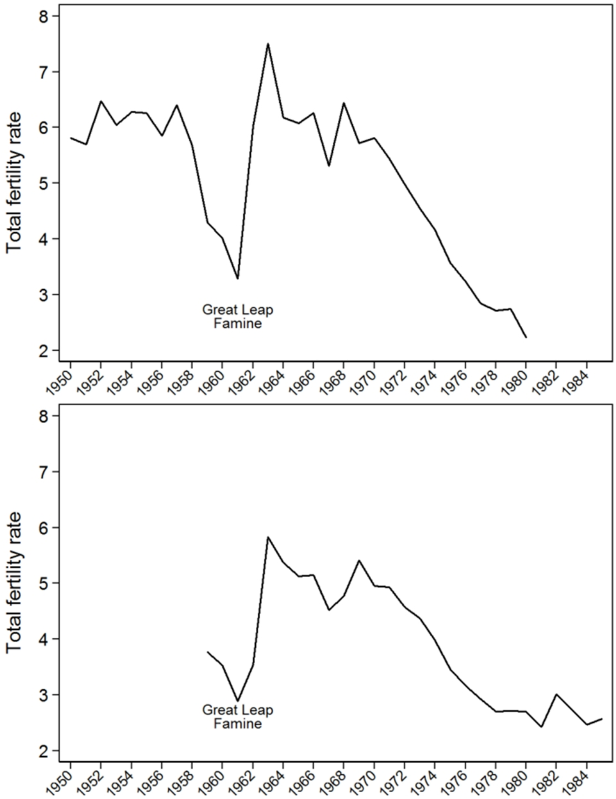
China's sex ratio at birth[5] rose dramatically throughout the 1980s (Banister 2004; Coale and Banister 1994). A large literature estimates that between 8.5 and 9.2 million females are missing from Chinese cohorts born between 1980 and 2000 – presumably the result of sex-selective abortion and childhood neglect (Almond, Li, and Zhang 2017; Cai and Lavely 2003; Chen, Li, and Meng 2013; Ebenstein 2014; Jiang et al. 2012; Tuljapurkar, Li, and Feldman 1995). Given parental preferences for sons,[6] the One-Child Policy (in 1980) and the rapid diffusion of ultrasound technologies capable of detecting fetal sex (during the early 1980s (Chen, Li, and Meng 2013)) are together generally considered responsible for China's 'missing women' that emerged during the 1980s and later (Banister 2004; Das Gupta 2005; Das Gupta and Li 1999; Ebenstein 2014; Ebenstein and Leung 2010; Johansson and Nygren 1991).[7]

However, demographic and economic theory suggest that as fertility declines in populations preferring sons, sex ratios will rise (Das Gupta and Bhat 1997; Das Gupta and Li 1999; Jayachandran 2017) – and strikingly, most of China's fertility decline occurred during the 1970s under China's first national birth planning policy, the "Later, Longer, Fewer" campaign. The campaign sought to reduce the country's birth rate by raising the minimum age at marriage ("later"), lengthening birth intervals ("longer"), and reducing fertility to a maximum of three births per couple ("fewer"). In rural areas (accounting for over 80% of the Chinese population in 1970), the total fertility rate (TFR) fell from 6.4 in 1970 to about 3 by 1979 (and from 3.2 to 1.4 during this period in urban areas; Figure 1) (Banister 2004; Cai and Lavely 2003; Coale and Banister 1994). This decline is among the most rapid documented reductions in fertility in global

[5] The sex ratio at birth is defined as the number of boys born per 100 girls.
[6] Scholars have linked China's skewed sex ratio to son preference and patriarchal traditions emphasizing the role of sons in elder care and lineage (Coale and Banister 1994; Ebenstein 2014; Ebenstein and Leung 2010; Jayachandran 2015).
[7] Coale and Banister (1994) find that average sex ratios over 5-year intervals among third and fourth order births rose as high as 109 prior to the One-Child Policy (see Table 4 in Coale and Banister 1994).

history.[8] Nonetheless, existing studies of China's fertility decline prior to 1980 are largely descriptive and report only modest sex imbalance (with unknown quantitative/statistical significance (Banister 2004; Cai and Lavely 2003; Coale and Banister 1994; Das Gupta and Li 1999; Jiang et al. 2012; Johansson and Nygren 1991).

Figure 1: Total fertility rate decline in China 1950–1984



Note: Total fertility rates (TFR) are calculated as the sum of age-specific fertility rates observed in a given calendar year. Panel A is reproduced using TFRs calculated in Chen (1984) from the 1982 National Sample Survey of Fertility and Contraception (the one-per-thousand survey) and Panel B shows highly consistent TFRs calculated using birth records in the 1988 National Sample Survey of Fertility and Contraception (the two-per-thousand survey).

[8] Other rapid declines in fertility have occurred in Iran (where the total fertility rate fell from about 6.5 in 1980 to 2.2 during the 1980s and 1990s), and South Korea (where the total fertility rate fell from 6.1 in 1960 to 2.8 during the 1960s and 1970s) (World Bank 2016).

Using historical fertility data (retrospective fertility histories in China's 1988 Two-per-Thousand National Sample Survey of Fertility and Contraception), this paper investigates the theoretical prediction that, despite efforts to prevent sex selection, sex ratios at birth began rising as China's total fertility rate plummeted during the 1970s – earlier and to higher levels than previously established. In doing so, it makes several important contributions to the existing literature. First, unlike previous work, it is able to isolate and study the subset of couples among whom the demand for sons is predicted to be the strongest: higher parity couples not yet having a boy (Arnold and Liu 1986; Das Gupta 2005; Ebenstein 2014).[9] Most previous studies use population census data in which births cannot be distinguished by parity and sex composition of older siblings.[10] Second, in estimating sex ratios by single year of age in each year (rather than pooling birth cohorts), it is able to distinguish differential rates of infant death during the first year of life (due either to differential neglect or possible infanticide) from more gradual neglect of girls as they age throughout childhood.[11] Third, it combines graphical and multivariate statistical analyses to test for meaningfully imbalanced sex ratios at birth using a sample restricted to years prior to the introduction of ultrasound technology in each province. This contribution is particularly important for analysis based on fertility histories (in which we know the ages at which girls become 'missing') rather than census counts. Fourth, it pays special attention to potential irregularities in the reporting of living girls, including the adoption of girls, generating new estimates of female under-reporting.

Consistent with theoretical predictions, we find that among third and higher parity births to couples without a surviving son,[12] sex ratios rose as high as 115–121 boys per 100 girls during the 1970s – higher than previously established. Our analysis implies that, although rare in absolute terms (accounting for less than 0.5% of births) and concentrated among a narrow subset of couples, approximately 840,000–1,100,000 girls are missing from Chinese birth cohorts during the 1970s. Moreover, we find that these missing girls are unlikely to be explained by systematic under-reporting of living children or instances of adoption.

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[9] We use the number of previous live births to define parity groups and birth orders.

[10] Coale and Banister (1994) show sex ratios by 5 year birth cohort among pooled first and second order births, and pooled third and fourth order births. The authors compare the sex ratios of births by within-sex birth order among babies born in 5 year birth cohorts, showing that girls are more likely to be missing when they have older sisters compared to boys when they have older brothers. However, the authors do not measure the absolute sex ratio among births in each year by birth order and sibship sex composition.

[11] Several studies using census counts show differential mortality rates over time or sex ratios at birth among pooled birth cohorts, neither approach permitting nuanced analysis of the age at which girls are missing from populations Cai and Lavely 2003; Coale and Banister 1994).

[12] Throughout this study, we characterize the sex composition of older siblings surviving at the time of the relevant birth. Results are robust to the inclusion of all previous births, including children who died earlier (available upon request).

This paper proceeds as follows. Section 2 provides a brief background on sex ratios over time in China and presents a stylized model of fertility decline and sex selection when there is son preference. Section 3 describes our data sources and methodology, and Section 4 presents our primary results. Section 5 then assesses data quality and investigates the possibility of differential under-reporting of female births, and Section 6 concludes.

2. Background and conceptual framework

2.1 Sex selection in China in recent history

Historical population research provides rich qualitative evidence of high sex ratios during China's Imperial period, extending into the first half of the 20^{th} century (King 2014; Mungello 2008; Wolf and Huang 1980). This work generally attributes the persistence of male-biased sex ratios to a strong preference for sons rooted in patriarchal traditions (Das Gupta and Li 1999; Ebenstein 2014; Ebenstein and Leung 2010; Greenhalgh 1988). Although there is evidence of advocacy against the practice (Mungello 2008), families of all social strata used female infanticide and abandonment to control family size and composition (Greenhalgh 1988; King 2014; Langer 1973; Lee and Wang 1999b; Mungello 2008). Some accounts suggest that 10% of female births may have ended this way, with rates as high as 40% reported among some subgroups (documented among Imperial families during specific periods, for example) (King 2014; Lee and Wang 1999a, 1999b).

The first population censuses conducted in the People's Republic of China (in 1953 and 1964) provide evidence that sex ratios were high during the early 20^{th} century, but then fell around the time of the 1949 communist revolution. The ratio of men to women born during the 1920s and 1930s appears to have ranged between 107.3 and 113.6, peaking during the 1940s at 112.7–117.7 in 1953 (Banister 1991).[13] Shortly after the communist revolution, China's sex ratio at birth then appears to have fallen to naturally occurring levels below 107 (Banister 1991).[14] Some research suggests this was due to efforts to promote gender equality and discourage 'feudal' attitudes toward

[13] Although it is not possible to distinguish if these high sex ratios emerged at birth or reflect differential mortality throughout childhood and early adulthood, qualitative records suggest that much of this imbalance may have emerged around the time of birth (Song 2012).

[14] The biologically 'natural' sex ratio at birth is generally believed to be 105–107 (Grech, Savona-Ventura, and Vassallo-Agius 2002; Johannson and Nygren 1991).

daughters, however son preference persisted as sons continued to ensure lineage continuation and provide greater economic security compared to daughters.[15]

From 'balanced' levels early in the communist era, existing studies then focus on the rapid resurgence of male-biased sex ratios beginning in 1980 under the One-Child Policy (and coincident with the diffusion of ultrasound technology across the country). A large body of research suggests that between 8.5 and 9.2 million girls are missing from cohorts born between 1980 and 2000 – generally at very young ages (Cai and Lavely 2003; Jiang et al. 2012). Specifically, in years 1990 and 1995, sex ratios at birth reached estimated levels of 111.8 and 116.6 (Banister 2004). Unlike earlier years, many studies argue that the diffusion of ultrasound technology (and the ability to detect fetal sex) during this period enabled families to selectively abort girls – effectively making sex selection 'easier' or less costly (Banister 2004; Cai and Lavely 2003; Chen, Li, and Meng 2013; Coale and Banister 1994).[16]

Little previous research reports evidence (or statistically significant evidence) of rising sex ratios during the 1970s – prior to the One-Child Policy and diffusion of ultrasound technology. Instead, many past studies are descriptive or focus on establishing 'baseline' population sex ratios to then quantify dramatic increases under the One-Child Policy (Banister 2004; Cai and Lavely 2003; Das Gupta 2005; Jiang et al. 2012; Johansson and Nygren 1991).[17] Nonetheless, modern China's most dramatic fertility decline occurred during this decade. Rural China's TFR fell by more than half, from approximately 6.4 in 1970 to about 3 in 1979 (Banister 2004; Cai and Lavely 2003; Coale and Banister 1994). Demographic and economic theory suggest that given preferences for sons, rapid fertility decline (whatever the cause – demand- or supply-driven) should be accompanied by rising sex ratios (Almond, Li, and Zhang 2017; Becker 1960, 1991; Das Gupta and Bhat 1997; Das Gupta and Li 1999; Jayachandran 2017; Jayachandran and Kuziemko 2010; Schultz 1985). Additionally, this increase in male-biased sex ratios should be concentrated among couples with the greatest demand for sons. Although individual and couple/household preferences are not observed, this group should disproportionately include higher parity couples without a surviving son.

[15] Despite the collective organization of agriculture, families remained patriarchal in nature. Property and lineage was passed through the male line and married couples lived with and cared for the husband's parents. Males were also awarded more work points and greater rations compared to female family members and offered greater opportunities for sociopolitical advancement through military or political careers (Arnold and Liu 1986; Ebenstein 2014; Ebenstein and Leung 2010; Greenhalgh 1988; Greenhalgh and Li 1993).

[16] Media archives and public policy statements discussing female infanticide in the early 1980s suggest that the practice was ongoing in some areas (Banister 1991; Greenhalgh and Winckler 2005; White 2000, 2009).

[17] Coale and Banister (1994) show sex ratios by 5 year birth cohort among pooled first and second order births, and pooled third and fourth order births. Further, the authors compare the sex ratios of births by within-sex birth order among babies born in 5 year birth cohorts, showing that girls are more likely to be missing when they have older sisters compared to boys when they have older brothers. However, the authors do not measure the absolute sex ratio among births in each year by birth order and sibship sex composition.

In this next section, we present a stylized conceptual framework for understanding decision-making about sex selection as fertility rates decline.

2.2 A simple conceptual framework for fertility decline and sex selection

In this section, we briefly summarize how fertility constraints theoretically affect sex selection behavior (as empirically documented by Das Gupta and Bhat (1997) and Guilmoto (2009), for example). As Das Gupta and Bhat (1997) note, the key necessary assumption is that when fertility declines, the desired family size falls more rapidly than the target number of sons (among couples with some form of son preference).[18] Put differently, the decline in the demand for daughters outpaces the decline in demand for sons, increasing sex selection pressure (Li, Feldman, and Tuljapurkar 2000). As fertility constraints tighten, the opportunity cost of having an unwanted girl then rises (because the probability of not having the target number of sons increases). Consequently, couples for whom the opportunity cost of a girl exceeds the cost of sex selection will sex select (prenatally or postnatally – in our case, because we only study province-years in which ultrasound technology was not available, postnatally) (Anukriti 2016; Ebenstein 2011; Lin, Liu, and Qian 2014). In the aggregate the consequences for sex ratios are unambiguous: an increase in the cost of children leads not only to smaller family sizes, but also greater neglect of daughters – and hence rising sex ratios. We provide a simple formal conceptualization of this decision in the online supplemental material.

3. Data and methods

3.1 Data

We use China's 1988 National Sample Survey of Fertility and Contraception (also known as the "two-per-thousand" fertility survey), which collected data from 417,518 women in 30 Chinese provinces and municipalities.[19] Conducted by the State Family

[18] Households are assumed to have a target number of children they would like to have taking the full cost of having and raising children into account – a target which is independent of their desire for sons. Importantly, we distinguish the "desired" or "target" number of children couples would like to have from standard demographic measures such as the 'ideal number of children,' 'desired total fertility,' and 'wanted total fertility.' The online supplemental material provides additional detail.

[19] Provinces and municipalities surveyed include: Anhui, Beijing, Fujian, Gansu, Guangdong, Guangxi, Guizhou, Hainan, Heilongjiang, Hebei, Henan, Hubei, Henan, Inner Mongolia, Jiangsu, Jiangxi, Jilin,

Planning Commission of China, the survey interviewed a representative sample of ever-married Chinese women ages 15–57 from approximately 14,000 sampling units (neighborhood small group for cities and village small group in rural areas) across the country. This survey was an expanded version of the 1982 National Sample Survey of Fertility and Contraception (the "one-per-thousand" fertility survey) and is generally believed to be accurate and good quality (Coale 1984; Coale and Banister 1994; Zhang and Zhao 2006).[20]

For the purpose of our study, the 1988 "two-per-thousand" fertility survey has several important advantages over its predecessor (the "one-per-thousand" fertility survey) as well as to China's national population censuses. The "one-per-thousand" survey also collected fertility histories, but this did not include information about infant and child deaths, preventing analysis of sex-specific mortality as children aged. China's 1982 and 1990 population censuses have been widely used to study population sex imbalance (Banister 2004; Cai and Lavely 2003; Das Gupta and Li 1999; Ebenstein 2014; Jiang et al. 2012), but they also lack information about child deaths – and importantly (given our focus), they do not contain sufficient information to estimate sex ratios by birth order, sex composition of previous births, and year. These features may be critical for understanding the early emergence of sex imbalance in the Chinese population (Muhuri and Preston 1991).

Each sampled woman in the 1988 "two-per-thousand" fertility survey provided detailed information about her complete fertility history, including all pregnancies ending in miscarriage, abortion, stillbirths, and live births. For each live birth, the survey collected the date of birth, sex of the child, whether or not the child was alive at the time of the survey, and if not, the date that the child died. The survey also recorded socioeconomic information about each mother, including her own date of birth, ethnicity, province of residence, and urban/rural residency status (*hukou*). Table 1 provides descriptive statistics of mothers having births in selected years during the period of study, and Appendix Table A-1 provides the sample size of birth records used.

Liaoning, Ningxia, Qinghai, Shaanxi, Shanghai, Shandong, Shanxi, Sichuan, Tianjin, Xinjiang, Yunnan, Zhejiang.

[20] The data used in our analyses is available upon request.

Table 1: Characteristics of mothers by year of delivery, 1967–1983

	Overall	1967	1970	1973	1976	1979	1982
Illiterate	54.7%	52.3%	46.1%	44.4%	41.3%	38.2%	30.7%
Semi-literate	11.1%	10.8%	10.4%	10.6%	11.1%	10.4%	8.6%
Primary education	23.9%	25.5%	29.0%	31.1%	31.8%	31.1%	26.2%
Secondary education or above	10.3%	11.4%	14.4%	13.9%	15.9%	20.4%	34.5%
Age at marriage	18.96	19.10	19.46	19.68	20.37	21.33	21.97
Maternal age	25.96	26.62	27.16	27.35	26.93	26.80	25.95
Rural household	87.6%	89.0%	88.9%	91.0%	91.8%	91.6%	89.0%

Source: State Family Planning Commission of China, 1988 Two-per-Thousand National Sample Survey of Fertility and Contraception.

3.2 Methods

3.2.1 Graphical analysis

We conduct both graphical and multivariate statistical analysis of population sex imbalance by year of birth, birth order, and sex composition of previous births (defined as the presence/absence of a living older male sibling at the time of the relevant birth – henceforth termed "sibship sex composition"). Our graphical analysis uses each birth, combined with birth records of all children born to the same mother, to determine both birth order and sibship sex composition.[21] We then calculate sex ratios (or the number of males for every 100 females) at birth and at each subsequent year of age up to age 5 by year of birth, birth order, and sibship sex composition between 1962 and 1987.[22] Finally, we plot these sex ratios (at birth and at each year of age up to 5) over time.

Given the lack of prenatal sex determination technology prior to the early 1980s (Chen, Li, and Meng 2013), we interpret sex ratios at birth above the naturally occurring level of 105–106 (Grech, Savona-Ventura, and Vassallo-Agius 2002; Johansson and Nygren 1991; Sen 1990) prior to the 1980s to reflect postnatal selection in favor of males – around the time of birth or in the first year of life, subject to reporting considerations examined in Section 5.

[21] As Footnote 11 describes, our results are insensitive to an alternate definition of sibship sex composition using indicators for whether or not any preceding birth was a boy, regardless of survival (results available on request).

[22] Appendix Table A-2 shows sex ratios at birth for each year, birth order, and sibship sex composition.

3.2.2 Multivariate statistical analysis

We also use Ordinary Least Squares (OLS) regression models to estimate the joint marginal relationship of birth year, birth order, and sibship sex composition with the marginal probability that a given birth is male (relative to 1967). Specifically, we estimate models of the following general form separately for birth order p births to mothers i in years y, grouping third- and higher-order births together:

$$Male_{ijyp} = \alpha + \beta_{ip}^{1}\,PriorSon_{ip} + \sum\nolimits_{y} \beta_{y}^{2} BirthYear_{y} \\ + \sum\nolimits_{y} \beta_{y}^{3} BirthYear_{y} \times PriorSon_{ip} + \sum\nolimits_{k} \beta_{k}^{4} X_{i}^{k} + \delta_{j} + \varepsilon_{ijyp}, \qquad (1)$$

where $male_{ijyp}$ is a dichotomous indicator variable for whether or not an order p child born to mother i living in province j in year y was male, $BirthYear_{y}$ is a dichotomous indicator for birth in year y, $PriorSon_{ip}$ is a dichotomous indicator for whether or not mother i had a previous son surviving to year y, and X_{i}^{k} is a$_{\text{vector}}$ of k maternal characteristics (indicators for residence in an urban area and mothers' educational attainment strata, as well as mother's age at marriage). In addition to main effects, we also include all two-way interactions between $BirthYear_{y}$ and $PriorSon_{ip}$. Equation 1 also includes provincial fixed effects (δ_{j}), which control for unobserved time-invariant differences across provinces. Appendix Tables A-3–A-4 show that our results are robust to the exclusion of urban and semi-urban populations.

We use linear probability models to allow for consistent fixed effects estimation while avoiding concerns about incidental parameters (Neyman and Scott 1948). Appendix Table A-5 reports results obtained using probit regressions, yielding comparable findings. We compute Huber–White robust standard errors clustered at the province level, relaxing the assumption that error terms are identical and independently distributed (i.i.d.) across provinces. The resulting estimates of $\beta_{y}^{1} - \beta_{yp}^{3}$ allow us to calculate the probability of being male in excess of biologically expected levels as a linear combination of year of birth, birth order, sibship sex composition, and all interactions among them. To draw conclusions about what these estimates imply about the prevalence of postnatal sex selection, we restrict our analyses to years before the introduction of ultrasound technology in each province (Chen, Li, and Meng 2013).[23]

[23] Data on ultrasound availability at the county-year level is from Chen and coauthors (2013), who digitize archival records of local histories as recorded in issues of the *Local Chronical* for over 1500 counties. In our analyses we adopt a conservative approach, restricting our sample to years before ultrasound technology became available in any county in each province. Including subsequent years increases the magnitude of our estimates.

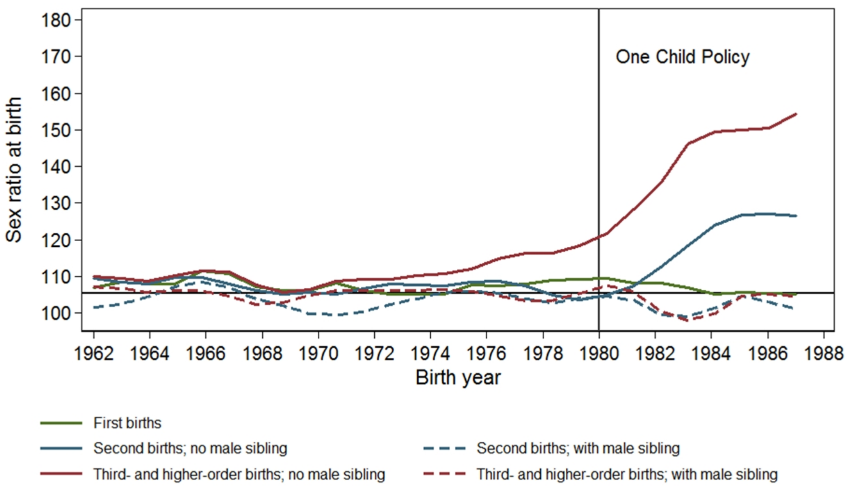
4. Results

4.1 Graphical analysis

4.1.1 Sex ratios at birth

Figure 2 plots sex ratios at birth by year, birth order, and sibship sex composition for each year between 1962 and 1987. At all parities, regardless of previous sons, sex ratios at birth were largely stable throughout the 1960s and early 1970s, oscillating around the natural rate. This is consistent with past analyses suggesting little sex imbalance at birth prior to the 1980s at this level of aggregation (Banister 2004; Coale and Banister 1994; Das Gupta and Li 1999; King 2014).

Figure 2: Sex ratio at birth by birth order and sibship sex composition in China 1962–1987



Note: Figure shows reported sex ratio at birth, by birth order and sibship sex composition. Sex ratio is calculated as the number of male births divided by the number of female births in each parity and sibship sex composition category.
Source: 1988 National Sample Survey of Fertility and Contraception.

However, consistent with theoretical predictions, as China's total fertility rate declined during the 1970s, the sex ratio of higher-order births depends critically on sibship sex composition.[24] Among mothers with at least one surviving son, this

[24] We conceptualize household fertility decision-making to depend on the sex composition of older surviving children (specifically, the presence of an older surviving son). In practice, all of our results are insensitive to

trajectory is also flat, remaining close to the natural rate throughout the 1970s. However, for mothers without sons, the sex ratio at birth rises rapidly throughout the decade – and does so earlier at higher parities. Among third- and higher-order births, the sex ratio among mothers without previous sons rises as high as 120.8 in 1977 (and 115.6 in 1979).[25] This increase in the probability of male births at higher parities among women without living sons then continues throughout the 1980s as fertility declines further, the One-Child Policy is introduced, and ultrasound technologies become available.

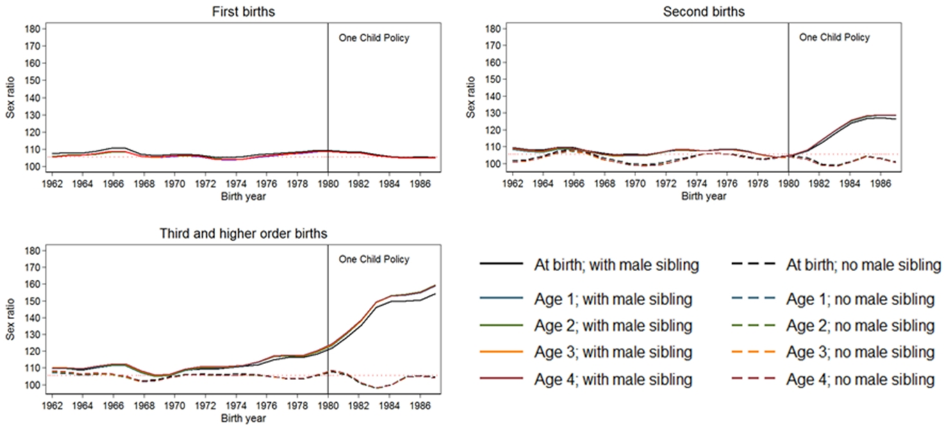
4.1.2 Sex ratios at ages 1–5

Figure 3 repeats the graphical analysis shown in Figure 2 for sex ratios at birth and each single-year age interval up to age 5. Births of each order are shown in separate panels. Relative to the sex ratio at birth within a given subgroup, sex ratios at subsequent ages up to age 5 generally change little. One exception is the modest increase at third- and higher-order births among households without a previous son. Within this subgroup, the sex ratio rises at age 1 by approximately 2–3 additional boys per 100 girls. The stability (and in some cases, increase) in age-specific sex ratios from birth up to age 5 contrasts with biologically higher rates of mortality among boys at all ages (Coale 1991). Overall, Figure 3 suggests that the majority of China's population sex imbalance during the 1970s occurred around the time of birth.

instead using the sex composition of previous births (specifically, the presence of a previous male birth), regardless of survival to the time of a fertility decision. These results are available upon request.

[25] Previous research reports average sex ratios over 5-year intervals at third and fourth parities, regardless of the sex composition of previous births, as high as 109 prior to the One-Child Policy (Coale and Banister 1994).

Figure 3: Sex ratio at age 1–4 by birth order and sibship sex composition in China 1962–1987



Note: Figure shows reported sex ratio at ages 1–4, by birth order and sibship sex composition. Sex ratio is calculates as the number of male children divided by the number of female children reaching each relative age in each birth order and sibship sex composition category.
Source: 1988 National Sample Survey of Fertility and Contraception.

4.2 Statistical analysis

The results of our statistical analysis (Table 2) are consistent with Figure 2.[26] We estimate the marginal probability that a child born is male by birth year, birth order, and sibship sex composition, calculating these probabilities using coefficient estimates obtained from Equation 1. Adjusting for observable maternal characteristics as well as provincial and birth year fixed effects, the probability that a first-born child is male does not rise during the 1970s relative to 1969 (the reference year, in which it is 51.5% – approximately the biologically expected rate). After a woman’s first child, sibship sex composition then becomes a key determinant of the probability of a male birth. With few exceptions, the probability that a child born to a woman with at least one older living son is statistically indistinguishable from the biologically expected probability throughout the 1970s. However, among women without surviving sons, a third- or higher-order child born after 1973 is 2–5 percentage points more likely to be a male (implying sex ratios at birth of 113–129).[27]

[26] Because we restrict our statistical analysis to province-year observations in which ultrasound technology was not available, Table 2 shows results for years 1965–1982.
[27] Estimates are statistically significant for years 1974–1981.

Table 2: Marginal probability of a male birth, conditional on live birth

Birth year	First birth: With surviving son	First birth: Without surviving son	Second birth: With surviving son	Second birth: Without surviving son	Third birth: With surviving son	Third birth: Without surviving son
1965		–0.56%	0.42%	1.15%	0.03%	0.59%
1966		1.93% **	1.37%	1.25%	1.32%*	2.87%***
1967				Reference year		
1968		–0.41%	–0.48%	–0.79%	–0.67%	0.78%
1969		–0.64%	–0.90%	0.92%	0.01%	0.03%
1970		–0.30%	–1.16%	–0.20%	0.66%	0.91%
1971		0.35%	–2.21%**	0.31%	0.33%	1.86%
1972		–1.09%	2.38%	0.68%	1.15%*	1.23%
1973		–0.21%	–1.57%	1.39%	0.16%	1.29%
1974		–1.09%	1.13%	0.27%	1.17%*	2.12%*
1975		–0.45%	–0.11%	0.79%	0.55%	1.69%**
1976		0.30%	0.82%	2.02%	0.77%	1.93%*
1977		–0.56%	1.49%	1.07%	0.64%	5.00%***
1978		0.90%	–1.18%	–0.21%	0.61%	2.68%**
1979		–0.39%	1.15%	0.33%	1.69%*	3.57%**
1980		0.95%	3.96%**	–0.84%	1.71%	5.11%***
1981		–0.46%	0.39%	2.47%	2.43%	4.84%**
1982		–0.35%	4.53%**	4.24%	4.16%	6.88%

Note: Each cell contains marginal probability that a birth occurring in each year, birth order, and sibling sex composition category is male. Marginal probabilities are calculated from coefficients estimated using OLS regressions (estimated separately for births of each order) of an indicator for a male birth on indicators for birth year and the sex composition of older siblings, as well as all two - way interactions between birth year and sibship sex composition. We control for residence in urban area, mother's educational attainment strata, mother's age at marriage, and time-invariant province fixed effects. Huber–White robust standard errors are clustered at the province level. * $p<0.10$, ** $p<0.05$, *** $p<0.001$.
Source: 1988 National Sample Survey of Fertility and Contraception.

5. Data quality and possible under-reporting of living girls

Because we consider deviations from the naturally occurring sex ratio at birth prior to the introduction of prenatal sex detection technology to reflect postnatal sex selection, a note about potential under-reporting of female births is warranted. Without the technological ability to identify and selectively abort female fetuses, high sex ratios at birth reflect either under-reporting of children born alive that died early in life or under-enumeration of children living at the time of the survey. The interpretation of these unreported female births is central to our paper. We interpret these ‘missing girls’ to reflect differential rates of infant/child female death, but if the majority of unreported girls were living but simply uncounted, then our interpretation would be incorrect.

Both true sex selection and under-enumeration of living children is well documented in the demography literature for Chinese birth cohorts from the 1980s and 1990s (Cai and Lavely 2003; Goodkind 2011; Merli and Raftery 2000; Zeng 1996; Zeng et al. 1993; Zhang and Zhao 2006). However, there is little existing evidence for birth cohorts from the 1970s. On one hand, penalties for violating fertility regulations were less severe prior to the One-Child Policy – and hence incentives for hiding unsanctioned births from enumerators were weaker. On the other hand, however, infant deaths may have been unreported in official registries at higher rates in earlier years for simple administrative reasons related to bureaucratic inefficiency (Coale and Banister 1994; Merli 1998). Some suggest that on balance, relative to later birth cohorts, the degree of under-enumeration of living children born during the 1970s was substantially less (Coale 1984; Coale and Banister 1994; Zeng et al. 1993). To the best of our knowledge, however, no study has empirically evaluated the degree of under-reporting of births from the 1970s in the 1988 "two-per-thousand" survey – including under-reporting by birth order and under-reporting of girls relative to boys.

We use several methods to investigate the extent to which unreported girls lived beyond infancy as unregistered and unenumerated children, which we present below.

5.1 Adoption and survey design

Before applying established demographic methods for assessing under-reporting of living girls, we first briefly consider how the design of the "two-per-thousand" survey (and enumerator instructions) handles adoption – a specific potential form under-reporting.[28] Survey enumerators were instructed to ensure that adopted children ("adopted-in") were not listed in pregnancy histories as 'own children' – and also to ensure that children given up for adoption ("adopted-out") were included in these histories. To accomplish this, the survey included cross-validation measures designed to explicitly handle adoptions in this way (SFPC 1988).[29] Although we are of course unable to verify how enumerators conducted fieldwork in practice, systematic under-

[28] If boys adopted into families are reported in our survey's fertility records, or if girls who were given up for adoption are not reported, then the sex ratios that we compute would be inflated.

[29] Specifically, before asking questions about each pregnancy, enumerators were instructed to ask how many "own children" were currently living with the respondent, how many were not living with the respondent, how many had been given up for adoption, and how many had died. Summing across these answers, enumerators were then to calculate the number of pregnancies resulting in live birth – including children "adopted-out." If this cross-validation exercise yielded discrepancies, "interviewer should probe for omissions, twins, and multiple births, or to see if adoptive children were listed as own children, etc." (SFPC 1988)

reporting of children adopted-out (along with other types of under-reporting) would be captured by our analyses below in Section 5.2.

5.2 Empirical assessment of under-reporting

We then test empirically for systematic under-reporting of living children who could have been adopted-out, or otherwise hidden from enumerators, using three approaches. The first two modify methods used to evaluate the quality of the 1982 "one-per-thousand" national fertility survey, and the third compares the 1988 "two-per-thousand" national fertility survey with the 1982 survey (which is generally considered good quality) (Banister 2004; Ní Bhrolcháin and Dyson 2007; Coale 1991; Coale and Banister 1994).

First, following Coale and Banister (1994), we investigate the extent to which possibly unreported female births in the 1988 "two-per-thousand" survey 'reappear' as adult women in China's population censuses, focusing on those births most likely to be underreported. We compare sex ratios at birth (number of male births for each 100 female births) for each birth cohort reported in the 1988 fertility survey with sex ratios for the same birth cohorts as reflected in the 1% micro samples of the 1982 and 1990 censuses. From cross-sectional census microsamples, we reconstruct sex ratios at birth by adjusting population counts for age- and sex-specific mortality rates, using a reverse survival method.[30] We find that sex ratios at birth in the 1988 fertility survey are consistent with mortality-adjusted sex ratios observed among the same birth cohorts in both the 1982 and 1990 population censuses (Appendix Figures A-1–A-2 and Appendix Table A-6).

To the extent possible, we also investigate the degree to which higher birth order girls (who may have been alive but disproportionately under-reported in fertility histories) are more likely to appear in later population censuses than higher birth order boys. Specifically, we use the same approach described above, stratifying by both birth order sibship sex composition.[31] Due to data requirements, we focus on birth cohorts

[30] Mortality rates are derived from three sets of life tables. These are: 1) life tables presented in Coale (1984), which interpolate between the 1964 and 1982 censuses; 2) life tables published in Banister (1991), which use China's Cancer Epidemiology Study of deaths between 1973–1975; and 3) life tables based directly on the 1982 population census (Jiang, Zhang, and Zhu 1984). For all mortality rate adjustments, we necessarily assume that age- and sex-specific morality rates were stable over the period of study.

[31] Because birth order is not directly reported in the population censuses, we reconstruct birth order and sibship sex composition for the subset of individuals still living with their parents in census years. Specifically, we use the total number of boys and girls ever born to a parent together with the sex and age of each child reported in the household roster, restricting our sample to households in which all children born to a mother still coresided with her at the time of the census (i.e., children for whom we know birth order and sex composition of older siblings with certainty). To maintain cells of adequate size (by birth year, birth

born between 1975 and 1979 in the 1990 census, adjusting for mortality using birth order-, age-, and sex- specific mortality rates derived from the 1988 "two-per-thousand" survey.[32] We find that girls born at higher parities and with no older brothers – precisely the circumstances under which sex selection is predicted to be strongest – are not more likely to reappear as adults in future censuses (i.e., we find no evidence of differential under-reporting by birth order and sex composition of previous births) (Appendix Figure A-3 and Appendix Table A-7).

Second, following Coale (1991), we use the 1988 "two-per-thousand" survey to calculate the age-specific rate at which women deliver male and female babies in each year. We then apply these fertility rates by maternal age and child sex (simultaneously) to age-specific population counts of women reported in population census microsamples (interpolated between the 1964 and 1982 censuses), yielding an estimate of the total number of boys and girls born in each calendar year. We then compare the estimated number of male and female births implied by these calculations to the actual number of individuals in each birth cohort in the 1982 and 1990 censuses to estimate the degree of underreporting for boys and girls by birth cohort in the 1988 fertility survey. We find that although females are slightly more likely to be unreported than males, the difference in rates remains relatively constant over time – and in fact decreases during the late 1970s (Appendix Figure A-4). For underreporting of surviving females to confound our main estimates of missing girls, they would need to increase relative to underreporting of males over time.

Third, we investigate the consistency of the 1988 "two-per-thousand" survey with its predecessor, the 1982 "one-per-thousand" survey (which others have shown to be good quality (Coale 1984)). To do so, we account for demographic changes between survey years by creating a matched sample of women across surveys. Specifically, for every woman in the "one-per-thousand" survey, we identify a woman in the "two-per-thousand" survey with the same characteristics,[33] pooling matched observations from both surveys together. We then regress, separately, (1) the reported number of children (male and female combined), (2) the reported number of male children, and (3) the reported number of female children on a dichotomous indicator variable for which survey the observation was drawn from. The results imply that the number of children (male and female combined) recorded in the 1988 "two-per-thousand" survey is 0.026

order, and sibship sex composition), we focus on birth cohorts from the second half of the 1970s (1975–1980).

32 Mortality rates from the 1988 "two-per-thousand" survey are shown to be consistent with life-table sources in Appendix Figures A-1–A-2 and Appendix Table A-6.

33 Individuals in each survey are matched using five individual-level characteristics: birth cohort, urban/rural residence, province of residence, educational attainment, and ethnicity. This means that the pooled data set includes an equal number of women from the "one-per-thousand" and "two-per-thousand" surveys, with each observation from the "one-per-thousand" survey matched to an observation in the "two-per-thousand" survey sharing exactly the same characteristics along these five dimensions.

fewer than in the 1982 survey [95% CI: 0.014 – 0.037]. Analogous estimates by sex imply 0.0164 fewer male births [95% CI: 0.010 – 0.023] and 0.009 fewer female births [95% CI: 0.003 – 0.016] in the 1988 survey. Overall, these results suggest a small degree of underreporting in the 1988 "two-per-thousand" survey relative to the 1982 survey. However, because underreporting is, to a small extent, more severe for male births than for female births, the implication is that our estimates of sex ratios at birth during the 1970s may be biased downwards.

6. Discussion

Guided by theory and previous literature, this paper uses fertility history data and a combination of graphical and statistical methods to investigate the possibility of previously undocumented sex imbalance at birth during China's sharp fertility decline throughout the 1970s. Analyzing sex ratios at birth by both birth order and sibship sex composition (simultaneously), we find that among couples predicted to have the greatest demand for sons (third and higher parity births to couples not yet having a boy), sex ratios rose as high as 115–121 boys per 100 girls during the 1970s. Importantly, this finding does not appear to be fully explained by under-reporting of female births or the adoption of girls. It also cannot be explained by prenatal sex selection because we study only births prior to the introduction of ultrasound technology.

Infanticide and postnatal neglect has a long and well-documented history in many cultures, including China during the late Imperial period and during times of political and economic instability in the 20th century (Das Gupta and Li 1999; Greenhalgh and Winckler 2005; King 2014; Langer 1973; Lee and Wang 1999b; Mungello 2008; White 2009; Wolf and Huang 1980). This paper provides quantitative measurements of the practice during the 1970s among a small, but quantitatively important, subset of couples.[34]

Overall, our statistical analysis implies that, although rare (accounting for about 0.5% of births) and concentrated among higher-order births to families without surviving sons, approximately 1.1 million girls are missing from Chinese cohorts born during the 1970s.[35] Our graphical analysis using unadjusted data (integrating under the curves shown for sex ratios at birth by birth order and sibship sex composition) yields a comparable estimate of 844,000 missing girls. Importantly, because these male-biased

[34] This is consistent with ethnographic reports of incidents of infanticide in the early 1980s (Greenhalgh and Winkler 2005; White 2006).
[35] See online supplemental material for detailed calculations of the estimated number of missing girls during the 1970s.

sex ratios at birth emerge prior to the introduction of ultrasound (Chen, Li, and Meng 2013), they cannot be explained by sex-selective abortion – and instead suggest disproportionately high female mortality rates shortly after birth, either through neglect, or in the extreme, infanticide. Differentially early discontinuation of breastfeeding after the birth of a girl among mothers without a son (and hoping to conceive again more quickly, for example) may explain only about 20–30% of these missing girls (Jayachandran and Kuziemko 2010).[36] More clearly than previous research, our results suggest a serious issue warranting further investigation.

More generally, given the coincidence between the Later, Longer, Fewer (LLF) birth control policy (the first large-scale population policy in China and predecessor to the One-Child Policy) and the rise in sex imbalance that we document, an important area for further research is the direct link between the two – and the unintended consequences of population policy generally. This concern is natural given the growing body of research establishing the relationship between fertility decline and population sex imbalance when there is an underlying preference for sons (Das Gupta and Bhat 1997; Jayachandran 2017; Li, Feldman, and Tuljapurkar 2000). Although family planning programs have not been shown to be the dominant factor responsible for global fertility decline (Babiarz and Miller 2016), both the intended and unintended consequences of many large-scale programs such as the LLF policy have not been studied quantitatively.

[36] Early discontinuation of breastfeeding has been reported to lead to higher infant mortality rates, with odds ratios ranging between 1.8 to 5.8, depending on when breastfeeding ended (WHO 2000). We find that mothers without previous sons are 3.3–8.5 percentage points more likely to discontinue breastfeeding daughters relative to sons (depending on parity). Accounting for differentially early weaning of girls may account for 20–30% of girls missing in the first year of life. Detailed calculations are available upon request.

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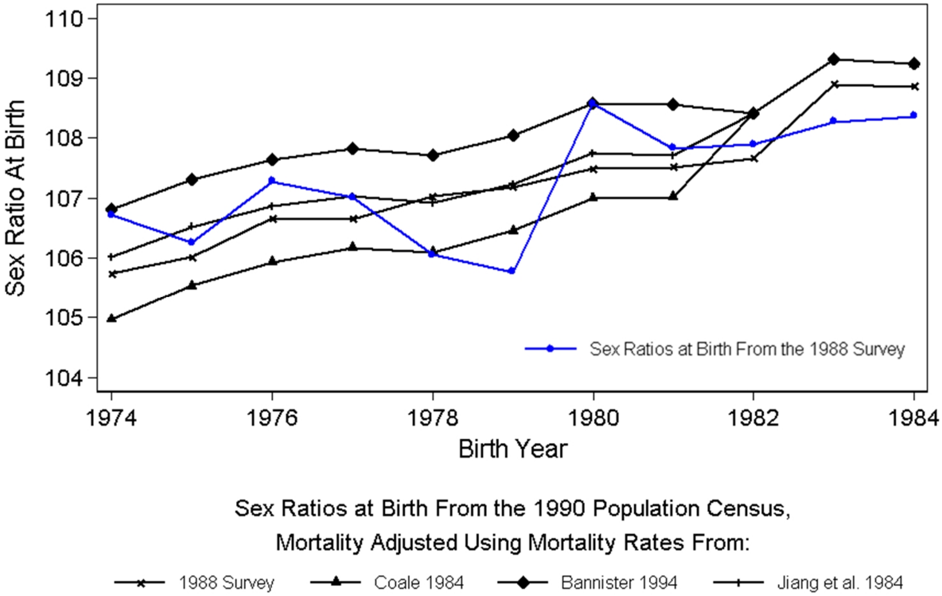
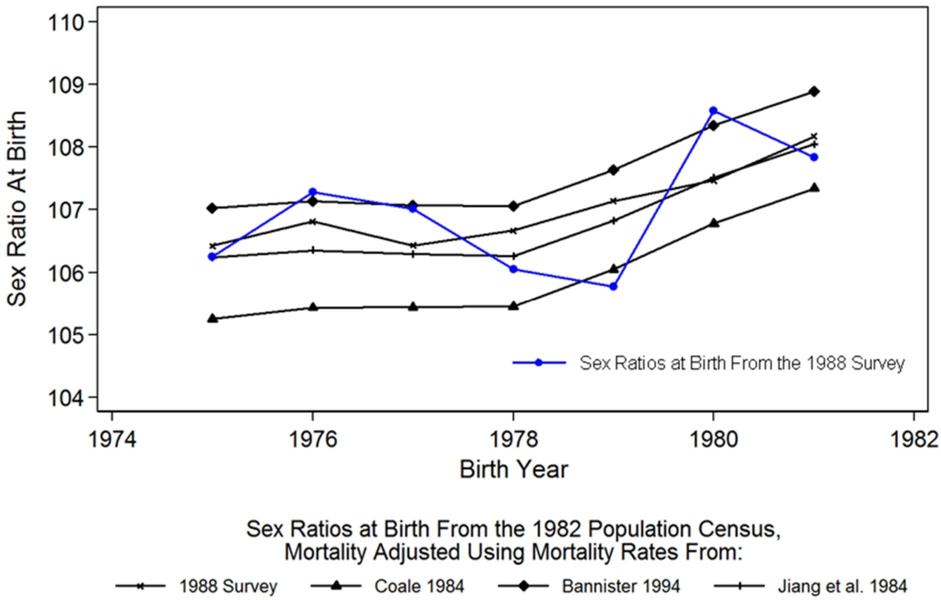
Appendix 1: Analysis of possible underreporting

Figure A-1: Sex ratio at birth calculated from 1988 Fertility Survey and implied by the 1990 Population Census



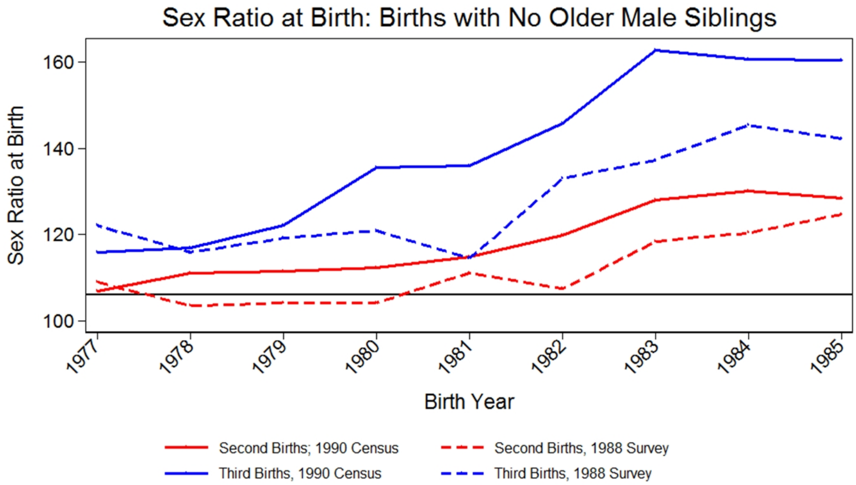
Note: Figure A-1 shows sex ratios calculated using the 1988 “two-per-thousand” survey and the 1% sample of the 1990 population census. Population census data adjusted for age- and sex-specific mortality rates using ‘reverse survival’ using mortality rates derived from four sources. These are: 1) mortality rates calculated using the deaths reported in the 1988 “two-per-thousand” survey, 2) life tables presented in Coale (1984), which interpolate between the 1964 and 1982 censuses; 2) life tables published in Banister (1994), which use China’s Cancer Epidemiology Study of deaths between 1973–1975; and life tables based directly on the 1982 population census (Jiang, Zhang, and Zhu 1984). For all mortality rate adjustments using life tables, we necessarily assume that age- and sex-specific mortality were stable over the period of study.

Figure A-2: Sex ratio at birth calculated from 1988 Fertility Survey and implied by the 1982 Population Census



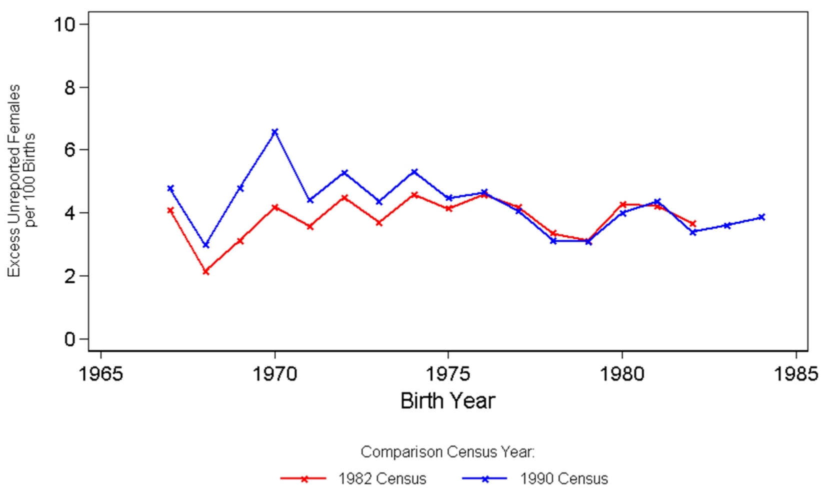
Note: Figure A-2 shows sex ratios calculated using the 1988 “two-per-thousand” survey and the 1982 population census. Population census data adjusted for age- and sex-specific mortality rates using ‘reverse survival’ using mortality rates derived from four sources. These are: 1) mortality rates calculated using the deaths reported in the 1988 “two-per-thousand” survey, 2) life tables presented in Coale (1984), which interpolate between the 1964 and 1982 censuses; 2) life tables published in Banister (1994), which use China’s Cancer Epidemiology Study of deaths between 1973–1975; and life tables based directly on the 1982 population census (Jiang, Zhang, and Zhu 1984). For all mortality rate adjustments using life tables, we necessarily assume that age- and sex-specific mortality were stable over the period of study.

Figure A-3: Sex ratio at birth by birth order calculated from the 1988 Fertility Survey and implied by 1990 Census Results



Note: Figure A-3 shows parity- specific sex ratios among children born to parents without a surviving male child, calculated using the 1988 “two-per-thousand” survey and the 1% sample of the 1990 population census. Population census data adjusted for age- and sex-specific mortality rates using ‘reverse survival’ using mortality rates derived from the 1988 “two-per-thousand” survey.

Figure A-4: Differential underreporting of female vs male births by birth year



Note: Figure A-4 shows the number of male and female births implied by birth rates reflected in the 1988 “two-per-thousand” survey and the number of individuals in each birth cohort enumerated in the 1982 and 1990 population censuses.

Table A-1: Sample size of birth records by birth order and sibship sex composition

Year	Birth order 1	Birth order 2		Birth order 3		Birth order 4–6	
		With son	With no son	With son	With no son	With son	With no son
1955	4645	1296	1821	767	509	273	125
1956	4827	1476	1998	1122	665	450	175
1957	5498	1910	2463	1639	902	955	310
1958	5260	1892	2309	1718	956	1293	395
1959	3845	1561	2017	1697	855	1537	438
1960	3797	1660	2090	1959	997	1961	590
1961	3660	1442	1772	1704	907	2073	510
1962	7896	2532	3087	3091	1584	4091	964
1963	10337	2983	3676	3767	1734	5756	1181
1964	8195	3177	3896	3269	1515	5622	1171
1965	6890	3604	4186	3579	1690	6366	1237
1966	6750	3871	4254	4183	1989	7174	1309
1967	6177	3196	3503	4061	1887	6726	1311
1968	8636	3825	4079	5490	2213	9481	1592
1969	8890	3259	3481	4960	1936	9333	1478
1970	9554	3719	3834	4732	1785	10088	1631
1971	9313	4057	4290	4614	1818	10471	1591
1972	8473	4122	4484	4585	1864	10087	1559
1973	8666	4321	4583	4637	1907	9001	1505
1974	9403	4052	4365	4517	2113	7842	1443
1975	8799	3962	4262	4256	1939	6663	1381
1976	9577	3966	4315	4040	1948	5694	1273
1977	9514	3803	4074	3693	1705	4617	1184
1978	10787	3901	4187	3586	1821	4389	1280
1979	12051	4212	4641	3690	1881	4118	1202
1980	12551	3615	4179	2379	1528	2523	919
1981	15827	3628	4452	2413	1513	2609	963
1982	18515	4057	4722	2593	1719	2527	920
1983	16592	3428	4417	2116	1455	1746	806
1984	16413	3731	4585	2091	1569	1583	811
1985	15899	4037	4724	1782	1423	1218	608
1986	18109	4890	5633	2129	1531	1287	765
1987	20970	5475	6467	2438	1767	1383	732
1988	8464	2067	2644	954	656	550	287

Source: 1988 “two-per-thousand” survey.

Table A-2: Unadjusted sex ratios by birth order and sibship sex composition

Year	Birth order 1	Birth order 2		Birth order 3		Birth order 4–6	
		With son	With no son	With son	With no son	With son	With no son
1955	107	110	99	120	109	101	105
1956	111	107	105	112	107	114	88
1957	104	104	114	97	119	101	85
1958	105	104	102	110	107	104	106
1959	103	93	101	104	101	105	112
1960	110	114	101	111	121	106	102
1961	107	104	104	110	113	109	102
1962	105	101	113	110	105	106	123
1963	110	102	104	109	111	105	104
1964	108	104	110	102	114	110	105
1965	106	107	110	102	108	105	108
1966	116	111	110	110	111	108	119
1967	108	106	110	100	113	107	108
1968	107	104	102	102	101	99	117
1969	106	101	109	102	104	104	103
1970	106	100	104	102	103	108	110
1971	110	96	106	104	113	106	110
1972	103	106	107	107	110	108	106
1973	107	98	111	99	103	107	117
1974	104	109	106	106	108	109	118
1975	107	105	107	106	107	106	114
1976	109	106	111	105	115	107	105
1977	106	106	108	104	121	104	120
1978	110	101	106	103	115	100	113
1979	108	102	102	101	118	108	112
1980	111	108	103	112	121	104	128
1981	108	102	109	106	121	114	122
1982	108	103	109	96	130	103	158
1983	109	94	120	93	140	98	142
1984	103	103	124	93	151	113	173
1985	107	105	128	99	142	108	154
1986	105	105	128	111	146	109	155
1987	105	98	125	98	158	105	159
1988	102	101	122	100	123	105	168

Source: 1988 “two-per-thousand” survey.

Table A-3: Marginal probability of a male birth, conditional on live birth; Rural and semi-rural subsample

Birth year	First birth		Second birth		Third birth	
	With surviving son	Without surviving son	With surviving son	Without surviving son	With surviving son	Without surviving son
1965		–0.22%	0.32%	1.05%	0.21%	0.02%
1966		2.39% **	2.27% **	1.02%	1.50% **	2.76% ***
1967			Reference year			
1968		–0.58%	–0.27%	–0.94%	–0.65%	0.65%
1969		–0.21%	–0.48%	0.88%	0.21%	–0.22%
1970		0.37%	–1.23%	0.08%	0.86%	0.96%
1971		0.99%	–1.91%	0.49%	0.48%	2.09% *
1972		–0.76%	12.29%	1.11%	1.24% *	1.60%
1973		0.23%	–1.70%	1.58%	0.45%	1.33%
1974		–0.51%	1.26%	0.52%	1.33% *	2.31% **
1975		0.25%	–0.13%	0.75%	0.72%	1.76% **
1976		1.59%	1.40%	2.70% *	0.95%	2.04%
1977		0.57%	2.68% *	1.51%	0.79%	4.94% ***
1978		1.80%	–1.37%	0.46%	0.76%	3.10% **
1979		0.45%	1.48%	42.36%	1.80% *	3.58% **
1980		1.82%	4.42% **	–0.45%	1.49%	5.19% ***
1981		1.28%	1.16%	3.27%	2.44%	4.15% **
1982		1.18%	4.52% ***	3.44%	0.01%	6.69%

Note: Each cell contains marginal probability that a birth occurring in each year, birth order, and sibling sex composition category is male. Marginal probabilities are calculated from coefficients estimated using OLS linear regressions (estimated separately for births of each order) of an indicator for a male birth on indicators for birth year and the sex composition of older siblings, as well as all two - way interactions between birth year and sibship sex composition. We control for residence in urban area, mother's educational attainment strata, mother's age at marriage, and time-invariant province fixed effects. Huber–White robust standard errors are clustered at the province level. * p<0.10, ** p<0.05, *** p<0.001.
Source: 1988 "two-per-thousand" survey.

Table A-4: Marginal probability of a male birth, conditional and live birth; rural subsample

Birth year	First birth		Second birth		Third birth	
	With surviving son	Without surviving son	With surviving son	Without surviving son	With surviving son	Without surviving son
1965		–0.98%	0.69%	0.54%	0.04%	–0.04%
1966		2.28% **	3.11% **	2.20%	1.79% *	2.53% **
1967			Reference year			
1968		–1.44%	–0.26%	–0.81%	–0.88%	1.33%
1969		–0.40%	0.60%	0.68%	0.36%	–0.34%
1970		0.32%	–1.39%	–0.19%	0.61%	1.76% *
1971		0.31%	–1.62%	1.06%	0.47%	2.44%
1972		–1.22%	–0.81%	2.14%	0.90%	1.59%
1973		–0.66%	–1.93%	1.85%	0.39%	1.11%
1974		–0.99%	1.27%	0.24%	1.39% *	3.48% **
1975		–1.02%	–0.01%	0.99%	0.89%	2.61% ***
1976		1.01%	0.73%	3.65% **	1.36%	2.73% *
1977		–0.77%	1.96% *	1.90%	1.16%	5.92% ***
1978		0.80%	–2.67%	0.90%	0.65%	4.73% ***
1979		–0.43%	1.46%	0.47%	1.89% *	5.55% **
1980		0.88%	4.80% **	1.24%	0.54%	4.99% ***
1981		–0.21%	–0.61%	3.41%	2.88%	5.38% **
1982		–0.15%	–1.06%	7.44% ***	4.76% **	7.37%

Note: Each cell contains marginal probability that a birth occurring in each year, birth order, and sibling sex composition category is male. Marginal probabilities are calculated from coefficients estimated using OLS linear regressions (estimated separately for births of each order) of an indicator for a male birth on indicators for birth year and the sex composition of older siblings, as well as all two - way interactions between birth year and sibship sex composition. We control for residence in urban area, mother's educational attainment strata, mother's age at marriage, and time-invariant province fixed effects. Huber–White robust standard errors are clustered at the province level. * p<0.10, ** p<0.05, *** p<0.001.
Source: 1988 "two-per-thousand" survey.

Table A-5: Marginal probability of a male birth relative among third parity births with no previously born son

	Full sample		Rural + semirural subsample		Rural subsample	
	LPM	**Probit**	**LPM**	**Probit**	**LPM**	**Probit**
No Son x Parity 3 x Year = 1965	–0.01386	–0.03480	–0.01820	–0.04567	–0.01257	–0.03153
	(0.013)	(0.033)	(0.014)	(0.035)	(0.015)	(0.037)
No Son x Parity 3 x Year = 1966	–0.00403	–0.01008	–0.00377	–0.00943	–0.00430	–0.01077
	(0.015)	(0.037)	(0.015)	(0.039)	(0.019)	(0.047)
No Son x Parity 3 x Year = 1967			Reference year			
No Son x Parity 3 x Year = 1968	–0.00493	–0.01241	–0.00330	–0.00829	0.01045	0.02620
	(0.014)	(0.036)	(0.015)	(0.037)	(0.016)	(0.040)
No Son x Parity 3 x Year = 1969	–0.01924	–0.04830	–0.02065	–0.05182	–0.01873	–0.04697
	(0.019)	(0.046)	(0.018)	(0.046)	(0.018)	(0.045)
No Son x Parity 3 x Year = 1970	–0.01699	–0.04266	–0.01533	–0.03847	–0.00023	–0.00058
	(0.014)	(0.035)	(0.016)	(0.039)	(0.018)	(0.044)
No Son x Parity 3 x Year = 1971	–0.00416	–0.01043	–0.00023	–0.00057	0.00804	0.02022
	(0.018)	(0.045)	(0.018)	(0.044)	(0.020)	(0.050)
No Son x Parity 3 x Year = 1972	–0.01873	–0.04700	–0.01272	–0.03190	–0.00474	–0.01186
	(0.016)	(0.041)	(0.017)	(0.041)	(0.017)	(0.044)
No Son x Parity 3 x Year = 1973	–0.00821	–0.02064	–0.00752	–0.01887	–0.00456	–0.01145
	(0.015)	(0.037)	(0.015)	(0.038)	(0.017)	(0.044)
No Son x Parity 3 x Year = 1974	–0.01006	–0.02524	–0.00649	–0.01628	0.00915	0.02304
	(0.016)	(0.041)	(0.016)	(0.040)	(0.017)	(0.044)
No Son x Parity 3 x Year = 1975	–0.00817	–0.02050	–0.00596	–0.01495	0.00548	0.01379
	(0.016)	(0.041)	(0.017)	(0.044)	(0.018)	(0.045)
No Son x Parity 3 x Year = 1976	–0.00787	–0.01977	–0.00544	–0.01365	0.00193	0.00487
	(0.011)	(0.027)	(0.011)	(0.027)	(0.016)	(0.039)
No Son x Parity 3 x Year = 1977	0.02415	0.06092	0.02518	0.06351	0.03592*	0.09060*
	(0.017)	(0.042)	(0.018)	(0.045)	(0.019)	(0.048)
No Son x Parity 3 x Year = 1978	0.00120	0.00302	0.00713	0.01796	0.02911	0.07328
	(0.021)	(0.052)	(0.021)	(0.053)	(0.018)	(0.045)
No Son x Parity 3 x Year = 1979	–0.00076	–0.00183	0.00144	0.00369	0.02485	0.06267
	(0.018)	(0.044)	(0.018)	(0.045)	(0.029)	(0.073)
No Son x Parity 3 x Year = 1980	0.01450	0.03666	0.02068	0.05218	0.03279	0.08248
	(0.024)	(0.061)	(0.025)	(0.062)	(0.025)	(0.064)
No Son x Parity 3 x Year = 1981	0.00460	0.01172	0.00081	0.00213	0.01332	0.03366
	(0.024)	(0.061)	(0.025)	(0.063)	(0.030)	(0.075)
No Son x Parity 3 x Year = 1982	0.04515*	0.11397*	0.05045*	0.12734*	0.01435	0.03679
	(0.025)	(0.063)	(0.026)	(0.065)	(0.054)	(0.138)

Note: Each cell contains marginal probability that a third- or higher-order birth occurring in each year to a couple with no previously born sons is male, relative to comparable births in the reference year 1969. Marginal probabilities are coefficients estimated using OLS linear regression or probit regression (as specified in column headers) of an indicator for birth year and the sex composition of older siblings, as well as all two-way interactions between birth year and sibship sex composition, estimated among third births. We control for residence in urban area, mother's educational attainment strata, mother's age at marriage, and time-invariant province fixed effects. Huber–White robust standard errors are clustered at the province level. * p<0.10, ** p<0.05, *** p<0.001.
Source: 1988 "two-per-thousand" survey.

Table A-6: Sex ratios at birth for all births, 1988 "Two-per-thousand" survey, 1990

Year	1988 "Two-per-thousand" survey	1990 Census				
		Not adjusted for mortality	Adjusted for mortality using:			
			1988 Survey	Coale 1984	Banister 1994	Jiang et al. 1984
1974	106.72	105.24	105.75	104.97	108.07	106.02
1975	106.25	105.74	106.03	105.54	108.67	106.53
1976	107.28	106.19	106.66	105.94	109.11	106.86
1977	107.02	106.49	106.66	106.17	109.37	107.03
1978	106.05	106.45	107.03	106.10	109.29	106.92
1979	105.77	106.84	107.19	106.45	109.64	107.23
1980	108.58	107.43	107.49	107.00	110.19	107.74
1981	107.83	107.37	107.52	107.02	110.19	107.71
1982	107.89	107.77	107.67	108.42	108.42	108.42
1983	108.28	108.69	108.90	109.31	109.31	109.31
1984	108.37	108.64	108.86	109.24	109.24	109.24
1985	110.87	108.65	108.88	109.22	109.22	109.22
Year	**1988 "Two-per-thousand" survey**	**1982 Census**				
		Not adjusted for mortality	**Adjusted for mortality using:**			
			1988 Survey	**Coale 1984**	**Banister 1994**	**Jiang et al. 1984**
1974	106.72	106.12	106.49	99.85	98.67	102.34
1975	106.25	106.18	106.43	105.26	107.03	106.24
1976	107.28	106.31	106.81	105.44	107.14	106.36
1977	107.02	106.27	106.43	105.44	107.08	106.29
1978	106.05	106.19	106.67	105.45	107.06	106.26
1979	105.77	106.71	107.14	106.05	107.64	106.83
1980	108.58	107.35	107.47	106.78	108.35	107.52
1981	107.83	107.83	108.17	107.34	108.89	108.04

Source: 1988 "two-per-thousand" survey. Mortality rates drawn from 1988 "two-per-thousand" survey, and life tables published in by Coale (1984), Banister (1994), and Jiang et al. (1984).

Table A-7: Sex ratios at birth for second and third order births with no older male sibling

	Second births with no older male sibling			Third births with no older male sibling		
	1988 "two-per-thousand" survey	1990 Census		1988 "two-per-thousand" survey	1990 Census	
Year		Not adjusted for mortality	Using survey-based mortality		Not adjusted for mortality	Using survey-based mortality
1977	107.54	106.82	105.69	120.87	116.00	114.21
1978	106.46	111.11	112.11	114.01	116.94	117.24
1979	102.31	111.41	111.10	115.59	122.17	120.45
1980	102.96	112.26	113.84	123.67	135.63	134.84
1981	108.62	114.92	114.46	121.67	135.93	131.88
1982	108.66	119.92	117.99	139.26	145.91	142.26
1983	120.19	128.13	125.21	140.53	162.85	154.30
1984	123.66	130.11	129.54	158.41	160.76	158.71
1985	127.88	128.52	128.55	145.29	160.54	156.53

Source: 1990 Population Census of China, 1% sample; 1988 "two-per-thousand" survey.

Appendix 2: A simple conceptual framework for fertility decline and sex selection

We formalize the intuition that as family size decreases, male-biased sex ratios increase. Our framework is adapted from Lin, Liu, and Qian (2014), who study the joint effect of legalized abortion and son preference on sex ratios at birth and relative female infant mortality rates in Taiwan. An important distinction is that the cost of sex selection in our model is an ex-post cost experienced through neglect; in contrast, most models treat this cost as an ex-ante cost (the cost of ultrasound, for example) (Anukriti 2016; Ebenstein 2011; Lin, Liu, and Qian 2014).

First, define the household utility function from giving births as:

$$\theta - \lambda \, , \tag{1}$$

where θ captures the utility of the child and varies with the child's gender. Specifically, if a son is born, then $\theta = 1 + \delta$; if a girl is born, then $\theta = 1$. $\delta > 0$ therefore captures the strength of son preference. $\lambda > 0$ is a cost parameter for raising children and can broadly be interpreted to include physical, financial, social, and psychological costs. Finally, without loss of generality, we assume households to have a reservation utility of zero from no children.

After a child is born, families can choose to neglect a child, which results in the child's death, with cost $c > 0$. This cost of neglect c broadly reflects the physical, financial, social, and psychological burden of neglect. Households with sufficiently high opportunity cost $\lambda > 1 + \max(\delta - c, \frac{\delta}{2})$ will never have any children. Likewise, households with sufficiently low opportunity cost $\lambda < 1 + \min(c, \frac{\delta}{2})$ will always become pregnant and not neglect. The remaining case includes families with intermediate values of λ, who become pregnant because of the option value of neglecting daughters:

$$1 + \min(c, \tfrac{\delta}{2}) < \lambda < 1 + \max\left(\delta - c, \tfrac{\delta}{2}\right) . \tag{2}$$

Once a daughter is born, households will prefer neglect to raising the daughter when:[37]

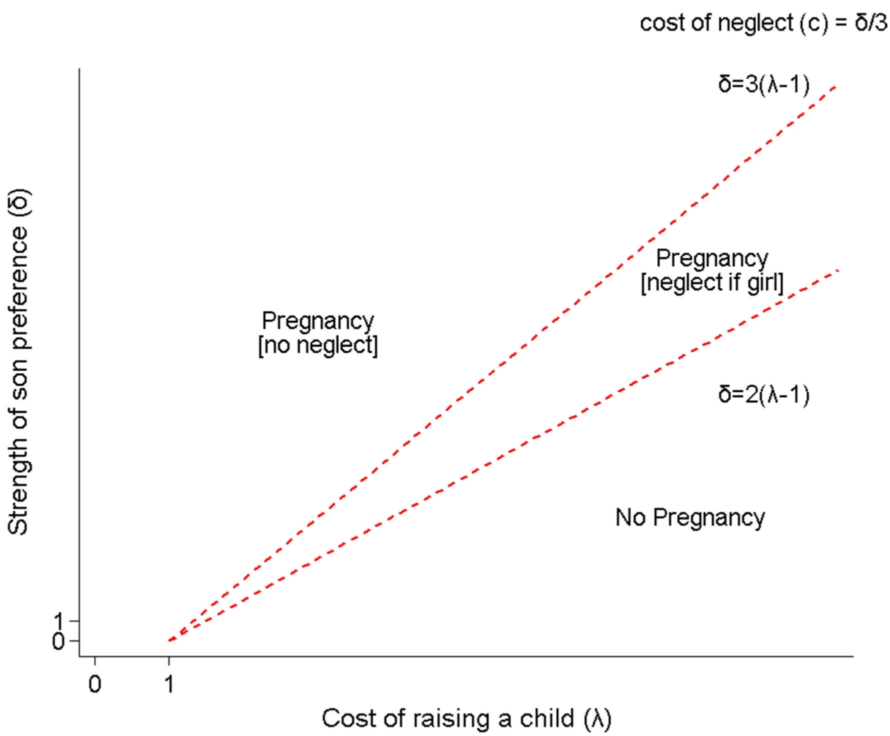
$$1 - \lambda < -c \, . \tag{3}$$

[37] In our framework, a son is never neglected because households with such preferences would never initiate a pregnancy.

Inspection of Equation 2 makes clear that as λ increases, a family's utility from having a daughter decreases – and hence the probability of neglect rises. Figure 2 shows these cases graphically.

In summary, as the cost of children λ increases, more families will choose to have no children, and some families that would otherwise become pregnant without neglect will now become pregnant – but instead neglect their daughters. In the aggregate, the impact on sex ratios is unambiguous: an increase in the cost of children (whatever the cause) leads not only to smaller family sizes, but also increased neglect of daughters and hence rising sex ratios.

Figure A-5: Graphical illustration of sex selection by strength of son preference and cost of raising a child



Note: Figure A-5 illustrates the conditions under which couples will choose to avoid pregnancy or become pregnant, and the conditions under which couples will neglect daughters following a birth. These outcomes are functions of the cost of raising a child (λ) and the strength of son preference (δ) as described in Appendix 2.

Appendix 3: Population-level estimates of missing girls from 1970–1979

Using the results of our statistical analysis, we calculate the implied number of missing girls as follows. First, we compute the implied sex ratio at birth for each year-parity-sibship sex composition cell for each statistically significant estimate (p<0.05) shown in Table 1:

$$\widehat{SRB}_{yps} = \frac{\widehat{ProbabilityMale}_{yps}}{(1-\widehat{ProbabilityMale}_{yps})} \times 100 ,$$

where $\widehat{ProbabilityMale}_{yps}$ is the estimated probability that a birth occurring in year y, at parity p, and with sibship sex composition p is male. Table A-7 shows the resulting sex ratios at birth ($\widehat{SRB}_{yps}$). We make the conservative assumption that statistically insignificant estimates imply a biologically expected sex ratio at birth. Because existing estimates of the biologically expected sex ratio at birth range between 105 and 106 boys per 100 girls (Grech, Savona-Ventura, and Vassallo-Agius 2002; Johansson and Nygren 1991; Sen 1990), we use 105.5 (results are not sensitive to choices in this range).

Next, we calculate the implied share of girls missing in each year-parity-sibship sex composition cell from the following identity:

$$\frac{SRB^e}{100} = \frac{\widehat{SRB}_{yps}}{100 + \, n^{\widehat{missing}}} ,$$

where SRB^e is the biologically expected ratio at birth and $n^{\widehat{missing}}$ is the count of missing girls for each 100 female births. Using this expression, we solve for $n^{\widehat{missing}}$ and divide by 100 to obtain the share of girls that are missing in each year-parity-sibship sex composition cell: $\left(\frac{\widehat{SRB}_{yps}}{SRB^e} - 1\right)$. Table A-8 shows these results, assuming SRB^e = 105.5.

Finally, we weight the cells in Table A-8 using the number of girls born in each cell as a share of the overall births observed in a given year (using counts shown in Table A-9). This yields year-specific rates of missing girls. We scale these year-specific rates by the total number of births across China in each year (Online 2016), summing across years 1970–1979 to obtain the total number of missing girls during the 1970s implied by our analysis (Table A-10).

To calculate the number of missing girls implied by Figure 1, we apply the same procedure described above using unadjusted sex ratios at birth in each year-parity-sibship sex composition cell in our sample.

Table A-8: Sex ratio at birth (boys per 100 girls) implied by Table 2

	First birth	Second birth		Third and higher birth	
Year		Male sibling	No male sibling	Male sibling	No male sibling
1970	105.50	105.5	105.50	105.50	105.50
1971	105.50	96.58	105.50	105.50	105.50
1972	105.50	105.5	105.50	110.48	105.50
1973	105.50	105.5	105.50	105.50	105.50
1974	105.50	105.5	105.50	110.57	114.84
1975	105.50	105.5	105.50	105.50	112.88
1976	105.50	105.5	105.50	105.50	113.98
1977	105.50	105.5	105.50	105.50	129.05
1978	105.50	105.5	105.50	105.50	117.49
1979	105.50	105.5	105.50	112.90	121.75

Table A-9: Missing girls as a share of female births in “two-per-thousand” survey

	First birth	Second birth		Third birth	
Year		Male sibling	No male sibling	Male sibling	No male sibling
1970	0.00%	0.00%	0.00%	0.00%	0.00%
1971	0.00%	–8.46%	0.00%	0.00%	0.00%
1972	0.00%	0.00%	0.00%	4.72%	0.00%
1973	0.00%	0.00%	0.00%	0.00%	0.00%
1974	0.00%	0.00%	0.00%	4.81%	8.85%
1975	0.00%	0.00%	0.00%	0.00%	6.99%
1976	0.00%	0.00%	0.00%	0.00%	8.04%
1977	0.00%	0.00%	0.00%	0.00%	22.32%
1978	0.00%	0.00%	0.00%	0.00%	11.36%
1979	0.00%	0.00%	0.00%	7.02%	15.40%

Table A-10: Number of births in "two-per-thousand" survey

Year	Female first birth	Female second birth		Female third birth		Total number of births observed (male and female)
		Male sibling	No male sibling	Male sibling	No male sibling	
1970	4629	1857	1884	7186	1654	35343
1971	4443	2072	2086	7338	1612	36154
1972	4167	2005	2171	7059	1645	35174
1973	4187	2183	2176	6687	1635	34620
1974	4617	1937	2122	5942	1676	33735
1975	4260	1935	2058	5306	1580	31262
1976	4590	1925	2044	4726	1524	30813
1977	4616	1844	1963	4076	1308	28590
1978	5138	1943	2028	3956	1449	29951
1979	5801	2081	2294	3814	1430	31795

Table A-11: Population-level estimates of missing girls

Year	Missing girls as share of total observed births		Total births occurring nation-wide		Population-level estimate of missing girls
1970	0.000%	×	27,877,013	=	0
1971	–0.485%	×	26,199,395	=	–127,003
1972	0.947%	×	26,083,358	=	247,017
1973	0.000%	×	25,041,528	=	0
1974	1.287%	×	22,669,321	=	291,665
1975	0.353%	×	21,376,746	=	75,527
1976	0.398%	×	18,752,772	=	74,567
1977	1.021%	×	18,073,552	=	184,575
1978	0.550%	×	17,567,268	=	96,552
1979	1.534%	×	17,381,984	=	266,721
					1,109,620