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## Social norms and the impact of early life events on gender inequality

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# SOCIAL NORMS AND THE IMPACT OF EARLY LIFE EVENTS ON GENDER INEQUALITY

*Wei Luo, Wei Huang, and Albert Park*

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## ABSTRACT

We study the influence of social norms in determining the impact of early life exposure to the Great Chinese Famine of 1959–1961 on gender inequality. We model how social norms interact with adverse shocks to affect male and female survival chances and influence subsequent human capital investments. We test these predictions empirically by using the Fifth National Population Census of the People's Republic of China in 2000 that has information on birthplace and estimate a difference-in-differences model that combines cohort and regional variation in exposure to the famine with regional variation in the culture of son preference. We find that son preference buffers the negative impact of intrauterine famine shocks on cohort male-to-female sex ratios and reduces famine's impact on gender inequality in health and education.

**Keywords:** famine, son preference, sex ratios, human capital investment

**JEL codes:** J13, J16, I24, I26

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## 1. Introduction

Previous literature has provided well-documented evidence on the long-run impact of prenatal events (Barker 1992, Case et al. 2005, Almond 2006, Maccini and Yang 2009, Almond and Currie 2011, Currie and Vogl 2013, Nilsson 2017). This link often is not gender-neutral, but empirical findings are mixed on which gender is more affected by prenatal events. Both biological and economic studies show that male fetuses and babies are more vulnerable to early life events (Kraemer 2000, Cameron 2004, Almond and Mazumder 2011, Dinkelman 2017, Nilsson 2017), while others find that girls are more affected (Pathania 2007, Maccini and Yang 2009, Shi 2011). Of interest, most research that finds girls are more responsive to early life events are in Asian settings (Pathania 2007, Maccini and Yang 2009, Shi 2011, Cui et al. 2020).

Abundant literature provides evidence that cultural preferences favoring sons contribute to gender disparities in health outcomes and mortality rates in Asian countries, such as the People's Republic of China (PRC) and India (Gupta 1987, Chen et al. 2007, Bhalotra et al. 2010, Jayachandran and Kuziemko 2011, Jayachandran and Pande 2017). Therefore, comparing gender differences in the long-run impact of prenatal events independent of the cultural setting may be problematic. Emerging literature points out that gender differences in the impact of early life events are context-specific. For example, Dinkelman (2017) mentioned that her finding that local environmental shocks have a more negative impact on males in South Africa is quite different from most findings in Asian settings, and it seems likely that differences in son preference across these continents could contribute to these differences. Besides, growing literature recognizes that shocks, investments, and interventions can interact in complex ways (Almond et al. 2018, Duque et al. 2018). In this paper, we study the extent to which social norms influence gender disparities in the long-run impact of early life events.

We study the long-run impact of early life events and their interaction with social norms in the context of perhaps the most severe famine in human history, the Great Chinese Famine of 1959–1961 (henceforth, the Great Famine), which struck provinces in the PRC unexpectedly and with large regional variations in death rates. During 1959–1961, agricultural production dropped sharply, and the estimated daily available food energy fell below the minimum food energy requirement (Ashton et al. 1992, Lin and Yang 2000). The prolonged famine caused an unprecedented number of deaths. National death rates were 14.6, 25.4, and 14.2 per thousand in 1959, 1960, and 1961, compared to 11.4, 10.8, and 12.0 per thousand in the previous 3 years

(1956–1958).<sup>1</sup> The Chinese central government eventually recognized the severity of the famine and moderated its policies, including reducing the transfer of grain from rural areas to urban areas and sending millions of people back to the countryside to boost agricultural production (Li and Yang 2005). By the end of 1961, death rates began to return to the pre-1959 level in over half of the provinces, and birth rates started to rebound.

Although the Great Famine was national in scope, the famine intensity varied significantly across provinces (Ashton et al. 1992, Lin and Yang 2000, Chen and Zhou 2007, Meng et al. 2015). Figure 1 displays the geographical variation of famine intensity by mapping the provincial death rate in 1960, the worst year of the famine in terms of fatalities. Provinces in central PRC, such as Henan and Anhui, and some southwestern provinces, were severely affected by the famine, while the northeastern provinces were less severely impacted. The Great Famine generated plausibly exogenous adverse shocks of varying magnitude to people living in different regions.

According to rich literature, the culture of son preference is deeply rooted and persistent in the PRC (Das Gupta et al. 2003, Guilimoto 2009). This tradition emphasizes the importance of continuing the family line through the male offspring, thereby reinforcing male dominance within a household (Murphy et al. 2011). The culture of son preference has profoundly shaped childbearing and child-rearing behavior in the PRC (Yi et al. 1993, Chen et al. 2007). One manifestation of son preference is sex-selection practices, which may be performed either prenatally or postnatally. For example, female infanticide, the neglect of baby girls, and the preferential allocation of household resources to sons can all be categorized as postnatal sex-selection strategies.

The cultural preference for sons is not only deeply rooted in the history of the PRC but also characterized by obvious and remarkable regional differences. Male-to-female sex ratios of newborn babies or young cohorts are often used as a proxy for the culture of son preference (Arnold and Liu 1986, Park and Cho 1995, Edlund 1999, Jayachandran and Kuziemko 2011). In Figure 2, we describe the geographic distribution of male-to-female sex ratios of cohorts aged 0–10 in the First National Population Census of the PRC in 1953 (henceforth, the 1953 population census). The figure shows that southern provinces, such as Guangdong and Fujian, and provinces in central PRC, such as Anhui and Jiangxi, have higher levels of son preference. Comparing this map with the geographic distribution of famine intensity reveals that the two are

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<sup>1</sup> The national death rate during the famine period is from the Statistical Yearbook published by the National Bureau of Statistics of China. The Great Chinese Famine is more severe compared to that happening in Greece (Neelsen and Stratmann 2011) and the Netherlands (Scholte et al. 2015). For example, the peak mortality rate of the Dutch famine was around 15 per thousand people, much lower than the peak mortality of 25.4 per thousand people during the Great Chinese Famine.

not significantly spatially correlated. In fact, the correlation coefficient between the provincial death rate in 1960 and sex ratios (males to females) of cohorts aged 0–10 in the 1953 population census is only 0.116.

Our theoretical model extends Bozzoli et al. (2009) and Valente (2015) and develops a nuanced and unified framework to: (i) show intrauterine adverse circumstances (famine exposure) and gender-specific postnatal investments (caused by son preference) generate mortality selection; (ii) link the selection effect to individuals' outcomes in adulthood, including height (expected health) and years of education (human capital); (iii) introduce the dimension of gender and investigate how gender inequality is shaped. Specifically, we posit that parents' unequal allocation of resources across children, as a result of son preference, especially under adverse circumstances, increases the relative survival chances of males versus females, leading to an increase in cohort male-to-female sex ratios. This phenomenon mitigates the effect of famine shocks on gender gaps in health and alters its influence on the gender gap in educational attainment.

To test our predictions about the long-run consequences of intrauterine famine shock and its interacting effect with gender norms, we assemble data from several sources. We collect cohort information on gender inequalities from the Fifth National Population Census of the PRC in 2000 (henceforth, the 2000 population census) and the 2010 China Family Panel Survey (CFPS). The advantage of using the 2000 population census is that it provides information on individuals' birthplace, enabling us to precisely identify famine severity received in utero and alleviate the concern about migration. In addition, we gather detailed province-level death rates from the Comprehensive Statistical Data and Materials on 50 Years of New China, which is compiled by the National Bureau of Statistics of China. At last, we collect information on the regional culture of son preference from pre-famine Census data (the 1953 population census). Our analysis adopts a generalized difference-in-differences approach to estimate the causal linkages. That is, our identification strategy is based on regional and cohort variation in famine severity received when in utero together with pre-famine regional variation in the culture of son preference.

We find that even though intrauterine famine shocks contribute to a reduction in male-to-female sex ratios in gender-neutral areas, the reduction in sex ratios induced by the famine shock in provinces with son preference is much smaller compared to that in gender-neutral provinces, consistent with our theoretical predictions on survival chances. The difference in the impact of famine on sex ratios in areas with son preference and gender-neutral areas associated with a one standard deviation increase in famine severity accounts for 1.89% of the mean cohort sex ratio (or 15.26% of the standard deviation). Moreover, we observe that intrauterine famine shocks



increase the gender gap in height in gender-neutral areas. However, the gender gap in height generated by the famine shock is narrower in areas with son preference, consistent with our theoretical predictions. Finally, we show that intrauterine famine shocks increase the gender gap in years of education in gender-neutral areas, but this is less true in areas with a culture of son preference. The change in the gender gap in years of education in response to a one standard deviation increase in famine severity in areas with son preference accounts for 6.41% of the mean (or 9.46% of the standard deviation). We also illustrate that our findings on education go beyond a mechanical change in survivors' health.

To address concerns related to selective fertility during the famine, we utilize birth month to capture the timing of pregnancy and show our results are not driven by selective fertility in response to the famine. Our analysis is also robust to the inclusion of rich pre-famine regional controls and rigorous fixed effects. We validate our measure of son preference by showing no correlation with pre-famine local conditions and cross-validate it across multiple alternative measures. Importantly, consistent findings persist when using alternative measures. Furthermore, our results pass several robustness tests, including using alternative sample restrictions and alternative measures of famine exposure, controlling for the impacts of the Chinese Civil War (1927-1949), the Cultural Revolution (1966-1977), and famine exposure received the first year of life, and a placebo test.

Our research contributes to the literature on the long-run impact of early life events.<sup>2</sup> Previous research studies on the long-run effect of early life events have discussed the selection effect to some extent (Neelsen and Stratmann 2011, Currie and Vogl 2013, Scholte et al. 2015, and Almond et al. 2018). However, the selection effect of early-life adverse shocks remained unproven. Valente (2015) linked mortality selection in utero and health at birth, but did not consider the long-run outcomes. In addition, as noted earlier, gender is an important dimension when considering health outcomes.<sup>3</sup> Prior research has emphasized that the lasting impact of early life events on survivors play out differently by gender (Almond et al. 2007, Dinkelman 2017, Nilsson 2017), especially in Asian contexts (Feeny et al. 2021, Sivadasan and Wu 2021, Wu et al. 2023). The first and most important contribution of our paper is to provide a formal investigation of the gender-specific selection effect of adverse life events related to social norms.

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<sup>2</sup> For a general review of early-life circumstances and adult outcomes, see Currie and Vogl (2013) and Almond et al. (2018).

<sup>3</sup> For example, Lei et al. (2012) demonstrate significant gender differences in cognitive ability among older Chinese people. In addition, evidence from developed societies also shows considerable gender disparities in mental and physical health conditions (McDonough and Walters 2001, Denton et al. 2004).

Specifically, our theoretical model extends Bozzoli et al. (2009) and Valente (2015) with a more intricate perspective and stands out by providing a comprehensive framework to understand the gender-specific selection effect of early-life adverse shocks to the long-run gender inequality in survivors' outcomes. Furthermore, our empirical findings not only corroborate our model but also align with research into the interplay between early-life adversities and subsequent human capital investments using quasi-natural experiments (Gunnsteinsson et al. 2014, Rossin-Slate and Wust 2020, Duque et al. 2021). Our findings also hold significant policy relevance to alleviate the enduring impacts of repercussions of events like the coronavirus disease (COVID-19) pandemic on newborns. The need for targeted interventions, including social safety nets addressing gender-specific vulnerabilities from early-life shocks and maternal and child health programs tailored to the well-being of pregnant women and female infants, becomes evident.

Our paper also relates to the literature on the lasting impact of the Great Famine. Previous research has documented significant effects of early life exposure to famine on adult outcomes of survivors, in terms of educational attainment, labor market outcomes, and health (Luo et al. 2006, Almond et al. 2007, Chen and Zhou 2007, Meng and Qian 2009, Shi 2011, Mu and Zhang 2011, Cui et al. 2020). However, very few studies focus on how these adverse shocks shaped inequality (Johnson and Jackson 2019), especially regarding the dimension of gender. Mu and Zhang (2011) find better health among female famine survivors. Cui et al. (2020) documented sizable gender disparities in cognitive abilities among the Chinese elderly who suffered from early-life hunger experience. But no papers systematically investigate the impact of early-life famine exposure on the gender gap. Our paper fills the gap by presenting causal evidence that intrauterine famine exposure shapes gender inequalities (gaps) in survival chances, health outcomes, and educational attainment in later life.

Our contribution can also be situated relative to the literature on son preference and the "missing girls" phenomenon (Gupta 1987, Sen 1990, Yi et al. 1993, Das Gupta et al. 2003, Ebenstein 2010). Our findings extend the literature on the impact of son preference on gender disparities in mortality (Chen et al. 2007, Bhalotra et al. 2010) and health outcomes (Jayachandran and Pande 2017) by showing that this impact is exacerbated by adverse shocks and results in unintended consequences for later-life gender inequality.

The rest of this paper is structured as follows. The next section presents the theoretical framework. Section 3 describes the data and variables. Section 4 introduces our identification strategy. In section 5, we present our empirical findings. Section 6 presents robustness checks. Finally, we conclude in section 7.

## 2. Theoretical Framework

### 2.1. Overview

We develop a simple model to show that adverse shocks negatively influence individuals' survival chances. Son preference mitigates the effect of adverse shocks on male survival chances by shifting the effective survival threshold. It alters the effect of adverse shocks on gender gaps in health and educational attainment in later life. Our model distinguishes the intrauterine period and the infancy and early childhood stages. As illustrated in Figure 3, children are exposed to famine shocks in utero. As sex selection technology is unavailable during that period (Chen et al. 2013), parents make gender differential investments only after observing the child's sex (during infancy and early childhood).

### 2.2. Setup of the Model

**Initial endowment distributions.** Initial health endowments of individual  $i$  ( $h_i$ ) are assumed to be distributed with a density function of  $f(h_i)$  and a cumulative distribution function (CDF henceforth) of  $F(h_i)$ . We suppose the CDF of health endowment  $F(h_i)$  satisfies that  $F''(h_i) > 0$  over the interval  $I$  (strictly convex on the interval  $I$ ). In addition, the density function of health endowment  $f(h_i)$  (or the first derivative of health endowment CDF  $F'(h_i)$ ) is always positive. Our model assumption holds broad applicability across a range of distributions commonly encountered in various fields.<sup>4</sup> For instance, rich empirical evidence in economics and biology supports the normal distribution of human height and weight (Tanner and Tanner 1981, Wachter and Trussell 1982, Steckel 1995), which serves as an example of the health distribution that meets the specified properties.

According to biological literature (Andersson and Bergstrom 1998, Kraemer 2000), boys are naturally weaker and more vulnerable than girls. We use subscripts to denote boys ( $b$ ) and girls ( $g$ ) and assume that the distribution of boys' endowments lies slightly to the left of that of girls (Catalano and Bruckner 2006, Mu and Zhang 2011).<sup>5</sup> In other words, we have  $h_{ig} = h_{ib} + \Delta h$ , where  $\Delta h > 0$ .

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<sup>4</sup> It is important to note that the convexity assumption in our model extends beyond just the normal distribution. For example, the chi-square distribution and Weibull distribution also satisfy these properties within specific interval. However, it's worth mentioning that certain distributions, such as the exponential and Pareto distributions, do not adhere to the convexity assumption.

<sup>5</sup> The phenomenon that the infant mortality of boys is higher than girls can be modeled in two ways: (i) the health endowment of boys and girls are different (weaker boys), and the survival threshold is uniform; or (ii) health distributions are uniform, but boys have a higher survival threshold. Previous studies in biology and economics document that females may have better health endowments than males (Catalano and Bruckner 2006, Mu and Zhang 2011) and support the first one.

**Postnatal investment and son preference.** Gender-specific postnatal investment is purely determined by son preference  $\theta_i$  and is assumed to decrease effective survival thresholds. We will define the effective survival threshold in the next two paragraphs. We use superscripts  $n$  and  $p$  to denote two social norms-gender-neutral ( $n$ ) and son preference ( $p$ ). We first define that in gender-neutral areas, preferences are equal for boys and girls, that is,  $\theta_g^n = \theta_b^n$ . For boys, we explicitly present the necessary assumption that son preference is greater in areas with son preference than in gender-neutral areas,  $\theta_b^p > \theta_b^n$ . Similarly, for girls, we present the assumption that willingness to invest in girls is greater in gender-neutral areas than in areas with son preference,  $\theta_g^n > \theta_g^p$ . It then follows that  $\theta_b^p > \theta_g^p$  in areas with son preference.<sup>6</sup>

**Famine shocks.** Adversity received in utero (negative health shock) may shift the entire health endowment distribution to the left, increase the mortality rate, and reduce the health parameters among survivors (the scarring effect). On the other hand, an intrauterine adverse environment may shift the survival threshold to the right and require a higher health endowment to survive. The increased survival threshold raises the mortality rate while resulting in a positive selection in survivors' health parameters, so-called the culling effect (see Catalano and Bruckner (2006) and Valente (2015) for a comprehensive discussion of both mechanisms). In our model, we assume intrauterine exposure to famine shocks increase the effective survival thresholds by  $\Delta\mu$  of both genders.<sup>7</sup>

**Survival chances.** We define that the biological survival threshold  $z$  is the survival threshold under natural conditions and is the same for both genders. As the distribution of boys' endowments lies slightly to the left of that of girls, boys are less likely to survive under natural conditions. In addition, we introduce the effective survival threshold that accounts for the effect of famine shock and postnatal investments. The effective survival threshold equals the biological survival threshold  $z$  plus the effect of intrauterine famine shock ( $\Delta\mu$ ) minus the effect of postnatal investment ( $\theta_i$ ). Individuals survive if the endowment draw is high enough, given the effective survival threshold  $Z$ , where  $Z \in I$ . For simplicity, we decompose the effective survival threshold into two parts: gender-specific survival thresholds accounting for postnatal investments  $z'_i$  (equal

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<sup>6</sup> Although we recognize that health endowments will influence health investments as well as educational investments, we do not include health investments in the model explicitly because doing so would not change the qualitative predictions of the model as long as gender differences in the negative impact of the famine on health endowments are not fully offset by compensating health investments, but adding health investments would make the model much less tractable.

<sup>7</sup> Assuming famine shocks reduce the health endowment or increase the survival threshold will not alter our predictions on gender gaps in survival chances, observed health, and education. However, it will change our predictions on the expected health and education of survivors.

to  $z-\theta_i$ ) and the effect of famine shock ( $\Delta\mu$ ). The survival chance ( $sr_i$ ) is defined in equation 1. The survival probabilities for non-exposed cohorts can be obtained by equating  $\Delta\mu$  to 0.

$$sr(h_i|h_i > Z) = 1 - F(Z) \quad (1)$$

**The health of survivors.** Survivors' expected health can be viewed as the expectation of a distribution of health endowment truncated by the effective survival threshold in equation 2. Here,  $g(\cdot)$  denotes the truncated density function of health,  $g(h_i) = f(h_i)\forall h_i > Z$  and  $g(h_i) = 0$  otherwise. For simplicity, we let  $\lambda(Z)$  represent  $\frac{\int_Z^\infty h_i g(h_i) dh_i}{1 - F(Z)}$ . We suppose  $\lambda'(Z) > 0$  and  $\lambda''(Z) > 0$  over the interval  $I$  for any  $Z \in I$ . For example, the expectation of a truncated normal distribution, as a function of the inverse mills ratio, satisfies those properties.

$$E(h_i|h_i > Z) = \frac{\int_Z^\infty h_i g(h_i) dh_i}{1 - F(Z)} = \lambda(Z) \quad (2)$$

**Schooling choice of survivors.** In a two-child family (one boy and one girl), parents choose children's schooling, consumption, and savings to maximize their utility (equation 3) subject to the life-time budget constraint (equation 4) and the determinants of children's earnings (equation 5). Parents' utility depends on their consumption in both periods,  $C_1$  and  $C_2$ , and the utility getting from their children's income.  $1 - \Delta\mu$  denotes parental response to children's famine experience,  $\beta$  is the discount factor, and  $\theta$  is the degree of altruism (preference) toward children of a specific gender.<sup>8</sup> Parental resources equal  $Y$  (earnings in period 1) as we assume that children do not transfer resources to their parents.<sup>9</sup> Parental consumption in period 2 is determined by the budget constraint  $C_2 = RX_1$ , where  $R$  denotes the rate of return on  $X_1$  and is assumed to be exogenously determined by the competitive market.  $X_1$  is saving in period 1 and  $p$  is the price of per unit schooling. We model children's earnings as the product of human capital that depends on schooling years  $s_i$  and surviving children's health  $Q_i$ , and  $w_i$  (the market-determined wage for a unit of human capital), shown in equation 5. The key assumption is that parents' investment in schooling is complementary to surviving children's health.

$$\text{Max } E\{\ln C_1 + \beta \ln C_2 + (1 - \Delta\mu)\beta\theta_b I_b + (1 - \Delta\mu)\beta\theta_g I_g\} \quad (3)$$

<sup>8</sup> As long as parents have some degree of altruism (or preference) toward their children, and  $\theta > 0$ , they invest a positive amount in children's schooling, since the marginal rate of these investments is very high for small investments, which satisfies the Inada conditions.

<sup>9</sup> The budget constraint of parents in period 1 is  $ps_b + ps_g + C_1 + X_1 = Y$ .

$$C_1 + \frac{C_2}{R} + ps_b + ps_g = Y \quad (4)$$

$$I_i = w_i Q_i s_i^\gamma, 0 < \gamma < 1 \quad (5)$$

## 2.3. Predictions of the Model

Our model predictions are based on the position of the effective survival line relative to the health endowment distributions. The effective survival thresholds are: (i) different to individuals who are exposed to the famine and those who are not (change in  $\Delta\mu$ ); and (ii) different to individuals who are in areas with son preference and those in gender-neutral areas (differences in  $\theta_i$ ). In addition, the health endowment distributions are different for boys and girls.<sup>10</sup>

### 2.3.1. Survival Chances

The impact of famine on survival chances is the difference in the survival probability of cohorts exposed to the famine and those who were not, equaling the difference between the two CDFs (equation 6). According to  $F'(h_i) > 0$ , we have  $F(z'_i) - F(z'_i + \Delta\mu) < 0$ , implying famine shocks reduce survival chances.

$$Sr(h_i|h_i > z'_i + \Delta\mu) - Sr(h_i|h_i > z'_i) = F(z'_i) - F(z'_i + \Delta\mu) \quad (6)$$

We show the impact of famine on survival probabilities for each gender (column 1 for boys and column 2 for girls) in gender-neutral areas (row 1) and areas with son preference (row 2) in Panel A of Table 1. In the last row in Panel A of Table 1, we show the sign of the difference in the impact of famine on survival chances for each gender between two areas (gender-neutral areas minus areas with son preference). We already know that the distribution of boys' endowments lies slightly to the left of girls, and preferences are equal in gender-neutral areas (thereby,  $z_b^n = z_g^n$ ). Hence, the position of the effective survival threshold relative to boys' health distribution lies right to that of girls (as illustrated in Figure 4). According to  $F''(h_i) > 0$  over the interval  $I$  for any  $Z \in I$ , we have that the negative impact of famine on boys' survival chance (row 1, column 1 in Panel A) is greater than that for girls (row 1, column 2 in Panel A) in gender-neutral areas, or  $F(z_b^n) - F(z_b^n + \Delta\mu) < F(z_g^n - \Delta h) - F(z_g^n - \Delta h + \Delta\mu)$ . We say that famine reduces the sex ratios among survivors in gender-neutral areas.

<sup>10</sup> Our model predictions only hold if the distribution of initial health endowments satisfies the prosperity described above.

As the preference for boys is higher in areas with son preference, we have  $z'_b{}^n > z'_b{}^p$ . By the same logic, the negative impact of famine on survival chances for boys in gender-neutral areas (row 1, column 1 in Panel A) is greater than that in areas with son preference (row 2, column 1 in Panel A). That is,  $F(z'_b{}^n) - F(z'_b{}^n + \Delta\mu) < F(z'_b{}^p) - F(z'_b{}^p + \Delta\mu)$ , and the sign of the difference in the impact of famine on boys' survival chances between these two areas is negative (row 3, column 1 in Panel A).

Similarly, the preference for girls is higher in gender-neutral areas, and the girls' survival threshold is greater in areas with son preference compared to gender-neutral areas ( $z'_g{}^p > z'_g{}^n$ ). Thereby, the negative impact of famine on girls' survival probabilities in gender-neutral areas (row 1, column 2 in Panel A) is smaller than in areas with son preference (row 2, column 2 in Panel A) which implies  $F(z'_g{}^n - \Delta h) - F(z'_g{}^n - \Delta h + \Delta\mu) > F(z'_g{}^p - \Delta h) - F(z'_g{}^p - \Delta h + \Delta\mu)$ , and the sign of the difference between the two areas is positive (row 3, column 2 in Panel A).

From these two inequalities, we know that the gender difference (boys minus girls) in the impact of famine on survival probability reduction is greater in gender-neutral areas compared to areas with son preference. As sex ratios capture the survival chances of boys relative to girls, we put forward Proposition 1.

*Proposition 1:* The negative impact of famine on survivors' sex ratios is greater in gender-neutral areas than in areas with son preference.

### 2.3.2. The Expected Health of Survivors

The impact of famine on survivors' expected health is the difference in the expected health of famine-exposed and non-exposed cohorts (equation 7). According to  $\lambda'(h_i) > 0$ , we have  $\lambda(z'_i + \Delta\mu) - \lambda(z'_i) > 0$ , suggesting that famine shocks increase the expected health among survivors (the culling effect).

$$E(h_i | h_i > z'_i + \Delta\mu) - E(h_i | h_i > z'_i) = \lambda(z'_i + \Delta\mu) - \lambda(z'_i) \quad (7)$$

Panel B of Table 1 shows the impact of famine on survivors' expected health for each gender in gender-neutral areas and areas with son preference. The last row of Panel B displays the sign of the difference in the impact of famine on survivors' expected health for each gender between the two areas (gender-neutral areas minus areas with son preference).

The distribution of boys' endowments lies slightly to the left of girls, and preferences are equal in gender-neutral areas. Hence, we get that the impact of famine on boys' expected health (row 1, column 1 in Panel B) is greater than that for girls (row 1, column 2 in Panel B) in gender-neutral

areas, based on  $\lambda''(Z) > 0$  over the interval  $I$ . That is,  $\lambda(z_b^m + \Delta\mu) - \lambda(z_b^m) > \lambda(z_g^m - \Delta h + \Delta\mu) - \lambda(z_g^m - \Delta h)$ . In other words, famine increases the gender gap in health in gender-neutral areas.

We now move to compare the effect of famine on expected health for each gender between two areas. As the preference for boys is higher in areas with son preference, we have  $z_b^m > z_b^p$ . Thereby, the impact of famine on the expected health for boys in gender-neutral areas (row 1, column 1 in Panel B) is greater than that in areas with son preference (row 2, column 1 in Panel B). That is,  $\lambda(z_b^m + \Delta\mu) - \lambda(z_b^m) > \lambda(z_b^p + \Delta\mu) - \lambda(z_b^p)$ , and the sign of the difference in the impact of famine on surviving boys' expected health between these two areas is positive (row 3, column 1 in Panel B).

Similarly, we get that the impact of famine on surviving girls' expected health in gender-neutral areas (row 1, column 2 in Panel B) is smaller than in areas with son preference (row 2, column 2 in Panel B), as the girls' survival threshold is greater in areas with son preference compared to gender-neutral areas ( $z_g^p > z_g^m$ ). Hence, we have that  $\lambda(z_g^m - \Delta h + \Delta\mu) - \lambda(z_g^m - \Delta h) < \lambda(z_g^p - \Delta h + \Delta\mu) - \lambda(z_g^p - \Delta h)$ , and the sign of the difference between the two areas is negative (row 3, column 2 in Panel B).

We know that the gender difference (boys versus girls) in the impact of the famine on expected health is greater in gender-neutral areas than in areas with son preference according to the above two inequalities. Hence, we put forward Proposition 2.

*Proposition 2:* The impact of famine on survivors' gender gap in health is larger in gender-neutral areas compared to areas with son preference.

### 2.3.3. Schooling Choice for Survivors

Solving for the optimal schooling choices for boys and girls based on equations 3–5 yields equation 8. Here, the Lagrangian multiplier for the lifetime budget constraint is  $\tau$ .

$$s_i = \left( \frac{(1 - \Delta\mu)\theta_i w_i Q_i \gamma \beta}{p\tau} \right)^{\frac{1}{1-\gamma}} \quad (8)$$

If we take the logarithm of schooling, the impact of famine on schooling can be expressed as the difference in schooling between famine-exposed and non-exposed cohorts (equation 9). It equals the human capital production parameter multiplied by the overall effect of famine on the survivors' expected health and the famine itself. Thereby, the effect of famine on survivors' education is ambiguous. It is determined by the relationship between  $\ln \lambda(z_i' + \Delta\mu) - \ln \lambda(z_i')$  and



$\ln(1 - \Delta\mu)$ , where  $\ln \lambda(z'_i + \Delta\mu) - \ln \lambda(z'_i)$  captures the ratio of famine survivors' expected health to non-exposed cohorts.

$$\ln s(h_i > z'_i + \Delta\mu) - \ln s(h_i > z'_i) = \frac{1}{1 - \gamma} (\ln \lambda(z'_i + \Delta\mu) + \ln(1 - \Delta\mu) - \ln \lambda(z'_i)) \quad (9)$$

Similar to Panels A and B, Panel C of Table 1 shows the impact of famine on survivors' education for each gender in gender-neutral areas and areas with son preference. The impact of famine on the gender gap in education in gender-neutral areas is determined by the relationship between the ratio of expected health of famine survivors to non-exposed cohorts for both genders (row 1, columns 1 and 2 in Panel C). As in Section 2.3.2, we already know that the impact of famine on boys' expected health is greater than that for girls in gender-neutral areas. Moreover, the expected health of non-exposed boys is smaller than girls due to the weak boy assumption. We know that  $\frac{\lambda(z'_b{}^n + \Delta\mu)}{\lambda(z'_b{}^n)} > \ln \frac{\lambda(z'_g{}^n + \Delta\mu)}{\lambda(z'_g{}^n)}$ , which implies that famine raises the gender gap in education in gender-neutral areas.

Furthermore, we consider the case in areas with son preference. As in Section 2.3.2, we already know that the impact of famine on boys' expected health is smaller than that for girls in areas with son preference. Moreover, the expected health of non-exposed boys is smaller than girls in areas with son preference. Hence, the relationship between  $\ln \frac{\lambda(z'_b{}^p + \Delta\mu)}{\lambda(z'_b{}^p)}$  and  $\ln \frac{\lambda(z'_g{}^p + \Delta\mu)}{\lambda(z'_g{}^p)}$  is ambiguous. The gender difference (boys versus girls) in the impact of the famine on education between these two areas (gender-neutral minus son preference) can be: (i) less than zero if  $\ln \frac{\lambda(z'_b{}^n + \Delta\mu)}{\lambda(z'_b{}^n)} - \ln \frac{\lambda(z'_g{}^n + \Delta\mu)}{\lambda(z'_g{}^n)} > \ln \frac{\lambda(z'_b{}^p + \Delta\mu)}{\lambda(z'_b{}^p)} - \ln \frac{\lambda(z'_g{}^p + \Delta\mu)}{\lambda(z'_g{}^p)}$ ; (ii) equal to zero if  $\ln \frac{\lambda(z'_b{}^n + \Delta\mu)}{\lambda(z'_b{}^n)} - \ln \frac{\lambda(z'_g{}^n + \Delta\mu)}{\lambda(z'_g{}^n)} = \ln \frac{\lambda(z'_b{}^p + \Delta\mu)}{\lambda(z'_b{}^p)} - \ln \frac{\lambda(z'_g{}^p + \Delta\mu)}{\lambda(z'_g{}^p)}$ ; and (iii) greater than zero if  $\ln \frac{\lambda(z'_b{}^n + \Delta\mu)}{\lambda(z'_b{}^n)} - \ln \frac{\lambda(z'_g{}^n + \Delta\mu)}{\lambda(z'_g{}^n)} > \ln \frac{\lambda(z'_b{}^p + \Delta\mu)}{\lambda(z'_b{}^p)} - \ln \frac{\lambda(z'_g{}^p + \Delta\mu)}{\lambda(z'_g{}^p)}$ .

**Proposition 3:** The relationship between the impact of famine on survivors' gender gap in education in gender-neutral areas and that in areas with son preference is ambiguous.

### **3. Data, Measures, and Descriptive Evidence**

#### **3.1. Data**

To empirically test our theoretical predictions, we combine data from several sources. First, we test our predictions about gender gaps in survival rates and years of education using data from the 2000 population census. Our main empirical analysis is based on a sample that includes cohorts who were born between 1954 and 1966.<sup>11</sup> The main advantage of using this data set rather than other household data is that the 2000 population census provides a large, representative sample across all 30 provinces, including a large number of persons who experienced the famine. This allows us to focus our analysis on cohorts with famine exposure in regions with different degrees of son preference while retaining a sufficient sample size in each birth cohort and birth province. In addition, the 2000 population census provides information on respondents' birth province and enables us to construct an exact measure of famine severity received in utero. Cross-province migration was not very prevalent in 2000. Based on our calculation, the inter-provincial migration rate was only 3.04% in 2000 for the whole population and 1.92% for the sample that we use. It also has basic demographic and education information, which can be used to construct two dependent variables of interest, male-to-female sex ratios and the gender gap in educational attainment. We combine the 2000 population census data with additional data on the provincial death rate by year and pre-famine sex ratios as a measure of son preference from the 1953 population census.<sup>12</sup>

To test our theoretical predictions on health outcomes, we use data from the 2010 China Family Panel Survey, a nationally representative dataset of Chinese households and individuals with a wealth of information on health and individual background variables. The CFPS sample in 2010 covers 25 provinces, municipalities, and autonomous regions, representing 95% of the Chinese population. The 2010 baseline survey interviewed a total of 14,960 households and 42,590 individuals.

#### **3.2. Measures**

Two key explanatory variables in our empirical analysis are intrauterine exposure to famine and regional level of son preference.

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<sup>11</sup> The 2000 population census only has information on current Hukou status and does not know the Hukou status when they were born or in utero. Here we do not divide the sample based on current Hukou status as the change of Hukou type during individuals' life courses is endogenous and highly selective.

<sup>12</sup> It comes from the Comprehensive Statistical Data and Materials on 50 Years of New China which is compiled by the Department of Comprehensive Statistics of the National Bureau of Statistics of China.

**Intrauterine famine exposure.** The construction of famine severity measure follows the approach in Almond et al. (2007) and is shown in equations 10 and 11. Famine severity received when in utero is constructed at the province birth year-month level and is defined as the province-level excess death rate in their birth year and the year before, weighted by the share of intrauterine months in these 2 years.  $EDR_j$  denotes the average excess death rate during the famine period, which is the difference between the average death rate during the famine (1958–1961) and the average death rate in normal years (1972–1976).

$$FS_{jmt} = [EDR_{jt} \times \max(9 - \text{birthm}, 0) + EDR_{j,t-1} \times \min(9, \text{birthm})]/9 \quad (10)$$

$$EDR_{jt} = \begin{cases} EDR_i, & 1957 < t < 1962 \\ 0, & \text{otherwise} \end{cases} \quad (11)$$

**Son preference measure.** We construct a proxy for the culture of son preference using the male-to-female sex ratios of young cohorts (aged 0–10) from pre-famine census data (the 1953 population census). This exercise enables our measure of son preference not to be influenced by shocks to sex ratios caused by the famine. As the biologically normal sex ratio (males-to-females) is 1.05, we define a dummy variable for son preference to equal 1 if the provincial male-to-female sex ratio of cohorts aged from 0 to 10 in the 1953 population census is above 1.07 following the method used by Mu and Zhang (2011). We also show our results remain robust to alternative cutoffs and continuous measures in the robustness check.

We further show the validity of the measure using various data sources. In Panels A and B of Figure 5, we plot the cohort sex ratios of individuals aged 0–10 in the 1953 census against sex ratios of the same age cohort in the 1982 and 1990 censuses. A significant correlation provides evidence that son preference culture is persistent. In addition, we construct three alternative measures of son preference using post-famine data and show that they are highly correlated with our son preference measure—sex ratios using the 1953 census.<sup>13</sup>

- (i) We follow the literature on the son-targeting fertility stopping rule that leads to the phenomenon of higher-order births being more likely to be sons (Yamaguchi 1989, Basu and De Jong 2010, Jayachandran and Pande 2017). We link parents and children in each household using data from the Third National Population Census of the PRC in 1982 and construct sex ratios of higher-order births at the province level.

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<sup>13</sup> It is worth noting that using these alternative measures calculated from post-famine data may be problematic as famine-exposed cohorts were at childbearing age when the census or survey was conducted. Prior studies show that famine may significantly influence individuals' marital and fertility decisions (Angrist 2002, Brandt et al. 2016), and mothers' exposure to adversity may influence the sex ratios of their offspring (Song 2014).

- (ii) We utilize information on the expected number of children from the 2001 Chinese National Family Planning and Reproductive Health Survey to calculate the expected share of children who are boys.
- (iii) We further exploit individuals' opinions about gender norms in the 2014 CFPS data to construct another measure of son preference. Panels C, D, and E of Figure 5 consistently show that all these alternative measures for boy preference are highly correlated with sex ratios in the 1953 census at the province level. Taken together, Figure 5 helps confirm the validity of our son preference measure. We also conduct regression analysis using these alternative boy preference measures in Panel B of Table 5. Overall, our results remain robust and statistically significant when using these measures.

**Pre-famine regional characteristics.** We control for gross domestic product (GDP) per capita, number of health facilities, and number of high school students in 1954 to capture pre-famine economic conditions, development of public health services, and educational institutions.<sup>14</sup> Due to the skewed distributions of these characteristics, we subtract the median values of these variables. We report the summary statistics of these variables in Table A1 in the Appendix.<sup>15</sup> We also check whether son preference culture is correlated with regional pre-famine conditions in Figure A1. We find that the correlation coefficients between the son preference measure and regional GDP per capita, number of health facilities, and number of high school students in 1954 are very low.

### 3.3. Descriptive Evidence

We first provide descriptive evidence on how the cultural preference for sons buffers the negative impact of famine shock on cohort sex ratios (males to females). In Figure 6, we plot male-to-female sex ratios of each birth year cohort by regional famine severity and the degree of preference for sons. It is clear that cohort sex ratios experienced a sharp decrease in 1960. This pattern is consistent with the argument in the biological literature that boys are physically more fragile and thus more likely to die in the face of severe adverse shocks (Kraemer 2000, Almond et al. 2007). What is interesting is the difference between these patterns in areas with a high degree of historical son preference and in areas with a low degree of historical son preference when controlling for famine severity. We find that cohort sex ratios dropped dramatically in areas with lower son preference.<sup>16</sup> In a society lacking access to prenatal sex selection technology, the

<sup>14</sup> The information in 1954 is used to construct pre-famine controls is because it is the earliest data available.

<sup>15</sup> The Appendix is available at <http://dx.doi.org/10.22617/WPS240387-2>.

<sup>16</sup> The impact of famine on cohort sex ratios is defined as the difference in cohort sex ratios between cohorts born between 1954 and 1957 and those born between 1958 and 1961.

culture of son preference would strongly affect the survival chances of young girls and alleviate the impact of adverse shocks on young boys. Taken together, Figure 6 suggests that son preference plays an important role in buffering the impact of the Great Famine on male survival chance (or cohort male-to-female sex ratios). However, this simple comparison in Figure 6 is naive, and we next present more rigorous causal evidence.

## 4. Identification Strategy

### 4.1. Average Treatment Effect Model

Our identification strategy is based on quasi-random variation in famine severity across provinces and birth cohorts together with regional variation in the culture of son preference. To be specific, we compare outcomes of interest across famine-exposed cohorts and non-exposed cohorts, across regions with high famine severity and low famine intensity, and across regions that have a culture of son preference and those that do not, controlling for pre-famine economic conditions, development of the public health system, and development of the education system. To fix ideas, our specification is based on a generalized difference-in-differences model:

$$Y_{jmt} = \alpha_0 + \alpha_1 FS_{jmt} + \alpha_2 FS_{jmt} \times SP_j + \alpha_3 FS_{jmt} \times X_j + \gamma_j + \delta_{mt} + \epsilon_{jmt} \quad (12)$$

Here,  $Y_{jmt}$  denotes the outcome variables of interest, which is the mean outcome for cohorts born in province  $j$  in month  $m$  of year  $t$ .  $FS_{jmt}$  measures intrauterine exposure to famine severity, constructed using province-level excess death in one's birth year ( $t$ ) and the year before ( $t - 1$ ) weighted by the share of intrauterine months in these 2 years. A continuous measure of famine intensity/treatment is used, and thereby, more variation in the data can be captured.  $SP_j$  is a dummy variable that equals 1 if provincial male-to-female sex ratios of cohorts aged 0–10 in the 1953 population census are above 1.07.

Our main concern is that regions with different pre-famine conditions may experience different trends in sex ratios or the gender gap in health or education. We next include variables that reflect pre-famine conditions ( $X_j$ ), including provincial GDP per capita, number of health institutions, and number of high school students in 1954. As  $X_j$  are cross-sectional and would be absorbed by the province fixed effects, we interact the  $FS_{jmt}$  with pre-famine regional characteristics. The estimates without pre-famine conditions can be interpreted as estimates when differences in pre-famine regional characteristics are not controlled for. As the distributions of values for these pre-famine variables are skewed, we subtract the median of these variables. This enables us to

interpret the empirical estimates as reflecting the impacts when pre-famine conditions are at median levels.

We also control for province fixed effects  $\gamma_j$  and birth year-month fixed effects  $\delta_{mt}$ . All models are weighted by the sample size of each cohort. We cluster standard errors at the province level.

In equation 12, the coefficient  $\alpha_1$  indicates the effect of famine in gender-neutral areas. The coefficient of interest,  $\alpha_2$ , indicates the difference in the impact of famine between areas with son preference and gender-neutral areas. The estimation strategy has all of the advantages and potential pitfalls of the standard DID approach. Province fixed effects control for all time-invariant factors that differ between provinces. Birth year-month fixed effects control for patterns in sex ratios or gender gaps in educational attainment or health that differ by cohorts. Our identification strategy relies on the assumption that there are no other contemporaneous shocks affecting gender gaps in outcome variables in gender-neutral areas and areas with son preference beyond the famine shocks (Bertrand et al. 2004). One potential destructive concern is that fertility decisions may be postponed during the famine, and selective fertility could influence our outcome variables of interest. We will address concerns about selective fertility in detail in the results section.

## 4.2. Flexible Estimates

A fully flexible specification is used to examine pre-trends of outcome variables of interest and to examine the timing of the impact of the famine shock interacted with regional son preference (equation 13). We restrict our sample to cohorts born between 1952 and 1966 in the 2000 population census.

$$Y_{jmt} = \beta_0 + \sum_{t=195}^{1965} (\beta_t \times EDR_j \times SP_j \times Year_t) + \sum_{t=195}^{1965} (EDR_j \times X_j \times Year_t) + \gamma_j + \delta_{mt} + \epsilon_{jt} \quad (13)$$

Here,  $Year_t$  denotes the birth year, which is equal to 1 if individuals are born in the year  $t$ . All other variables are defined as before. We interact regional excess death rate during the famine with regional son preference and each of the birth year dummies. The estimated vectors of  $\beta_t$ s reveal the correlation between famine severity, regional son preference, and the outcome variables of interest for each cohort. We would expect the estimated  $\beta_t$ s to be close to zero and statistically insignificant over time before the famine and for the magnitude of the coefficients on the triple interaction term to be larger for the famine-exposed cohorts.

## 5. Empirical Results

### 5.1. The Impact on Cohort Sex Ratios

In Table 2, we investigate the impact of intrauterine famine exposure and its interacting effect with regional son preference on cohort male-to-female sex ratios. We multiply sex ratios by 100 to make it feasible to interpret the estimated coefficients. In columns 1–3, the dependent variable is cohort male-to-female sex ratios at province birth year and month level. Column 1 shows that intrauterine famine exposure reduces sex ratios, but the interaction term between famine exposure and son preference is positive, indicating the negative effect of famine exposure on cohort male-to-female sex ratios is mitigated by regional son preference. Results in column 2 hold to including pre-famine regional controls, even though the estimates of the famine term become less precise. Column 2 shows, on average, that the difference in the impact of famine on sex ratios in areas with son preference and gender-neutral areas associated with a one standard deviation increase in famine severity by 1.98 percentage points, accounting for 1.89% of the mean cohort sex ratio (or 15.26% of the standard deviation), consistent with Proposition 1.

One potential concern about our identifying assumption is that fertility behavior may be selective during the famine, which will confound our analysis. We address this concern by focusing on those born in months, implying that they were conceived before the famine. To be specific, we divide famine-exposed cohorts into two groups: cohorts who were conceived at least 5 months before the famine (cohorts who were born between January 1958 and June 1959) and cohorts who were conceived and born during the famine (cohorts born from July 1959 to the end of famine). Cohorts who were conceived and born during the famine could be a selected sample affected by fertility response to the famine. We re-estimate the model when excluding those conceived and born during the famine in column 3 and find our estimates reassuring. We also report estimates using alternative ways to aggregate the dependent variable. Columns 4–5 report the results when cohort sex ratios are aggregated at the province birth year level. The remaining two columns report the results using the moving average of cohort sex ratios (6 months before and after a particular birth year and month). Overall, our results remain robust to alternative ways to aggregate the dependent variable.

We present the estimates for the fully flexible specification in panel A of Figure 7.<sup>17</sup> The point estimates of coefficients of the triple interaction of famine severity, son preference, and birth year in equation 19 are plotted, and a clear pattern emerges. Before the Great Famine, the coefficients

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<sup>17</sup> We also divide the sample by the level of son preference and conduct the flexible estimates for both groups in Figure A2 which demonstrate the same patterns.

were constant over time and small in magnitude, suggesting no substantial effect of future famine severity on cohorts born before the Great Famine. In addition, the point estimates steadily increase among cohorts born during the famine, reaching a peak for cohorts born in 1961 and conceived in 1960. In summary, our estimates in Table 2 and Panel A of Figure 7 verify our predictions on cohort sex ratios (Propositions 1) and suggest that the culture of son preference buffers the negative impact of famine shock on male survival chances.

## 5.2. The Impact on Adult Height

To test the prediction about the gender gap in survivors' health, we use adult height to proxy for health, which is from the 2010 CFPS. One caveat is that adult height may not be a perfect proxy for survivors' health parameters. Therefore, the interpretations of the following results should be cautious. Due to the small sample size of CFPS data, we aggregate the dependent variable at the province birth year level. Column 1 of Table 3 presents estimates without pre-famine regional controls. In column 2, we further include these pre-famine controls. It shows that the impact of a one standard deviation increase in famine exposure raises the gender gap in height in gender-neutral areas by 2.45 centimeters. Meanwhile, the effect of one standard deviation increase in famine exposure on the gender gap in height in areas with son preference is 2.90 centimeters lower than that in gender-neutral areas, consistent with Proposition 2. In the last column, we re-estimate the model when excluding those conceived and born during the famine and find estimates hold.

In addition, Panel A of Appendix Table A8 reports the estimated impact on survivors' height for each gender. We find that famine exposure significantly increases height for males (Columns 1–2), consistent with our model prediction and the findings of Gørgens et al. (2012). In addition, the coefficient of the interaction term is negative for males while positive for females. This finding indicates that the impact of famine on height is more prominent for males in gender-neutral areas than in areas with son preference. Conversely, for females, the effect of famine is smaller in gender-neutral areas than in areas in son preference. These observed patterns align with our predictions in Section 2.3.2 that son preference will buffer the impact of famine on health for males but not in the case of females.

## 5.3. The Impact of the Gender Gap on Education

Table 4 presents the estimated effect of intrauterine famine exposure and its interacting effect with regional son preference on the gender gap in years of education. The specifications reported in each column of Table 4 are the same as those reported in Table 2. In columns 1–3, the dependent variable is the gender gap in years of education at the province birth year and month



level. Column 1 reports the estimates without pre-famine regional conditions. We find that intrauterine famine exposure significantly increases the gender gap in years of education. Moreover, we find that the interaction term between famine exposure and son preference is negative. These results remain robust when additionally including pre-famine regional controls in column 2. Column 2 shows that the change in the gender gap in years of education in response to a one standard deviation increase in famine severity in areas with son preference is 0.087 lower than in gender-neutral areas, equal to 6.41% of the mean (or 9.46% of the standard deviation). This means that famine shocks are expected to reduce the gender gap in education in areas with son preference, consistent with Proposition 3. In column 3, we re-estimate the model when excluding those conceived and born during the famine and observe robust estimates. We also report estimates using alternative ways to aggregate the dependent variable. Columns 4–5 report the results when the gender gap in years of education is aggregated at the province birth year level. The remaining two columns report the results when using the moving average of the gender gap in years of education (6 months before and after a particular birth year and month). Overall, our results remain robust to alternative ways to aggregate the dependent variable.

Next, we present the results of the fully flexible specification for the determinants of the gender gap in years of education. We plot the coefficients of the triple interaction terms in panel B of Figure 7. The patterns suggest that the relationship between famine shock interacting with regional son preference remains small in magnitude among pre-famine born cohorts and experiences a considerable decrease for cohorts born during the famine. It is worth noting that the largest impact of famine severity interacting with son preference is for those born in 1961 and conceived in 1960, the most severe year of the famine. Taken together, our findings suggest that the culture of son preference generates unintended consequences for the gender gap in years of education under adverse circumstances. Our findings are close to Yi et al. (2015), who find that household resource allocation responses exacerbate the adverse impact of adverse shocks to children on their educational attainment. In our context, household resource allocation (for example, schooling investment) between boys and girls may be responsive to famine experience.

Our findings on educational outcomes exhibit a nuanced relationship, which is far from a mechanical change in survivors' health. In Table A7, we document that our results on the gender gap in education remain statistically significant after controlling for the gender gap in health (height), albeit with a reduced magnitude. Moreover, Panel B of Table A8 reveals that famine exerts no discernible influence on schooling years for males, while it does lead to a reduction in schooling years for females. This pattern resonates cohesively with Section 2.3.3, emphasizing the ambiguity in the impact of famine on survivors' education, contingent on the ratio of famine

survivors' expected health to non-exposed cohorts and the parental response to children's famine experience. The negative coefficient for girls indicates that parents' unfavorable responses to children's famine experience dominate the change in expected health induced by famine shocks for females. In addition, the positive coefficient of the interaction term for males (despite insignificant) and females indicates that the negative impact of famine on years of education is more significant in gender-neutral areas compared to areas with son preference. The difference in the impact of famine on education between gender-neutral areas and areas with son preference for both gender hinges on the ratio of famine survivors' expected health to non-exposed cohorts between these two regions, which can be either positive or negative.

## 6. Robustness Checks

We conduct several exercises to ensure the robustness of the analysis. These robustness checks include: (i) validating the son preference measure, (ii) using alternative measures of key explanatory variables, (iii) ruling out the potential confounding effects of other policies or events, (iv) using alternative sample restrictions and including famine exposure during age 0–1, (v) a placebo test, and (vi) addressing the selection issue.

**Son preference measure.** We address several concerns related to our main son preference measure.

- (i) One possible issue is that our son preference measure may be influenced by the Chinese Civil War, probably reflecting the change in economic and social circumstances around the 1940s instead of local preference for sons.<sup>18</sup> To mitigate this, we construct an alternative measure using the sex ratios of younger cohorts aged 0-3, less likely to be influenced by the war. Columns 1 and 5 in Table 5 show our results are robust, even though the estimated impacts of famine become less precise.
- (ii) Another concern is raised related to the construction of the son preference dummy variable. We show that our results hold when using the continuous measure (columns 2 and 6) and an alternative cutoff of 1.06 to construct the dummy variable (columns 3 and 7) in Panel A of Table 5.

Furthermore, we demonstrate that the son preference measure exhibits no correlation with pre-famine local healthcare supply, economic development, and population size in Appendix Table A2. In addition, our results remain robust and significant when controlling for these conditions in Table A3. Finally, we switch to post-famine data to construct alternative measures of son

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<sup>18</sup> The Chinese Civil War was a civil war between the Chinese Communist Party (CCP) and Kuomintang (KMT) between 1929 and 1949.

preference in Panel B of Table 5: (i) according to the literature on the son-targeted fertility-stopping rule, we measure son preference using the sex ratios of higher-order birth and find our results reassuring (columns 1 and 4); (ii) our results hold when using the fraction of boys in expected children to proxy son preference (columns 2 and 5); and (iii) we find our results remain robust when using the opinions about gender norms to capture son preference (columns 3 and 6).

**Famine exposure measure.** In Panel A of Table 6, we show that our results are insensitive to using the average death rate in 1954–1957 (pre-famine death rate) as the death rate in normal years to construct a new measure of famine severity. We also use province-level actual death rates to construct famine exposure to capture the different timing of the famine across provinces. Panel B of Table 6 shows that our results remain robust and statistically significant when using provincial actual death rates to construct famine exposure.

**Other social changes.** We acknowledge the concern about the possibly confounding impacts of two major social changes: the Chinese Civil War and the Cultural Revolution. The Chinese Civil War is considered to affect individuals' survival chances (or cohort sex ratios). We use the province-level duration of civil war to capture its intensity. We include the interaction term between famine exposure and civil war duration in Table A4 (columns 1–2 and 5–6), and we find our results are still here. The Cultural Revolution may influence individuals' educational attainment. We get the measure of reform intensity from Walder (2014) based on recorded history from county gazetteers. The remaining columns of Table A4 show that our results remain significant and similar in magnitude compared to the main tables when including the controls of the Cultural Revolution.<sup>19</sup>

**Alternative sample restrictions and famine exposure received in the first year of life.** We further examine the robustness of our estimates by including cohorts born between 1967 and 1970. Appendix Table A5 shows that our results are robust to including cohorts who were born after the famine. We also show that our results are robust to adding famine exposure during age 0–1. We construct a measure of famine exposure received in the first year of life. We include both famine exposure variables and their interactions with son preference and pre-famine regional conditions in Appendix Table A6. Overall, our results hold when adding this control.

**Placebo test and selection during the famine.** Changes in provincial excess death rates should not be systematically correlated to other omitted factors that affect cohort sex ratios or educational attainment to ensure that difference-in-differences estimators remain free from bias.

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<sup>19</sup> The measure of the Cultural Revolution (CR) intensity is the normalized total number of abnormal deaths of each province during 1966–1976. One SD increase in CR intensity leads to a 0.04-percentage point decrease in sex ratios and a 0.002 in the gender gap in years of education.

We conduct a placebo test using a younger subsample (those born between 1966 and 1978). Since none of these birth cohorts were ever directly exposed to the famine, we expect that the Great Famine, as well as the interaction between famine severity and regional son preference, would not produce any substantial effects. Table 7 shows that this placebo test works well. We examine overall selection effects during the famine using intrauterine famine severity to predict individual characteristics, specifically family background and sibship size, using 2010 CFPS. Overall, Table A9 implies that the gender difference in overall selection effects during the famine is not a serious concern.

## **7. Discussion and Conclusions**

In this paper, we study the long-run impact of early life events and their interaction with social norms, in the setting of the Great Famine. We first put forward a theoretical framework to show how social norms can interact with adverse shocks to affect male and female survival chances that shape gender inequalities in later life. We empirically test our theoretical predictions using data from the 2000 population census and the 2010 CFPS.

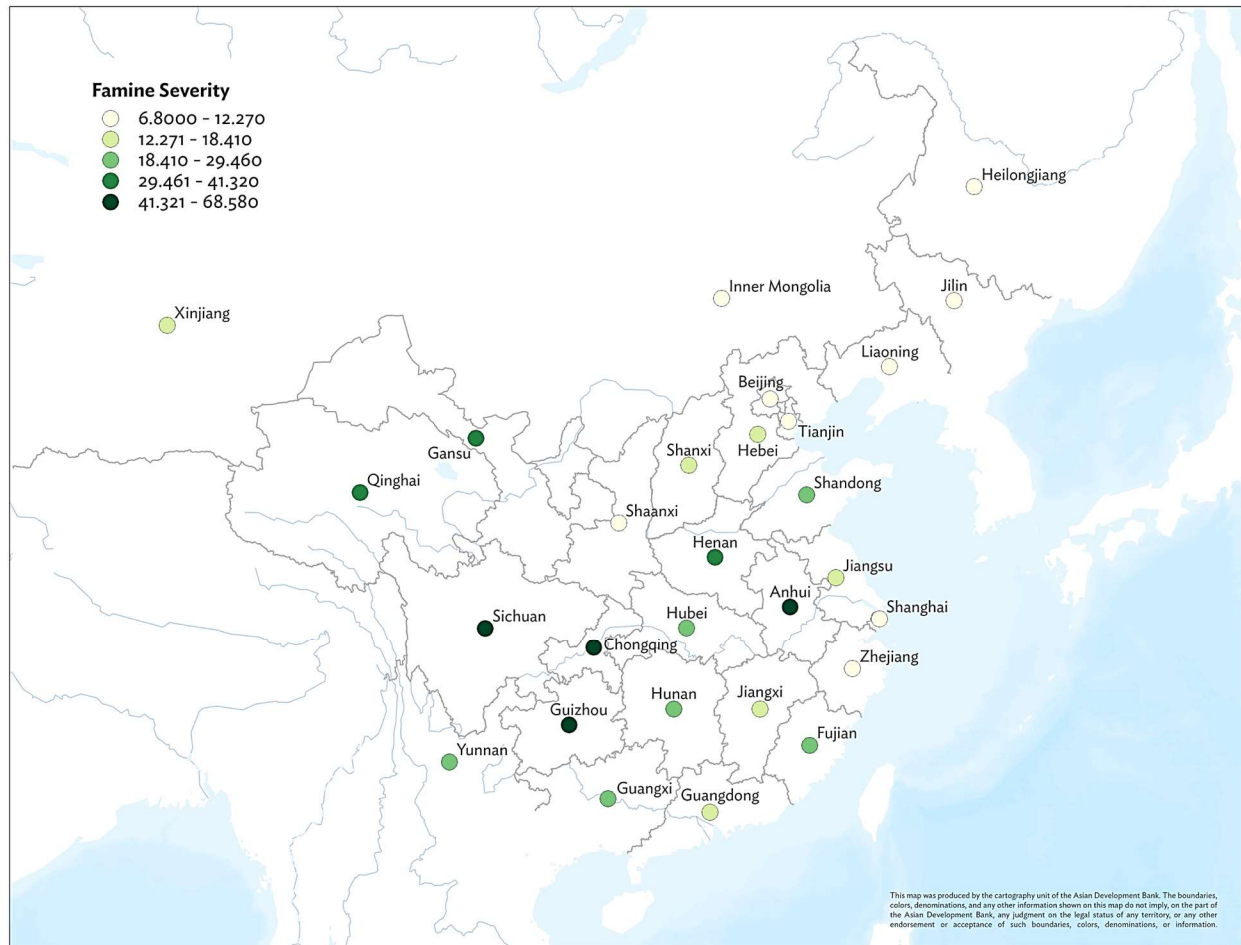
We find that even though intrauterine famine shocks contribute to a reduction in sex ratios (males to females), the change in male-to-female sex ratios induced by the famine shocks in provinces with son preference is much smaller compared to that in gender-neutral provinces, implying that the culture of son preference buffers the negative impact of famine shock on male survival chances and worsens the survival chances of girls during the famine. In addition, we find that even though intrauterine famine shocks increase the gender gap in height in gender-neutral areas, the interaction effect with son preference more than fully offsets the increase in the gender gap in height caused by the famine. Finally, we show that intrauterine famine shocks increase the gender gap in years of education, but this is less true in areas with a culture of son preference. Taken together, these findings suggest that the culture of son preference buffers the negative impact of famine shock on male-to-female sex ratios (survival chances) and also changes the gender gap in health and education among survivors. These results are proven not to be driven by selective fertility and pass several robustness tests, including alternative sample restriction, alternative measures of treatment variables, controlling for the impact of the Chinese Civil War and the Cultural Revolution, and a placebo test.

Our findings link two pieces of literature on the consequences of the cultural preference for sons and the lasting impact of early life events. We show that the culture of son preference can interact with early-life adverse shocks to shape gender inequalities in later life. Our theoretical framework and empirical findings, taken together, suggest that context is an important mediator

of how local environmental shocks play out with respect to gender. In societies with strong son preference, especially in East Asia, people may discriminate against girls, especially under adverse circumstances. However, the long-term consequences may be different.

## FIGURES AND TABLES

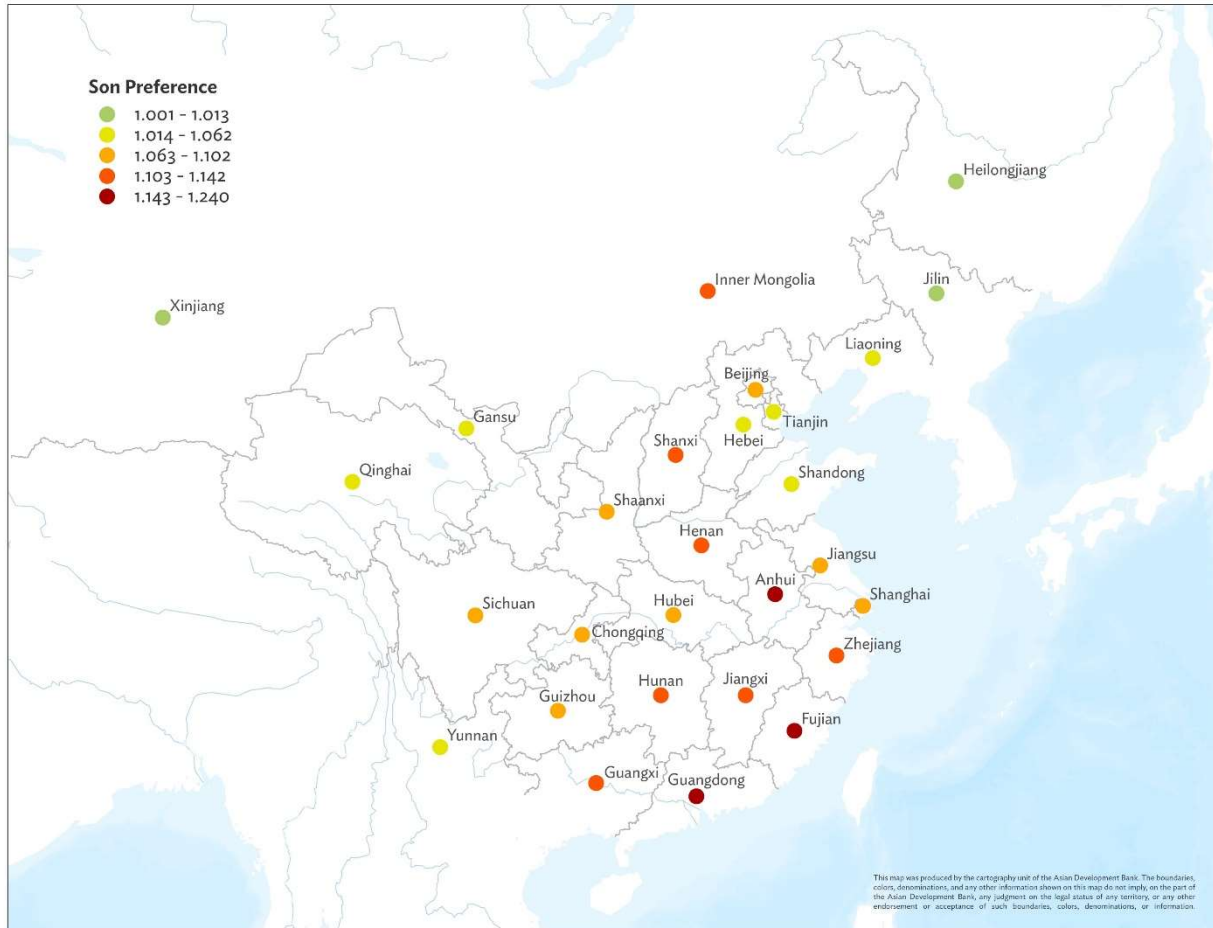
**Figure 1: Geographic Distribution of Famine Severity**



Note: This figure shows the geographic distribution of famine severity measured by the number of dead people (per thousand) in 1960.

Source: Statistical Yearbook from the National Bureau of Statistics of China.

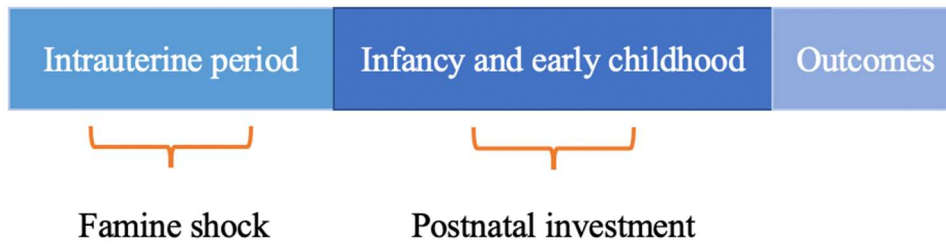
**Figure 2: Geographic Distribution of Son Preference**



Note: This figure shows the geographic distribution of the culture of son preference measured by male-to-female sex ratios of cohorts aged from 0–10 in the 1953 census.

Source: First National Population Census of the People's Republic of China, 1953.

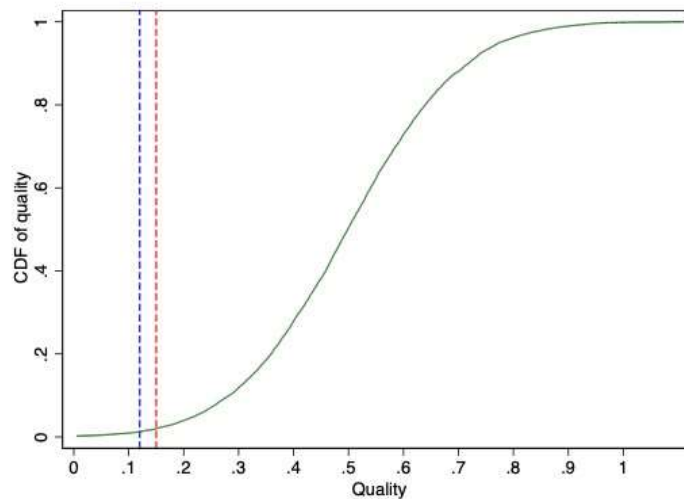
**Figure 3: Illustration of Model Period**



Note: The figure shows the timing in our theoretical model.

Source: Authors.

**Figure 4: Relative Position of Survival Thresholds to Health Distributions**

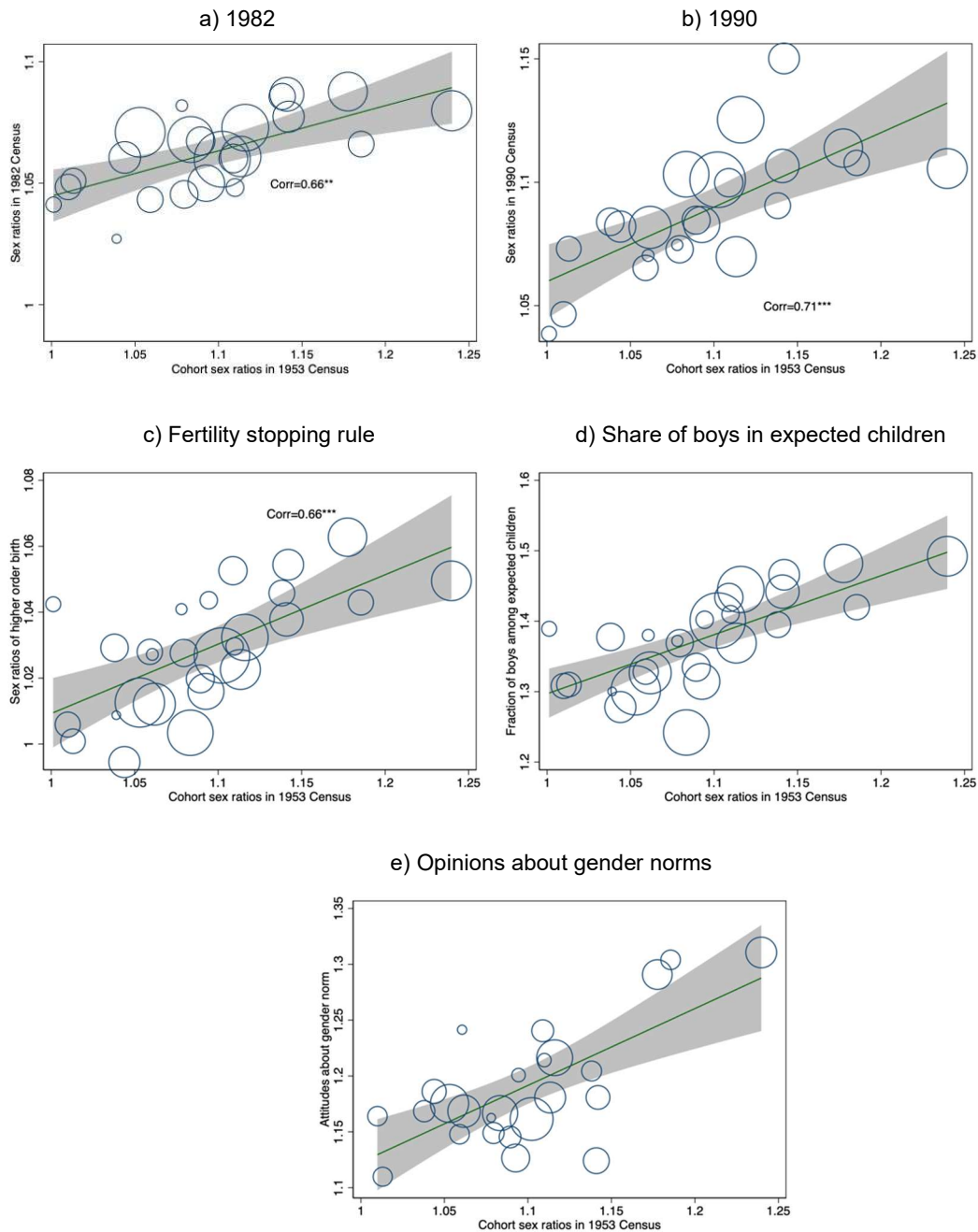


Notes: This figure plots the relative position of effective survival thresholds for both genders with respect to the cumulative distribution function of male health. The left blue line denotes the threshold for females, and the right red line represents that for males.

Source: Authors' estimates.



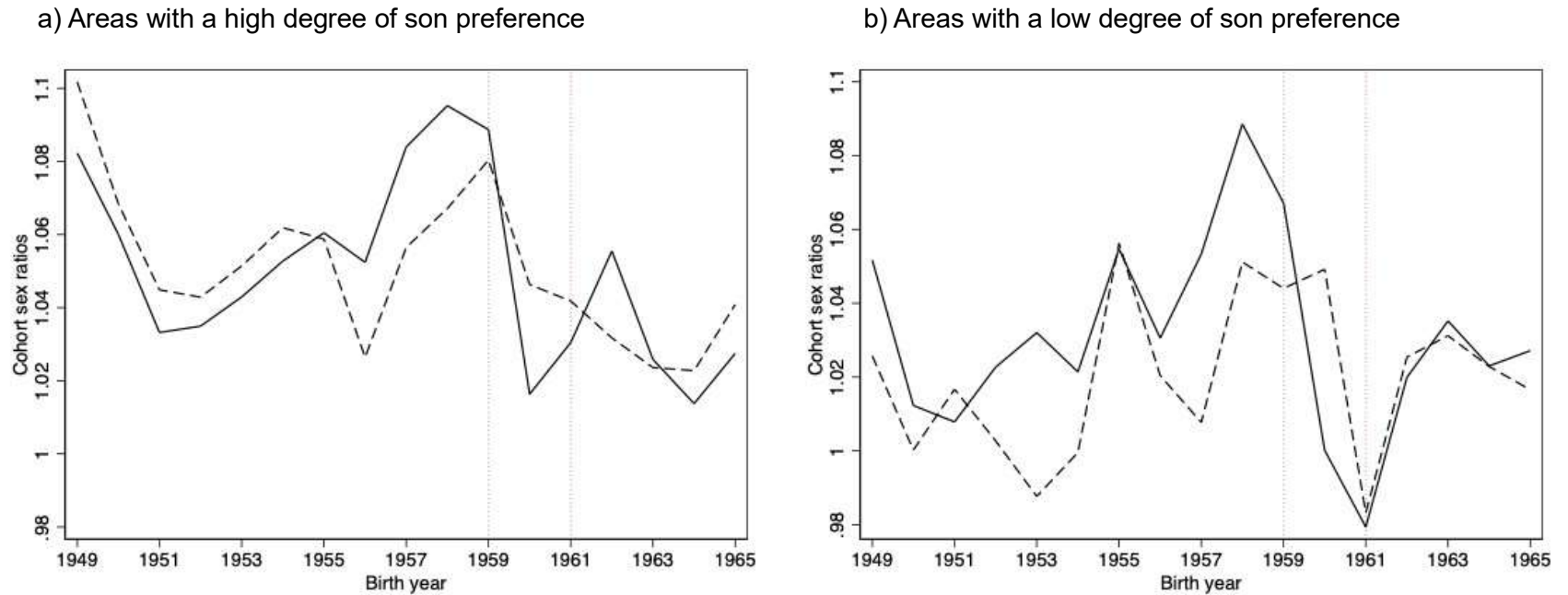
**Figure 5: Correlations Between Main Son Preference Measure and Alternative Measures**



Notes: The figure plots the correlations between cohort sex ratios in 1953 (aged 0–10) and sex ratios in the 1982 census (Panel A) and the 1990 Census (Panel B), sex ratios of higher-order births based on son targeted fertility stopping rule (Panel C), the share of boys in expected children in the 2001 Chinese National Family Planning and Reproductive Health Survey (Panel D), and opinions about gender norms in the 2014 China Family Panel Study (Panel E). Each circle represents a province. Circle size presents the size of the population in the 1953 Census.

Sources: Authors' estimates; China Family Panel Study, 2014; Chinese National Family Planning and Reproductive Health Survey, 2001; First National Population Census of the People's Republic of China, 1953; Third National Population Census of the People's Republic of China, 1982; Fifth National Population Census of the People's Republic of China, 1990.

**Figure 6: Sex Ratios by Famine Intensity and Regional Son Preference**

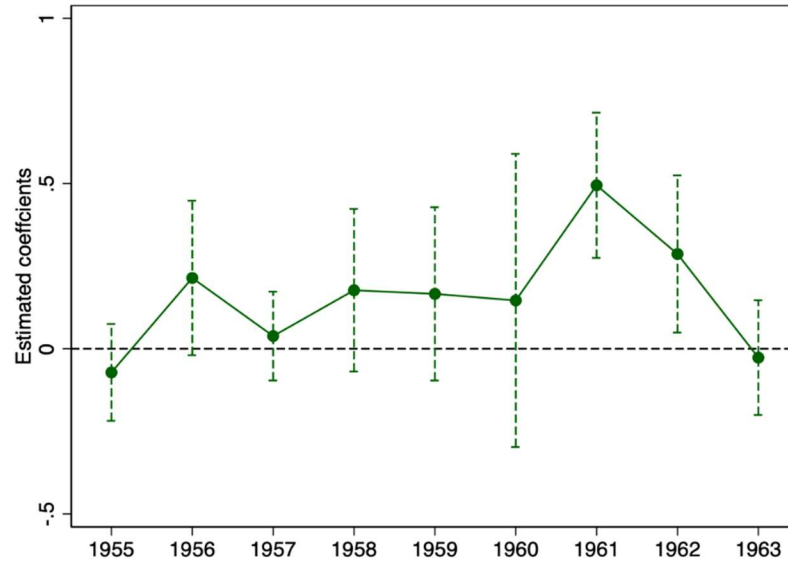


Notes: This figure shows cohort male-to-female sex ratios by famine severity and regional culture of son preference. The black solid lines denote cohort sex ratios in high famine intensity areas, and the black dashed lines indicate cohort sex ratios in low famine intensity areas. Cohorts born between 1959 and 1961 (within two red vertical lines) are famine-born cohorts.

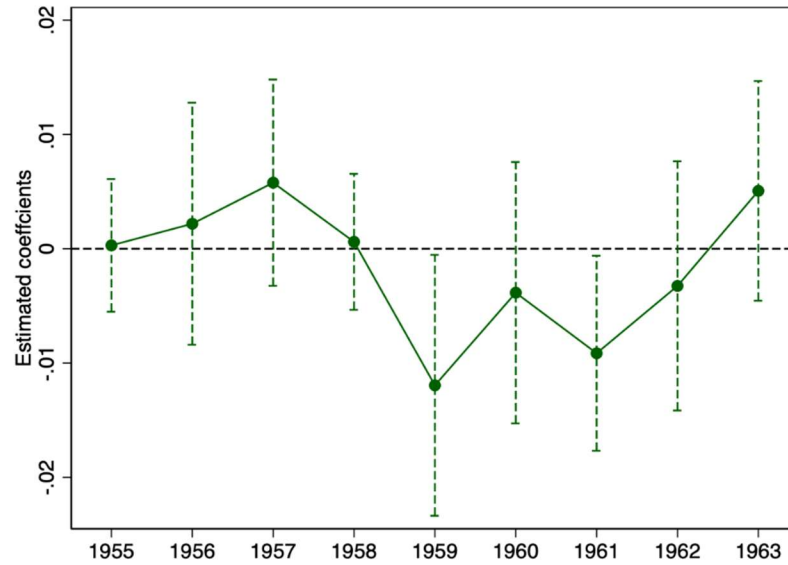
Source: Authors' estimates; Fifth National Population Census of the People's Republic of China, 2000; Statistical Yearbook from the National Bureau of Statistics of China.

**Figure 7: Flexible Estimates**

a) Cohort male-to-female sex ratios



b) Gender gap in years of education



Notes: The figure presents parameter estimates on interactions between birth year fixed effects, regional excess death rate during the famine and regional son preference on outcome variables, with 95% confidence intervals reported. We multiply sex ratios by 100. We include controls for pre-famine regional characteristics interacted with birth year fixed effects. Standard errors clustered at the province level are used to construct confidence intervals.

Source: Authors' estimates.

**Table 1: Effect of Famine on Survivors' Outcomes—Theoretical Predictions**

	Boys (1)	Girls (2)
Panel A: Effects of Famine on Survival Chances		
Gender-neutral	$F(z'_b{}^n) - F(z'_b{}^n + \Delta\mu)$	$F(z'_g{}^n - \Delta h) - F(z'_g{}^n - \Delta h + \Delta\mu)$
Son preference	$F(z'_b{}^p) - F(z'_b{}^p + \Delta\mu)$	$F(z'_g{}^p - \Delta h) - F(z'_g{}^p - \Delta h + \Delta\mu)$
Difference	<0	>0
Panel B: Effects of Famine on Survivors' Expected Health		
Gender-neutral	$\lambda(z'_b{}^n + \Delta\mu) - \lambda(z'_b{}^n)$	$\lambda(z'_g{}^n - \Delta h + \Delta\mu) - \lambda(z'_g{}^n - \Delta h)$
Son preference	$\lambda(z'_b{}^p + \Delta\mu) - \lambda(z'_b{}^p)$	$\lambda(z'_g{}^p - \Delta h + \Delta\mu) - \lambda(z'_g{}^p - \Delta h)$
Difference	>0	<0
Panel C: Effects of Famine on Survivors' Education		
Gender-neutral	$\ln \lambda(z'_b{}^n + \Delta\mu) + \ln(1 - \Delta\mu) - \ln \lambda(z'_b{}^n)$	$\ln \lambda(z'_g{}^n - \Delta h + \Delta\mu) + \ln(1 - \Delta\mu) - \ln \lambda(z'_g{}^n - \Delta h)$
Son preference	$\ln \lambda(z'_b{}^p + \Delta\mu) + \ln(1 - \Delta\mu) - \ln \lambda(z'_b{}^p)$	$\ln \lambda(z'_g{}^p - \Delta h + \Delta\mu) + \ln(1 - \Delta\mu) - \ln \lambda(z'_g{}^p - \Delta h)$

Source: Authors.

**Table 2: Estimated Impacts on Cohort Male-to-Female Sex Ratios**

	Male-to-Female Sex Ratios						
	Birth Year and Month Level			Birth Year Level		Moving Average	
	(1)	(2)	(3)	(4)	(5)	(6)	(7)
Famine	-0.2364*** (0.084)	-0.1064 (0.097)	-0.0935 (0.1353)	-0.1724*** (0.054)	-0.1094** (0.047)	-0.2273*** (0.081)	-0.0941 (0.091)
Famine*son preference	0.2640*** (0.078)	0.3367*** (0.059)	0.2556** (0.1126)	0.1881*** (0.055)	0.2489*** (0.052)	0.2523*** (0.073)	0.3219*** (0.057)
Pre-famine controls	No	Yes	No	No	Yes	No	Yes
Observations	4212	4212	3402	4212	4212	4212	4212
R-squared	0.165	0.166	0.178	0.466	0.470	0.497	0.502

Notes: This table reports the effect of intrauterine famine exposure and its interaction with son preference on cohort sex ratios. The unit of observation is at the province birth year and month level. In columns 1 through 3, the dependent variable is the male-to-female sex ratios of each cell at the province birth year and month level. In columns 4–5, we use dependent variables aggregated at province and birth year level. In columns 6–7, we use a moving average of the dependent variable (6 months before and after). We multiply sex ratios (dependent variable) by 100. Pre-famine regional control variables are provincial gross domestic product per capita, number of health institutions, and number of high school students in 1954 (subtracted from the median values). Province-fixed effects and birth year and month-fixed effects are controlled for. All models are weighted by the population size of each unit. Standard errors are clustered at the province level and reported in parentheses. \*\*\*p<0.01, \*\*p<0.05, \*p<0.1.

Sources: Authors' estimates; Comprehensive Statistical Data and Materials on 50 Years of New China; Fifth National Population Census of the People's Republic of China, 2000.

**Table 3: Estimated Impacts on the Gender Gap in Adult Height**

	Gender Gap in Height		
	(1)	(2)	(3)
Famine	0.2672*** (0.0784)	0.4167*** (0.0702)	0.4649*** (0.0914)
Famine*son preference	-0.3383*** (0.0697)	-0.4927*** (0.0750)	-0.5285*** (0.0858)
Pre-famine controls	No	Yes	No
Observations	265	265	229
R-squared	0.301	0.342	0.271

Notes: This table reports the effect of intrauterine famine exposure and its interaction with son preference on the gender gap in adult height. The unit of observation is at the province birth year level. Pre-famine regional control variables are provincial gross domestic product per capita, number of health institutions, and number of high school students in 1954 (subtracted from the median values). Province-fixed effects and birth year effects are controlled for. All models are weighted by population size of each unit. Standard errors are clustered at the province level and reported in parentheses. \*\*\*p<0.01, \*\*p<0.05, \*p<0.1.

Sources: Authors' estimates; China Family Panel Studies, 2010; Comprehensive Statistical Data and Materials on 50 Years of New China.

**Table 4: Estimated Impacts on the Gender Gap in Years of Education**

	Gender Gap in Education						
	Birth Year and Month Level			Birth Year Level		Moving Average	
	(1)	(2)	(3)	(4)	(5)	(6)	(7)
Famine	0.0117** (0.004)	0.0110*** (0.004)	0.0126** (0.0046)	0.0092** (0.003)	0.0100*** (0.003)	0.0109** (0.004)	0.0101** (0.004)
Famine*son preference	-0.0127*** (0.003)	-0.0148*** (0.003)	-0.0124*** (0.0034)	-0.0111*** (0.003)	-0.0132*** (0.003)	-0.0120*** (0.004)	-0.0143*** (0.003)
Pre-famine controls	No	Yes	No	No	Yes	No	Yes
Observations	4212	4212	3402	4212	4212	4212	4212
R-squared	0.821	0.823	0.831	0.955	0.957	0.955	0.957

Notes: This table reports the effect of intrauterine famine exposure and its interaction with son preference on the gender gap in years of education. The unit of observation is at the province birth year and month level. In columns 1 through 3, the dependent variable is the gender gap in years of education of each cohort at the province birth year and month level. In columns 4–5, we use dependent variable aggregated at province and birth year level. In columns 6–7, we use a moving average of the dependent variable (6 months before and after). Pre-famine regional control variables are provincial gross domestic product per capita, number of health institutions, and number of high school students in 1954 (subtracted from the median values). Province fixed effects and birth year and month fixed effects are controlled for. All models are weighted by population size of each unit. Standard errors are clustered at province level and reported in parentheses. \*\*\*p<0.01, \*\*p<0.05, \*p<0.1.

Sources: Authors' estimates; Fifth National Population Census of the People's Republic of China, 2000; Comprehensive Statistical Data and Materials on 50 Years of New China.

**Table 5: Robustness to Different Measures of Son Preference**

	(1)	(2)	(3)	(4)	(5)	(6)
	Male-to-Female Sex Ratios			Gender Gap in Education		
Panel A: Pre-famine measure						
Famine	0.0016	-2.4372***	-0.0825	0.0066*	0.1120***	0.0105***
	(0.097)	(0.8061)	(0.0995)	(0.004)	(0.0380)	(0.0035)
Famine*measurey	0.3034***			-0.0149***		
	(0.063)			(0.002)		
Famine*measurec		2.3171***			-0.1005***	
		(0.7470)			(0.0350)	
Famine*measure06			0.3357***			-0.0160***
			(0.0547)			(0.0024)
Observations	4212	4212	4212	4212	4212	4212
R-squared	0.166	0.165	0.166	0.823	0.822	0.823
Panel B: post-famine measure						
Famine	-6.2429***	-2.9176***	-1.7588***	0.3176***	0.1343***	0.0686***
	(2.205)	(0.9338)	(0.6231)	(0.101)	(0.0447)	(0.0184)
Famine*measure1	6.1028***			-0.3039***		
	(2.121)			(0.097)		

*Continued on the next page*



	(1)	(2)	(3)	(4)	(5)	(6)
	Male-to-Female Sex Ratios			Gender Gap in Education		
Famine*measure2		2.1415*** (0.6614)			-0.0940*** (0.0314)	
Famine*measure3			1.4855*** (0.5193)			-0.0539*** (0.0143)
Observations	4212	4212	3900	4212	4212	3900
R-squared	0.165	0.165	0.174	0.822	0.822	0.833

Notes: In Panel A, the following measures are considered: (i) *measurey* represents the sex ratios of younger cohorts aged 0–3 in the 1953 census who were less likely to be influenced by the war; (ii) *measurec* represents a continuous measure of sex ratios of cohorts aged 0–10 in the 1953 census; and (iii) *measure06* represents using an alternative cutoff of 1.06 to construct the dummy variable based on sex ratios in the 1953 census. In Panel B, we consider three sets of son preference measures using alternative data sources: (i) *measure1* denotes the sex ratios of higher-order birth based on the 1982 census; (ii) *measure2* represents the fraction of boys in expected children based on the Chinese National Family Planning and Reproductive Health Survey conducted in 2001; and (iii) *measure3* represents individuals' gender attitudes (or norms) using the CFPS data, with a higher value indicating more male-biased gender norms. We multiply sex ratios (dependent variable) by 100. The unit of observation is at the province birth year and month level. Pre-famine regional control variables are provincial gross domestic product per capita, number of health institutions, and number of high school students in 1954 (subtracted from the median values). Province-fixed effects and birth year and month-fixed effects are controlled for. All models are weighted by the population size of each unit. Standard errors are clustered at the province level and reported in parentheses. \*\*\*p<0.01, \*\*p<0.05, \*p<0.1.

Source: Authors' estimates; China Family Panel Studies, 2010; Chinese National Family Planning and Reproductive Health Survey, 2001; First National Population Census of the People's Republic of China, 1953; Third National Population Census of the People's Republic of China, 1982.

**Table 6: Robustness to Different Measures of Famine Exposure**

	Male-to-Female Sex Ratios		Gender Gap in Education	
	(1)	(2)	(3)	(4)
Panel A: alternative excess death rate				
Famine_excess	-0.3566** (0.1507)	-0.3282** (0.1298)	0.0205*** (0.0050)	0.0223*** (0.0039)
Famine_excess*son preference	0.3998** (0.1488)	0.4966*** (0.1464)	-0.0227*** (0.0047)	-0.0352*** (0.0039)
Pre-famine controls	No	Yes	No	Yes
Observations	4212	4212	4212	4212
R-squared	0.165	0.165	0.821	0.823
Panel B: actual death rate				
Famine_actual	-0.3566* (0.184)	-0.3576** (0.150)	0.0127** (0.006)	0.0127* (0.007)
Famine_actual*son preference	0.3600* (0.178)	0.4666*** (0.156)	-0.0129** (0.006)	-0.0111 (0.009)
Pre-famine controls	No	Yes	No	Yes

*Continued on the next page*

	Male-to-Female Sex Ratios		Gender Gap in Education	
	(1)	(2)	(3)	(4)
Observations	4212	4212	4212	4212
R-squared	0.165	0.167	0.820	0.821

Notes: Panel A reports the estimates using an alternative definition of excess death rate. *Famine\_excess* is constructed using the average death rate in 1954 and 1957 (pre-famine death rate) as the death rate in normal years. Panel B reports the estimates using the actual provincial death rate to construct famine exposure in utero. *Famine\_actual* denotes the intrauterine famine exposure based on the actual provincial death rate. Province-fixed effects and birth year and month-fixed effects are controlled for. All models are weighted by the population size of each unit. Standard errors are clustered at province level and reported in parentheses. \*\*\*p<0.01, \*\*p<0.05, \*p<0.1.

Source: Authors' estimates.

**Table 7: Placebo Tests**

	Male-to-Female Sex Ratios		Gender Gap in Education	
	(1)	(2)	(3)	(4)
Famine	−0.0010 (0.0751)	0.0246 (0.0921)	−0.0037 (0.0035)	−0.0014 (0.0031)
Famine*son preference	0.0298 (0.0628)	0.0486 (0.0709)	0.0034 (0.0031)	0.0051 (0.0032)
Pre-famine controls	No	Yes	No	Yes
Sample	1966–1978	1966–1978	1966–1978	1966–1978
Observations	4212	4212	4212	4212
R-squared	0.187	0.187	0.777	0.777

Notes: We construct a fake (placebo) famine shock that happened 13 years later (between 1971 and 1974) and conduct our analysis on cohorts who were born between 1966 and 1978. We multiply sex ratios (dependent variable) by 100. The unit of observation is at the province birth year and month level. Pre-famine regional control variables are provincial GDP per capita, number of health institutions, and number of high school students in 1954 (subtracted from the median values). Province-fixed effects and birth year and month-fixed effects are controlled for. All models are weighted by the population size of each unit. Standard errors are clustered at the province level and reported in parentheses. \*\*\*p<0.01, \*\*p<0.05, \*p<0.1.

Sources: Authors' estimates; Comprehensive Statistical Data and Materials on 50 Years of New China; Fifth National Population Census of the People's Republic of China, 2000.

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## **Social Norms and the Impact of Early Life Events on Gender Inequality**

This working paper analyzes the influence of social norms in determining the impact of early life exposure to the Great Chinese Famine of 1959–1961 on gender inequality. It uses a model to explain how social norms interact with adverse shocks to affect male and female survival chances and influence subsequent human capital investments using the 2000 population census of the People’s Republic of China. Results show that son preference buffers the negative impact of intrauterine famine shocks on cohort male-to-female sex ratios and reduces famine’s impact on gender inequality in health and education.

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