Does Buddhism Mitigate Sex Imbalance in Crisis? Evidence from China's Great Famine

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Abstract

This study examines the role of Buddhism in mitigating sex imbalance during and after China's Great Famine (1958-1966). Using the historical distribution of Buddhist temples as a proxy, we employ a double-difference strategy that leverages both temporal variation from the famine and cross-regional variation in Buddhist presence. A triple-differences model further explores the interactive effects between Buddhism and Confucian clans. Results indicate that the increase in male-biased sex ratios was significantly attenuated in areas with stronger Buddhist presence. Moreover, this mitigating effect was considerably stronger in areas with dense Confucian clan structures. These findings suggest that core Buddhist values helped counteract heightened clan influence and intensified son preference during the crisis, thereby promoting more balanced sex ratios.

Keywords: Buddhism, sex imbalance, Confucian clan, famine, China
JEL: J13, J16, Z12

1. Introduction

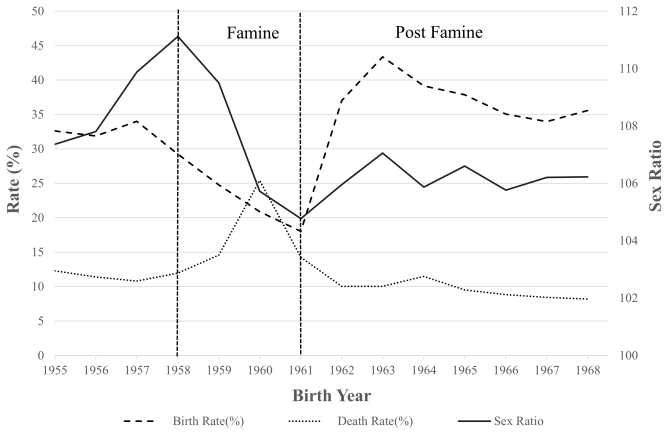
The adaptive sex ratio adjustment hypothesis proposed by Trivers and Willard (1973) in evolutionary biology posits that mothers in good physical or socioeconomic condition are more likely to give birth to sons, whereas

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those in poorer condition tend to have daughters. Although this hypothesis has been extensively tested in various animal species, empirical evidence for humans remains mixed (Keller et al., 2001; Lazarus, 2002; Stein AD, 2004; Song, 2012; Zhongwei Zhao, 2013). Figure 1 illustrates that China's national sex ratio declined sharply during the Great Famine (1958-1961) but rose steadily in the post-famine years (1962-1966), coinciding with a rapid rebound in birth rates—patterns often interpreted as supporting the Trivers-Willard hypothesis.[1] However, this hypothesis alone cannot account for the substantial regional variation in sex ratios observed across China during these periods. This paper presents evidence that Buddhism, a long-standing and influential religion in China, played a role in mitigating sex imbalance during and after China's Great Famine (1958-1966).[2]

Figure 1: National Sex Ratio, Birth Rate, and Death Rate of China



Data Source: Sex ratios are calculated using the 1990 Chinese population census. Birth and death rates are derived from (Lu and Zhai, 2009).

China's deep-rooted cultural preference for sons, strongly shaped by Confucianism, is widely recognized as a primary driver of the male-biased sex

[1]The sex ratio is defined as the number of males per 100 females in the reference population.

[2]We chose 1966 as the cutoff to avoid the influence of Chinas Cultural Revolution, which began that year and sought to suppress many traditional cultural practices, including Buddhism.

ratios. This preference has been embodied through Confucian clans, which have perpetuated Confucian ideals at the grassroots level since the Song dynasty (Chen et al., 2022). Son preference manifests in both prenatal and postnatal forms. Before technologies such as B-scan ultrasonography became available in the early 1980s, prenatal sex-selective abortion was virtually infeasible (Chen et al., 2013; Kashyap, 2018). In households with a strong preference for sons, families often employed the "stop rule"—continuing to have children until the desired number of sons was achieved—or resorted to postnatal sex-selective practices, such as the neglect, abandonment, or infanticide of daughters (Das Gupta, 1987; Miller, 1987; Jayachandran and Kuziemko, 2011). The "stop rule" required sufficient resources to support additional children, whereas postnatal methods were more likely to occur during periods of severe resource constraints, such as those caused by floods, earthquakes, or famines (Lee, 1981; Jimmerson, 1990).

While Confucianism centers the family as society's basic unit, Buddhism prioritizes individual salvation through ethical conduct and adherence to the Five Precepts.[3] Buddhism may help alleviate sex imbalance through at least two distinct channels. First, the principle of karma—the law of cause and effect—directly discourages practices such as sex-selective abortion and infanticide, thereby fostering a more balanced sex ratio. Second, its teachings on universal compassion and the equality of all beings may foster gender attitudes that diverge from Confucian patriarchy, improving female child survival over time. Establishing a convincing causal link, however, is challenging due to the potential endogeneity of religious influence and its complex interaction with prevailing Confucian norms.

Three main challenges arise as follows. First, the influences of Buddhism and Confucian clans are difficult to measure. Earlier studies have used affiliation rates, attendance, missionary activity, and self-reported religiosity from surveys to capture religious influence (Iannaccone, 1998). These measures, however, may not accurately reflect the influence of Buddhism in China due to its non-exclusive nature and the state's official secular stance. We follow a more recent strand of literature that constructs local-level measures of religious influence using historical data (Mehmood and Seror, 2023; Overstreet, 2024). Specifically, we use the number of Buddhist temples as a proxy for the

[3]The Five Precepts are commitments to abstain from killing, stealing, sexual misconduct, lying, and intoxication, forming the core moral code for Buddhist followers.

local strength of Buddhist influence. Similiarly, following Chen et al. (2022), we use the number of genealogies that compiled before the famine as a proxy for the influence of Confucian clans.

The second major challenge is identification. Buddhist influence evolves slowly with minimal temporal variation, making it hard to use time-based changes for causal inference. Thus, its primary variation is cross-sectional, and identification relies on the selection-on-observables assumption. To alleviate this concern, we adopt a double-difference strategy that compares the change in sex ratios (during and after the famine versus normal years) between prefectures with high and low Buddhist influence. Prefecture and year fixed effects are included to account for time-invariant unobservables and aggregate time shocks. This strategy mitigates endogeneity concerns by focusing on changes in differences rather than cross-sectional levels. Furthermore, to explore how Buddhist influence interacts with Confucian clan effects, we employ a triple-differences model. This specification captures the heterogeneous regional responses to famine by exploiting variations in both the extent of Buddhist influence and Confucian clan presence.

The third challenge involves adequately controlling for local famine intensity. A main threat to our identification strategie is bias arising from selection into treatment on observables. Prefectures with stronger Buddhist presence may systematically differ from others in their exposure to the famine. For example, such regions may might have benefited from more robust community networks or more favorable local policies, leading to lower mortality. However, to the best of our knowledge, systematic local level mortality data for the early years of the famine are unavailable. To mitigate this concern, we follow Yi et al. (2022) and construct a proxy for regional famine intensity based on cohort-specific population losses driven by plausibly exogenous weather shocks. This method offers two main advantages over prior indices. First, it nets out prefecture fixed effects to capture only exogenous weather-driven variation. Second, it circumvents the well-documented issue of misreporting in official sources (Chen, 2010; Walder, 2015).

Our findings reveal that Buddhism exerted a persistent moderating effect on sex ratios during and after the famine. Relative to the pre-famine baseline, prefectures with stronger Buddhist influence exhibited post-famine sex ratios that were 6.98 percentage points lower than those of prefectures with weaker Buddhist presence. Furthermore, in prefectures where both Buddhist and Confucian clan influences were strong, sex ratios were 17.42 percentage points lower than in prefectures dominated by Confucian clans but lacking

Buddhist presence. These results suggest that while Confucian clans intensified male bias during and after the famine, Buddhist influence mitigated it by discouraging sex-selective behavior and fostering more gender-equitable cultural norms.

We next examine potential mechanisms underlying our findings. First, we test whether Buddhism mitigated sex selection by alleviating famine severity during the famine years. Second, using a provincial-level survey data, we show that Buddhist affiliation is associated with smaller gender gaps in early-life resource access during the famine, suggesting a mitigating effect on adverse shocks for female children. Third, among individuals exposed to the Great Famine during early childhood, we find that Buddhist affiliation is associated with attenuated son preference, particularly among women, consistent with a moderation of gender-biased postnatal behavior under extreme resource constraints. Fourth, census evidence indicates that, relative to prefectures dominated by Confucian clan institutions alone, regions where Buddhism coexisted with strong clan influence experienced higher post-famine female survival rates. Together, these results suggest that Buddhist values moderated the demographic consequences of patriarchal kinship systems by promoting more gender-inclusive attitudes and behaviors during and after the famine periods.

This paper makes three main contributions to the literature. First, this paper contributes to the literature on resource shocks and gender disparities by introducing a cultural perspective. Using China's 1982 One-Per-Thousand Fertility Survey, Song (2012) documented an abrupt decline in the sex ratio at birth during the Great Famine, followed by a compensatory rise in the post-famine period, consistent with the adaptive sex ratio adjustment hypothesis. However, that study overlooks regional cultural heterogeneity, thereby masking important variation across local contexts. Recent work underscores this point: Malein and Tapia (2024) finds that Russia's 1891-1892 famine increased sex ratios in districts characterized by patriarchal oral traditions, while Luo et al. (2024) shows that son preference in China's famine buffered boys but harmed girls. In contrast, Pitkänen (1993) reports no comparable effect during the Great Finnish Famine. Our study extends this literature by demonstrating that cultural heterogeneity fundamentally shapes demographic responses to resource shocks.

Second, this study emphasizes that incorporating religion is essential for a more comprehensive understanding of China's long-standing sex imbalance. Prior research attributes the post-1980s rise in sex ratios primarily to fac-

tors such as gender wage disparities (Qian, 2008), the one-child policy (Li et al., 2011), access to prenatal ultrasound technology (Chen et al., 2013), and the patriarchal lineage structures embedded in Confucian clans (Zhang and Ma, 2017). However, these studies largely focus on the national trend and underexplore the spatial heterogeneity underlying earlier periods. In particular, the causes of pronounced sex imbalances observed in many regions during the 1960s remain insufficiently understood, as birth control policies were relatively lenient and sex-selective technologies were not yet available. By incorporating a religious dimension, our analysis sheds light on how culturally embedded moral norms moderated sex selection behavior well before the diffusion of modern reproductive technologies.

Third, this paper engages with the surprisingly limited research attention devoted to Buddhism. Despite its vast global following, empirical studies on Buddhism remain far fewer than those on Hinduism and are markedly less prevalent than analyses of Christianity and Islam (Iyer, 2016). This imbalance may partly reflect the enduring dominance of Confucianism as the prevailing ideological and ethical framework in East Asian societies such as China, Japan, and South Korea. In addition, through the long process of Sinicization, Buddhism has often been absorbed into and perceived as an intrinsic component of traditional Chinese culture, rather than as an independent religious system. Consequently, scholars have tended to prioritize Confucianism while overlooking Buddhisms autonomous role in shaping social and demographic outcomes in China (Root, 2024). Likewise, the limited literature examining Buddhism and sex ratios in East Asia often omits, or insufficiently distinguishes, the influence of Confucianism, thereby contributing to inconsistent empirical findings (Kim and Song, 2007; Skirbekk et al., 2015). By explicitly accounting for Confucian influence through measures of clan strength, this paper isolates the specific contribution of Buddhism to observed sex-ratio patterns and, in doing so, addresses an important gap in the existing literature.

The remainder of the paper proceeds as follows. Section 2 provides historical background on Buddhism, son perference culture, and the Great Famine. Section 3 details the data sources and key variable measurements. Section 4 outlines our identification strategies. Sections 5 and 6 present the main findings and robustness checks. Section 7 examines the underlying mechanisms, and the final section concludes.

2. Background

2.1. Buddhism in China

Buddhism, the most widespread and influential religion in China's history, has profoundly shaped nearly every aspect of Chinese intellectual, moral, and social life (Chen, 1973). Introduced during the Han dynasty (206 BCE-220 CE), Buddhism gradually permeated both elite Confucian circles and the broader populace.[4] Alongside Confucianism and Taoism, it is considered one of the three pillars of traditional Chinese culture. In China, Buddhism is primarily categorized into three main branches: Han Buddhism (also known as Chinese Buddhism), Tibetan Buddhism, and Theravada Buddhism. Among these, Han Buddhism is the most widespread, as reflected in the number and geographic distribution of registered temples across the country. Tibetan Buddhism and Theravada Buddhism are predominantly practiced by ethnic minorities in regions such as the Tibetan Plateau and Inner Mongolia. In everyday Chinese usage, the term "Buddhism" generally refers to Han Buddhism, which is the focus of this paper.

When Buddhism was first introduced to China, it encountered fundamental tensions with the hierarchical and family-centered values of Confucianism (Chen, 1973). Over time, however, both Chinese scholars and foreign monks adapted Buddhist doctrines to fit prevailing cultural norms, enabling the religion to gain legitimacy and become deeply embedded within Chinese society. This process of Sinicization gave rise to Han Buddhism, a distinct tradition that integrated elements of Confucian ethics and local customs while retaining Buddhisms core spiritual principles. Several of these core values may have played a crucial role in curbing sex-selective behaviors during periods of crisis, when moral and social structures were under severe strain.

A central tenet of Buddhism is the principle of karma, which holds that an individual's actions determine their future states of existence.[5] Life is viewed as an endless cycle of rebirths shaped by karma accumulated in past lives,

[4]The influence of Buddhism is often underestimated. Since its arrival during the Han dynasty, it surpassed Taoism in philosophical sophistication. Without Buddhism's intellectual challenge, Taoism might not have achieved comparable prominence, as rulers frequently employed it to balance Buddhisms growing influence (Denis and eds., 1986; Ji, 2007). Additionally, Buddhism played a crucial role in the theoretical development of Neo-Confucianism during the Song dynasty (Li, 2009).

[5]There are five states of existence: deity, human, animal, hungry ghost, and denizen of hell. The first two are considered good states, while the last three are evil.

with ones actions in the present determining the form of future existence. The ultimate goal is to escape this cycle by attaining Nirvana, the extinction of desire and liberation from rebirth.[6] To achieve this, adherents must avoid generating negative karma by following the Five Precepts.

Another core set of Buddhist doctrines centers on Karuna (universal compassion) and sattva-samata (the equality of all beings). Karuna embodies the aspiration to alleviate the suffering of all sentient beings, transcending distinctions of kinship, caste, gender, or species.[7] It is grounded in the belief that all life forms are interconnected through the continuous cycle of birth, death, and moral causation. The doctrine of sattva-samat holds that every sentient being possesses Buddha-nature, implying that men and women share an equal capacity for enlightenment. Together, these principles prohibit actions that inflict harm—such as killing, exploitation, or discrimination—and render practices like infanticide or sex-selective abortion morally indefensible, as they generate negative karma and hinder spiritual progress.

The influence of Buddhism remains significant today. According to the 2018 Chinese General Social Survey (CGSS), Buddhism is the most common religious identity in China, with 4% of Chinese adults (approximately 42 million people) formally identifying as Buddhist, and a third of Chinese adults engaging in Buddhist beliefs or practices.[8]

2.2. Confucianism, family clans and son preference culture

Confucianism, a philosophical and ethical system founded by Confucius, became the state ideology during the Han dynasty and remained dominant until the end of the Qing dynasty (1644-1911). Serving as the foundation of education in China for over two millennia, Confucianism is deeply intertwined with family clans and the culture of son preference, which can be understood from three key perspectives.

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[6]Nirvana is the term used in Buddhism to describe the ultimate soteriological goal: the extinguishing of passions and liberation from the cycle of rebirth (Meister, 2009).

[7]In Buddhism, sentient beings or living beings are beings with consciousness, sentience, or in some contexts life itself (Buswell, 2004).

[8]Formal affiliation does not fully capture the extent of Buddhist belief and practice. In China, many individuals who engage in Buddhist beliefs and practices do not formally identify as Buddhist because they practice Buddhism alongside Taoism or folk religion, believe only monks or nuns have the right to call themselves Buddhist, or because Buddhism does not require formal affiliation.

First, Confucianism institutionalizes patriarchy by emphasizing male authority within both family and society (Tai, 2021). Confucian philosophy asserts that the smooth functioning and harmony of society depend on five cardinal relationships: ruler-subject, father-son, husband-wife, elder brother-younger brother, and friend-friend. These relationships underscore a hierarchical and patriarchal structure, positioning men in dominant roles over women. Canonical texts such as the Book of Rites prescribe womens obedience and modesty as virtues, while practices like primogeniture ensured that property and lineage remained within the male line.

Second, Confucianism regards filial piety as a cornerstone of its ethics. In Hsiao-ching, or the Classic on Filial Piety, it is stated, "the filial son serving his parents should be most respectful to them while living, most joyful in supporting them, greatly worried at their illness, deeply grieved at their death, and utterly solemn at the sacrifices" and "of the 3,000 offenses included under the five punishments, none is greater than unfilial conduct." This sentiment is echoed in Chinese proverbs, such as: "There are three kinds of unfilial piety, the greatest is not having sons." These teachings reinforced the moral necessity of male offspring as bearers of family continuity.

Third, ancestor worship further entrenched son preference. In Chinese families, filial obligations extended beyond the living to include the veneration of ancestors (Watts, 1989). These rituals embodied the Confucian virtue of reciprocity, expressing gratitude for the care and sacrifices of previous generations. Because the continuation of ancestral rites depended on male heirs, the absence of sons was perceived not only as a personal shortcoming but also as a moral and social failure that threatened the continuity of lineage and family honor.

2.3. China's Great Famine

China's Great Famine was a devastating period of mass starvation from 1958 to 1961, regarded as one of the most severe famines in human history.[9] Estimates of excess deaths range from 16.5 to 30 million (Coale, 1981, 1991; Aird, 1982; Banister, 1984; Lin and Yang, 2000).

[9]The Chinese government designates 1959-1961 as the "Three Years of Natural Disasters," but scholars often include 1958 as part of the famine period, as the mortality rate had already started to rise that year (Ashton et al., 1984; Chang and Wen, 1997; Cao et al., 2022).

Many scholars argue that the Great Famine was closely linked to the Great Leap Forward, a radical socio-economic campaign launched by the Chinese Communist Party in 1958 (Peng, 1987; Kung and Lin, 2003; Kung and Chen, 2011). Intended to accelerate China's transformation from an agrarian economy to an industrialized socialist state, the campaign introduced sweeping institutional changes that severely disrupted agricultural production. The establishment of people's communes abolished private land ownership and centralized all farming activities under state control, while exaggerated grain reporting led to excessive state procurement and grain shortages in rural areas (Li and Yang, 2005; Meng et al., 2015). The crisis was further aggravated by adverse weather conditions, including floods and droughts, which compounded the decline in agricultural output (Ashton et al., 1984). Collectively, these political, institutional, and environmental failures converged to produce the most devastating famine in modern Chinese history.

The famine not only caused widespread starvation but also generated pronounced gender imbalances, as households with a strong preference for sons often prioritized the survival of male children (Luo et al., 2024). Under conditions of extreme scarcity, this intensified son preference compelled families to make devastating choices about resource allocation. Many directed limited food and care toward boys, viewing sons as essential to household survival and future security. Consequently, female infants and young girls suffered disproportionately high mortality rates, leading to a markedly male-biased sex ratio among survivors (Zhao and Reimondos, 2012).

3. Data and Measurement

This section presents the principal data sources and core variable constructions. The empirical analysis integrates three complementary datasets, producing a final sample of 249 prefectures and more than two million observations. This extensive coverage affords substantial geographic and demographic variation, enabling a rigorous examination of regional disparities in sex ratios during and after the Great Famine.

3.1. Buddhist temples

Our primary source for religious data is the Chinese National Religious Affairs Administration, which documents 42,438 registered religious temples nationwide, including 34,089 Buddhist and 8,349 Taoist temples. The dataset

provides detailed information on temple names, locations, and religious affiliations. Although the construction dates of these temples are unavailable, making it difficult to exclude those built after the famine, the establishment of new temples in China is tightly regulated. Under current religious administration rules, any new temple must demonstrate a historical foundation, such as cultural relics, ruins, or documented lineage. Hence, this dataset likely serves as a reasonable proxy for the spatial distribution of Buddhist temples prior to the famine. To validate robustness, we also use data from the 2004 Economic Census, which records 72,887 sites across the five officially recognized religions (Buddhism, Taoism, Islam, Protestantism, and Catholicism), covering all 31 provinces in China.

We use the number of Buddhist temples as the primary measure of local Buddhist influence. Beyond serving as places of worship, temples function as community centers where religious practice intertwines with local customs, festivals, and social life, thereby reinforcing the persistence of Buddhist culture. To quantify this influence at the prefecture level, we construct two indicators: (1) the logarithm of the number of Buddhist temples per 10,000 residents, and (2) a binary variable indicating whether this value is above or below the national median. Given our focus on Han Buddhism, we exclude Tibet, Inner Mongolia, and Xinjiang, where Tibetan Buddhism or Islam predominates.

3.2. Genealogies

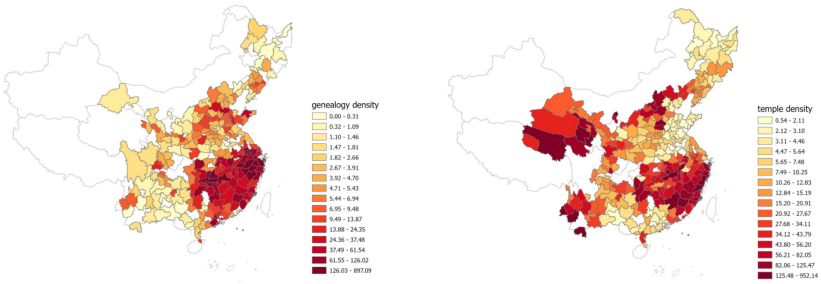
Following Chen et al. (2022) and Cao et al. (2022), we use genealogical data to assess the strength of Confucian clans across prefectures. This data comes from the Clan Culture Database, a specialized repository that digitizes records from the General Catalogue of Chinese Genealogy, compiled by the Shanghai Library. The database contains over 50,000 genealogical records dating up to the end of 2004, covering 608 surnames across China. Each record includes key details such as surname, location, starting year, and data source. With the 100 most common surnames accounting for over 85% of Chinas 1.3 billion population, this dataset serves as a valuable tool for analyzing the regional influence of Confucian clans.

We measure the influence of Confucian clans at the prefecture level by analyzing the total number of genealogies compiled in each region. To mitigate potential endogeneity concerns, we focus on genealogies compiled before 1960. We construct two measures of clan strength: (1) the logarithm of the number of pre-1960 genealogies, normalized by the prefectures population

size, and (2) a binary variable indicating whether this measure is above or below the national median. Given the highly left-skewed distribution of the first measure, as noted by Cao et al. (2022), we primarily rely on the second measure in our analysis.

Figure 2 illustrates the geographic distribution of Buddhist temples and genealogies based on the first measure. Notably, prefectures with high temple density often overlap with areas characterized by extensive genealogical records, underscoring the need to account for Confucian clan influence when estimating the effect of Buddhism on sex ratios.

Figure 2: Geographical description of genealogies and Buddhist temples



(a) Genealogies (b) Buddhist temples

Note: Xinjiang, Tibet and Inner Mongolia are excluded.
Data source: General Catalogue of Chinese Genealogy and Chinese National Religious Affairs Administrations.

3.3. Sex Ratios in China, 1954-1966

The analysis of sex ratios draws on the 1% sample of the 1990 China Population Census (1990 Census).[10] The dataset provides information on individuals sex, birth year, ethnicity, registration type, education, and relationship to the household head. Our sample comprises individuals born

[10]We opt for the 1990 Census over earlier and later counterparts for two key reasons. First, the 1982 census suffered from substantial underreporting of births and deaths, partly driven by efforts to retain ration allocations and evade penalties under the newly implemented One-Child Policy (Kennedy and Shi, 2019). Second, the 1990 Census is preferred to its 2000 counterpart because migration was more restricted in the earlier period, and its reliability is widely recognized (Lavely, 2001; Zhang and Zhao, 2006).

between 1954 and 1966 who, according to the 1990 Census, had resided in the prefecture of their hukou registration for at least five years. Given the strict hukou system prior to the 1990s, internal migration was highly constrained, thereby minimizing concerns about selective migration bias in our estimates (Qian, 2008; Li et al., 2011; Meng et al., 2015).

Table 1 presents summary statistics used in our empirical analysis. Prefectures are divided into two groups based on the national median of Buddhist temples per 10,000 residents. Mean values for key variables are shown separately for the periods before and after 1958, with standard errors in brackets. The dataset contains roughly 1.1 million individual observations for each group across 249 prefectures. Prefectures with higher Buddhist density tend to have slightly younger populations, a higher proportion of Han ethnicity, a greater share of residents with rural hukou status, and more likely to have completed only primary education. These demographic characteristics remained relatively stable after 1958, indicating no major compositional shifts associated with Buddhist density. However, the fraction of male births (our main outcome variable) varied over time across prefectures. The difference in the male birth share between high- and low-Buddhist prefectures changed from 0.1 percentage points before the famine to -0.5 percentage points afterward, corresponding to a shift in sex ratios from 0.44 to -2.13 percentage points. This pattern provides preliminary evidence that Buddhism may have mitigated male-biased sex ratios in the early post-famine years.

4. Identification Strategies

4.1. Effect of Buddhism on Sex Ratios

We begin by employing a double-difference strategy that exploits the timing of the famine as well as the cross-sectional differences in the measure of Buddhist influence. Specifically, we estimate the following model:

$$S_{ict} = \beta BH_c \times FM_t + \gamma Intensity_c \times FM_t + \theta W_{ict} + \delta_c + \lambda_t + \varepsilon_{ict} \quad (1)$$

where S_{ict} is a child's gender indicator, equal to 1 for boys and 0 for girls, born in prefecture c in year t. BH_c is a binary variable indicating whether the density of Buddhist temples in a prefecture is above the national median, and FM_t is a dummy for the famine period of 1958-1966. $Intensity_c$ denotes famine intensity in prefecture c, which is constructed based on cohort-specific

population losses attributable to plausibly exogenous weather shocks.[11] W_{ict} is a vector of individual characteristics, including age, ethnicity, hukou type, and education level.[12] Prefecture and year fixed effects (δ_c and λ_t) account for time-invariant prefecture characteristics and common temporal shocks, respectively. ε_{ict} denotes the idiosyncratic error term. Our key estimand, β, identifies the differential effect of the famine. It quantifies the extent to which prefectures with strong Buddhist influence experienced a different change in the proportion of male births, compared to those with weak influence prefectures, from the pre- to the famine and post-famine periods. A statistically negative β suggests that a stronger Buddhist presence was associated with a mitigated increase in sex ratios in response to the famine.

To separately estimate the effects of Buddhism on sex ratios during the famine (1958-1961) and post-famine (1962-1966), we estimate the following specification:

$$\begin{aligned} S_{ict} =& \beta_1 BH_c \times FM_{1t} + \beta_2 BH_c \times FM_{2t} + \gamma Intensity_c \times FM_t \\ &+ \theta W_{ict} + \delta_c + \lambda_t + \varepsilon_{ict} \end{aligned} \tag{2}$$

where FM_{1t} and FM_{2t} are dummy variables for the famine and post-famine periods, respectively. In this specification, β_1 measures the differential change in male birth shares during the famine years in high- versus low-Buddhist prefectures, while β_2 captures the corresponding effect in the post-famine period. Comparing the magnitudes of these two coefficients allows us to assess whether Buddhism's moderating effect on gender imbalance persisted beyond the immediate crisis period.

The estimation strategy in Equation (1) controls for time-invariant and prefecture-specific unobservables and common time shocks via fixed effects. However, it may still be subject to two sources of bias. The first arises from time-varying confounders—factors that evolve over time and correlate with both the outcome and the key interaction term. For instance, attitudes toward son preference may diverge across prefectures with differing Buddhist temple densities and could thus shape male birth shares during and in the aftermath of the famine. The second source of bias stems from cross-sectional

[11] Details on the construction methodology are provided in the Appendix A.

[12] Although education level is uncorrelated with gender at birth, it is correlated with survival chances during the famine, as only families with resources beyond basic necessities could afford to educate their children.

heterogeneity: time-invariant prefecture characteristics that are correlated with our measures of Buddhist influence and induce differential responses to the famine. Such heterogeneity is captured by interaction terms of the form $Z_c \times Famine_t$, where Z_c denotes a vector of unobserved, time-invariant characteristics of prefecture c. Notably, this second bias constitutes a special case of the first, as it can be framed as a time-variant confounder. Our identification strategy departs from the classical difference-in-differences framework in a key respect: the Great Famine constituted a universal shock across all prefectures, while the treatment, Buddhist influence, exhibits only cross-sectional variation.

To rule out the first source of biases concern, we estimate the following event-study model that allows for dynamic treatment effects:

$$\mathrm{S}_{ict} = \sum_{t=1954}^{1966} \beta_t BH_c \times D_t + \gamma Intensity_c \times FM_t + \theta W_{ict} + \delta_c + \lambda_t + \varepsilon_{ict} \quad (3)$$

where D_t denotes a series of year dummies for 1954-1966. This specification enables us to test whether the pre-famine coefficients ($\beta_{t<1958}$), are statistically indistinguishable from zero, as implied that no differential pre-trends exist between high- and low-Buddhist prefectures. To further address this concern, we employ a group interactive fixed effect model, which includes the unobservable interaction term $\delta_c \times \lambda_t$, in robustness check section.

4.2. Effects of Buddhism and Confucian Clans on Sex Ratios

We further extend the analysis by employing a triple-difference regression model to investigate the interactive effect of Buddhism and Confucian clans. As described earlier, the density of Buddhist temples serves as a proxy for Buddhist cultural strength, while the density of genealogies captures the influence of Confucian clans. To distinguish these two cultural dimensions, we classify prefectures into four categories based on the density rankings of temples and genealogies: strong Buddhist influence, strong clan influence, weak Buddhist influence, and weak clan influence. Specifically, we use the one-third and two-thirds percentiles of each distribution as cutoffs and exclude prefectures within the middle range to sharpen contrasts. The indicator BH_c equals 1 if a prefectures Buddhist temple density lies above the two-thirds percentile, indicating strong Buddhist influence, and 0 otherwise. Similarly, CH_c equals 1 for prefectures above the two-thirds percentile in genealogy

density, representing strong clan influence, and 0 otherwise. We then estimate the model as follows:

$$\begin{aligned} S_{ict} =& \beta_1 BH_c \times FM_t + \beta_2 CH_c \times FM_t + \beta_3 BH_c \times CH_c \times FM_t \\ &+ \gamma Intensity_c \times FM_t + \theta W_{ict} + \delta_c + \lambda_t + \varepsilon_{ict}, \end{aligned} \tag{4}$$

where BH_c and CH_c are the top-quantile indicators defined above. All remaining variables are defined as before. Our primary coefficient of interest is β_3. A negative β_3 indicates that, from the pre- to post-famine period, the increase in sex ratios was attenuated in prefectures with strong Buddhist influence relative to those with weak Buddhist influence, both within the subset of areas dominated by Confucian clans. In effect, β_3 captures the differential effect of Buddhist influence on sex ratio changes during and after the famine, conditional on strong Confucian clan presence.

To estimate the separate effects of Buddhism on sex ratios during and after the famine, while controlling for Confucian clan influence, we estimate the following specification:

$$\begin{aligned} S_{ict} =& \beta_1 BH_c \times FM_{1t} + \beta_2 BH_c \times FM_{2t} + \beta_3 BH_c \times CH_c \times FM_{1t} \\ &+ \beta_4 CH_c \times FM_{1t} + \beta_5 CH_c \times FM_{2t} + \beta_6 BH_c \times CH_c \times FM_{2t} \\ &+ \gamma Intensity_c \times FM_t + \theta W_{ict} + \delta_c + \lambda_t + \varepsilon_{ict}, \end{aligned} \tag{5}$$

where β_3 and β_6 capture the additional moderating effect of Buddhism on sex ratios in regions with strong Confucian clan influence during each period. A negative β_3 suggests that Buddhism helped mitigate the patriarchal influence of Confucian clans during the famine, resulting in a smaller increase in the fraction of male birth. Similarly, a negative β_6 indicates that this moderating effect persisted during the post-famine, implying a sustained role for Buddhism in promoting gender balance.

To address potential bias from unobserved time-varying confounders, we estimate the dynamic treatment effects analog of Equation (3):

$$\begin{aligned} S_{ict} =& \beta_1 BH_c \times FM_t + \beta_2 CH_c \times FM_t + \sum_{t=1954}^{1966} \beta_t BH_c \times CH_c \times D_t \\ &+ \gamma Intensity_c \times FM_t + \theta W_{ict} + \delta_c + \lambda_t + \varepsilon_{ict} \end{aligned} \tag{6}$$

where β_t captures the dynamic treatment effect of Buddhist influence on the proportion of male births, while accounting for the role of Confucian clans.

We use this specification to verify that the coefficients for the pre-famine years ($\beta_{t<1958}$) are statistically indistinguishable from zero, consistent with the absence of time-varying confounders and supporting the validity of the parallel trends assumption.

5. Main Results

5.1. Estimates of Buddhism on Sex Ratios

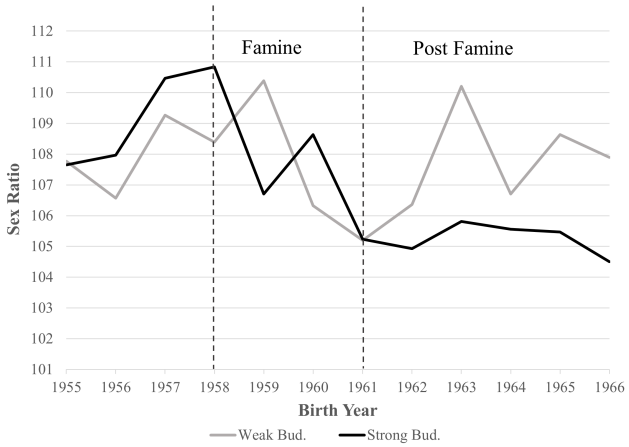
In this section, we present the main findings. Before turning to the regression results, we begin with a descriptive analysis of the raw data to examine the relationship between Buddhism and sex ratio trends, abstracting from the influence of Confucian clans. For this purpose, we divide the sample into two groups based on whether the number of Buddhist temples per capita exceeds the national median.

Figure 3 shows that prior to the famine (pre-1958), sex ratios rose gradually in both groups but declined sharply during the famine years (1958-1961), a pattern consistent with evidence that women generally exhibit higher survival rates under adverse conditions (Kraemer, 2000; Zarulli et al., 2018). In the post-famine period (after 1961), as birth rates recovered, sex ratios rebounded sharply in prefectures with weak Buddhist influence but remained low in those with strong Buddhist presence. This pattern suggests that Buddhism may have helped curb postnatal sex-selection in the aftermath of the famine.

We next formally examine the effect of Buddhism on sex ratios during and after the famine by estimating Equation (1). To distinguish short-term from longer-term effects, we also estimate Equation (2), which separates the famine (1958-1961) from the post-famine (1962-1966). Table 2 presents the results. As discussed earlier, Buddhist influence is measured using both a binary indicator and a continuous temple-density variable. Columns (1)-(4), and (5)-(8) employ these two indicators, respectively. All standard errors are clustered at the prefecture level.

Column (1) reports the estimation results from Equation (1) without controling for the famine intensity. The findings indicate that, relative to the pre-famine years, a greater Buddhist presence is associated with an average annual reduction of 1.64 percentage points in the proportion of male births during and after the famine period. This implies that, after the famine, sex ratios in high-Buddhist prefectures were 6.72 percentage points lower than

Figure 3: Sex Ratios Trends Shaped by Buddhism



Note: The two groups are categorized based on whether the number of Buddhist temples per capita is above or below the national median.

in low-Buddhist prefectures, compared to pre-famine levels.[13] In Column (2), we introduce interactions between the famine intensity and the famine indicator. This adjustment helps address the concern that factors (i.e., grain production) other than Buddhist influence, which may have mitigated famine severity, could be driving the differential changes in sex ratios. The coefficient on the key interaction term becomes larger, corresponding to a 1.78 percentage points reduction in male birth share, and remains highly statistically significant.

Columns (3)-(4) assess the impact of Buddhism on the proportion of male births during the famine (1958-1961) and the post-famine (1962-1966), without and with controlling for the famine intensity, respectively. The findings reveal that, relative to the pre-famine years, stronger Buddhist influence is associated with an average annual reduction of 0.77 percentage points in male birth share during the famine period and 1.91 percentage points during the post-famine period. This implies that, on average, sex ratios in prefec-

[13]Table 1 shows that the average proportion of male births before 1958 was 0.520, corresponding to a sex ratio of $\frac{0.520}{1-0.520} \times 100 = 108.33$. Thus, the reduction in the proportion of male births is calculated as $0.520 - 0.016 = 0.504$. The corresponding sex ratio is $\frac{0.504}{1-0.504} \times 100 = 101.61$, which is 6.72 points lower than the pre-famine sex ratio.

tures with strong Buddhist influence were 3.59 percentage points lower than in those with weak Buddhist influence during the famine, and 8.15 percentage points lower in the subsequent post-famine years. After controlling for famine intensity, the corresponding estimates increase in magnitude and are statistically significant at the 5% and 1% levels, respectively. These findings remain robust when the using the continuous measure of Buddhist influence, as shown in Columns (5)-(8).

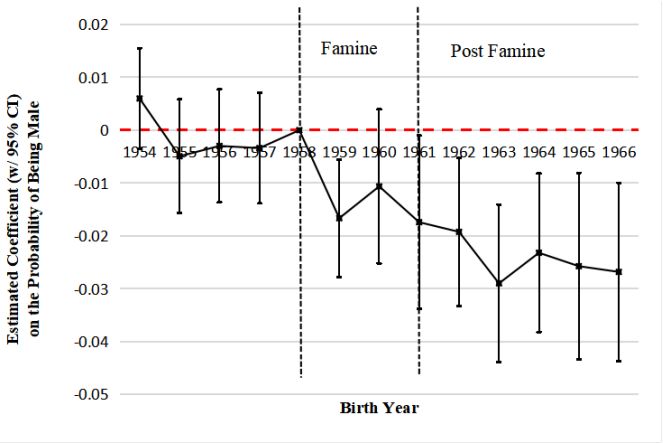
Table 3 presents the heterogeneous estimation results regarding the effect of Buddhism on the male birth shares across various sub-samples, controlling for famine intensity. Panel A reports results for the full 1958-1966 period. Across all sub-samples, Buddhist influence is associated with a lower male birth shares, and all estimates are statistically significant. This indicates that, relative to the pre-famine baseline, sex ratios were consistently lower in high-Buddhist prefectures throughout this period. Panel B further examines the famine and post-famine periods separately. During the famine, the negative effect remains significant for the Han majority and urban hukou groups. In the post-famine period, the effect is consistently negative and significant across all subsamples. Taken together, these results suggest that Buddhist presence is associated with an annual reduction in the male birth shares in both periods and across all population subgroups.

Finally, we estimate the dynamic treatment effects of Buddhism using the regression specified in Equation (3). The estimated coefficients, β_t, are plotted in Figure 4, using 1958 as the baseline. As shown in Figure 4, the differences in male birth shares between prefectures with strong and weak Buddhist influences remained stable during the pre-famine years, with all estimates close to zero. However, during the famine, these differences began to decline in prefectures with strong Buddhist influence and continued to decrease in the post-famine years. In other words, prefectures with strong Buddhist influence exhibited significantly lower sex ratios than those with weak Buddhist influence during these periods.

5.2. Estimates of Buddhism and Confucian Clans on Sex Ratios

We begin by presenting the raw data to illustrate the effects of Buddhism and Confucian clans on sex ratio trends, as depicted in Figure 5. The density of Buddhist temples is used as a proxy for the influence of Buddhism, while the density of genealogies represents the influence of Confucian clans. Prefectures are classified into four groups based on whether the density of Buddhist temples or genealogies falls within the top quantile: weak Buddhist and weak

Figure 4: Effect of Buddhism on the Proportion of Male Births by Year



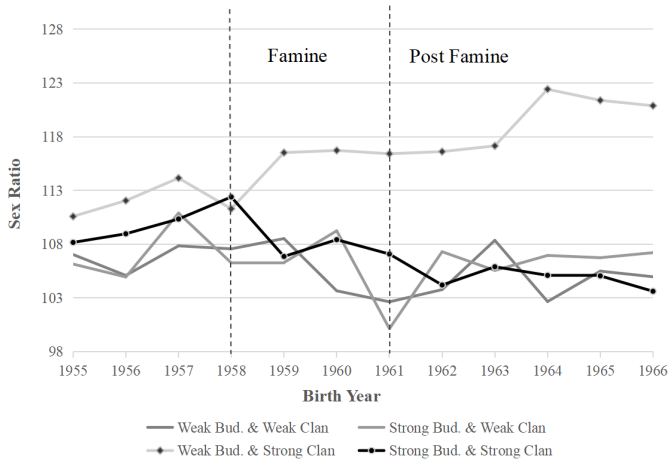
Notes: Coefficients of the interaction terms between BH_c and year dummies are plotted in Equation (3). Year 1958 is set as the reference group.

clan influence, strong Buddhist and weak clan influence, weak Buddhist and strong clan influence, and strong Buddhist and strong clan influence.

Figure 5 illustrates that prior to the famine (pre-1958), sex ratio trends across the four groups were relatively similar. However, during the famine period, sex ratios rose sharply in prefectures with strong clan influence but weak Buddhist presence, while the group with both strong clan and strong Buddhist influence saw a decrease. This trend persisted into the post-famine period (1962-1966), with sex ratios continuing to increase in the group characterized by strong clan influence but weak Buddhist presence, while the group with both strong clan and strong Buddhist influence maintained lower and more stable levels. The other two groups, characterized by weak clan influence, also maintained lower and stable sex ratios during these two periods. These patterns suggest that regions with strong clan influence but a lack of Buddhist presence experienced significant sex imbalances during the famine and the post-famine periods, likely driven by postnatal sex-selective practices. In contrast, regions with strong Buddhist influence appear to have mitigated this imbalance, potentially by discouraging such practices. This highlights the contrasting roles of Buddhism and Confucian clans in shaping sex ratios during these periods.

To formally examine the interactive effect of Buddhism and Confucian clans, we estimate the regression model specified in Equation (4). We further

Figure 5: Sex Ratio Trends Shaped by Buddhism and Confucian Clans



Note: Prefectures are categorized into four groups based on whether the density of Buddhist temples or genealogies falls within the top quantile: (1) weak Buddhist and weak clan influence, (2) strong Buddhist and weak clan influence, (3) weak Buddhist and strong clan influence, and (4) strong Buddhist and strong clan influence.

estimate separate effects for the famine and post-famine periods using the model in Equation (5). The results are presented in Table 4. As previously outlined, we measure the influence of Buddhism and Confucian clans using both binary and continuous variables. Columns (1)-(4) of Table 4 present the results based on binary variables, while Columns (5)-(8) report the findings using continuous measures.

Specifically, Column (1) of Table 4 reports the estimation results from Equation (4) without controlling for famine intensity. The results remain largely unchanged when famine intensity is included as a control in Column (2). The estimates in Column (2) indicate that, relative to prefectures with strong clan but weak Buddhist influence, prefectures characterized by both strong Buddhist and strong clan influence experienced an average annual reduction in male birth shares of 4.07 percentage points during the 1958-1966 period, compared to the pre-famine years. This estimate is statistically significant at the 1% level, implying that after the famine, the sex ratio in prefectures with both strong Buddhist and clan influences was 16.28 percentage points lower than in prefectures with strong clan but weak Buddhist

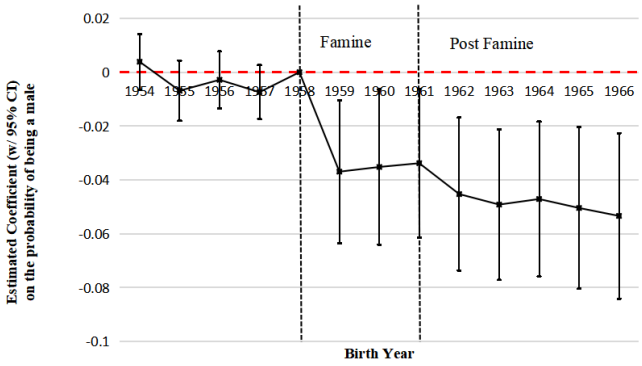
influence.[14] Notably, the estimated interactive effect of -0.407 is larger than the -0.0161 result from our earlier double-difference analysis. This discrepancy can be explained by the fact that, on average, regions with strong clan influence exhibited a 2.69 percentage point higher proportion of male births after the famine, equivalent to a 10.89 percentage point increase in the sex ratio relative to regions with weak clan influence. These findings suggest that while Confucian clans exacerbated sex imbalances during this period, a strong Buddhist presence appears to have counteracted this effect.

Columns (3)-(4) report the estimated separate effects of Buddhism and Confucian clans during the famine (1958-1961) and post-famine (1962-1966) periods, without and with controls for famine intensity, respectively. The results in Column (4) indicate that, relative to prefectures with strong clan but weak Buddhist influence, prefectures characterized by strong influences of both Buddhism and Confucian clans saw an average annual decline in male birth shares of 2.27 percentage points during the famine and 4.65 percentage points in the post-famine period. These estimates are statistically significant at the 5% and 1% levels, respectively. Columns (5)-(8) assess the robustness of these findings using continuous measures of Buddhist and clan influence. Estimates for the full 1958-1966 period remain robust in Columns (5)-(6). However, Columns (7)-(8) show that while the results for the post-famine period remain robust, the corresponding estimates for the famine period are not statistically significant.

Table 5 presents the estimation results from Equations (4) and (5) across various sub-samples. Panel A shows that, when controlling for the influence of Confucian clans, Buddhism is consistently associated with a reduction in the proportion of male births across all sub-samples during the 1958-1966 period. Notably, the results for the Han majority, rural hukou, and less-educated sub-samples are statistically significant. These findings align with the existing literature, which suggests that the Han majority, rural hukou, and less-educated groups are more susceptible to the influence of clan (Cao et al., 2022). Panel B further examines the interactive effects separately for the famine (1958-1961) and post-famine (1962-1966) periods. The results

[14] Table 1 shows that the average proportion of male births before 1958 was 0.520, corresponding to a sex ratio of $\frac{0.520}{1-0.520} \times 100 = 108.33$. Thus, the reduction in the proportion of male births is calculated as $0.520 - 0.0407 = 0.4793$. The corresponding sex ratio is $\frac{0.4793}{1-0.4793} \times 100 = 92.05$, which is 16.28 points lower than the pre-famine sex ratio.

Figure 6: Effects of Buddhism and Confucian Clans on the Proportion of Male Births



Notes: Coefficients of the interaction terms between BH_c, CH_c and year dummies in Equation (6) are plotted. Year 1958 is set as the reference group.

remain consistently negative for all sub-samples and statistically significant at the 1% level for the Han majority, rural hukou, and less-educated sub-samples in the post-famine period, indicating that Buddhism helped mitigate sex imbalances within these groups during this period.

Finally, we estimate the dynamic treatment effects of Buddhism and Confucian clans using the regression specified in Equation (6). The estimated coefficients of the interaction term $BH_c \times CH_c \times D_t$, β_t, are plotted in Figure 6, using 1958 as the baseline. Figure 6 shows that, prior to the famine, the differences in the proportion of male births between prefectures with strong and weak Buddhist influence—while controlling for clan influence—remained relatively stable around zero. However, during the famine years, these differences began to decrease in prefectures with both strong Buddhist and clan influence, and continued to decrease in the post-famine years. This pattern indicates that sex ratios were significantly lower in prefectures with both strong Buddhist and clan influence than in those with strong clan but weak Buddhist influence during these periods.

5.3. Dynamic interactive effects of Buddhism and Confusion clans

Owing to the long process of Sinicization, Buddhism in China has been deeply embedded within an indigenous Confucian cultural environment, giving rise to persistent and potentially non-linear interaction between Buddhist influence and Confucian clan structures. To examine how Buddhist presence

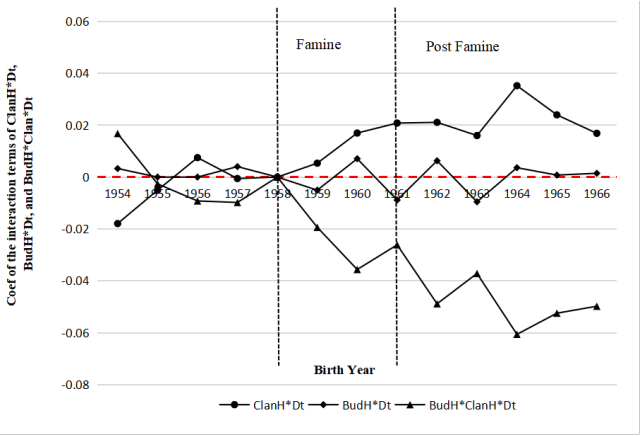
interacts with Confucian clans over time, we estimate the following cohort-specific specification:

$$\begin{aligned} S_{ict} = & \sum_{t=1954}^{1966} \beta_{1t} CH_c * D_t + \sum_{t=1954}^{1966} \beta_{2t} BH_c * D_t + \gamma Intensity_c \times FM \\ & + \sum_{t=1954}^{1966} \beta_{3t} BH_c * CH_c * D_t + \gamma W_{ict} + \delta_c + \lambda_t + \varepsilon_{ict} \end{aligned} \tag{7}$$

where β_{1t} capture the cohort-specific effect of Confucian clans on male birth shares, β_{2t} measure the corresponding effect of Buddhist influence, and β_{3t} characterize their interaction effects. The dummy variable for the 1958 birth cohort, along with all its interactions, is excluded from the analysis to serve as the baseline. If Confucian clans exacerbated female survival disadvantages, we would expect β_{1t} to be relatively stable in the pre-famine cohorts and to become positive during the famine years, reflecting an increase in male birth shares. Conversely, if Buddhism enhanced female survival, β_{2t} should be flat prior to the famine and turn negative during the famine period, indicating a relative decline in male births. Finally, if Buddhism dampened the adverse effects of Confucian clans on female survival, conditional on clan strength, the interaction coefficients β_{3t} should exhibit a shift toward negative values around the famine cohorts, capturing a mitigating effect on sex ratios.

We present the estimated coefficients β_{1t}, β_{2t}, and β_{3t} in Figure 7. For cohorts born prior to the famine, the estimated effects of Confucian clans and Buddhism on male birth shares are small and relatively stable across cohorts, indicating no systematic pre-trends. In contrast, pronounced divergences emerge for cohorts exposed to the famine. Specifically, stronger Confucian clan influence is associated with a higher proportion of male births during the famine period, as reflected in a positive shift in β_{1t}. This pattern persists into the post-famine cohorts, suggesting that heightened clan influence and intensified son preference under conditions of extreme scarcity had lasting effects on sex imbalance. By contrast, the interaction term β_{3t} is negative during the famine period, indicating that Buddhist influence mitigated the effect of Confucian clans on male-biased birth outcomes. This finding is consistent with the interpretation that Buddhism attenuated sex selection pressures precisely where clan-based patriarchal norms were strongest. Finally, the coefficient β_{2t}—capturing the effect of Buddhism in prefectures with relatively weak clan influence—exhibits little variation across cohorts

Figure 7: Effects of Buddhism and Clan on the Proportion of Male Births



Notes: Coefficients of the interaction terms of $CH_c * D_t$, $BH_c * D_t$, and $BH_c * CH_c * D_t$ in Equation (7) are plotted. Year 1958 is set as the reference group.

before and after the famine. This stability suggests that Buddhism alone did not substantially alter sex ratios in the absence of strong clan structures. Rather, its influence appears to operate primarily through moderating clan-driven responses to crisis.

6. Robustness Check

We conduct a series of robustness checks to assess the stability and credibility of our core findings. Detailed results are reported in Appendix B. First, the main results are robust to alternative constructions of famine intensity. Second, we re-estimate Equation (1) using a group interactive fixed effects (GIFE) specification to address potential bias from unobserved time-varying confounders. The estimates remain qualitatively and quantitatively similar to the baseline results in Table 2. Third, drawing on data from the 2004 Economic Census of China, we show that the observed relationship between Buddhist presence and sex ratios cannot be explained by the influence of other religions, including Taoism, Islam, Protestantism, or Catholicism. Finally, we implement a leave-one-province-out exercise, sequentially excluding each of the 22 provinces from the estimation sample. The results remain substantively unchanged, indicating that the findings are not driven by any single region.

7. Mechanisms

In this section, we examine four mechanisms through which Buddhism may have contributed to the attenuation of sex ratio imbalances during and after the Great Famine (1958-1966). First, Buddhist presence may have mitigated famine severity, thereby dampening the intensification of sex selection under extreme resource constraints. Second, Buddhism may have promoted more gender-neutral inter-household resource allocation, reducing postnatal discrimination against female children during periods of scarcity. Third, exposure to Buddhist norms may have shaped son-preference attitudes among individuals who experienced the famine in early childhood. Finally, Buddhist influence may have improved the long-run survival prospects of female children.

7.1. Famine Relief

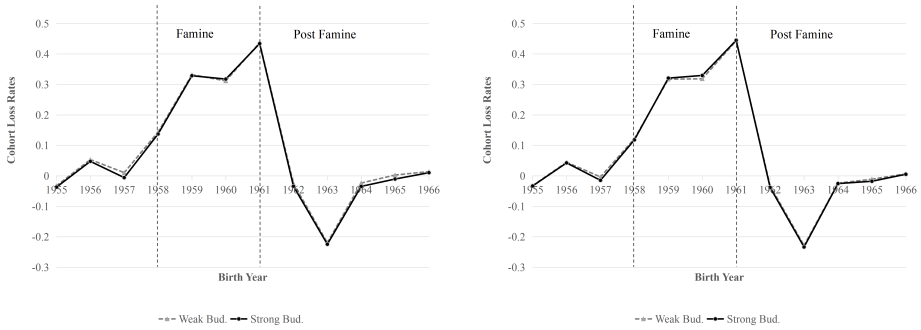
Religious belief systems have long shaped how societies interpret, respond to, and recover from natural disasters, including floods, earthquakes, and famines. (McCleary and Barro, 2006; Dehejia et al., 2007; Islam, 2012; Bush et al., 2015; Bentzen, 2019; Jenkins, 2021). In the Chinese context, historical evidence indicates that Buddhism played a substantive role in famine relief, often complementing state-led interventions (Gernet, 1995; Shiue, 2005; Fuller, 2013). In particular, Fuller (2013) documents that the Buddhist Relief Society emerged as the largest provider of relief grain during the 1920-1921 Great North China Famine, distributing over 14 million kilograms of grain to the southern Zhili region.

We examine whether Buddhism mitigated sex selection by alleviating famine severity during the peak famine years (1958-1961). A central challenge in studying this mechanism is data availability: systematic and reliable mortality records from the famine period do not exist. Following Meng et al. (2015), we proxy mortality rate using the birth cohort loss rate, defined as the proportional difference between a counterfactual cohort size—extrapolated from pre-famine and post-famine demographic trends—and the observed cohort size reconstructed from the 1990 Census.

Figure 8(a) plots the evolution of predicted female birth cohort loss rates for prefectures with strong versus weak Buddhist influence. Specifically, we compute annual averages of cohort loss rates within each group. The two series track each other closely throughout the famine period, indicating that

prefectures with stronger Buddhist influence did not experience systematically lower female birth cohort loss. Figure 8(b) presents analogous patterns for male cohorts, yielding the same conclusion. Taken together, these results suggest that Buddhist presence was not associated with a meaningful reduction in aggregate famine severity for either gender.

Figure 8: Birth Cohort Loss: By the Strength of Buddhism and Gender



(a) Female (b) Male

Notes: Figure (a) and (b) plot estimated birth cohort loss rates for females and males, respectively, in prefectures with strong and weak Buddhist influence. Cohort sizes are reconstructed using individuals observed in the 1990 Census. Buddhist influence is classified as strong (weak) if the number of Buddhist temples per capita lies above the top tercile (below the bottom tercile) of the prefecture-level distribution.

7.2. Inter-household Resource Allocation

We next study whether Buddhism promoted more gender-neutral inter-household resource allocation. The influence of religion on demographic behavior remains contested, operating through a range of mechanisms including doctrinal norms, gender preferences, contextual factors, and moral prescriptions (Dharmalingam and Morgan, 2004; Borooah and Iyer, 2005; Guillot and Allendorf, 2010). Building on earlier results, we find suggestive evidence that Buddhism may attenuate postnatal sex selection—such as differential access to nutrition and healthcare—thereby reducing female disadvantage during periods of scarcity.

To investigate this channel, we draw on data from the China Longitudinal Ageing Social Survey (CLASS). It is a nationally representative, biannual survey that collects detailed information on personal and household characteristics, religious affiliation, health status, and economic conditions of older

adults in China. The CLASS sample comprises approximately 12,000 individuals aged 60 and above, spanning 134 counties across 28 provinces. For our analysis, we utilize the third national wave, conducted between 2018 and 2019. Thus, all respondents in this wave were born prior to 1959, ensuring that they directly experienced the Great Famine of 1959-1961 during their formative years.

This dataset contains two retrospective questions particularly relevant to our empirical strategy: (1) "Did you often go to bed hungry when you were a child?" and (2) "Were you able to receive adequate medical services when you were ill during your childhood?" Drawing on the premise that Buddhist beliefs are typically transmitted intergenerationally and shaped by a strong local Buddhist cultural influence, these response variables allow us to infer the gender-differentiated impacts of Buddhism on childhood outcomes.[15] Specifically, we test whether girls with Buddhist affiliations were less likely to experience childhood hunger and more likely to access adequate medical care during the famine period. These two metrics serve as proxies for postnatal investment, enabling us to assess whether Buddhist norms moderated gender-biased resource allocation in the face of extreme scarcity.

Table 6 presents the corresponding estimates. In Panel A, Columns (1)-(2) and (5)-(6) demonstrate that female respondents with Buddhist affiliations were statistically significantly less likely to report having experienced childhood hunger and more likely to report having accessed adequate medical care during their childhood (not statistically significant). By contrast, Columns (3)-(4) and (7)-(8) show no such effects for female respondents with non-Buddhist religious affiliations.

To address the concern that religious affiliation adopted in later life may confound the empirical results—driven by idiosyncratic life events rather than intergenerational transmission from parents—Panel B restricts the sample to female respondents who identified with their religion before the age of 20. The results remain consistent with those in Panel A, supporting the interpretation that parental Buddhist affiliation is associated with reduced gender gaps in early-life resource access.

[15]Based on the 2018 CGSS, 47.4% of Buddhist respondents reported that at least one parent identified as Buddhist. Disaggregated by gender, this figure is 46.4% for males and 50.3% for females. Detailed descriptive statistics are provided in the Appendix.

7.3. Shaping Son Preference Attitudes

We next investigate if Buddhism shapes son preference attitudes among individuals exposed to the famine during early childhood—specifically, those born after 1955. Son preference in the Chinese context is commonly understood to arise from norms of filial piety and concerns about old-age support. To proxy for its intensity, we draw on two attitudinal measures from the 2018 CLASS. Respondents are asked to evaluate the following statements: (1) "Children should make sacrifices for their parents when necessary," and (2) "Raising sons for old age." Responses are recorded on a five-point Likert scale, with lower values indicating stronger agreement and, hence, a higher intensity of son preference.

Table 7 reports the regression results. Panel A shows that, in the full sample, Buddhist affiliation is on average associated with a higher intensity of son preference. This pattern is consistent with the historical Sinicization of Buddhism, whereby Buddhist practices in China became intertwined with Confucian norms emphasizing patrilineality and filial obligation. Panel B, however, reveals substantial heterogeneity when restricting the sample to cohorts exposed to the famine during early childhood. Within this subsample, the association between Buddhism and son preference diverges sharply by gender: Buddhist affiliation is associated with an attenuation of son preference among women, while the positive association persists among men. This gender-specific divergence sheds light on the underlying mechanism. Under conditions of extreme scarcity, Buddhist ethical norms may have constrained discriminatory postnatal responses to resource shortages, a margin of adjustment over which mothers exert greater influence. In contrast, in non-crisis contexts, institutionalized Buddhism appears to coexist with prevailing patriarchal norms.

7.4. Rising Long-term Survival Rates for Girls

Finally, we study whether Buddhism improves long-term survival rates for girls. Rooted in the core tenets of compassion and equality, Buddhist communities frequently develop dedicated support systems and charitable initiatives, which deliver critical resources and practical assistance to girls and their families (Huang and Weller, 1998; Laliberté, 2012; Weller, 2018). In turn, the embrace of such Buddhist beliefs encourages households to extend equal care and respect to girls and boys alike, thereby enhancing girls survival prospects and overall well-being in the long run.

To investigate this mechanism, we analyze data from the 1990 Census, with a focus on the number of surviving female and male children. While the dataset does not contain specific information on gender-based child survival during the famine period, we utilize data on children born to mothers who were born during 1958-1966. This analysis is predicated on the assumption that the influence of Buddhism is intergenerationally transmitted from mothers to their children. Specifically, we test the hypothesis that girls are more likely to achieve long-term survival if their mothers were born and survived the famine in regions with a strong Buddhist influence, relative to those in regions with a weak Buddhist influence. We construct a novel outcome variable, Y_{ict}, to measure the disparity in surviving children:

$$Y_{ict} = BoySurvived_{ict} - GirlSurvived_{ict} \tag{8}$$

where $BoySurvived_{ict}$ and $GirlSurvived_{ict}$ represent the number of surviving boys and girls, respectively, born to mother i in prefecture c and birth year t. Biologically, Y_{ict} should be close to zero on average in the absence of gender-based selection. Thus, a positive Y_{ict} indicates that the mother had more surviving male children than female children, while a negative value suggests the opposite. We apply Equations (4) and (5) in our analysis. As before, we control for the famine intensity and demographic characteristics of the mothers, as well as birth year and prefecture fixed effects. Additionally, we control for the total number of births per mother to ensure that the observed differences in quantity are not driven by the total number of children.

The estimation results are presented in Table 8. Column (1) shows that, compared to regions with strong clan influence but weak Buddhist presence, mothers in prefectures with strong clan and strong Buddhist influence had, on average, 8.82 percentage points more surviving female children than male children. This result is statistically significant at the 1% level. Moreover, we find that, on average, mothers born during and after the famine in prefectures with strong clan influence tend to have 8.15 percentage points fewer surviving female children than those in prefectures with weak clan influence before the famine. This is consistent with earlier findings and is statistically significant at the 1% level. These results remain largely unchanged when we include famine intensity as a control in Column (2). Columns (3)-(4) further reveal that the impact emerged during the famine and intensified in the subsequent years. Columns (5)-(8) confirm the robustness of these results when using continuous measures. Together, these findings provide additional evidence

supporting the balancing effect of Buddhism in counteracting the influence of Confucian clans.

8. Conclusions

This paper examines Buddhisms role in alleviating sex imbalances during and after China's Great Famine (1958-1966). Using the number of Buddhist temples per capita as a proxy, we find that regions with a strong Buddhist presence exhibited significantly lower sex ratios both during the famine years (1958-1961) and in the immediate aftermath (1962-1966). Controlling for the influence of Confucian clans, we further show that prefectures characterized by both strong Buddhist and clan traditions maintained more balanced sex ratios compared to those dominated by clan influence alone. We argue that the culture of son preference intensified in clan-dense regions following the famine, exacerbating gender imbalances. In contrast, Buddhist beliefs may have alleviated this imbalance by suppressing sex-selective practices and fostering more balanced demographic outcomes.

These findings highlight the substantial influence of cultural and religious values in shaping gender dynamics. In crisis contexts—where discriminatory norms often intensify—policymakers should consider leveraging indigenous belief systems that impose moral constraints and promote compassion and gender equity. Integrating such ethical frameworks into gender equity initiatives may enhance both their effectiveness and cultural legitimacy, particularly when institutional interventions face inherent limitations.

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CRediT authorship contribution statement

Zhenhao Gong: Writing —original draft, Writing —review & editing, Resources, Methodology, Investigation, Formal analysis, Conceptualization. **Ziqing Tian**: Validation, Methodology, Conceptualization. **Heli Xu**: Writing review & editing, Visualization, Software, Methodology, Formal analysis, Data curation.

Data availability

Data will be made available on request.

Declaration of competing interest

The authors declare that they have no known competing financial interests or personal relationships that could have appeared to influence the work reported in this paper.

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Table 1: Summary Statistics

	Before 1958		After 1958	
	Low Buddhist	High Buddhist	Low Buddhist	High Buddhist
# Buddhist temples	35.41	229.02	33.78	230.32
	(29.34)	(245.1)	(27.94)	(246.73)
# Genealogies	62.24	268.57	56.63	269.04
	(112.35)	(479.88)	(106.03)	(479.76)
Population	66,668	46,804	65,450	46,655
	(33949)	(21827)	(33052)	(21916)
Buddhist density	5.48	49.77	5.37	50.49
	(3.31)	(50.37)	(3.27)	(55.09)
Clan density	9.51	59.11	9.20	59.46
	(23.29)	(117.91)	(23.43)	(118.99)
Fraction of male births	0.520	0.521	0.518	0.513
	(0.5)	(0.5)	(0.5)	(0.5)
Sex ratios	108.33	108.77	107.46	105.33
	(0.5)	(0.5)	(0.5)	(0.5)
Age	33.54	33.50	26.46	26.41
	(1.49)	(1.49)	(2.19)	(2.2)
Ethnicity: Han	0.94	0.96	0.93	0.95
	(0.24)	(0.2)	(0.25)	(0.21)
Rural	0.73	0.77	0.75	0.79
	(0.44)	(0.42)	(0.43)	(0.41)
Less educated	0.44	0.53	0.31	0.37
	(0.5)	(0.5)	(0.46)	(0.48)
Observations	413,177	396,736	712,340	710,085

Notes: We categorize the 249 prefectures into two groups by the national median of Buddhist temples per 10,000 population: "Low Buddhist" (below the median) and "High Buddhist" (above the median). "Less educated" refers to individuals whose education level is at or below primary school. We report the mean values of each variable for the pre-1958 and post-1958 periods separately, with standard errors in square brackets.

Table 2: Effects of Buddhism on the Proportion of Male Births

	Outcome variable: a dummy variable for male birth.							
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
Outcome mean	0.5177	0.5177	0.5177	0.5177	0.5177	0.5177	0.5177	0.5177
BH×FM	−0.0164***	−0.0178***						
	(0.0050)	(0.0057)						
BH×FM$_1$			−0.00766**	−0.00837**				
			(0.0035)	(0.0040)				
BH×FM$_2$			−0.0191***	−0.0207***				
			(0.0057)	(0.0066)				
Log(BH)×FM					−0.00665***	−0.00754***		
					(0.0019)	(0.0021)		
Log(BH)× FM$_1$							−0.00290**	−0.00367**
							(0.0014)	(0.0016)
Log(BH)× FM$_2$							−0.00784***	−0.00877***
							(0.0021)	(0.0024)
Observations	2,232,338	1,695,810	2,232,338	1,695,810	2,232,338	1,695,810	2,232,338	1,695,810
Prefectures	249	249	249	249	249	249	249	249
R-squared	0.075	0.078	0.075	0.078	0.075	0.078	0.075	0.078
Intensity×FM	No	Yes	No	Yes	No	Yes	No	Yes
Demo controls	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Prefecture FE	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Year FE	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes

Notes: In Columns (1)-(4), the variable BH is a dummy indicating whether the number of Buddhist temples in a prefecture, adjusted for population size, is above the national median. In Columns (5)-(8), a continuous measure of Buddhist influence is used, where $Log(BH) = Log[(\#Buddhist\ temples + 1)/pop]$. The variable FM is defined as 1958–1966, with FM_1 covering 1958–1961 and FM_2 covering 1962–1966. Demographic controls include age, education level, ethnicity, and hukou type. Standard errors, clustered at the prefecture level, are reported in parentheses. $^{***}p < 0.01$, $^{**}p < 0.05$, $^{*}p < 0.1$.

Table 3: Effects of Buddhism on the Proportion of Male Births Across Sub-samples

	Outcome variable: a dummy variable for male birth.						
	(1)	(2)	(3)	(4)	(5)	(6)	(7)
Sub-samples	All	Not Han	Han	Urban	Rural	Less Edu.	More Edu.
Outcome mean	0.5177	0.5160	0.5178	0.5399	0.5126	0.5685	0.4543
Panel A:							
BH×FM	-0.0178^{***}	-0.0314^{***}	-0.0170^{***}	-0.208^{**}	-0.0127^{***}	-0.0256^{***}	-0.0147^{*}
	(0.0057)	(0.0095)	(0.0060)	(0.0092)	(0.0040)	(0.0073)	(0.0087)
Panel B:							
BH×FM_1	-0.00837^{**}	-0.0214^{*}	-0.0078^{*}	-0.0144^{***}	-0.00184	-0.00626	-0.00931
	(0.0040)	(0.0112)	(0.0042)	(0.0056)	(0.0040)	(0.0056)	(0.0060)
BH×FM_2	-0.0207^{***}	-0.0341^{***}	-0.0199^{***}	-0.0234^{**}	-0.0158^{***}	-0.0318^{***}	-0.0164^{*}
	(0.0066)	(0.0103)	(0.0069)	(0.0116)	(0.0045)	(0.0083)	(0.0099)
Observations	1,695,810	103,490	1,592,320	422,473	1,273,337	659,788	1,036,102
Prefectures	249	249	249	249	249	249	249
R-squared	0.078	0.113	0.077	0.027	0.102	0.063	0.027
Intensity×FM	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Demo controls	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Prefecture FE	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Year FE	Yes	Yes	Yes	Yes	Yes	Yes	Yes

Notes: The variable *BH* is a dummy indicating whether the number of Buddhist temples in a prefecture, adjusted for population size, is above the national median. The variable FM is defined as 1958-1966, with FM_1 covering 1958-1961 and FM_2 covering 1962-1966. Demographic controls include age, education level, ethnicity, and hukou type. "Less Edu." refers to individuals whose education level is at or below primary school, while "More Edu." refers to those whose education level is above primary school. Standard errors, clustered at the prefecture level, are reported in parentheses. $^{***}p < 0.01$, $^{**}p < 0.05$, $^{*}p < 0.1$.

Table 4: Effects of Buddhism and Confucian clans on Proportion of Male Births

	Outcome variable: a dummy variable for male birth.							
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
Outcome mean	0.5177	0.5177	0.5177	0.5177	0.5177	0.5177	0.5177	0.5177
BH×CH×FM	−0.0430***	−0.0407**						
	(0.0146)	(0.0168)						
BH×FM	−0.0018	−0.0055						
	(0.0076)	(0.0087)						
CH×FM	0.0236*	0.0269*						
	(0.0130)	(0.0152)						
BH×CH×FM_1			−0.0262**	−0.0227*				
			(0.0114)	(0.0137)				
BH×CH×FM_2			−0.0484***	−0.0465**				
			(0.0163)	(0.0186)				
Log(BH)×Log(CH)×FM					−0.0030***	−0.0027**		
					(0.0011)	(0.0012)		
Log(BH)×Log(CH)×FM_1							−0.0012	−0.0007
							(0.0009)	(0.0011)
Log(BH)×Log(CH)×FM_2							−0.0036***	−0.0033**
							(0.0012)	(0.0014)
Observations	1,081,336	844,332	1,081,336	844,332	1,081,336	844,332	1,081,336	844,332
Prefectures	249	249	249	249	249	249	249	249
R-squared	0.085	0.091	0.085	0.091	0.085	0.091	0.085	0.091
Intensity×FM	No	Yes	No	Yes	No	Yes	No	Yes
Demo controls	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Prefecture FE	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Year FE	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes

Notes: In Columns (1)-(4), the variables BH and CH are the top-quantile indicators defined in Equation (4). In Columns (5)-(8), $Log(BH) = Log[(\#Buddhist\ temples + 1)/pop]$ and $Log(CH) = Log[(\#Genealogies + 1)/pop]$. The variable FM is defined as 1958-1966, with FM_1 covering 1958-1961 and FM_2 covering 1962-1966. Demographic controls include age, education level, ethnicity, and hukou type. Standard errors, clustered at the prefecture level, are reported in parentheses. $^{***}p < 0.01$, $^{**}p < 0.05$, $^{*}p < 0.1$.

Table 5: Effects of Buddhism and Confucian Clans Across Sub-samples

	Outcome variable: a dummy variable for male birth.						
	(1)	(2)	(3)	(4)	(5)	(6)	(7)
Sub-samples	All	Not Han	Han	Urban	Rural	Less Edu.	More Edu.
Outcome mean	0.5177	0.5160	0.5178	0.5399	0.5126	0.5685	0.4543
Panel A:							
BH×CH×FM	−0.0407**	−0.0137	−0.0448**	−0.0225	−0.0338***	−0.0605**	−0.0278
	(0.0168)	(0.0253)	(0.0173)	(0.0244)	(0.0123)	(0.0249)	(0.0223)
Panel B:							
BH×CH×FM$_1$	−0.0227*	−0.0172	−0.0260*	−0.0167	−0.0166	0.00106	−0.0286
	(0.0137)	(0.0333)	(0.0145)	(0.0199)	(0.0134)	(0.0222)	(0.0177)
BH×CH×FM$_2$	−0.0465**	−0.0126	−0.0510***	−0.0252	−0.0389***	−0.0811***	−0.0276
	(0.0186)	(0.0304)	(0.0190)	(0.0278)	(0.0131)	(0.0273)	(0.0246)
Observations	844,332	64,056	780,276	199,646	644,686	359,816	484,516
Prefectures	249	249	249	249	249	249	249
R-squared	0.091	0.121	0.090	0.031	0.120	0.071	0.033
Intensity×FM	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Demo controls	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Year FE	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Prefecture FE	Yes	Yes	Yes	Yes	Yes	Yes	Yes

Notes: The variables BH and CH are the top-quantile indicators defined in Equation (4). The variable FM is defined as 1958-1966, with FM_1 covering 1958-1961 and FM_2 covering 1962-1966. "Less Edu." refers to individuals whose education level is at or below primary school, while "More Edu." refers to those whose education level is above primary school. Demographic controls include age, education level, ethnicity, and hukou type. Standard errors, clustered at the prefecture level, are reported in parentheses. $^{***}p < 0.01$, $^{**}p < 0.05$, $^{*}p < 0.1$.

Table 6: Effects of Buddhism on Inter-household Resource Allocation

	Hunger Experience				Medical Service			
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
Outcome mean	0.6748	0.6748	0.6748	0.6748	0.1895	0.1895	0.1895	0.1895
Panel A: Start having religion beliefs across all ages								
Buddhist affiliation×Female	−0.0750**	−0.0630**			0.0417	0.0363		
	(0.0319)	(0.0312)			(0.0267)	(0.0262)		
Other affiliations×Female			0.0299	−0.00219			−0.0256	−0.0132
			(0.0337)	(0.0336)			(0.0282)	(0.0283)
Control Variables	No	Yes	No	Yes	No	Yes	No	Yes
Observations	11419	11419	11419	11419	11419	11419	11419	11419
R-squared	0.001	0.047	0.000	0.047	0.001	0.039	0.001	0.039
Panel B: Start having religion beliefs in ages less than 20								
Buddhist affiliation×Female	−0.344**	−0.297**			0.0569	0.0268		
	(0.1353)	(0.1323)			(0.1131)	(0.1111)		
Other affiliations×Female			0.179**	0.127			−0.0560	−0.0661
			(0.0793)	(0.0831)			(0.0664)	(0.0698)
Control Variables	No	Yes	No	Yes	No	Yes	No	Yes
Observations	11419	11419	11419	11419	11419	11419	11419	11419
R-squared	0.001	0.047	0.001	0.047	0.001	0.039	0.001	0.039

Notes: Data are from CLASS 2018. Other religious affiliations include Taoism, Christianity, Catholicism, and Islam. Control variables comprise age, ethnicity, household registration status, educational attainment, and marital status. Standard errors, clustered at the provincial level, are reported in parentheses.$^{***}p < 0.01$, $^{**}p < 0.05$, $^{*}p < 0.1$.

Table 7: Effects of Buddhism on Son Preference Attitudes

	Female				Male			
	Filial piety		Elderly care		Filial piety		Elderly care	
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
Outcome mean	2.7003	2.7003	2.1756	2.1756	2.7089	2.7089	2.1370	2.1370
Panel A: All sample								
Buddhist affliliation	−0.109	−0.104	−0.191	−0.168	−0.320***	−0.289***	−0.239*	−0.190
	(0.1014)	(0.1009)	(0.1231)	(0.1233)	(0.1055)	(0.1039)	(0.1331)	(0.1323)
Control Variables	No	Yes	No	Yes	No	Yes	No	Yes
Observations	5682	5682	5682	5682	5737	5737	5737	5737
R-squared	0.000	0.022	0.007	0.015	0.003	0.040	0.007	0.027
Panel B: Experienced famine in their childhood (born after 1955)								
Buddhist affliliation	0.120	0.158	0.0577	0.135	−0.650*	−0.564*	−0.302	−0.224
	(0.2453)	(0.2384)	(0.3048)	(0.2978)	(0.3393)	(0.3396)	(0.3739)	(0.3704)
Control Variables	No	Yes	No	Yes	No	Yes	No	Yes
Observations	1061	1061	1061	1061	1000	1000	1000	1000
R-squared	0.002	0.076	0.007	0.070	0.004	0.073	0.012	0.046

Notes: Data are from the CLASS 2018 survey. Control variables include age, ethnicity, household registration status, educational attainment, and marital status. Standard errors, clustered at the provincial level, are reported in parentheses. $^{***}p < 0.01$, $^{**}p < 0.05$, $^{*}p < 0.1$.

Table 8: Effects of Buddhism and Clans on Gender Gaps in Child Survival

	Outcome variable: Gender gap in surviving children (male-to-female count).							
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
Outcome mean	0.0481	0.0481	0.0481	0.0481	0.0481	0.0481	0.0481	0.0481
BH×CH×FM	−0.0882***	−0.0769**						
	(0.0310)	(0.0323)						
BH×FM	0.0292	0.0131						
	(0.0255)	(0.0267)						
CH×FM	0.0815***	0.0781***						
	(0.0233)	(0.0231)						
BH×CH×FM_1			−0.0636***	−0.0469*				
			(0.0222)	(0.0242)				
BH×CH×FM_2			−0.0966***	−0.0872**				
			(0.0353)	(0.0368)				
Log(BH)×Log(CH)×FM					−0.00979***	−0.00731**		
					(0.0034)	(0.0034)		
Log(BH)×Log(CH)×FM_1							−0.0059**	−0.0059**
							(0.0028)	(0.0028)
Log(BH)×Log(CH)×FM_2							−0.0078**	−0.0078**
							(0.0039)	(0.0039)
Observations	1,077,267	815,636	1,077,267	815,636	1,077,267	815,636	815,636	815,636
Prefectures	249	249	249	249	249	249	249	249
R-squared	0.034	0.035	0.034	0.035	0.034	0.035	0.036	0.036
Intensity×FM	No	Yes	No	Yes	No	Yes	No	Yes
Demo controls	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Prefecture FE	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Year FE	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes

Notes: In Columns (1)-(4), the variables BH and CH are the top-quantile indicators defined in Equation (4). In Columns (5)-(8), $Log(BH) = Log[(\#Buddhist\ temples + 1)/pop]$ and $Log(CH) = Log[(\#Genealogies + 1)/pop]$. The variable FM is defined as 1958-1966, with FM_1 covering 1958-1961 and FM_2 covering 1962-1966. Demographic controls include age, education level, ethnicity, and hukou type. Standard errors, clustered at the prefecture level, are reported in parentheses. $^{***}p < 0.01$, $^{**}p < 0.05$, $^{*}p < 0.1$.

Appendix A. Constructing Famine Intensity

Following Yi (2022), we construct an exogenous measure of famine severity that purges prefecture fixed effects and varies exclusively with random weather shocks. This measure is motivated by the institutional mechanism underlying the famine. During the Great Leap Forward, local procurement targets were determined largely by reported agricultural output in the preceding year. Agricultural output, in turn, depended on two components: time-invariant prefecture-specific agricultural endowments and transitory weather realizations. While the former is absorbed by prefecture fixed effects, the latter constitutes idiosyncratic shocks that are orthogonal to persistent prefecture characteristics and other unobserved confounders.

We therefore isolate the component of famine intensity attributable to random weather fluctuations. Specifically, we construct cross-prefecture variation in famine severity induced by weather shocks in four steps:

Step 1: Construct Cohort Loss Rate.

Using the 1990 Census, we define the birth cohort loss rate for prefecture c and birth year t as:

$$L_{ct} = \frac{\hat{P}_{ct} - P_{ct}}{\hat{P}_{ct}} = 1 - \frac{P_{ct}}{\hat{P}_{ct}}, \tag{A.1}$$

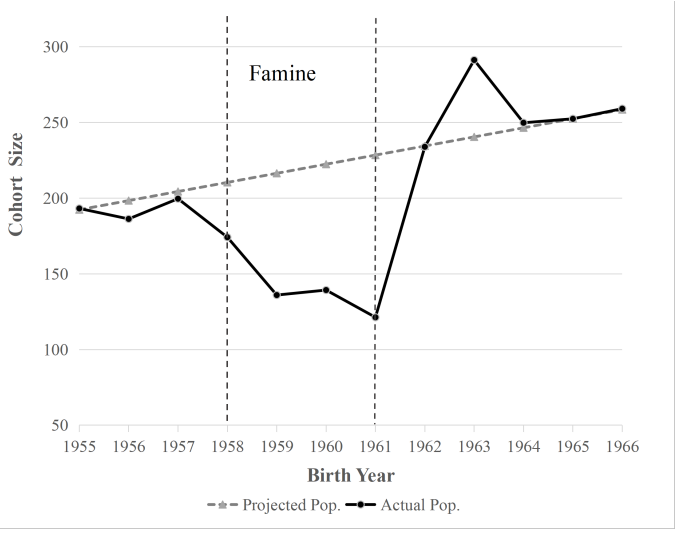
where $\hat{P}_{ct}$ denotes the linearly projected counterfactual cohort size and and P_{ct} denotes the observed cohort size reconstructed from the 1990 Census. The counterfactual $\hat{P}_{ct}$ is obtained by extrapolating pre- and post-famine demographic trends, thereby capturing expected cohort size in the absence of famine-related mortality.

Figure A.1 illustrates the projected and observed cohort sizes for China as a whole. In normal years, the observed cohort size closely tracks the projected trend. During the famine, however, actual cohort size falls sharply below the projection, generating substantial cohort loss. This divergence provides the basis for our prefecture-level measure of famine intensity.

Step 2: Thermal Agricultural Productivity.

Cohort loss rates may differ systematically across prefectures for reasons correlated with local Buddhist presence, including persistent differences in agricultural endowments, infrastructure, or baseline public health conditions. To isolate the component of cohort loss attributable to exogenous weather variation, we construct a measure of agricultural productivity based solely on thermal exposure.

Figure A.1: Prefecture-level Birth Cohort Sizes (Mean)



Note: The soild line plots mean birth cohort size by birth year t across prefectures (e.g., 50 on the y-axis indicates $50,000$ individuals). Individuals are observed in 1990. The dashed line is a linear trend based on normal years.

Specifically, we define thermal agricultural productivity (TAP) for prefecture c and year t the cumulative number of growing-degree days during the main agricultural season (April 1 to September 30):

$$\text{TAP}_{ct} = \sum_{d \in \mathcal{G}_t} g(T_{ctd}), \tag{A.2}$$

where $\mathcal{G}_t$ indexes days in the growing season and $g(T_{ctd})$ is defined as:

$$g(T_{ctd}) = \begin{cases} 0 & \text{if} \quad T_{ctd} < 8, \\ T_{ctd} - 8 & \text{if} \quad 8 \leq T_{ctd} < 33, \\ \frac{25}{8}\left[41 - T_{ctd}\right] & \text{if} \quad 33 \leq T_{ctd} < 41, \\ 0 & \text{if} \quad T_{ctd} \geq 41. \end{cases} \tag{A.3}$$

This specification is grounded in agronomic principles: Crops accumulate effective heat only when temperature exceeds a lower physiological threshold (here set at 8°C). Yields increase with temperature up to a critical ceiling (33°C), beyond which yields decline. T_{ctd} denotes the daily mean instrumental temperature for prefecture c on day d of year t, which is matched to the nearest weather station using Euclidean distance.

Step 3: Random Element of Cohort Loss.

We estimate the following regression to isolate the component of cohort loss attributable to transitory weather-induced agricultural shocks, net of time-invariant prefecture characteristics:

$$L_{ct} = \beta_0 + \beta_1 \text{ TAP }_{c,t-1} + \sum_{y=1953}^{1970} \lambda_y \text{ TAP }_{c,t-1} D_y + \sum_{y=1953}^{1970} \eta_y D_y + \nu_c + \varepsilon_{ct}, \quad \text{(A.4)}$$

where D_y is a year indicator, ν_c are prefecture fixed effects, and ε_{ct} is an idiosyncratic error term. In this specification, β_1 captures the average relationship between lagged TAP and cohort loss rate across years. The interaction terms, λ_y, allow this relationship to vary flexibly by year, thereby accommodating changes in institutional responses over time (i.e., procurement intensity). η_y absorb aggregate shocks common to all prefectures in a given year that are unrelated to local agricultural conditions.

Step 4: Exogenous Measure of Famine Intensity.

In the final step, we use the estimated coefficients $\hat{\beta}_1$, $\hat{\lambda}_t$, and $\hat{\eta}_t$ from Equation (A.4) to construct the predicted cohort loss rate attributable to lagged TAP:

$$\hat{L}_{ct} = \hat{\beta}_1 \text{ TAP }_{c,t-1} + \sum_{y=1953}^{1970} \hat{\lambda}_y \text{ TAP }_{c,t-1} D_y + \sum_{y=1953}^{1970} \hat{\eta}_y D_y. \quad \text{(A.5)}$$

We then define the exogenous measure of famine intensity for prefecture c as the difference between predicted cohort loss during the famine years and predicted loss during non-famine years:

$$Famine_c = \frac{1}{3} \sum_{t=1959}^{1961} \hat{L}_{ct} - \frac{1}{15} \left[\sum_{t=1953}^{1958} \hat{L}_{ct} + \sum_{t=1962}^{1970} \hat{L}_{ct} \right]. \quad \text{(A.6)}$$

This measure captures the excess predicted cohort loss rate during 1959-1961 relative to normal periods. By construction, it excludes prefecture fixed effects and is driven solely by realized temperature shocks operating through agricultural productivity. As shown in Yi (2022), this predicted measure is strongly correlated with observed cohort loss rate and exhibits substantial cross-prefecture variation in famine intensity.

Appendix B. Robustness Checks

Appendix B.1. Alternative Constructions of Famine Intensity

Our identification strategy would be biased if prefectures with a stronger Buddhist presence differ systematically from other prefectures in their exposure to the famine. To better account for heterogeneity in famine intensity, we construct four alternative measures of famine intensity, as outlined in the four steps of Appendix A. We then assess the robustness of our baseline results by re-estimating Equation (1) using these four alternative famine intensity measures. Table B.1 reports the corresponding results. In Columns (1)-(2) and (3)-(4), we employ quadratic and exponentially projected counterfactual populations to compute cohort loss rates in Equation (A.1), as implemented in Step 1 of our famine intensity construction. In Columns (5)-(6), we adopt the weighted average of the three nearest weather stations to construct T_{ctd} as defined in Equation (A.3) of Step 2. In Columns (7)-(8), we utilize temperatures across all months to construct TAP_{ct} as specified in Equation (A.2) of Step 2. As shown in Table B.1, our baseline findings remain robust across all four alternative measures of famine intensity.

Table B.1: Effects of Buddhism on the Proportion of Male Births

	Outcome variable: a dummy variable for male birth.							
FM Intensity Constructions	Quadratic Projection		Exponential Projection		Weighted Average of Three Stations		Temperatures in All Months	
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
BH×FM	−0.0178***		−0.0178***		−0.0176***		−0.0176***	
	(0.0057)		(0.0057)		(0.0057)		(0.0058)	
BH×FM$_1$		−0.00837**		−0.00837**		−0.00820**		−0.00821**
		(0.0040)		(0.0040)		(0.0040)		(0.0040)
BH×FM$_2$		−0.0207***		−0.0207***		−0.0206***		−0.0206***
		(0.0066)		(0.0066)		(0.0066)		(0.0066)
Observations	1,695,810	1,695,810	1,695,810	1,695,810	1,695,810	1,695,810	1,695,810	1,695,810
Prefectures	249	249	249	249	249	249	249	249
R-squared	0.078	0.078	0.078	0.078	0.078	0.078	0.078	0.078
Intensity×FM	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Demo controls	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Prefecture FE	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Year FE	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes

Notes: In Columns (1)-(2) and (3)-(4), we use quadratic and exponentially projected counterfactual populations to calculate the cohort loss rate in Step 1 of constructing famine intensity, respectively. In Columns (5)-(6) and (7)-(8), we adopt the weighted average of the three nearest weather stations to construct T_{ctd} and temperatures in all months to construct TAP_{ct} in Step 2 of the famine intensity measure, respectively.

Appendix B.2. Implementing a Group Interactive Fixed Effect Model

As discussed in Section 4.1, our baseline two-way fixed effects (TWFE) specification in Equation (1) may be susceptible to bias arising from time-varying unobserved confounders. This reflects a standard model misspecification problem. The TWFE framework implicitly assumes that the famine shock affected all prefectures homogeneously, while prefecture-specific unobservables remained constant over time. This assumption may be untenable if latent determinants—such as evolving son-preference norms or other cultural factors—change over time and induce heterogeneous responses to the famine across prefectures.

To address this problem, we implement the group interactive fixed effects (GIFE) model proposed by Gong and Kim (2024). The specification corresponding to Equation (1) can be written as:

$$S_{ict} = \beta BH_c \times FM_t + \theta W_{ict} + \omega_c' f_t + \varepsilon_{ict}, \tag{B.1}$$

where f_t is an $(r \times 1)$ vector of unobserved common shocks, and ω_c is an $(r \times 1)$ vector of prefecture-specific factor loadings that capture heterogeneous responses to those shocks. The interactive term $\omega_c' f_t$ therefore allows for non-additive, time-varying unobserved heterogeneity across prefectures. If $\omega_c = (\delta_c, 1)'$ and $f_t = (1, \lambda_t)'$, then the interactive structure reduces to additive prefecture and year fixed effects, and Equation (B.1) collapses to the TWFE model. The analogue of Equation (2) under the GIFE framework is:

$$S_{ict} = \beta_1 BH_c \times FM_{1t} + \beta_2 BH_c \times FM_{2t} + \theta W_{ict} + \omega_c' f_t + \varepsilon_{ict}. \tag{B.2}$$

We estimate Equations (B.1) and (B.2) using the group iterative least squares procedure proposed by Gong and Kim (2024). Table B.2 reports the corresponding results. Columns (1) and (3) present estimates from the baseline TWFE specifications in Equations (1) and (2), respectively. Columns (2) and (4) report the corresponding estimates from the GIFE specifications in Equations (B.1) and (B.2). Columns (5)-(8) repeat the analysis using a continuous measure of temple-density. The qualitative conclusions remain unchanged across specifications, indicating that the main findings are robust to both alternative functional forms and more flexible controls for time-varying unobserved heterogeneity.

Table B.2: Effects of Buddhism on the Proportion of Male Births

	Outcome variable: a dummy variable for male birth.							
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
Model Specifications	TWFE	GIFE	TWFE	GIFE	TWFE	GIFE	TWFE	GIFE
BH×FM	-0.0164^{***}	-0.0168^{***}						
	(0.0050)	(0.0025)						
BH×FM_1			-0.00766^{**}	-0.0158^{***}				
			(0.0035)	(0.0029)				
BH×FM_2			-0.0191^{***}	-0.0177^{***}				
			(0.0057)	(0.0030)				
Log(BH)×FM					-0.0067^{***}	-0.0063^{***}		
					(0.0019)	(0.0009)		
Log(BH)× FM_1							-0.0029^{**}	-0.0061^{***}
							(0.0014)	(0.0009)
Log(BH)× FM_2							-0.0078^{***}	-0.0069^{***}
							(0.0021)	(0.0011)
Observations	1,695,810	1,695,810	1,695,810	1,695,810	1,695,810	1,695,810	1,695,810	1,695,810
Prefectures	249	249	249	249	249	249	249	249
R-squared	0.078	—	0.078	—	0.078	—	0.078	—
Intensity×FM	No	No	No	No	No	No	No	No
Demo controls	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Prefecture FE	Yes	No	Yes	No	Yes	No	Yes	No
Year FE	Yes	No	Yes	No	Yes	No	Yes	No
Interactive FE	No	Yes	No	Yes	No	Yes	No	Yes

Notes: Columns (1) and (3) present estimates from the TWFE specifications in Equations (1) and (2), respectively. Columns (2) and (4) report the corresponding estimates from the GIFE specifications in Equations (B.1) and (B.2). Columns (5)-(8) repeat the analysis using a continuous measure of temple-density. All of the variables are defined the same as before. Demographic controls include age, education level, ethnicity, and hukou type. Standard errors, clustered at the prefecture level, are reported in parentheses. $^{***}p < 0.01$, $^{**}p < 0.05$, $^{*}p < 0.1$.

Appendix B.3. Other Religious Beliefs

One potential concern is that domestic Taoism or other officially recognized religions (Islam, Protestantism, or Catholicism) rather than Buddhism, may have mitigated sex imbalance during and after the Great Famine. To address this possibility, we draw on data from China's 2004 Economic Census, which provides the first comprehensive and publicly available official enumeration of religious sites nationwide.

The census recorded a total of 72,887 religious sites across China's 31 provinces in 2004, including 16,676 Buddhist temples, 4,939 Taoist temples, 34,305 Islamic temples, 14,509 Protestant churches, and 2,419 Catholic churches. For each site, the census reports detailed information on name, location, scale, and annual income. We obtain this data set from Yang et al. (2019) and construct prefecture-level measures of religious presence analogous to our Buddhist intensity variable.

Figure B.1 presents the raw patterns underlying the relationship between Buddhist presence and sex ratio dynamics. Consistent with Figure 3, sex ratios increase markedly in prefectures where other religions are more prevalent, while remaining comparatively lower in prefectures where Buddhism is the dominant religious presence during the post-famine period.[16] This suggests that Buddhism, rather than other religions, played a crucial role in mitigating sex imbalances during this time.

We then re-estimate Equations (1) and (2) by replacing the variable BH_c with an alternative indicator BD_c, which equals one if the number of Buddhist sites in a prefecture exceeds the number of sites affiliated with each of the other four religions, thereby capturing Buddhisms dominant religious presence. The results, reported in Table B.3, show that prefectures classified as Buddhist-dominant experienced a statistically significant reduction in the male birth ratio during both the famine period (1958-1961) and the post-famine period (1962-1966). The magnitude of the estimated coefficients implies that sex ratios at birth in Buddhist-dominant prefectures were systematically lower than in prefectures where another religion constituted the dominant religious presence. These findings reinforce the interpretation that the mitigating effect on sex imbalance is specific to Buddhist influence rather than reflecting broader patterns of religious activity.

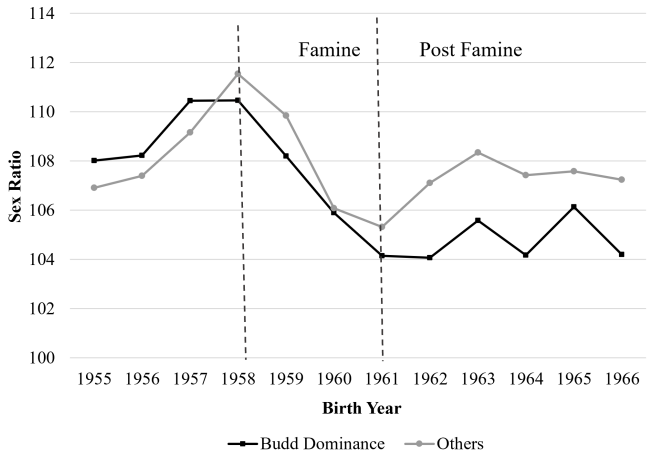
[16]The two groups are classified based on whether the number of Buddhist sites exceeds that of the other four religious sites in a given prefecture.

Table B.3: Effects of Buddhism on the Proportion of Male Births

	Outcome variable: a dummy variable for male birth.							
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
Outcome mean	0.5177	0.5177	0.5177	0.5177	0.5177	0.5177	0.5177	0.5177
BD×FM	−0.0129***	−0.0135**						
	(0.0049)	(0.0055)						
BD×FM_1			−0.0062*	−0.0041				
			(0.0035)	(0.0039)				
BD×FM_2			−0.0149***	−0.0164***				
			(0.0055)	(0.0062)				
Log(BD)×FM					−0.00334**	−0.0034**		
					(0.0014)	(0.0016)		
Log(BD)×FM_1							−0.0009	−0.0008
							(0.0011)	(0.0012)
Log(BD) ×FM_2							−0.0041***	−0.0042**
							(0.0016)	(0.0018)
Observations	2,232,338	1,695,810	2,232,338	1,695,810	2,232,338	1,690,056	2,217,652	1,690,056
Prefectures	249	249	249	249	249	249	249	249
R-squared	0.075	0.078	0.075	0.078	0.075	0.078	0.075	0.078
Intensity×FM	No	Yes	No	Yes	No	Yes	No	Yes
Demo controls	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Prefecture FE	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Year FE	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes

Notes: In Columns (1)-(4), the variable BD is a dummy indicating whether the number of Buddhist temples in a prefecture exceeds that of the other four religious sites. In Columns (5)-(8), a continuous measure of Buddhist influence is used, where $Log(BD) = Log[(\#\ Buddhist\ temples + 1)/populations]$. The variable FM is defined as 1958–1966, with FM_1 covering 1958–1961 and $FM2$ covering 1962–1966. Demographic controls include age, education level, ethnicity, and hukou type. Standard errors, clustered at the prefecture level, are reported in parentheses. $^{***}p < 0.01$, $^{**}p < 0.05$, $^{*}p < 0.1$.

Figure B.1: Sex Ratios Trends Shaped by the Dominant Influence of Religions

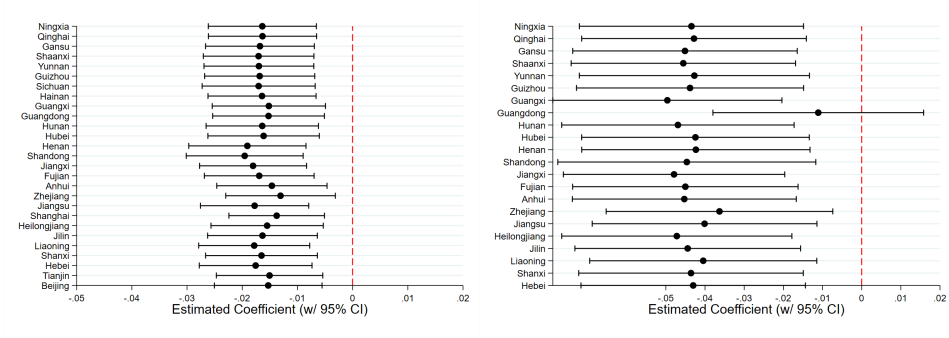


Note: Prefectures classified based on whether the number of Buddhist sites exceeds that of the other four religious sites.

Appendix B.4. Leave-one-province-out Exercise

It is possible that having only a small number of provinces drives the results. We address this concern with a simple robustness exercise: we repeatedly re-estimate Equation (1) and Equation (4) while excluding one province at a time. Figure B.2(a) plots the coefficients on the interaction term between Buddhist influence and the famine period dummy from Equation (1), and Figure B.2(b) plots the coefficients on the three-way interaction term among Buddhist influence, clan influence, and the famine period dummy from Equation (4). The figure shows that the coefficients for the key interaction terms remain strongly negative and large in magnitude regardless of which province is omitted.

Figure B.2: Estimates with One Province Excluded at a Time



(a) Buddhism×Famine (b) Buddhism×Clan×Famine

Notes: Figure (a) plots the coefficients on the interaction term between Buddhist influence and the famine period dummy in Equation (1). Figure (b) plots the coefficients on the threeway interaction term among Buddhist influence, clan influence, and the famine period dummy in Equation (4). Each coefficient is estimated on a subsample from which one province is sequentially excluded from the original dataset.