Culture, Economic Stress, and Missing Girls[1]

Viktor Malein[2]

Tamar Matiashvili[3]

Francisco J. Beltrán Tapia[4]

Abstract

Cultural norms play a pivotal role in shaping how societies respond to crises. This study examines the gender-asymmetric effects of a resource shock on mortality. Studying the 1891-1892 Russian famine, we compare cohorts born before and after the famine in districts differentially affected by the famine and with diverse gender norms, expressed in folklore. Our findings reveal that areas where women were depicted more negatively suffered a more skewed sex ratio favouring male survival. The study demonstrates the importance of the cultural channel in driving these results.

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[2] Higher School of Economics (email: vmalein@hse.ru)
[3] Stanford University (email: tamrim@stanford.edu)
[4] Norwegian University of Science and Technology (email: francisco.beltran.tapia@ntnu.no)

A daughter is someone else's treasure. To take care and feed, teach and guard, and give in marriage.
(Russian proverb)[5]

Introduction

Women typically live longer than men due to biological differences (Austad 2006; Bardford et al. 2006; Zarulli et al. 2018).[6] This biological advantage also helps women endure the harsh conditions of famines more effectively (Watkins and Menken 1985; Henry 1990; Dyson 1991; Janetta 1992; Macintyre 2002; Hionidou 2006; Zarulli et al. 2018).[7] However, this advantage can nonetheless be enhanced or offset by gender differences in levels of care or exposure to mortality crises (Dyson 1991; Macantyre 2002; Edgerton-Tarpley 2008).[8] Cultural and social norms can dictate how food and care are allocated within families when resources are scarce. Crises can also expose pre-existing gender biases as they hide in the background during normal times. Existing gender-discriminatory practices in terms of food or health care also imply that women and girls face famines in weaker conditions, increasing their mortality risks.

This article studies the aftermath of a resource crisis to assess how cultural gender norms shape gender-discriminatory practices and subsequently affect sex-specific mortality rates. Focusing on the 1891/92 Russian Famine, we document the gender-asymmetric effect of a resource shock on mortality.[9] On average, exposure to famine resulted in relatively higher male-

[5] The original version: "Дочь — чужое сокровище. Холь да корми, учи да стереги, да в люди отдай." (Source)
[6] The biological foundations of sex differences in human mortality include hormonal and chromosomal genetic differences, as well as a female advantage in immune protection against infectious diseases (Zarulli et al. 2018, 7).
[7] Women are smaller and store more body fat, thus increasing their chances of surviving acute deprivation, especially during the first months of scarcity (Rivers 1982; Henry 1990; Livi Bacci 1991; Collins 1995). Women are also more resistant to most epidemic diseases (Lord et al. 1998; Bouman et al. 2005; Marie et al. 2012), which is also a major driving of mortality during famines due to the interaction between malnutrition and the spread of diseases due to nutritional deficiencies, worse hygienic conditions and increased migratory flows (Pitkänen and Mielke 1993; Macintyre 2002; Speakman 2013). The evidence is not so clear for particular diseases such as malaria, measles, or pulmonary diseases (Macyntire 2002, 248). Likewise, the decline in fertility that is frequently associated with famines reduces the number of female deaths during pregnancy and childbirth (Ó Gráda 1995; 1999).
[8] Differences in coping mechanisms, such as access to the cooking pot, the knowledge of wild foods, the possibility to turn to prostitution, or the willingness to seek assistance, can benefit women. Males are more likely to resort to the possibility of escaping from the crisis area, but this coping mechanism may result in a greater risk of contracting deadly infectious diseases or running into other hazards (Pitkänen and Mielke 1993). Likewise, famines tend to be followed by epidemics, and women are usually in charge of caring for the sick, which in turn increases their likelihood of falling ill and lowers their chances of survival. Relief programs can also especially target women and children, thus increasing their chances of survival (El Bushra and Piza Lopez 1994; Ó Gráda 1999). See also Rivers (1982) or Neumayer and Plümper (2007).
[9] Section "Interpretation of the empirical findings" discusses conditions under which our empirical estimates can be interpreted as causal.

to-female child sex ratios among the affected cohorts. By leveraging unique folklore data collected by anthropologists (Michalopoulos and Xue 2021), we show heterogeneity of the famine effect – the sex ratios are skewed more in those districts where the oral traditions of the population tend to depict women more negatively.[10] We thus tie entrenched gender norms to gender-biased mortality, demonstrating through various exercises that gender norms played a primary role in the asymmetric response of mortality to the famine.

The 1891/92 Famine and the subsequent Cholera outbreak in the Russian Empire contributed to approximately half a million deaths (Charnysh 2022), affecting one of the most ethnically diverse parts of the Empire – the Volga Region. Twenty different ethnic groups resided in the Volga Region at this time. In 1891, crop failure combined with depleted reserves from previous years led to a severe food shortage. Families faced difficult intrahousehold resource allocation decisions – and unlike adults, infants and children were wholly reliant on their parents' decisions.

Each ethnic group varied significantly in its view of women. We utilize oral motifs data from Michalopoulos and Xue (2021), who use Berezkin's *Folklore and Mythology Catalog* (2015), which collects oral motifs from 958 ethnic groups. These authors take the *Catalog* and group various motifs from a given ethnic group by themes. Using Folklore data offers us several advantages. First, it contains the most comprehensive information on the distribution of different groups of motifs. Of particular interest to us are the motifs depicting men or women as submissive or stupid.[11] Second, this dataset covers a large number of ethnic groups that we can directly link to the ethnic population shares in different districts that we digitized from the 1897 Population census. Third, unlike other measures of gender discrimination, the distribution of motifs originated within the conditions of traditional societies many centuries ago and thus cannot be affected by economic factors during the period of study.[12]

To calculate female negative bias for each ethnic group, we calculate the difference between the number of submissive-stupid motifs depicting women and men as such, divided by

[10] The original data was collected by Berezkin 2015. Then, Michalopoulos and Xue (2021) demonstrated a wide range of possibilities for using these data in economic research (for example, studying the economic status of women).

[11] Some examples of descriptions of motifs found in the Catalog that Michalopolous and Xue (2021) rate as likely depicting women as submissive and dependent include: "a magic medicine demonstrates that the only chaste woman is a servant girl, orphan, etc. King chooses the chaste one," or "a task-giver asks the hero to get a woman." Online Appendix G includes additional information about motifs.

[12] The late 19th century Russian Empire represented a rapidly industrializing country.

the total number of motifs. We then create a population-weighted average of this measure for each district of the Russian Empire. The 5-95 percentile range shows sizable variation of 5.5-10 percent.[13] Other themes suggesting discrimination against women also appear in the database. Prominent themes include role in domestic affairs, violence, physical activity, and sexuality. For studying resource allocation decisions, we believe the measure of submissiveness and stupidity to be most relevant, and we verify this assumption statistically.

We validate our female negative bias measure against other proxies of gender equity. We find that the share of young single women was negatively correlated with this measure. Additionally, we find that places with higher bias had lower female labor force participation rates and lower women's employment share in high-skill professions *relative to men* in 1897. Thus, we have confidence that this measure truly captured ingrained gender biases that were correlated with women's life outcomes.

To identify the causal effect of famine on sex-biased mortality, we apply the difference-in-differences methodology by comparing cohorts born before and after the famine in the areas with high versus low famine intensity. We leverage information on the provision of food relief from the central government to the population in the affected areas to construct a district-level measure of famine intensity. Additionally, we leverage historical climate data to predict food provision from the government. Our predicted measure reflects the deviation of 1891 precipitation and temperature levels from long-term mean values. Thus, it allows us to overcome potential biases if the food distribution was affected by factors correlated with our measure of gender bias. In general, our findings are consistent for both measures of famine intensity. Our results are also robust to alternative measures such as the per capita size of food loans provided by the government and the duration of the food relief program.

Ultimately, we are interested in the interaction of the negative female bias with our famine measure. The cohorts exposed to both higher gender bias and higher famine intensity indeed exhibit relatively higher sex ratios, thus indicating relatively higher survival chances for boys. We also demonstrate that accounting for economic factors does not explain away the effect of culture. This result stresses the importance of deep-rooted gender stereotypes independently of economic considerations. Importantly, we demonstrate that accounting for religion does not

[13] In comparison, the well-known measure of historical gender norms (plow adoption) from *Ethnographic Atlas* offers us only a tiny variation as most of the societies within the region of study adopted plow.

remove the significance of the effects associated with cultural norms. Hence, aspects of cultural identity seem to matter regardless of religion.

We strengthen the interpretation of our regression analysis by comparing the effects of exposure to famine with exposure to cholera outbreak caused by the spread of pandemic disease from Asia that hit both the famine-stricken and other areas in 1892. We demonstrate that the coefficient on the interaction between the measures of cholera exposure and gender bias has low magnitude and statistical precision. While parents had agency in food distribution and thus survival outcomes during the famine, they had limited agency over survival outcomes of contagious diseases. We further validate our findings by performing several sensitivity checks: accounting for outliers, spatial autocorrelation in residuals, and an alternative definition of the outcome variable. Additionally, our findings hold after incorporating gender-biased motifs other than submissiveness into regression analysis.

The effect of patriarchal norms on girls' differential survival likelihood after the famine may have manifested in several different ways. In particular, the gender bias may have resulted in a direct withholding of resources from female children relative to male children. However, the effect may also have operated through the son-stopping rule in societies with son preference (for example, Basu and De Jong 2010). If parents continue having children until they reach their desired number of sons, the girls would be more likely to grow up in larger families. As a result the girls would be more vulnerable to famine making them disproportionately exposed to its adverse effects. While we cannot test these explanations directly, we hypothesize that both of them explain the differential survival of girls.[14]

In its essence, our paper connects pre-existing gender norms, a resource shock, and populations' gender-biased mortality response. It, therefore, directly contributes to the literature tying resource shocks and gender-biased consequences. First, our results confirm a well-documented pattern that women generally have higher survival chances in harsh environmental conditions (for example, Zarulli et al. 2018).[15] However, this general pattern masks a high heterogeneity evident in region-specific studies. On the one hand, some studies document

[14] In section 2, we discuss existing studies providing qualitative evidence on the differential treatment of boys and girls in the Russian Empire. Further, in Online Appendix C we provide suggestive empirical evidence on the possibility of son-stopping rule behavior.

[15] We demonstrate that aggregated sex ratios at age 0-4 tend to be lower in the areas with higher mortality rates, not caused by famine.

worsening excess female mortality following harsh economic crises in societies characterized by strong son preference (Das Gupta and Shuzhuo 1999; Aldashev and Guirkinger 2012[16]). By contrast, other works studying different severe 19th and 20th-century famines (such as the great Finnish famine of the 1860s, the Madras 1876–1878 famine, the 1945 Dutch Hunger Winter, or the 1959-1961 Chinese Famine) do not find consistent results (Pitkänen 1993; Maharatna 1996; Sami 2011; Song 2012), pointing to underlying heterogeneities of sex ratio response to resource shocks.[17] Likewise, Bhalotra (2010) finds that infant mortality in India is strongly counter-cyclical – it is much higher during economic downturns – and the increased mortality comes entirely from girls. What can cause such differences across regions? Our paper shows that heterogeneities in cultural norms can mediate the effect of resource shocks on gender-biased investments, suggesting that these relationships should always be viewed in their specific contexts.

Our paper also contributes to understanding the "missing girls" phenomenon.[18] One strand of the literature directly connects women's relative economic value and their survival chances (Alesina, Giuliano, and Nunn 2013, 2018; Xue 2016; Qian 2008; Becker 2022).[19] An alternative explanation underlines the role played by cultural dimensions, such as the dowry and patrilocal and patrilineality system, as well as other social norms making sons responsible for old-age support and religious rituals (Jayachandran and Kuziemko 2011; Jayachandran 2015; Bhalotra et al. 2020).[20] We connect existing gender norms and sex-biased vital outcomes using a "natural

[16] While Das Gupta and Shuzhuo (1999) study the famine that devastated China between 1959-61, Aldashev and Guirkinger (2012) explore the events that followed the Russian army's entry into Kazakhstan between 1898 and 1908.

[17] Relatedly, in Indonesia, a financial crisis did not lead to worse outcomes for girls in health, education, or mortality (Levine and Ames 2003).

[18] For example, Sen 1992, Das Gupta and Shuzhuo 1999, Hesketh and Xing 2006; Bhaskar and Gupta 2007; Anderson and Ray 2010; Chao et al. 2019; Fenske et al. 2022; Becker 2022 in the context of Asian countries and Beltrán Tapia and Gallego-Martínez 2017; Szoltysek et al. 2022; Malein and Beltrán Tapia 2022; Beltrán Tapia and Cappelli 2023 for the historical Europe.

[19] Alesina, Giuliano, and Nunn (2013, 2018) find that plow-based farming favored men because it required physical strength, leading to men having higher economic value in societies that relied on plows. Xue (2016) finds that improvements in women's economic role caused by the Cotton Revolution in Imperial China improved the social status of women. In a more modern context, Qian (2008) explored how specific agricultural reforms, which favored crops that women were particularly skilled at cultivating, led to increased income for women. This increase in female income, in turn, is directly correlated with higher survival rates for girls in China. Furthermore, recent findings by Becker (2022) demonstrate that younger ages at marriage, for example, child marriage, are particularly prevalent among descendants of historically pastoral societies.

[20] Daughters in India are weaned sooner than sons in order to regain fertility and achieve the number of desired sons (Jayachandran and Kuziemko 2011). Likewise, Bhalotra et al. (2020) show how an exogenous shift in the price of dowry increases the cost of raising girls and reduces their survival chances.

experiment" of a famine that hit ethnicities with different beliefs about women's values and induced a resource constraint, resulting in sex-biased mortality. We, therefore, stress the role of culture over economic factors in explaining the survival chances of girls. In this regard, our contribution also relates to the literature showing the persistence of gender norms even after dramatic changes in the economic environment (for example, Grosjean and Khattar 2019; Baranov et al. 2023; Galor et al. 2020).

The rest of the paper is organized as follows. First, we provide background information on the Russian Empire's demographic landscape and prevalent gender norms and discuss the causes of the 1891/92 Famine. Next, we discuss data sources and the construction of the main variables, formulate the testing hypothesis, and describe our methodology. Next, we present our main results and robustness checks, followed by concluding remarks.

Historical context

Although the Russian Empire was one of the largest producers of cereal in the world, famines occurred regularly throughout the 19th and the first half of the 20th centuries (Wheatcroft 1983; Livi-Bacci 1993; Adamets 2002, p. 158).[21] According to some historians, one of the major features separating 19th-century Russia from Western Europe was its inability to eliminate these catastrophes (Hoch 1989, p. 55). Although all regions were equally subjected to the risk of famines up to 1870, these crises became increasingly more localized, especially in the Volga region, the northern Caucasus, southern Ukraine, and Kirghizia (Adamets 2002, p. 160).[22] Contemporaries were well aware of the importance of this problem, as well as of rural poverty in general, and aimed various vital reforms at tackling these (such as the abolition of serfdom in 1861 or the agrarian reform of 1907).

[21] Terrible famines occurred during the Civil War (1921-23), the collectivization period (1932-33), and 1946-47.

[22] One reason for famine localization was the uneven development of transport infrastructure, especially railroads, which left distant and more isolated parts of the Empire more exposed to poor harvests (for example, in Asia). At the same time, market integration encouraged certain regions to specialize in wheat production – main export crop in Imperial Russia. As a result, these regions became more dependent on agriculture and vulnerable to adverse climate shocks. This interpretation is consistent with evidence from Mironov (1985), who documents convergence in regional grain prices and their alignment with world prices in the second half of the nineteenth century. Such price convergence reflects growing market integration, which enabled provinces to exploit their comparative advantage through agricultural specialization.

The 1891/92 Famine was the last large mortality crisis of the *ancien régime* in Russia (Adamets 2002, p. 163). According to various estimates, around half a million people died, making this episode equivalent to the 1846 Irish and 1877 Madras famine (Long 1988).[23] The famine especially hit the Volga basin, a region where different ethnic and religious groups lived together (Orthodox Russians and Ukrainians, Protestant and Catholic Germans, and various groups of Muslim populations – see Figure A1).

The extremely adverse climate conditions were the main trigger of this crisis. Soil scientist P. Zemyatchenskii provided a historical account of the detrimental role of climatic factors:

The dry autumn […], the snowless winter and, finally, the dry spring turned the top layer of […] earth partly into a dry dust, [and] partly into a fine-grained, crumbly, powder, which, with the onset of strong storms in April, lost their hold, and were raised up in whole clouds, concealing the sun's rays and turning day into night. Witnesses unanimously testified that the phenomenon had such a dreadful and frightening character that everyone expected 'the end of the world.'[24]

The majority of the participants of the debates organized by the Imperial Free Economic Society in November 1891 regarding the famine agreed that climate factors played a principal role in harvest failure. As one of the speakers A. S. Ermolov argued: *"In this context I must repeat what was already said by others, that is, that this year no efforts by the farmer, no kind of technical improvement, no kind of treatment and fertilization of the soil would have made any noticeable difference in the localities stricken by the crop failure."*[25]

The crop failure in 1891 was the most significant calamity over a thirty-year span (Figure 1). Accordingly, it caused a significant negative income shock for peasantry: grain production and real wages of agricultural workers reached their two-decade trough (Figure 2).

The Russian government attempted to reduce the negative consequences of crop failure by providing food loans to the population in the affected areas. However, the results of this relief campaign were often delayed due to bureaucratic issues and limited access to transport

[23] It is important to stress that Famine mostly contributed to the population losses indirectly by weakening people's immune systems and making them more vulnerable to infectious diseases such as typhus and cholera (Wheatcroft 1992; Henze 2010).
[24] Piotr Zemyatchenskii, 'Velikoanadol´skii uchastok', Trudy Ekspeditsii, snaryazhennoi Lesnym Departmentom, pod rukovodstvom professora Dokuchaeva, Nauchnii otdel, 1, 1894, 3, p. 15, cited in Johnson (2015).
[25] Cited from Simms (1982, p. 240).

infrastructure (Charnysh, 2022)[26]. Figure A2 shows that most government relief was distributed only in 1892, while the food shortage was already present in 1891. This suggests that the conditions of many households during 1891 were particularly vulnerable.

The famine hit an impoverished society with extremely low standards of living.[27] In 1875, life expectancy was about 30 and 33 years for men and women, respectively (Adamets 2002, 162; Hoch 1998). Infant and child mortality rates were extremely high, probably the highest in Europe, due to mass poverty (Ransel 1991). Broadly speaking, in the European part of the Russian Empire, only about half of the children survived to age five, but this average masks significant regional differences (Patterson 1995; Glavatskaya et al. 2017; Natkhov and Vasilenok 2023).[28] Religious and ethnic dimensions influenced childcare practices that, in turn, shaped infant and child mortality rates (Patterson 1995; Glavatskaya et al. 2017; Bonneuil and Fursa 2017). Orthodox Russians suffered by far the highest mortality rates, compared to Catholics, Protestants, Muslims, and Jews (Natkhov and Vasilenok 2023). Under conditions of high mortality, resignation and fatalism about the fate of their offspring were common among peasants (Frieden 1978, 246). This could additionally contribute to malnutrition and neglect of the children and increase mortality, especially of less valued children.

It is therefore plausible to hypothesize that, under the harsh circumstances caused by a large famine, families treated their sons and daughters differently, especially considering the strong son preference that characterized Russian society at that time (Malein and Beltrán Tapia 2022). Strong patriarchal institutions subordinated women to their fathers and husbands (Evans Clements 1991, 5-7; Glickman 1991, 148-150).[29] The system of apportioning land and strict patrilocal rules favored boys, so parents had incentives to treat their sons and daughters differently.[30] For example, Worobec (1984) discusses how, in the rural customary law of

[26] Almost 12 million people were receiving assistance in the form of food and seed loans or employment on public works at the height of the relief effort in early 1892 (Charnysh 2022).
[27] Although the view that living standards were stagnant in late Imperial Russia has been recently challenged (see, for instance, Mironov 2012, Markevich and Zhuravskaya 2018 or Natkhov and Vasilenok 2023), Russian peasants nonetheless lived at the verge of subsistence as infant mortality rates and the very famine itself testify (see also Dennison and Nafziger 2012). The decades before the famine, the Empire witnessed explosive population growth and a decline in the average peasant's land allotment (Engel 1994, p. 2; Moon 1999, p. 32).
[28] Infant mortality in Perm province in 1896/97 was as high as 437 deaths per thousand live births (Bakharev and Glavatskaya 2019, p. 206).
[29] A more detailed explanation of the Russian patriarchal system during this period can be found in Evans Clement (1991, pp. 5-7), Glickman (1991, pp. 148-150), Worobec (1991, 175-216) and Engel (1994, pp. 8-25).
[30] Several studies provide supportive anecdotal evidence (Semionova 1973, pp. 8-9; Ransel 1988, p. 130; Moon 1999, pp. 182-184, 192; Avdeev et al. 2004, p. 726).

prerevolutionary Russia, property tended to be passed along male lines, with male heirs (sons) commonly preferred under traditional inheritance practices. Likewise, the evidence of infanticide, and especially of child abandonment, also suggests the presence of son preference.[31]

It is also plausible that places with more negative gender norms predisposed girls to worse outcomes, even if explicit withholding of resources was not present. As discussed earlier, in societies with stronger son preference, girls may systematically grow up in larger families due to the son-stopping rule, making them more vulnerable to resource constraints caused by famine.[32] If boys were preferred to girls and thus held a higher social value, girls would have experienced greater peril either through direct resource withholding or a higher likelihood of living in high-risk, large families.

The only existing evidence about gender differential mortality following a crisis in Imperial Russia is Aldashev and Guirkinger (2012). The authors argue that worsening economic circumstances increased excess female mortality in Kazakhstan at the turn of the 20th century, especially in poorer households. However, it remains unclear whether such patterns were confined to this specific context or were more widespread across the Empire. This article addresses this question by turning to the wealth of information provided by the 1897 Russian Imperial Census and other data sources.

Data

Child sex ratios

We construct our primary outcome variable – the ratio between boys and girls - by leveraging the 1897 population Census. This was the first and only true population census designed to fully register the population and its socioeconomic and demographic characteristics (Rowney and Stockwell 1978). Following contemporary international statistical practices, qualified enumerators visited every household in each village, wrote down the number of individuals living there, and classified a rich set of demographic characteristics: age (birth year) and gender, literacy, occupational status, and ethnic and religious denominations. The 1897 Census was not likely to suffer from crisis-related underreporting since it was taken several years

[31] See Malein and Beltrán Tapia (2022) for a longer discussion of these issues.
[32] In Online Appendix C, we document how, in the Russian Empire, the *variance* in household size was highly correlated with the negative gender bias, suggesting that the male-stopping rule could have existed in this society, although we do not have conclusive evidence.

after the 1891/92 Famine and was not plagued by any major famine, disease, or war in a relatively regular year. Crucially, the 1897 Census contains information about the size of the population by birth year and gender at the city and district (rural) level. Since the 1891/92 Famine affected rural areas, we use the population of the latter to construct the sex ratios, defined as the number of boys per girl. We primarily focus on cohorts born after 1887, who were either infants or young children during the 1891/92 Famine, or had not yet been born at the time of the famine. This allows us to focus on the effects of Famine on the groups of the population entirely dependent on parents. Note that sex ratios among older cohorts were likely affected by the famine as women's survival strategies could differ from those of men. Since we focus on parental agency in explaining gender-specific mortality, we exclude older cohorts in most of the regressions throughout the analysis.

Thus, we construct a panel of the gender-specific cohorts born between 1887 and 1897 and calculate the ratio of boys to girls (sex ratios) for the cohorts born before and after the 1891 Famine event. Therefore, those born before 1892 were exposed to the famine, while those born after were never exposed. Figure A3 shows a histogram of the distribution of sex ratios in the estimation sample.[33]

One of the common concerns regarding census data is the presence of age heaping, which is well documented in the literature (see Charnysh 2022). Specifically, in our setting, the concern is that age heaping could depend on gender, for example, census respondents could report their sons' ages more accurately than their daughters'. Potentially, it could create imbalances in the sex ratios and bias our results. Reassuringly, we find no correlation between the measure of age heaping (Myers index) and constructed sex ratios, which strengthens the validity of our data (Figure A4).[34]

Folklore-based measure of female discrimination.

The main challenge in studying the effects of gender discrimination practices is that such practices are hidden, and it is difficult to convert them into a quantitative measure. However, as Michalopoulos and Xue (2021) demonstrate, the differences in oral traditions (folklore) explain

[33] The extended version of the sample also includes the following age groups by 1897: 10-19, 20-29, 30-39, and 40-49.

[34] Beltrán Tapia and Malein (2022) provide a comprehension discussion of the 1897 census data and further validity checks for constructed sex ratios.

the variation in gender disparities across modern societies (for example, countries, groups of migrants). More specifically, the authors explore the dataset of oral traditions of 958 societies across the globe, originally compiled by anthropologist and folklorist Yuri Berezkin, which reflects various topics, such as environmental events (for example, volcano eruption), trust, fraudulent/risk-taking behavior, and gender roles. To our knowledge, this is a unique dataset spanning ethnicities from all over the globe that allows for the direct measurement of historical gender norms. The study identified 1,073 gender-related motifs within these oral traditions. Among these, 30 percent depict women as submissive and dependent, while only 12 percent describe men in such terms. The authors find that countries and ethnic groups with a higher prevalence of motifs favoring men have lower female labor force participation rates today. The results by Michalopoulos and Xue (2021) demonstrate that folklore data reflect deep-rooted gender attitudes that originated from the historical past and persisted over time. These data are free from many confounding factors associated with analyzing facets of ethnic or religious groups recorded contemporaneously with the studied event. These gender norms had existed for centuries and were not likely to be related to the causes or consequences of the 1891/92 Famine.

We use these data in the following way: first, we link each ethnic group from Michalopoulos and Xue's (2021) sample to the corresponding group in the 1897 Census, giving us 68 unique ethnic groups.[35] Next, within each ethnic group g's set of motifs, we calculate the proportion of the motifs containing the image of women as submissive and stupid relative to men:

$$Share\ motifs_g = \frac{\#\ female\ submissive\ motifs_g - \#male\ submissive\ motifs_g}{\#\ total\ motifs_g}$$

We take the difference between the female and male motifs to account for such instances where certain groups may have had a multitude of motifs related to submissiveness or stupidity of both men and women. Counting only female or male motifs may overestimate gender-specific negative bias.

We then aggregate these group-specific measures at the district level using the group's population shares as weights to obtain a population-weighted average of our bias measure:

[35] Note that 1897 Census reports mother tongue that we use as an indicator of ethnicity.

$$Female\ negative\ bias\ _i = \sum_{g=1}^{G} Share\ motifs_g \times Share\ pop_{gi}$$

Figure 3 shows the spatial distribution of our female negative bias measure. Variation was relatively large in the Volga region, enabling us to conduct meaningful statistical analysis using this variable. To validate our measure of gender stereotypes based on folklore, we turn to the results by demographers (see Szołtysek et al. 2017 and Beltrán Tapia and Szołtysek 2022). In these studies, the authors construct a Patriarchy Index—a composite measure of gender inequality that includes, among other components, the share of young brides. They show that this index is positively and statistically significantly associated with child sex ratios across Europe. Accordingly, we assume that a lower share of single women among those aged 15-29 would indicate a lower female bargaining position and their disadvantageous status in society. Indeed, the data from the 1897 Census shows that the share of singles among young women tend to be lower in the areas with a higher prevalence of gender bias (Figure 4).[36]

Furthermore, we reveal a negative correlation between the folklore-based measure of gender bias and female labor force participation and women's employment share in high-skill occupations (Figure 4).[37] Additionally, in Table A3, we demonstrate that the prevalence of other motifs related to domestic affairs, violence, physical activities and sexuality do not predict women's economic status, further validating our preferred measure of gender bias. Taken together, these results indicate that women in the areas with higher gender bias had lower agency over their marriage and limited opportunities in the labor market, showing that folklore-based measure of gender discrimination indeed reflects the disadvantaged economic and social status of women in society.

The geography of the famine

This article aims to assess how the 1891/92 Famine shaped the survival chances of boys and girls. As explained above, the famine mainly affected the Volga basin, a highly diverse region where different ethnic and religious groups lived together. It then spread further, affecting provinces in the northern and central parts of the Empire. The central government supplied food

[36] We collected data on age and marriage status only for a subsample of the provinces affected by the famine.
[37] To construct these measures we applied data from the 1897 Population Census and coded high-skill/human capital occupations such as public administration, science, literature, art, teaching, and medicine.

relief in the affected areas. Figure 5 shows the actual shares of the population receiving food support in 1891/92. Hence, our food relief measure allows us to measure the intensive margin of the exposure to the famine. We validate this measure by showing its high correlation with the 1891 harvest productivity (Figure A6)[38].

One of the concerns regarding this measure of famine intensity is that food provision in the affected districts can be correlated with various factors, which can bias the results of our analysis. To mitigate this concern, we first demonstrate that our measure of gender bias is orthogonal to the provision of food support (Figure A7). Additionally, we construct an alternative measure that exploits districts' exposure to climate shock – adverse climate conditions that triggered the 1891/92 Famine. We use historical climate data to explain variation in the distribution of food relief by climate shocks observed during the winter, spring, and summer seasons of 1891. To construct the measure, we compute the average temperature and precipitation for each month and climate station using daily observations from 1892-1901 as a reference period.[39] Then, we take monthly averages of temperature and precipitation in 1891 and compute the climate shocks as follows:

$$z_{im,1891} = \frac{x_{im,1891} - \underline{x_{im}}}{\sigma_{im}}$$

Given that index *i* denotes a climate station, and m stands for a month, $\underline{x_{im}}$ is an average of precipitation/temperature observed in the month *m* between 1892 and 1901 at climate station *i*. σ_{im} is a standard deviation of monthly temperature/precipitation observed in 1892-1901. As we observe 1891 climate shocks at the climate station level, we create a spatial grid covering the Russian Empire's territory. Then, for each empty grid cell (unobserved climate data), we compute values using the *Inverse-Probability-Weighting* (IPW) method and aggregate grid values to the district level. We utilize information about 14 climate shocks that affected the harvesting period in 1891 (between January and July of 1891) alongside exogenous soil

[38] To compute a measure of harvest we take an average of rye, summer wheat, winter wheat and potato yields weighted by the crop's sown area. The yield = total harvest – seeds / total sq. sown land.

[39] Our data includes historical daily temperature and precipitation observations from 52 climate stations in the European part of the Russian Empire. We obtain the data from National Oceanic and Atmospheric Administration (NOAA) https://www.ncdc.noaa.gov/cdo-web/

productivity[40] to predict the proportion of the district's population receiving food relief from the government. Then, we use predicted values to measure famine intensity in our subsequent analysis.

Table A2 shows the estimated coefficients of the climate shock on food relief. Our estimates are consistent with a general narrative provided by historians emphasizing the detrimental impact of the dry early spring that left grain seeds unprotected from frost, followed by a hot summer. This is consistent with Figure A8 showing that unusually high temperatures in July 1891 hit Low Volga and Southern Provinces in the European part of Russia.

By construction, the predicted food relief measure reflects only the differences in climate conditions in 1891 and cannot be biased due to the government's preferences and local economic conditions. Later, we provide the results based on both actual and predicted measures of food provision.

Additional variables

We draw data on births and infant mortality from the Vital Statistics volumes collected by the Ministry of Internal Affairs. The vital statistics in the Russian Empire in the second half of the 19th and the beginning of the 20th century were primarily based on metric books by clergymen. In 1865, the government introduced the procedure of reporting data to the state authorities, who then aggregated data at the district level and sent it to the Central Statistical Committee (CSC). The Central Statistical Committee reported data in tabular format in annual statistical volumes *"Dvizhenie Naselenija" (Population Movement)*. Novosel'skij (1916) reports that statisticians followed a universal and reliable data collection program. The data prepared by CSC demonstrated consistency with alternative sources, such as the State Medical Department's mortality indicators, that signal the data's suitability for statistical analysis (Rashin 1956). We leverage this data to construct infant mortality rates and birth ratios that we include as control variables in our analysis.[41]

[40] We apply the Caloric Suitability Index constructed by Galor and Özak (2016).

[41] An important caveat is that administrative mortality records report combined data for both rural areas and cities located in these districts (the separate data is only available for large cities). Since the Famine was predominantly a rural phenomenon, the measures based on administrative data are less accurate. Still, it reflects useful information we utilize in our regression analysis as control variables.

Additionally, we augment our dataset with exogenous variables that can affect the formation of gender norms and intensity of the 1891/92 Famine: soil productivity, temperature, precipitation, ruggedness, distances from the district's centroid to a navigable river, coastline, and capitals (Moscow and Saint-Petersburg)[42]. Table A1 provides summary statistics of the main variables.

Methodology

Hypotheses

Absent gender-specific preferences, the biological disadvantage of boys, which other studies demonstrate (for example, Zarulli et al. 2018) and we also observe in our data (see Figure A9), should theoretically be aggravated by the extreme living conditions during the famine. Then, exposure to famine should result in relatively *lower* sex ratios among the surviving cohorts born before 1892 and thus affected by the Famine. These children would have been between the ages of 0-4 when the famine hit in 1891. However, if families somewhat prioritized boys, the results could be different. If cultural norms favoring boys matter, the joint impact of famine intensity and gender bias should offset the female biological advantage. Thus, our main testing hypothesis is that in the presence of cultural bias against women, the famine effect results in relatively *higher* sex ratios among exposed cohorts.

Empirical strategy

We investigate the joint effect of gender bias and exposure to the 1891/92 Famine on the survival chances of girls relative to boys by leveraging district-level variation in exposure to the 1891/92 Famine and prevalence of the cultural norms against women with cohort-level variation in the birth year (before and after the 1891/92 Famine). Equations 1.1 and 1.2 formalize our difference-in-differences approach. Equation 1.1 tests the general impact of famine intensity on the survival chances of girls, whereas equation 1.2 (triple difference) tests the heterogeneity of the famine effect on sex ratios depending on the prevalence of the gender bias:

[42] As a measure of exogenous soil productivity, we use the Caloric Suitability Index (Galor and Özak 2016). GAEZ provides climate data, and the computation of the distances is based on the Russian Empire shape file supplied by Kessler, 2017: "Maps," https://hdl.handle.net/10622/DN9QDM, IISH Data Collection, V2. Ruggedness data is provided by D. Puga https://diegopuga.org/data/rugged/

$$Sex\ ratio_{it} = \gamma_i + \mu_t + \beta_1 1(Year < 1892)_t \times Famine_{i,1891} + \delta X_{it} + \varepsilon_{it} \textbf{ (1.1)}$$

$$Sex\ ratio_{it} = \gamma_i + \mu_t + \beta_1 1(Year < 1892)_t \times Famine_{i,1891} + \beta_2 1(Year < 1892)_t \times Famine_{i,1891} \times Bias_i + \beta_3 1(Year < 1892)_t \times Bias_i + \delta X_{it} + \varepsilon_{it} \textbf{ (1.2)}$$

$Sex\ ratio_{it}$ represents the cohort-specific ratio of boys relative to girls. This measure reflects the relative survival chances of girls. $1(Year < 1892)_t$ denotes a binary indicator that is equal to one for the cohorts born between 1887-1891 – before or during the famine; and equal to zero for the cohorts born between 1893-1897.[43] The 1892 cohort is excluded in our main estimation because the supplied food relief mainly reached the target population in 1892 – the year following the harvest failure (see Figure A2). Hence, the harsh resource constraints created by harvest failure were partially alleviated by 1892. The 1892 cohort is also excluded because it consisted of the children born in the second half of 1892 when the situation with food availability significantly improved due to a new better harvest.[44] $Famine_{i,1891}$ measures the proportion of the district's population that received food support from the central government during the relief campaign in 1891/92. $Bias_i$ is a measure of gender bias, discussed in the previous section. It denotes the population-weighted average of the motifs depicting women as submissive or stupid. γ_i and μ_t capture all time-invariant district-level characteristics as well as the factors associated with particular cohorts.

Vector X_{it} includes additional controls derived from administrative mortality and birth records. Firstly, we include the ratio of boys to girls at birth to account for the negative effect of nutritional stress of mother on the probability of giving birth to a son (Myers, 1978). Secondly, we account for the mortality factor. It is worth stressing that the mortality rate is negatively associated with sex ratios as shown by Figure A9. Although district fixed effects control for time-invariant differences in baseline mortality across districts, famine shocks may still have heterogeneous effects. That is, districts with persistently high (or low) mortality may differ in

[43] To analyze the sex ratios, we limit our sample to European provinces and exclude the Caucasus region. We do it for three main reasons. Firstly, the administrative mortality data are available only for European provinces. Secondly, the sex ratios in Caucasus provinces seem to be highly distorted (in some cases, we observe values above 1.5-2), and we need to investigate these cases further. Thirdly, the Caucasus region differed from the European part in many dimensions and can be considered an outlier. It is worth noting that the Caucasus region was outside the 1891/92 Famine area.

[44] In Online Appendix F, we elaborate more on the reasoning behind exclusion of the 1892 birth cohort. Our results nonetheless do not change when we include 1892 in the treated group.

their vulnerability to famine and in how famine translates into excess mortality.[45] To account for this possibility, we include a triple interaction between the cohort indicator (year < 1892), baseline infant mortality rates (not affected by famine) and the famine-intensity measure in our regression specification. As an additional proxy for mortality, we also include cohort size normalized by the district's average cohort size in 1887–1897. Additionally, we include triple interactions with other geographical characteristics of the districts. We motivate the inclusion of these controls by the fact that some of them are correlated with the measure of gender bias as shown by Figure 6. Together, these controls capture whether areas with various observables (for example, geographical location, climate) responded differently to famine shock. The more restrictive version of equation (1) also includes the interaction of province indicators with cohort dummies, allowing us to account for all province-specific factors varying across cohorts.[46] Finally, we consider specification with district-specific linear trends as yet another check for the robustness of our findings.

The sign and significance of the coefficient β_1 in equation 1.1 reflects the two opposite effects discussed above: biological factors ("-") and gender bias ("+"). If positive, it would indicate the stronger impact of gender bias. The coefficient β_2 in equation 1.2 reflects the joint importance of gender cultural norms and resource constraints created by famine for the relative survival chances of girls. Positive and significant β_2 would support our testing hypothesis.

Results

1891/92 Famine and population losses

First, we apply a simple cross-sectional regression approach and show that the 1891/92 Famine produced a significant death toll on the population (see Figure 7).[47] Moving from the 5th to the 95th percentile of the famine intensity increases the infant mortality rate in 1891 by 3.8 deaths per 100 births, corresponding to an 18 percent increase relative to the sample mean of 21.13. These estimates confirm that the 1891/92 Famine was probably the most adverse demographic shock in the late 19th-century Russian Empire. Population losses resulting from the Famine are also evident in population data from the 1897 Census. The relative sizes of the

[45] We provide an additional discussion in section "Interpretation of the empirical findings".
[46] Provinces formed the first unit of administrative division in the Russian Empire. District ("uezd") and county ("volost") constituted second and third units. The 1897 Census data is available mostly at the district level.
[47] See also spatial variation in 1891 infant mortality rates on Figure A5.

cohorts born before 1892 in the affected districts are significantly lower than those in the non-famine districts. In particular, our estimation shows that the cohorts in the districts with high famine intensity (95th) were 8-10.2 percent smaller, a result consistent with the infant mortality data and with the dramatic effect of this demographic shock. (Table 1).

Gender bias and survival chances of girls

The estimation results of the main empirical specification. Regression analysis based on equation 1.1 provides results consistent with the proposed hypothesis. It is important to emphasize that in the case of neutral gender norms, exposure to famine shock should theoretically produce a negative impact on sex ratios. However, one can observe that the coefficient β_1 on famine intensity is positive and significant (Table 2, column 1). This means that the survival advantages of girls in the areas hit by famine were offset by other factors.

We then estimate equation 1.2 with an additional interaction term $1(Year < 1892)_t \times Famine_{i,1891} \times Bias_i$. The coefficient β_2 appears with a positive sign, although it is not statistically significant (Table 2, column 2). However, the addition of province-by-cohort fixed effects and then additional controls interacted with the bias measure improves the magnitude and precision of the coefficient β_2 (Table 2, columns 3-6). It is also worth mentioning that both cohort-varying controls – sex ratio at birth and cohort size have significant coefficients with expected signs. Our estimation results, presented in Table 2 Panel B, demonstrate that switching to the predicted measure of famine intensity reveals the same pattern – the cohort sex ratios increase with a joint increase of predicted famine intensity and gender bias. Since the estimates of famine intensity measures based on government food relief and the predicted measure of famine intensity are quite close to each other, we apply actual value of food relief in our subsequent analysis. In Online Appendix D, we also replicate our main results with the predicted measure based on the climate shock. In Online Appendix E, we replicate our findings using alternative measures such as food loans per capita and months on food relief. We also consider the 1891 harvest as an alternative measure.

Our results indicate that the combination of famine shock and cultural bias against women accounted for the decreased survival chances of girls. Quantitatively, our estimates imply that moving from the 5th to the 95th percentile of famine intensity with an average value of bias against women would decrease the relative survival chances of girls born before 1892 by

approximately 9 percent compared to the cohorts born after the famine [1.8*0.63*0.08]. Given the sample average sex ratio [0.985] and the corresponding size of the cohorts of boys and girls [2262 and 2290], the obtained estimate translates to the decrease in the number of girls per average cohort by 186 [2262 / (0.985+0.09) – 2290].[48]

Alternative regression specification and the sample composition. Further, we estimate a flexible regression specification by interacting the measure of gender bias with cohort indicators and splitting the sample into two parts – districts with high famine intensity (> 75th percentile) and not affected by famine.

$$Sex\ ratio_{it} = \gamma_i + \mu_t + \sum_{t=1887;t\neq1893}^{1887} \beta_t Year_t \times Bias_i + \delta X_{it} + \varepsilon_{it} \mid Famine_{i,1891} > 75^{th} \ (\mathbf{2.1})$$

$$Sex\ ratio_{it} = \gamma_i + \mu_t + \sum_{t=1887;t\neq1893}^{1887} \beta_t Year_t \times Bias_i + \delta X_{it} + \varepsilon_{it} \mid Famine_{i,1891} = 0 \ (\mathbf{2.2})$$

This specification represents a difference-in-differences model where the treatment indicator is a continuous measure of cultural bias against women interacted with a full set of cohort dummies. We measure the effect relative to the 1893 cohort, which serves as a reference category. Essentially, equation 2.2. serves as an empirical test of the assumption that in the absence of the resources shock, the gender bias would not affect the *relative* survival chances of girls born before and after the famine. In this flexible difference-in-differences model, we expand our sample by including all available cohorts, including adults. While our primary focus is on the survival chances of children, this approach enables us to assess the heterogeneity of the effects across different age groups.

The obtained results correspond to the estimates reported in Table 2. As Figure 8a shows, the cohorts born in or before 1892 have consistently higher sex ratios in the famine districts with higher gender bias. Reassuringly, we do not find any consistent patterns in the sex ratios of the cohorts in the districts not affected by the famine (Figure 8b). One can also observe that exposure to famine and gender bias affect sex ratios among children rather than adults. The pattern of coefficients for adults is similar in both samples.

[48] We use estimates of Table 2, column 4 for this calculation.

Accounting for the economic value of female labor

To test whether the effect of cultural norms arises from economic factors affecting the relative value of girls, we extend our analysis by including interactions of famine exposure with measures reflecting the economic value of female labor and general economic development: M/F ratio in high-skilled occupations, share of industrial workers in the labor force and urbanization rate. Additionally, we account for the differences in agricultural specialization (crop farming vs. cattle breeding) by including the interaction of famine intensity with a proportion of workers employed in cattle breeding in the regression. Crop farming in the Russian Empire was associated with intensive female labor, unlike cattle breeding. The purpose of this exercise is to test whether adding the proxies for the value of female labor can explain away the effect associated with gender bias.

As Table 3 demonstrates, the coefficient on women's relative employment in high-skilled occupations appears with the expected sign, while other variables are zero or positive. The inclusion of these variables does not alter the magnitude and significance of the gender bias coefficient. It stresses the importance of cultural factors in driving gender-discriminatory practices under harsh conditions.

Accounting for religion and different groups of motifs

In Table 4, we test whether the joint importance of famine and gender bias on sex ratios holds after accounting for religious factors. We augment our regression specification by adding interactions between famine intensity and district-level population shares of the main religious groups (Muslims, Protestants, Catholics, Jews, and Old Believers). Interestingly, one can observe that adding religious groups does not alter the sign and significance of the coefficient on gender bias. The coefficients on a few religious groups (Protestants, Catholics, Jewish) remain significant with the inclusion of the measure of gender bias, indicating that religion played an important role in explaining the survival chances of girls independently of cultural norms expressed in folklore.

Robustness and additional results

One of the potential concerns regarding the interpretation of our findings is the spread of infectious diseases that could trigger differential mortality rates among boys and girls in the areas

hit by famine. It is a documented fact that the areas affected by the 1891/92 Famine were also hit by cholera in the subsequent year (see Figure 9), although the cholera outbreak was wider in geographic scope. To test the significance of exposure to infectious disease, we collect data on registered cholera cases in each year between 1887 and 1895.[49] One can see the large spike in the number of cholera cases in 1892 (Figure A11) that was primarily caused by the pandemic originating in Asia and then brought to the southern provinces of the Russian Empire via trade routes and then to the areas on the North through the Volga river (Henze 2010). Leveraging this data, we construct the measure of exposure to cholera, which varies by cohort and district. Then, we augment regression equation (1) by including an additional interaction term: $Cholera_{it} \times Bias_i$.

$$Y_{it} = \gamma_i + \mu_t + \beta_1 1(Year < 1892)_t \times Famine_{i,1891} + \beta_2 1(Year < 1892)_t \times Famine_{i,1891} \times Bias_i + \beta_3 Cholera_{it} \times Bias_i + \beta_4 1(Year < 1892)_t \times Bias_i + \beta_5 Cholera_{it} + \delta X_{it} + \varepsilon_{it} \quad \textbf{(3)}$$

It allows us to disentangle the effect of the drastic resource shock created by famine from the effects of the spread of infectious diseases. The results of our analysis presented in Table 5 indicate that the joint effect of famine and gender bias is not affected by accounting for the spread of cholera. We interpret this in favor of the proposed hypothesis. Since parents have limited or no agency over the impact of highly contagious diseases (cholera), gender bias should not influence the relative survival chances of girls.[50]

To provide further support for our findings, we estimate our regression model for the sample of urban areas (cities and towns). Since the famine hit rural areas, one should reasonably expect null effects of famine intensity on the sex ratios of the cohorts in urban areas.[51] This is confirmed by the estimation results presented in Table 6.[52]

We provide several additional exercises to test: 1) the sensitivity of the results to the alternative computation of the standard errors (columns 2-4 in Table A4); 2) sensitivity to

[49] Unfortunately, the information on the spread of cholera is not available in 1896 and 1897. We impute missing values using observations from 1895.
[50] This is a plausible assumption given that most of the rural population was poor with a general lack of knowledge of hygiene and limited access to healthcare facilities.
[51] The validity of this exercise relies on the assumption that the population did not move outside the areas of their permanent living which we discuss and justify in section "Interpretation of the empirical findings".
[52] See further discussion in section "Interpretation of the empirical findings".

outliers (columns 5-8 in Table A4 and Figure A12); 3) an alternative definition of the outcome when we apply the percentage of boys in cohort instead of sex ratios (Table A5); 4) sensitivity to the addition of other groups of motifs (Table A6). Additionally, we compare our baseline estimates with the estimates obtained after reshuffling the distribution of gender motifs across ethnic groups (permutation test, Figure A13). On top of that, we provide additional tests showing robustness of our findings to binary treatment (Online Appendix B) and applying alternative measures of the famine intensity (Online Appendix E).

Interpretation of the empirical findings

In this section, we discuss the conditions under which our empirical findings can be interpreted as causal.

First, the causal interpretation of our findings requires that the famine intensity is properly measured. This assumption would be violated if the government applied a discriminative approach, providing fewer or lower food loans to ethnic/religious minorities. In that case, we would underestimate the true intensity of famine and introduce a bias in our estimates. Although we do not observe a significant empirical link between major groups of non-Orthodox minorities and the provision of food relief (Figure A14), we support our findings by applying the predicted measure of famine intensity as explained above. Since the predicted measure reflects variation in climate indicators, it should be free from potential biases due to discriminative approach in food supply.

Second, famine intensity could trigger migration from the affected to non-affected regions after 1892. Since we observe sex ratios in 1897, it could introduce measurement error and bias our estimation results. To address this concern, we argue that the likelihood of permanent migration among children and other dependent family members was lower compared to male adults due to institutional constraints in Imperial Russia (for example, tax obligations, passport regulation). In addition, we show empirically that famine intensity does not significantly increase the likelihood of migration measured by the share of local-born residents in the district population (Figure A15). Thus, families were unlikely to have moved to nearby districts to flee the famine and thus reassures us that population movements during the famine are unlikely to confound our results. Therefore, we are confident that our findings are not biased due to

improperly measured famine intensity. This implies that the estimates of the famine effect can be interpreted as causal.

The causal interpretation of the gender-biased norms as the mechanism of the famine's impact requires further assumptions. At the baseline, our regression model should be correctly specified and include all covariates of gender bias that have significant effects on sex ratios when interacted with famine intensity. This way, we can, to the best of our ability, isolate the effects of gender norms on the gender-asymmetric mortality response to the famine. Specifically, ethnic groups can be associated with differences in gender norms and general mortality rates that both affect the relative survival chances of girls. For example, it is well documented in the literature that Muslim minorities practiced breast-feeding for a longer period, which improved mortality and hence pushed sex ratios upward. (Natkhov and Vasilenok 2023). Demographers also indicated that the lack of daycare, especially during harvests, contributed to infant and child mortality in 19th-century Imperial Russia (Tezjakov 1904). Hence, specialization in less female-labor-intensive activities (for example, cattle breeding) observed among specific ethnic groups could increase the amount of time available for daycare and thus indirectly affect mortality and sex ratios.

Therefore, the main threat to our identification strategy is that omitted covariates associated with high mortality could produce stronger or weaker effects during the famine. Without accounting for these interactive effects, our estimate of the gender bias effect can be biased and lack precision. This is a valid concern given that the infant mortality rate is indeed positively correlated with gender bias (Figure 6). We do our best to account for this by incorporating mortality rate and other covariates into regression analysis. Further, we implement Hainmueller (2012) approach that calibrates observation weights to impose balance of covariates depicted by Figure 6. This balance improvement makes the results of our difference-in-difference analysis more credible (see Online Appendix B).

Additionally, our estimates by cohort ("event-study" Figure 8a) show no visible trends in sex ratios for the cohorts born after the famine. This suggests that areas with high famine intensity and negative female bias were not on different trends before the famine. Importantly, we also find no systematic pre- and post-event trends for the cohorts born in the districts unaffected by the famine (Figure 8b). The estimation results for urban areas further support the validity of our findings (Table 6). Alongside other exercises, it mitigates possible concerns that

the differential response of gender mortality was caused by the interaction of gender bias with factors other than famine. If such factors were common for urban and rural parts of the districts, they would have been captured by estimates in Table 6.

Concluding remarks

Adverse economic shocks place families under severe resource constraints. The need to ration the scant available resources can, in turn, exacerbate existing gender inequalities, especially in contexts where the perceived relative value of raising boys and girls is different. Studying the 1891/92 Russian Famine, one of the most devastating catastrophes in the history of the Russian Empire, this article sheds light on how culture shapes the allocation of resources and care within families under harsh economic circumstances. Leveraging the relative intensity with which women are depicted as submissive or stupid in the oral traditions of different ethnic groups, our identification strategy relies on a difference-in-differences estimation that compares the cohorts born before and after the famine in the areas with different intensity of famine. Our results show that girls suffered higher mortality rates than boys in areas hit by the famine where cultural norms penalized women.

In order to avoid potential endogeneity in measuring the intensity of the famine using the provision of food relief, we construct an exogenous measure that links the intensity of the famine to the deviation in rainfall and temperature levels experienced during the shock. The cholera outbreak that hit the Russian Empire after the pandemic provides a placebo test since parents had limited influence over who survived contagious diseases. In contrast to their role in allocating resources under the severe scarcity characterized by the famine, the gender-bias measure employed here shows no effect on the sex-specific mortality outcomes arising from the cholera epidemic.

As well as robust to different specifications, our exercise rules out that other mechanisms are driving the results and, therefore, stresses the crucial importance of cultural dimensions in driving gender discrimination. For example, including variables capturing woman's economic status in the regressions does not affect the cultural channel. Moreover, the cultural mechanism stressed here, namely the belief in women's submissiveness and stupidity, remains significant after accounting for religious factors. Even though our estimates demonstrate the importance of religion, the cultural channel is mainly independent of it.

The main limitation of our approach is that we are unable to precisely identify whether excess female mortality is driven by outright neglect, by prioritising boys when allocating the scarce resources that are needed for survival in such an extreme crisis, or through pre-existing differences in living conditions and family size for girls and boys due to the son-stopping rule. In addition, our main explanatory variable measures gender bias according to how women are depicted in the oral tradition of different ethnicities. However, we do not observe sex ratios at the ethnicity-district levels, so we rely on a population-weighted measure of the relative share of each ethnicity in the districts' population. Further research based on more detailed census or administrative data can overcome this limitation. Even though our findings refer to a specific historical period and country – late 19th century Imperial Russia – we believe they can provide potentially generalizable insights into the "missing girls" phenomenon observed in other countries with a predominantly agrarian economy and strong patriarchal norms.

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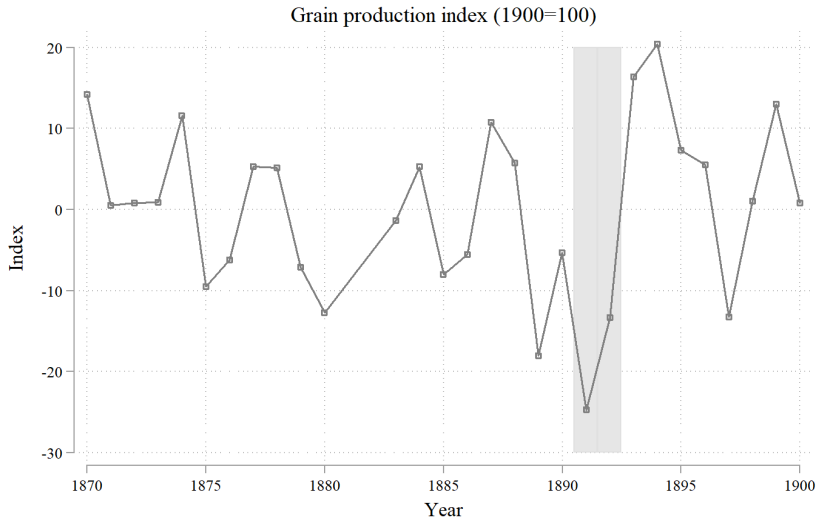
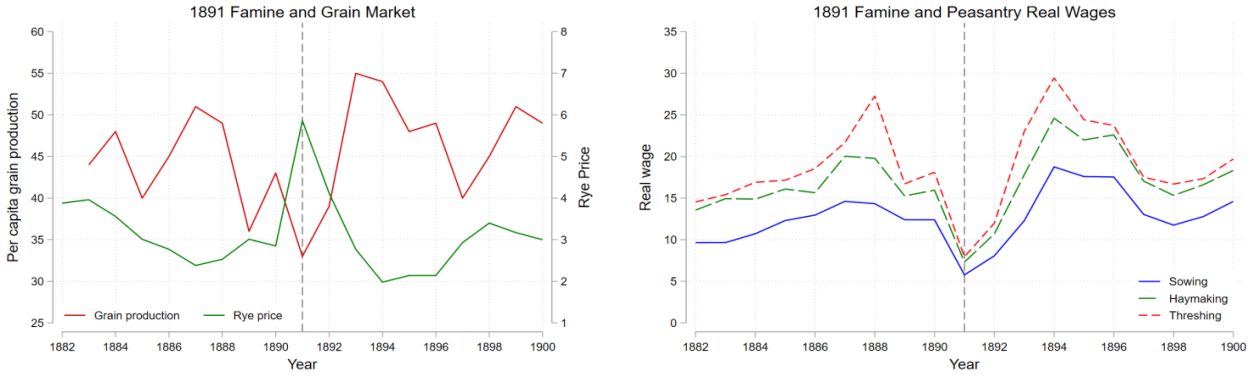
Figures and Tables

Figure 1. Russian Empire grain production index (de-trended) in 1870-1900



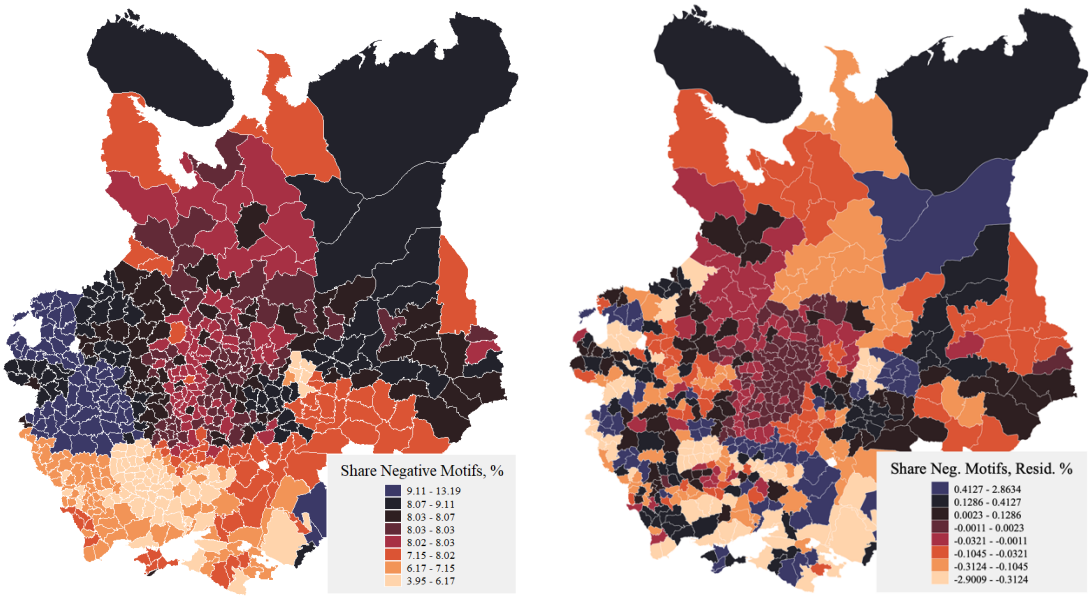
Sources: Raymond W. Goldsmith, "The Economic Growth of Tsarist Russia 1860 1913," Economic Development and Cultural Change, 9 (April 1961): 446.

Figure 2. Relationship between exposure to 1891/92 Famine, Agricultural Prices and Income



Sources: V.M. Obukhov "Dvizhenie urozhaev v evrop. Rossii v period 1883-1915 gg." and S.G. Strumilin "Dinamika posennoi platy za 1883-1916 gg." The data is reported by Wheatcroft in Edmondson, L., & Waldron, P. (Eds.). (1992). *Notes:* This figure shows the evolution of grain production and real wages of agricultural workers (by job category) between 1882 and 1900. Vertical dashed line represents 1891/92 Famine.

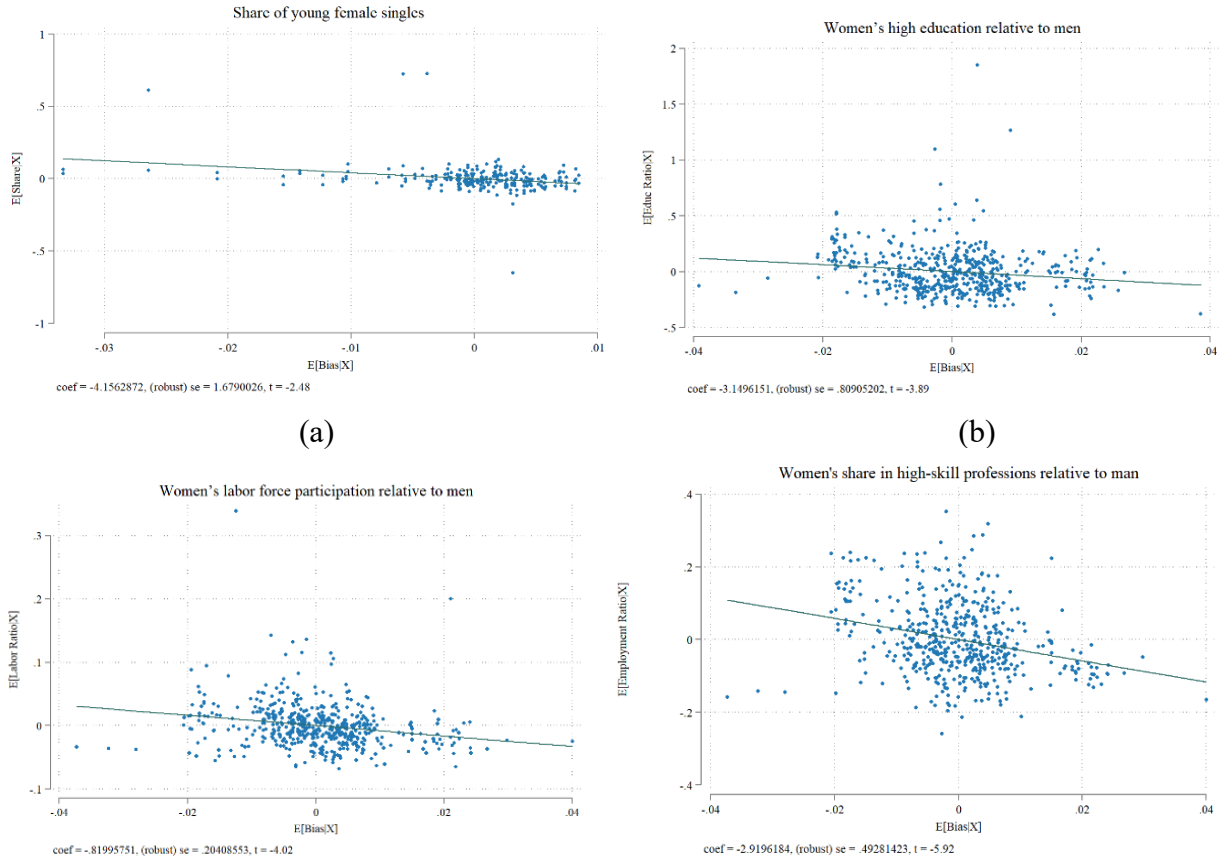
Figure 3. Distribution of the Female Negative Bias in the European part of the Russian Empire.



(a) (b)

Sources: Shapefiles from Kessler and Markevich (2020). *Notes:* This figure shows the spatial distribution of gender bias (Folklore-based) across districts in the European part of the Russian Empire and Caucasus. Figure (a) shows the raw distribution of the (net) percent of the negative female motifs of all motifs. Figure (b) plots the residuals, after controlling for the province fixed effects, showing that between district variation remains, especially in the Volga region.

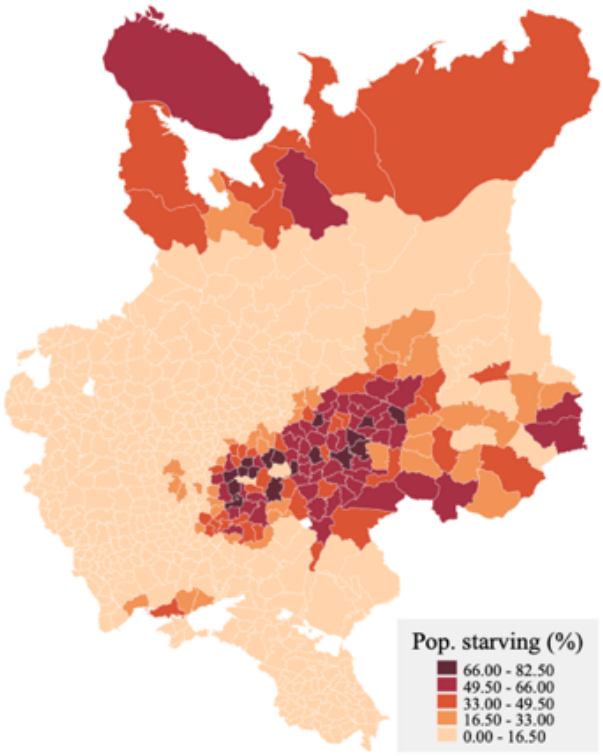
Figure 4. Woman's education and labor force outcomes relative to man



Notes: This figure shows relationship between women's socio-economic status on the measure of gender bias. Panels (a)-(d) shows residual scatterplots between the measures of woman economic status and gender discrimination (folklore bias). These measures include: (a) the share of singles among young women (age 15-29)[53], (b) F/M ratio of district inhabitants with high school degree; (c) with employment; (d) with employment in high-skilled occupations (public servants, doctors, teachers and scientists). We apply the following controls in the regressions: literacy rate, latitude, longitude, exogenous soil productivity, average temperature and precipitation (GAEZ), ruggedness, distances to the railroad, to the nearest navigable river, coastline and capitals (Moscow and Saint Petersburg).

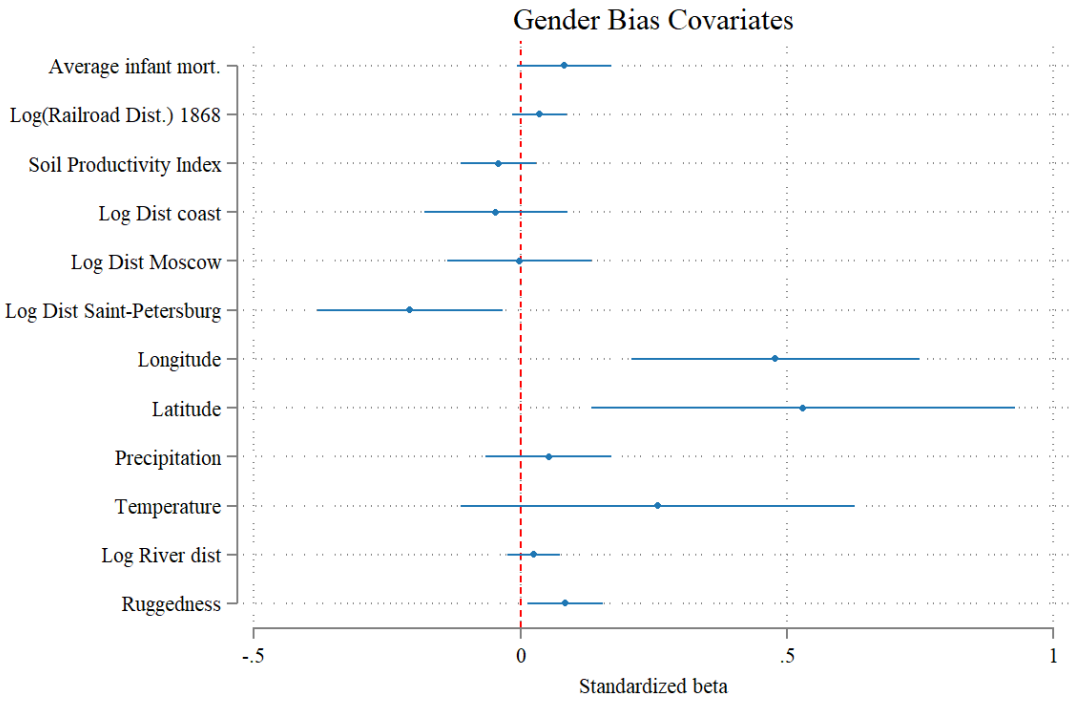
[53] This measure is only available for the provinces affected by the Famine.

Figure 5. The geographic impact of the 1891/92 famine.



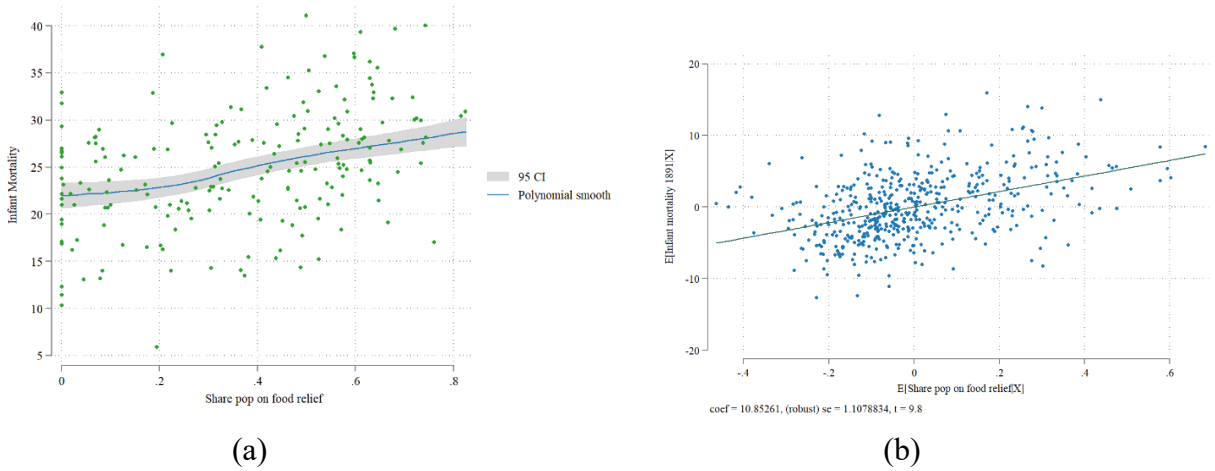
Notes: This figure shows the spatial distribution of the share of the district population receiving food relief at the district ("uezd") level (European part of the Russian Empire).

Figure 6. Female negative bias district-level covariates



Notes: This figure plots standardized coefficients after regressing the folklore-based gender bias measure on the set of district-level controls: average infant mortality rate (excluding famine years), latitude, longitude, exogenous soil productivity, average temperature and precipitation (GAEZ), ruggedness, distances to the railroad, to the nearest navigable river, coastline and capitals (Moscow and Saint Petersburg), and province dummies.

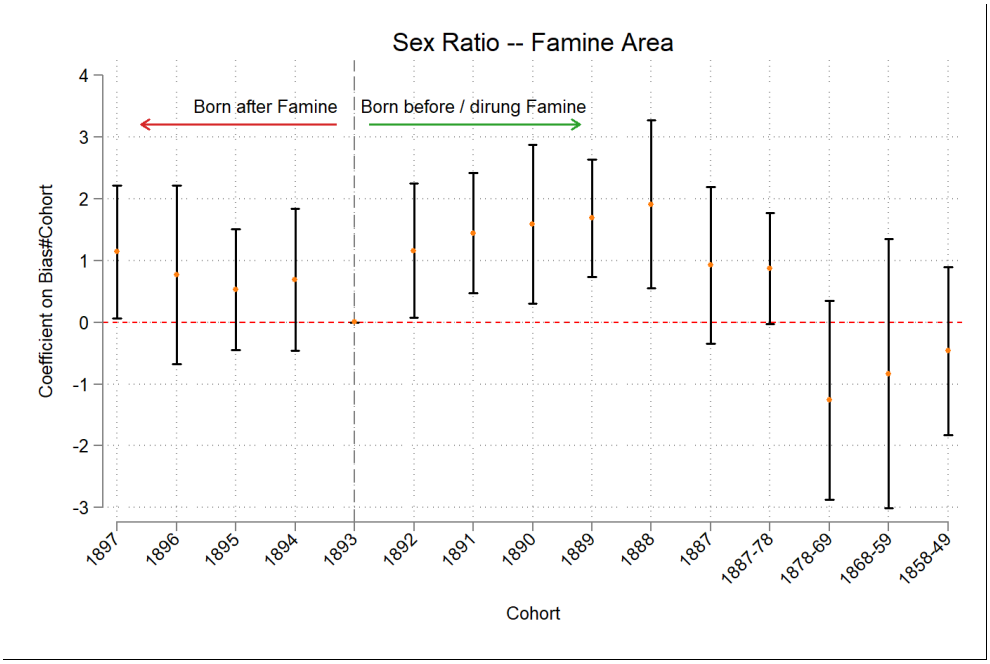
Figure 7. 1891/92 Famine and infant mortality



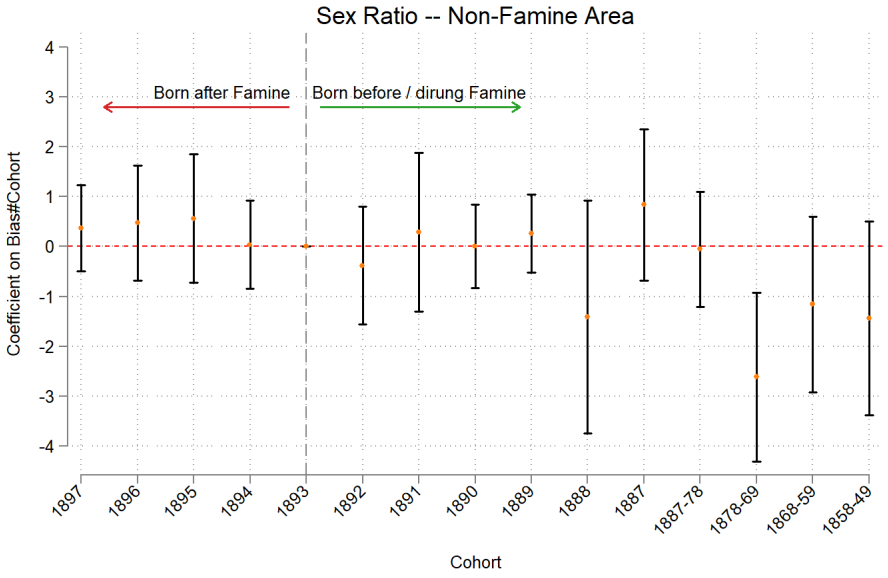
(a) (b)

Notes: This figure shows (a) unconditional and (b) conditional relationship between famine intensity and the 1891 infant mortality rate. A residual scatterplot is obtained after regressing 1891 infant mortality on famine intensity – a share of the district's population on food relief. Other controls include literacy rate, latitude, longitude, exogenous soil productivity, average temperature and precipitation (GAEZ), ruggedness, distances to the railroad, to the nearest navigable river, coastline and capitals (Moscow and Saint Petersburg), and province dummies.

Figure 8. Effect of famine and gender bias on sex ratios (by cohort)



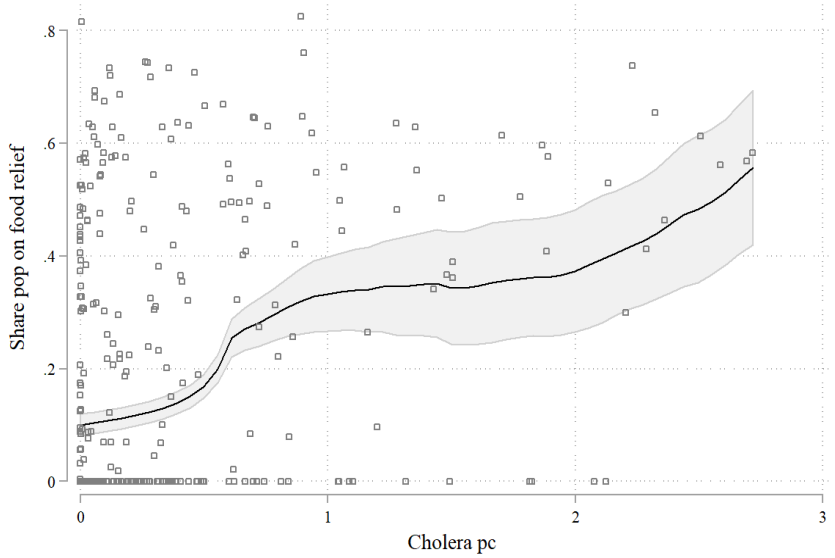
(a)



(b)

Notes: This figure plots coefficient on the interaction term between the measure of gender bias and cohort indicators. (a) plots estimates using the sample of districts with the measure of famine intensity above 75th percentile (about 1/3 of the district's population receive food relief). (b) plots estimates using the sample of districts with zero famine intensity. The coefficients are normalized to the 1893 cohort (omitted category). The regression specification includes cohort dummies, district fixed effects, cohort-by-province dummies, sex ratio at birth, cohort size and interaction of cohort dummies with average infant mortality (1887-1890, 1893-1897), latitude, longitude, exogenous soil productivity, average temperature and precipitation (GAEZ), ruggedness, distances to the railroad, to the nearest navigable river, coastline and capitals (Moscow and Saint Petersburg). Standard errors are clustered at the district level.

Figure 9. 1892 Cholera vs 1891 Famine



Notes: This figure shows the relationship (local polynomial smooth) between 1891 Famine intensity and 1892 Cholera incidence (# of cases divided by district's population).

Table 1. 1891/92 Famine effect on the cohort size

	(1)	(2)
	Dep variable: Cohort size	
Famine intensity × Year < 1892	-0.146***	
	(0.036)	
Famine intensity predicted × Year < 1892		-0.188***
		(0.045)
Outcome mean / s.d	1.005 / 0.261	1.005 / 0.261
Famine intensity mean / s.d	0.145 / 0.232	0.145 / 0.232
Cohort FE	✓	✓
District FE	✓	✓
Observations	5,006	5,006
R-squared	0.716	0.720

Notes: This table shows the effect of the 1891/92 Famine on the size of the cohort (normalized by the average size of the cohorts born in 1887-1897). We apply a simple difference-in-difference approach by regressing cohort size on the treatment indicator (equals one for the cohorts born before 1892) and famine intensity measure alongside cohort and district fixed effects: $Cohort\ size_{it} = \gamma_i + \mu_t + \beta 1(Year < 1892)_t \times Famine_{i,1891} + \varepsilon_{it}$. Standard errors are clustered at the district level. *** p<0.01, ** p<0.05, * p<0.1

Table 2. Effect of famine and gender bias on sex ratios

	(1)	(2)	(3)	(4)	(5)	(6)	(7)
	Dependent variable: M / F Ratio						
	Panel A: Actual famine exposure						
Famine intensity × Year < 1892	0.013**	0.017***	-0.047	-0.127	2.075*	1.140	-0.202*
	(0.006)	(0.006)	(0.049)	(0.085)	(1.174)	(1.764)	(0.103)
Bias × Famine × Year < 1892			0.816	1.756*	1.983***	2.962***	2.461*
			(0.635)	(1.049)	(0.597)	(1.047)	(1.334)
Bias × Year < 1892			-0.247**	-0.390	-0.122	-0.541	0.131
			(0.123)	(0.361)	(0.153)	(0.419)	(0.275)
Sex ratio at birth		0.101***	0.101***	0.116***	0.100***	0.112***	0.079***
		(0.027)	(0.027)	(0.028)	(0.027)	(0.028)	(0.028)
Cohort size		0.022*	0.025*	0.047	0.035**	0.051	0.041**
		(0.013)	(0.013)	(0.032)	(0.015)	(0.035)	(0.018)
Observations	5,006	5,006	5,006	5,006	5,006	5,006	5,006
R-squared	0.260	0.265	0.266	0.358	0.276	0.364	0.355
	Panel B: Predicted famine exposure						
Famine intensity × Year < 1892	0.013*	0.018**	-0.052	-0.130	1.760	1.475	-0.179*
	(0.007)	(0.008)	(0.047)	(0.084)	(1.209)	(2.035)	(0.092)
Bias × Famine × Year < 1892			0.914	1.843**	1.705***	2.811***	2.130*
			(0.623)	(0.932)	(0.621)	(1.069)	(1.206)
Bias × Year < 1892			-0.261**	-0.421	-0.089	-0.513	0.120
			(0.127)	(0.372)	(0.161)	(0.424)	(0.282)
Sex ratio at birth		0.101***	0.101***	0.116***	0.100***	0.112***	0.079***
		(0.027)	(0.027)	(0.028)	(0.027)	(0.028)	(0.028)
Cohort size		0.022*	0.025*	0.047	0.035**	0.051	0.041**
		(0.013)	(0.013)	(0.032)	(0.015)	(0.035)	(0.018)
Cohort FE		✓	✓	✓	✓	✓	✓
District FE		✓	✓	✓	✓	✓	✓
Province × Cohort FE				✓		✓	
Controls+					✓	✓	
District specific linear trends							✓
Observations	5,006	5,006	5,006	5,006	5,006	5,006	5,006
R-squared	0.260	0.265	0.266	0.358	0.276	0.365	0.355

Notes: This table shows estimates of the joint effect of the 1891/92 Famine and gender bias on sex ratios. In Panel A, Famine intensity is the share of the district population receiving food relief from the government. In Panel B, we apply predicted food relief as described in Section "Empirical strategy". Year < 1892 is a binary indicator switching on for the cohorts, born before 1892. Bias is the measure of gender bias against women based on folklore data. Outcome mean / s.d = 1.005 / 0.261; Famine intensity mean / s.d = 0.145 / 0.232; Bias mean / s.d. = 0.078 / 0.012. Cohort size is normalized by the average size of the cohorts born in 1887-1897. Columns (1)-(2) shows the estimate of famine exposure on relative survival chances of boys (eq.1). Controls+ include the term – Famine intensity × Year < 1892 interacted with average infant mortality (1887-1890, 1893-1897), latitude, longitude, exogenous soil productivity, average temperature and precipitation (GAEZ), ruggedness, distances to the railroad, to the nearest navigable river, coastline and capitals (Moscow and Saint Petersburg). Standard errors are clustered at the district level. *** p<0.01, ** p<0.05, * p<0.1.

Table 3. Effect of famine and gender bias on sex ratios: accounting for the impact of economic factors

	(1)	(2)	(3)	(4)	(5)	(6)	(7)
	Dependent variable: M / F Ratio						
Bias × Famine × Year < 1892	2.962***	2.947***	3.011***	2.874***	2.958***	3.033***	3.033***
	(1.047)	(1.082)	(1.048)	(1.058)	(1.044)	(1.051)	(1.095)
High-skill employment F/M ratio × Famine × Year < 1892		-0.037					-0.066
		(0.078)					(0.085)
High education F/M ratio × Famine × Year < 1892			0.032				0.040
			(0.033)				(0.037)
Industry employment × Famine × Year < 1892				0.070			0.050
				(0.140)			(0.142)
Urbanization × Famine × Year < 1892					0.018		0.017
					(0.085)		(0.082)
Share cattle producers × Famine × Year < 1892						-1.151	-1.157
						(0.943)	(0.911)
Cohort FE	✓	✓	✓	✓	✓	✓	✓
District FE	✓	✓	✓	✓	✓	✓	✓
Province × Cohort FE	✓	✓	✓	✓	✓	✓	✓
Controls+	✓	✓	✓	✓	✓	✓	✓
Observations	5,006	5,006	5,006	5,006	5,006	5,006	5,006
R-squared	0.364	0.364	0.364	0.364	0.364	0.364	0.365

Notes: This table shows estimates of the joint effect of the 1891/92 Famine and gender bias on sex ratios. Additional controls include the interaction of famine intensity and cohort dummy with the indicators of women's economic status: women employment rate in high skilled occupations relative to man, and women high school education relative to man. In addition, we include general economic indicators: industrial employment rate, urbanization rate, and share of workers involved in cattle breeding. Controls+ include the term – Famine intensity × Year < 1892 interacted with average infant mortality (1887-1890, 1893-1897), latitude, longitude, exogenous soil productivity, average temperature and precipitation (GAEZ), ruggedness, distances to the railroad, to the nearest navigable river, coastline and capitals (Moscow and Saint Petersburg). We do not report lower order interaction terms for better exposition. Standard errors are clustered at the district level. *** p<0.01, ** p<0.05, * p<0.

Table 4. Effect of famine and gender bias on sex ratios: accounting for religion

	(1)	(2)	(3)	(4)	(5)	(6)
	Dependent variable: M / F Ratio					
Folklore bias × Famine intensity × Year < 1892	2.962***	2.201*	3.184***	3.248***	2.605**	2.910***
	(1.047)	(1.136)	(1.036)	(1.062)	(1.045)	(1.066)
Famine intensity × Share Muslims × Year < 1892		0.118				
		(0.114)				
Famine intensity × Share Catholics × Year < 1892			-1.382**			
			(0.601)			
Famine intensity × Share Protestants × Year < 1892				-0.482**		
				(0.238)		
Famine intensity × Share Jewish × Year < 1892					-5.499*	
					(3.175)	
Famine intensity × Share Old-Believers × Year < 1892						0.246
						(0.289)
Cohort FE	✓	✓	✓	✓	✓	✓	✓	✓
District FE	✓	✓	✓	✓	✓	✓	✓	✓
Province × Cohort FE	✓	✓	✓	✓	✓	✓	✓	✓
Controls+	✓	✓	✓	✓	✓	✓	✓	✓
Observations	5,006	5,006	5,006	5,006	5,006	5,006
R-squared	0.364	0.364	0.365	0.365	0.364	0.364

Notes: This table shows the sensitivity of the coefficient on Bias × Famine × Year < 1892 to additional interactions of famine intensity with shares of religious minorities in population: Muslims, Catholics, Protestants and Old-Believers (subdivision within Orthodox Church). Controls+ include the term – Famine intensity × Year < 1892 interacted with average infant mortality (1887-1890, 1893-1897), latitude, longitude, exogenous soil productivity, average temperature and precipitation (GAEZ), ruggedness, distances to the railroad, to the nearest navigable river, coastline and capitals (Moscow and Saint Petersburg). We do not report lower order interaction terms for better exposition. Standard errors are clustered at the district level. *** p<0.01, ** p<0.05, * p<0.1

Table 5. Effect of famine, gender bias and cholera on sex ratios

	(1)	(2)
	Dep var: M/F Sex Ratio	
Cholera cases pc	0.001	-0.006
	(0.002)	(0.006)
Folklore Bias × Cholera cases pc		0.101
		(0.107)
Famine × Year < 1892	0.017***	1.822
	(0.006)	(1.168)
Bias × Famine × Year < 1892		1.932***
		(0.609)
Famine × Year < 1892		0.174
		(0.155)
Cohort FE	✓	✓
District FE	✓	✓
Controls+		✓
Observations	4,982	4,982
R-squared	0.274	0.284

Notes: This table shows the robustness of the main findings to adding the measure of Cholera exposure. Cholera cases pc denotes the number of cholera cases registered in each year between 1887 and 1897, normalized by 1890 population. Controls+ include the term – Famine intensity × Year < 1892 interacted with average infant mortality (1887-1890, 1893-1897), latitude, longitude, exogenous soil productivity, average temperature and precipitation (GAEZ), ruggedness, distances to the railroad, to the nearest navigable river, coastline and capitals (Moscow and Saint Petersburg). Standard errors are clustered at the district level. *** p<0.01, ** p<0.05, * p<0.1.

Table 6. Effect of famine and gender bias on sex ratios in urban areas (placebo treatment)

	(1)	(2)	(3)	(4)	(5)	(6)
	Dependent variable: M / F Ratio					
Famine intensity × Year < 1892	-0.023	-0.020	-0.138	-0.067	9.360*	0.341
	(0.027)	(0.027)	(0.302)	(0.399)	(4.822)	(9.970)
Bias × Famine × Year < 1892			1.509	0.983	0.938	-1.770
			(3.902)	(5.038)	(3.935)	(4.877)
Bias × Year < 1892			-0.499	-1.042	-1.369*	-0.485
			(0.427)	(1.077)	(0.698)	(1.156)
Cohort FE	✓	✓	✓	✓	✓	✓
District FE	✓	✓	✓	✓	✓	✓
Province × Cohort FE				✓		✓
Controls+					✓	✓
Observations	5,000	5,000	5,000	5,000	5,000	5,000
R-squared	0.104	0.105	0.105	0.204	0.116	0.215

Notes: This table replicates Table 2 for the sample of *urban areas* (cities and towns). Due to the small size of the cohorts in some towns and resulting extreme sex ratios, the values in the tails of the distribution (1th and 99th percentiles) are winsorized. Controls+ include the term – Famine intensity × Year < 1892 interacted with average infant mortality (1887-1890, 1893-1897), latitude, longitude, exogenous soil productivity, average temperature and precipitation (GAEZ), ruggedness, distances to the railroad, to the nearest navigable river, coastline and capitals (Moscow and Saint Petersburg). Standard errors are clustered at the district level. *** p<0.01, ** p<0.05, * p<0.1.

Online Appendices

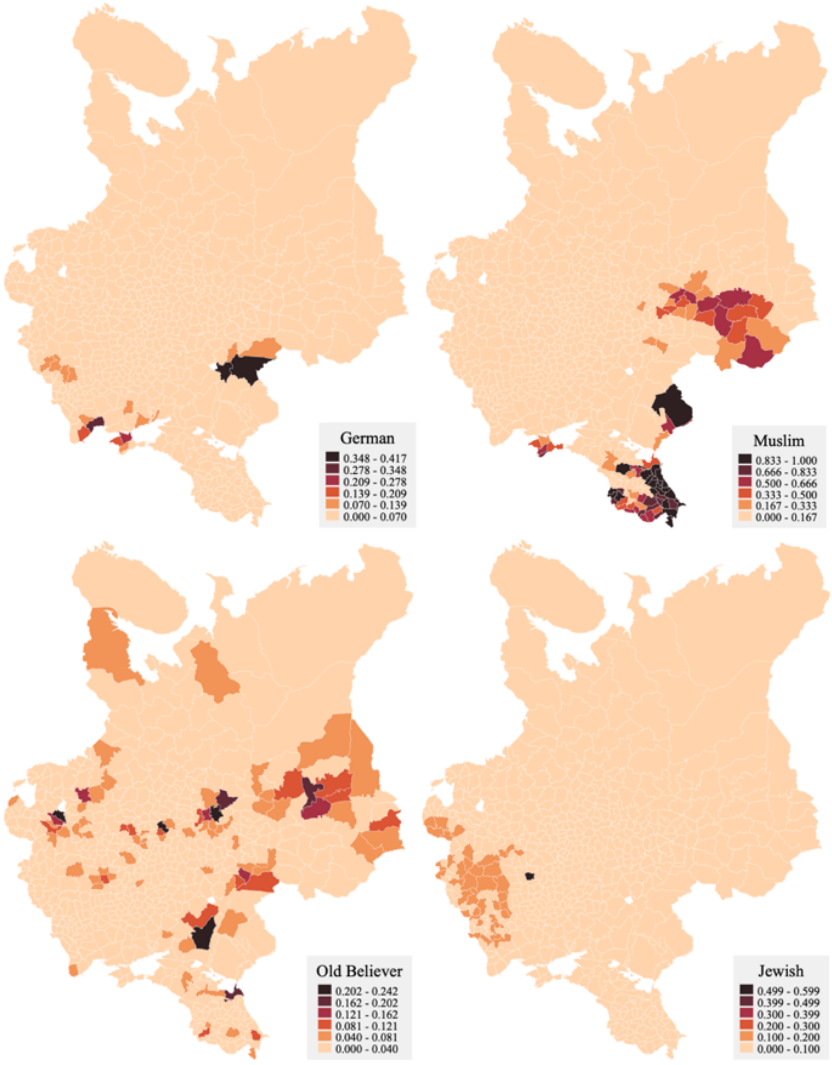
Appendix A. Additional Tables and Figures Covered in the Paper

Table A1. Descriptive Statistics

Variable	Obs	Mean	Std. Dev.	Min	Max
Sex Ratio	5,006	0.985	0.053	0.473	1.554
Sex ratio at birth	5,006	1.056	0.038	0.455	1.541
Cohort size (normalized)	5,006	1.005	0.161	0.627	1.766
Female negative bias	501	0.078	0.012	0.039	0.113
Share pop starving	501	0.145	0.232	0.000	0.825
Share pop starving, predicted	501	0.151	0.205	0.000	0.689
Mortality rate (age 0-1) 1891	501	20.79	5.845	9.899	39.37
Cholera cases per capita	498	0.372	1.332	0.000	26.42
Log railroad distance	501	3.523	1.406	-1.153	6.701
Log soil productivity index	501	8.260	0.519	0.000	8.689
Log distance to coastline	501	6.000	0.844	2.419	7.076
Log distance to Moscow	501	6.292	0.668	2.141	7.384
F / M labor force participation	501	0.154	0.054	0.043	0.460
F / M high-skilled employment	501	0.242	0.119	0.000	0.642
F / M high school education	501	0.488	0.220	0.107	2.297
Av. Precipitation	501	48.593	5.930	16.76	65.39
Av. Temperature	501	5.30	2.130	-3.960	11.63
Log distance to river	501	4.20	1.064	0.089	6.576
Ruggedness	501	29.05	19.45	7.224	288.3

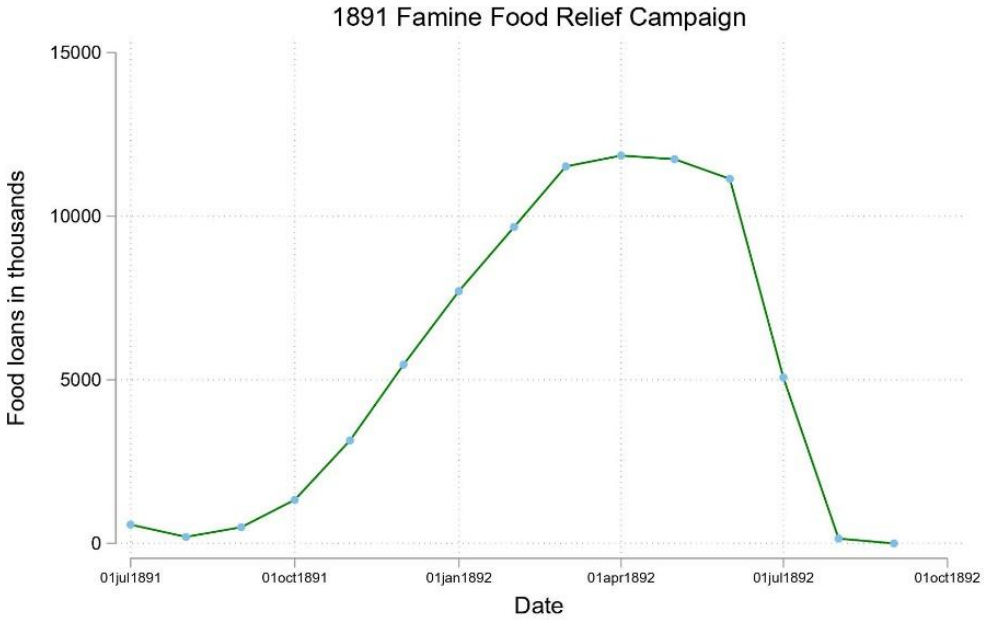
Notes: Controls that vary by cohort and district include sex ratio, sex ratio at birth, cohort size, and cholera cases per capita. Other reported controls vary at the district level.

Figure A1. Ethnic and religious groups in 1897.



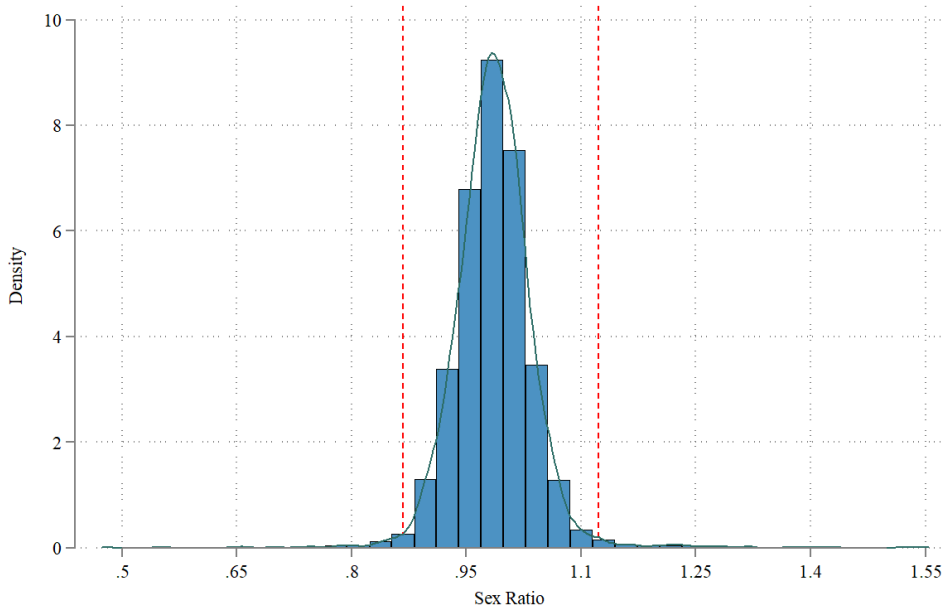
Notes: This figure shows the spatial distribution of the ethnic and religious minorities in European Russia and the Caucasus based on the 1897 Census. Germans almost exclusively represented the Catholic and Protestant populations within the Volga basin.

Figure A2. Disbursement of state-sponsored food relief campaign over time



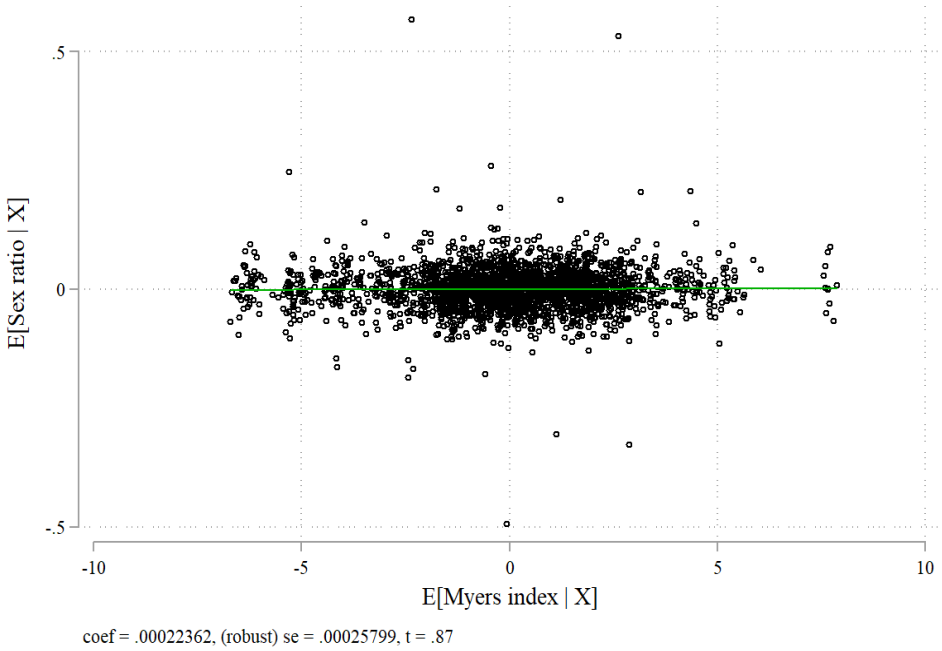
Sources: Statisticheskie dannye po vydahe ssud na obsemenenie I prodovolstvie naseleniyu postraadavshemu ot neurozhaya v 1891-1892 gg. Reported by Whitecroft

Figure A3. Distribution of sex ratios



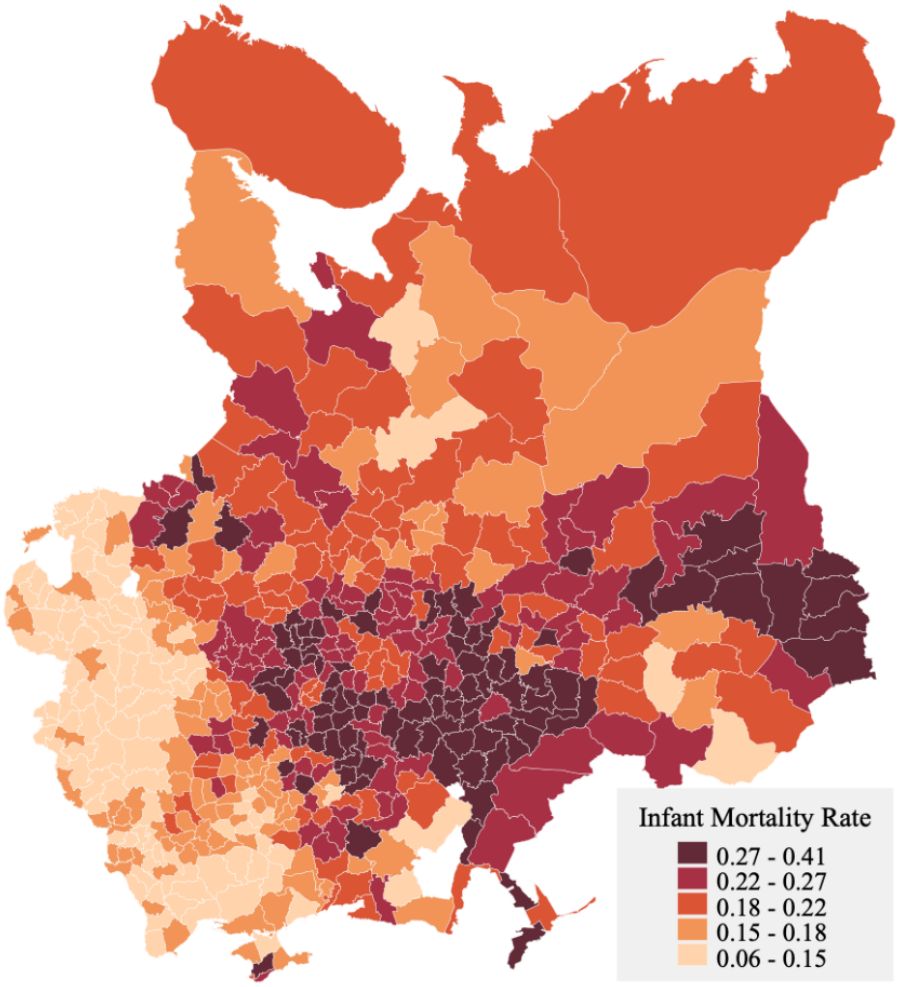
Sources: Census 1897. *Notes:* This figure shows the distribution of sex ratios – the ratio between boys and girls among cohorts born between 1887 and 1897 (excluding 1892). Red lines indicate the 1st and 99th percentiles of the sample distribution.

Figure A4. Age heaping (ages 15-74) and sex ratios



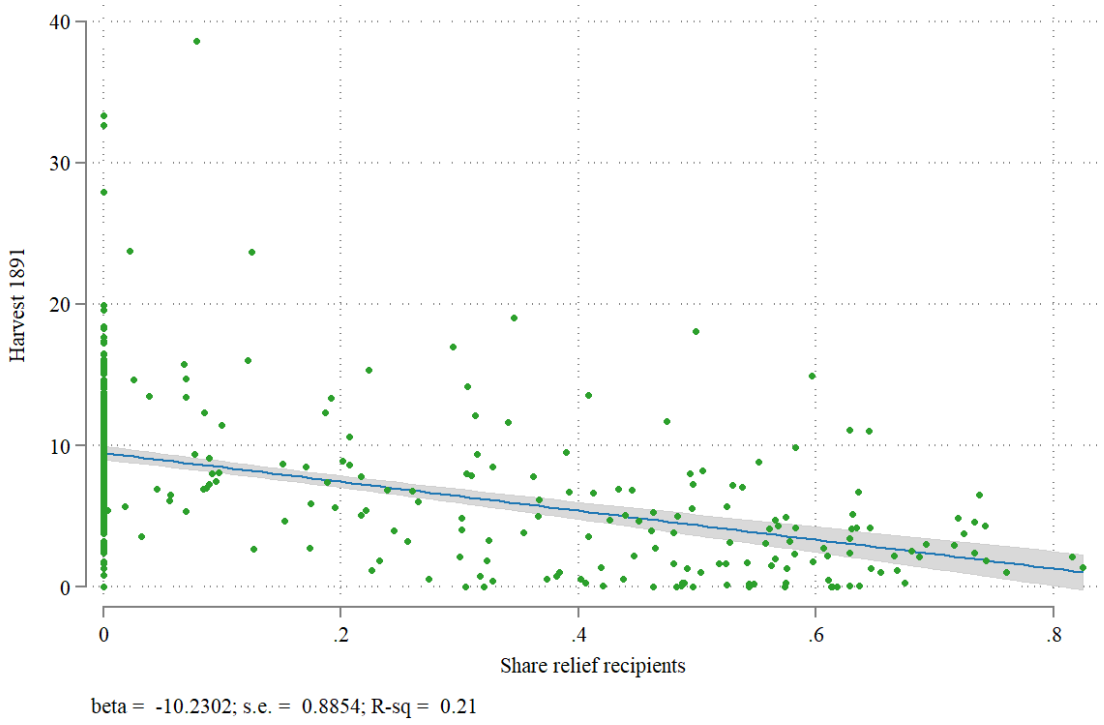
Notes: This figure shows a conditional scatterplot obtained after regressing the sex ratio on the measure of age-hyping – Myers index for ages 15-74 (The index is constructed by Charnysh 2022 using Census 1897).

Figure A5. Infant mortality rates, 1891



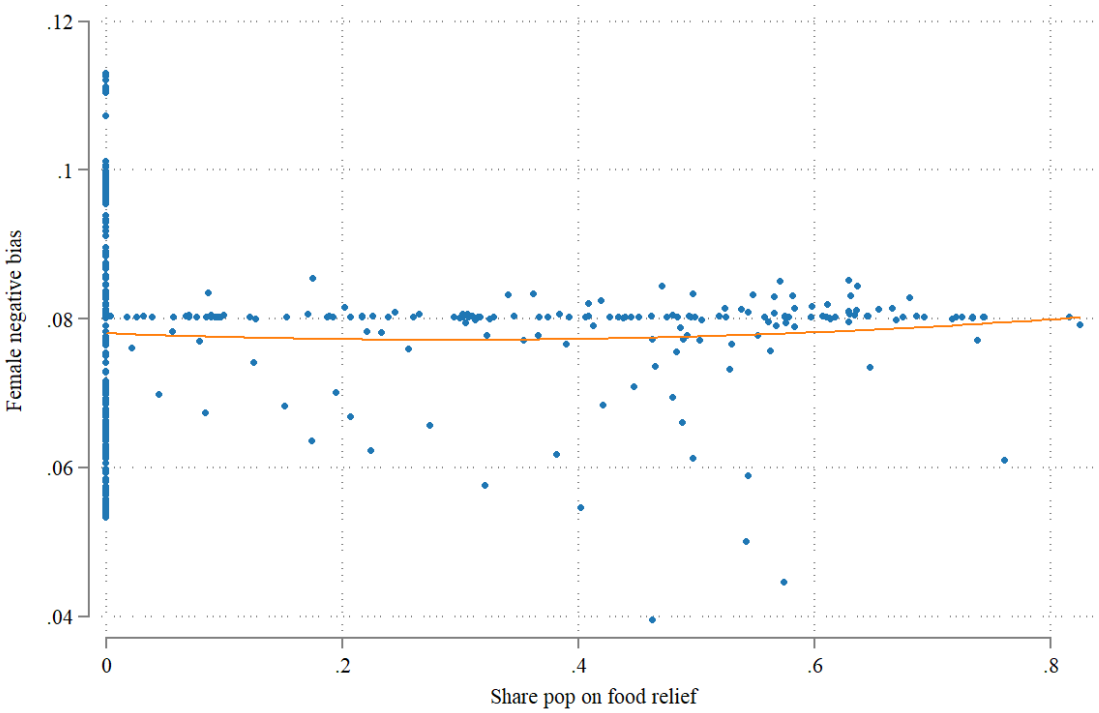
Notes: This figure shows the spatial distribution of the 1891 infant mortality rate across districts in the European part of the Russian Empire.

Figure A6. Harvest per capita, 1891 and share of food relief recipients.



Notes: This figure shows an unconditional relationship between the 1891 grain harvest and famine intensity – the share of the district's population receiving food relief from the government.

Figure A7. Female bias vs. share of food relief recipients across districts



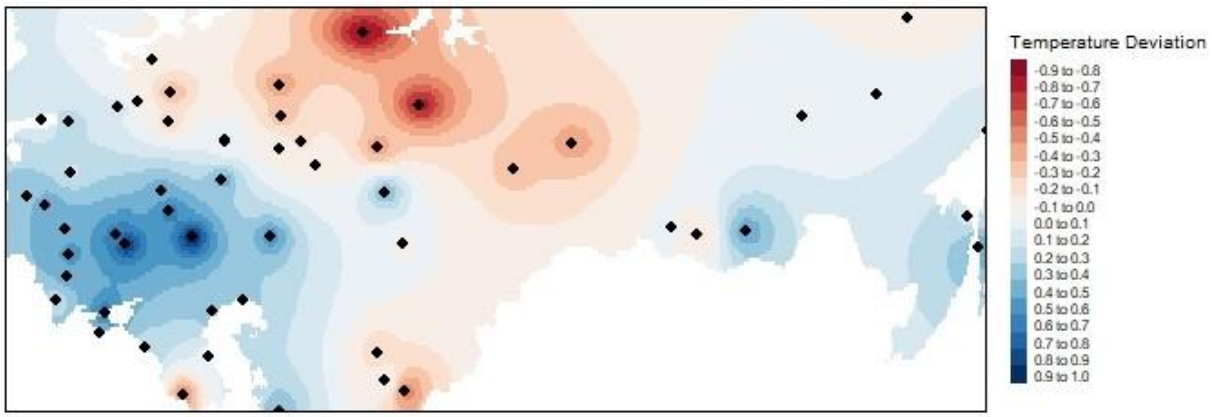
Notes: This figure shows an unconditional relationship between Female negative bias and the share of the district population receiving food relief from the government.

Table A2. Climate shock as predictor of famine food relief

	OLS coefficient
January precipitation	-12.526
	(11.368)
February precipitation	2.799
	(18.218)
March precipitation	**-23.326*** **
	(8.003)
April precipitation	3.187
	(7.664)
May precipitation	5.575
	(9.065)
June precipitation	-9.279
	(12.339)
July precipitation	-1.011
	(7.682)
January temperature	-0.990
	(11.997)
February temperature	-9.318
	(6.820)
March temperature	-18.470
	(12.002)
April temperature	21.632
	(15.721)
May temperature	-7.029
	(17.802)
June temperature	1.922
	(12.211)
July temperature	**42.570*** **
	(11.396)
Soil productivity	-2.133
	(1.411)
Observations	576
R-squared	0.834

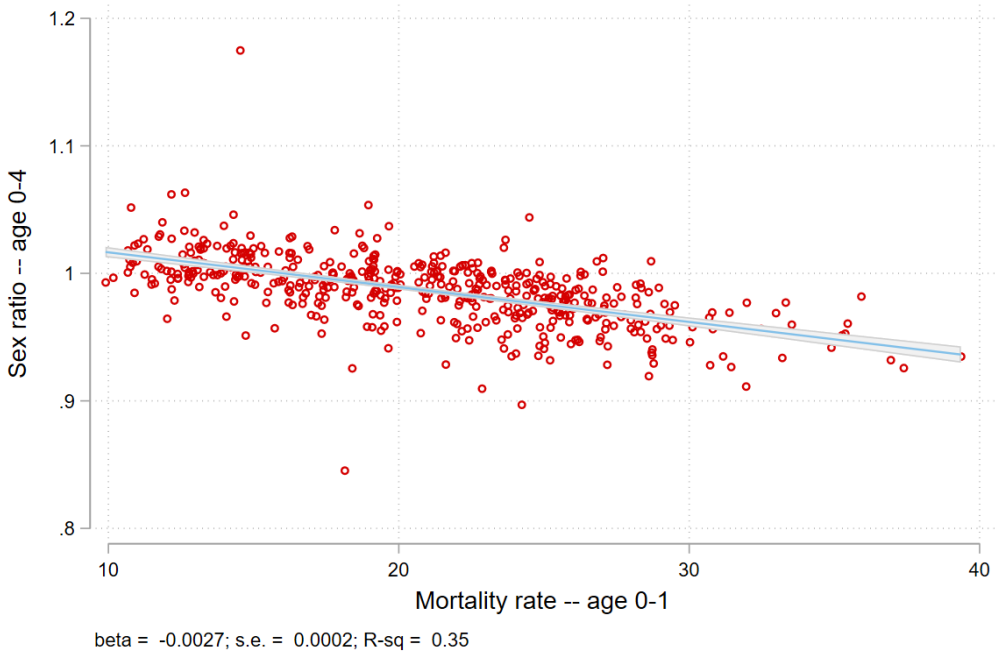
Notes: This table shows estimates of the coefficient from the linear regression model based on climate observations. The monthly temperature and precipitation shocks are standardized deviations of 1891 levels from long-term mean values. Other controls include province dummies.

Figure A8. Deviation of temperature levels in July 1891 from that month's historical mean values



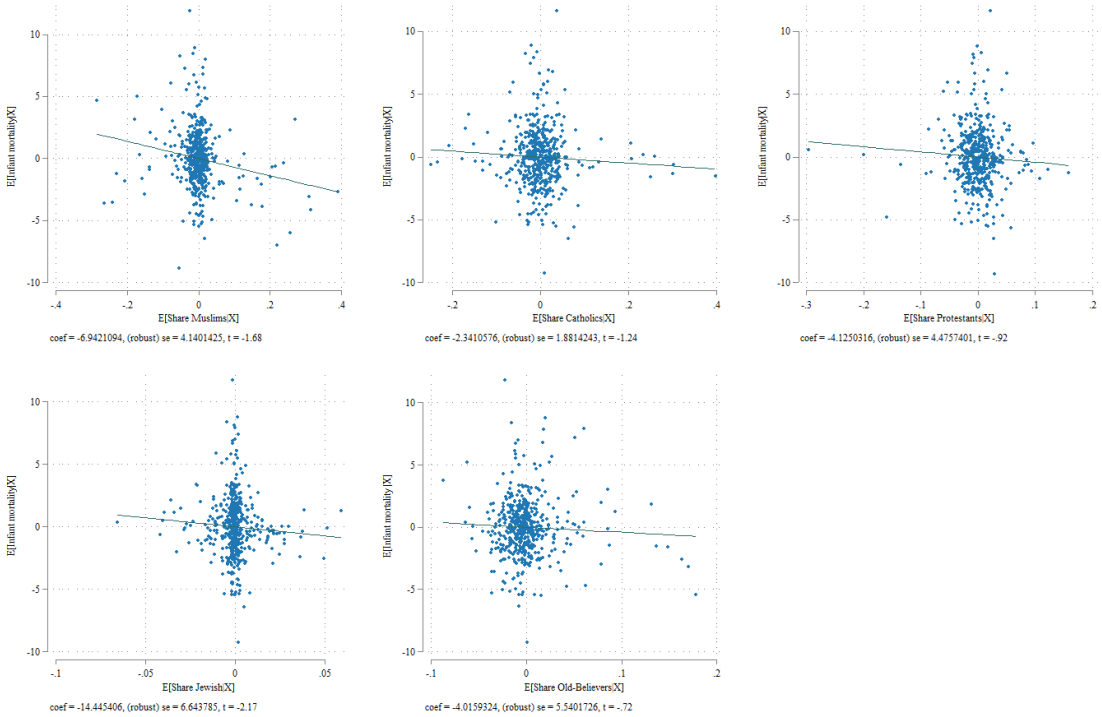
Notes: This figure shows the spatial distribution of climate shock (temperature). Dots denote the location of climate stations.

Figure A9. Relationship between infant mortality and sex ratios



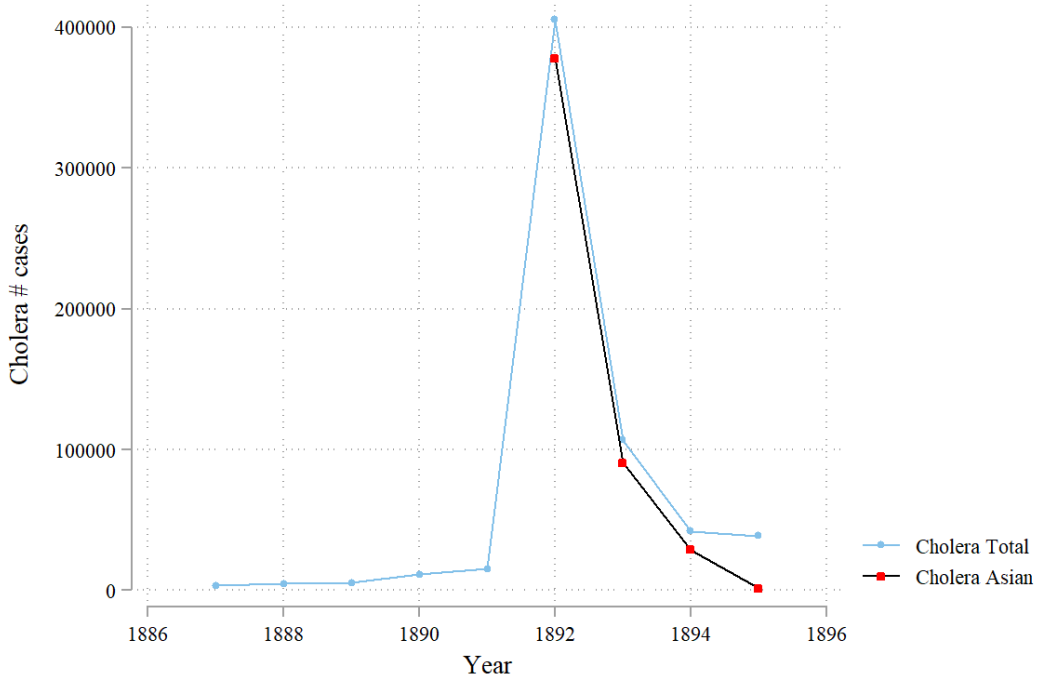
Notes: This figure shows an unconditional relationship between sex ratios at age 0-4 (Source: Census 1897) and average infant mortality rate for the period between 1887 and 1897 excluding famine years (Source: Vital Statistics).

Figure A10. Religious minorities and infant mortality



Notes: This figure plots conditional scatterplots of average infant mortality in 1887-1897 (excluding famine years) vs. shares of religious minorities in the population. Other controls include birth ratio, literacy rate, latitude, longitude, exogenous soil productivity, average temperature and precipitation (GAEZ), ruggedness, distances to the railroad, to a nearest navigable river, coastline and capitals, and province dummies.

Figure A11. Cholera incidence in 1887-1895



Notes: This figure shows the incidence of Cholera over time in the Russian Empire. Source: Medical Department Report.

Table A3. Different groups of motifs as predictors of women's economic status

	(1)	(2)	(3)	(4)	(5)	(6)
	Panel A: F/M Employment ratio in high-skilled professions					
Submissive	-0.820***	-0.674**	-0.940***	-0.847***	-0.774***	-0.747***
	(0.204)	(0.281)	(0.219)	(0.207)	(0.209)	(0.213)
Domestic affairs		-0.825				
		(1.193)				
Violent			-0.484			
			(0.336)			
Physically active				0.553		
				(0.508)		
Sexual					1.508	
					(1.251)	
Other						-0.541*
						(0.327)
Observations	515	515	515	515	515	515
R-squared	0.595	0.596	0.597	0.596	0.596	0.598
	Panel B: F/M Employment ratio (all occupations)					
Submissive	-2.920***	-5.925***	-2.972***	-2.870***	-3.063***	-2.908***
	(0.493)	(0.529)	(0.540)	(0.511)	(0.519)	(0.501)
Domestic affairs		16.960***				
		(2.050)				
Violent			-0.212			
			(0.777)			
Physically active				-1.009		
				(1.770)		
Sexual					-4.678	
					(3.753)	
Other						-0.084
						(0.847)
Observations	515	515	515	515	515	515
	0.349	0.417	0.349	0.350	0.352	0.349
R-squared						

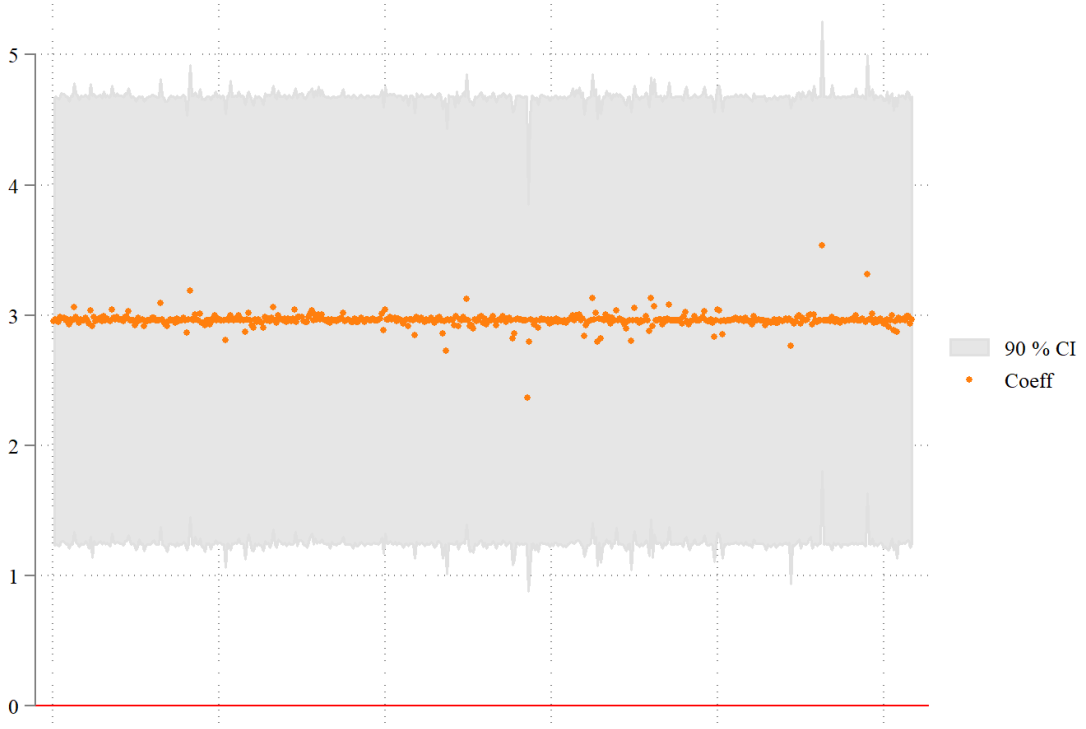
Note: This table reports regression estimates on the relationship between different groups of motifs (folklore) and the measures of woman economic status. Other controls include literacy rate, latitude, longitude, exogenous soil productivity, average temperature and precipitation (GAEZ), ruggedness, distances to the railroad, to the nearest navigable river, coastline and capitals (Moscow and Saint Petersburg).

Table A4. Alternative Standard Errors and Robustness to Outliers

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
	Panel A: Actual Famine exposure. Dep Var M/F Ratio:							
Bias × Famine × Year < 1892	2.962***	2.962***	2.962***	2.962***	1.831**	2.218***	2.962***	2.962***
	(1.047)	(0.953)	(0.979)	(1.066)	(0.735)	(0.818)	(1.047)	(1.047)
	Baseline	Conley Adj Spatial cut-off 500 km	Conley Adj Spatial cut-off 1000 km	Conley Adj Spatial cut-off 1500 km	Omit Cook's Distance outliers	Winsorize Sex ratio 1st/99th	Winsorize Bias 1st/99th	Winsorize Famine 1st/99th
Observations	5,006	5,006	5,006	5,006	4,765	5,006	5,006	5,006
R-squared	0.364	0.364	0.364	0.364	0.548	0.427	0.364	0.364

Notes: This table tests the robustness of the main findings to alternative standard errors (Conley Adjustment) and potential outliers. Regression specification corresponds to Table 2, column 6 (panel A). We do not report lower order interaction terms for better exposition. *** p<0.01, ** p<0.05, * p<0.1.

Figure A12. Robustness to outliers



Notes: This figure plots the distribution of the coefficients on *Bias* × *Famine intensity* × *Year* < *1892* after excluding one of the districts from the sample. Regression specification corresponds to Table 2, column 6 (panel A).

Table A5**.** Alternative outcome variable

	(1)	(2)	(3)	(4)	(5)	(6)
	Dependent variable: % Boys in cohort					
Famine intensity × Year < 1892	0.295**	0.405***	-1.052	-2.568	44.749	20.705
	(0.149)	(0.156)	(1.214)	(2.084)	(29.319)	(44.687)
Bias × Famine × Year < 1892			18.670	36.125	48.239***	67.812***
			(15.774)	(25.740)	(14.907)	(25.935)
Bias × Year < 1892			-5.939**	-6.465	-2.744	-10.190
			(3.008)	(8.980)	(3.796)	(10.425)
Cohort FE	✓	✓	✓	✓	✓	✓
District FE	✓	✓	✓	✓	✓	✓
Province × Cohort FE			✓		✓	✓
Controls+				✓	✓	✓
Observations	5,006	5,006	5,006	5,006	5,006	5,006
R-squared	0.260	0.266	0.266	0.355	0.276	0.361

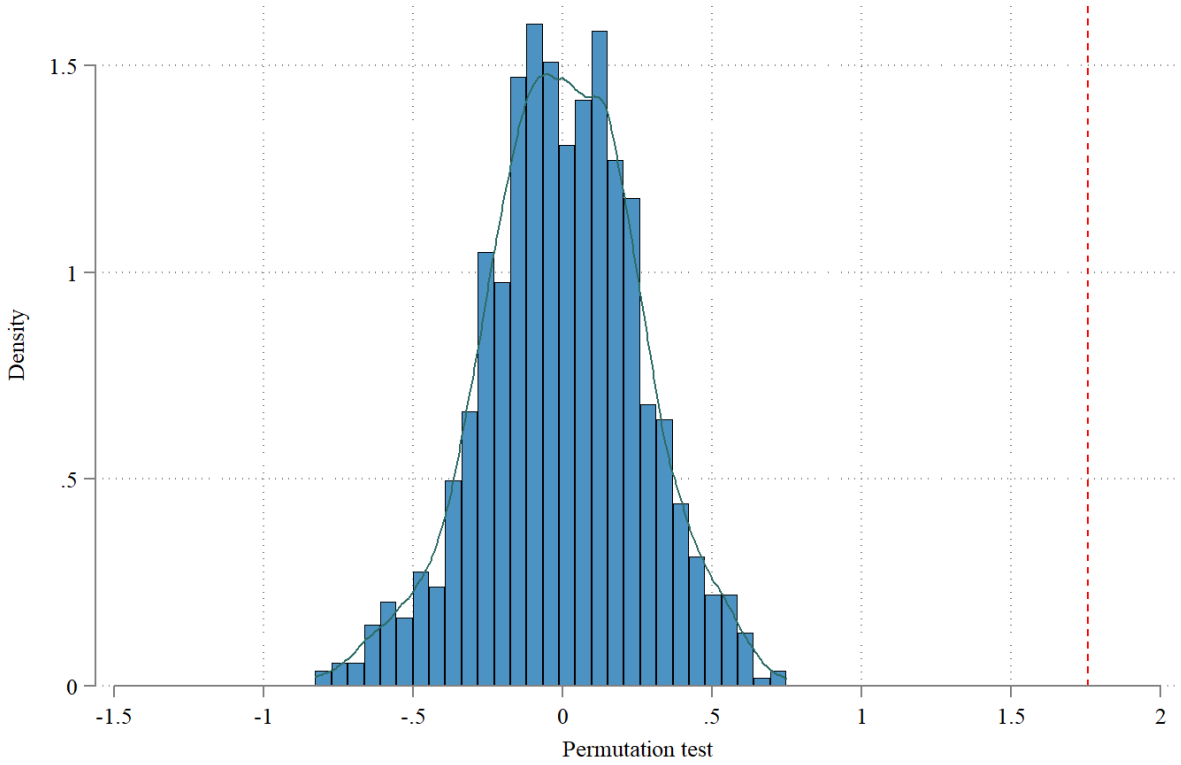
Notes: This table replicates the findings of Table 2 (panel A) with an alternative outcome variable – % of boys per cohort born in 1887-1897. Controls+ include the term – Famine intensity × Year < 1892 interacted with average infant mortality (1887-1890, 1893-1897), latitude, longitude, exogenous soil productivity, average temperature and precipitation (GAEZ), ruggedness, distances to the railroad, to the nearest navigable river, coastline and capitals (Moscow and Saint Petersburg).

Table A6. Effect of famine and gender bias on sex ratios: by group of motifs

	(1)	(2)	(3)	(4)	(5)	(6)
	Dependent variable: M / F Ratio					
Submissive × Famine intensity × Year < 1892	2.962***					
	(1.047)					
Domestic affairs × Famine intensity × Year < 1892		3.257				
		(4.304)				
Violent × Famine intensity × Year < 1892			-2.834			
			(1.824)			
Physically active × Famine intensity × Year < 1892				8.414		
				(5.325)		
Sexual × Famine intensity × Year < 1892					17.836***	
					(6.420)	
Other × Famine intensity × Year < 1892						3.172
						(2.350)
Cohort FE	✓	✓	✓	✓	✓	✓
District FE	✓	✓	✓	✓	✓	✓
Province × Cohort FE	✓	✓	✓	✓	✓	✓
Controls+	✓	✓	✓	✓	✓	✓
Observations	5,006	5,006	5,006	5,006	5,006	5,006
R-squared	0.364	0.363	0.363	0.364	0.364	0.363

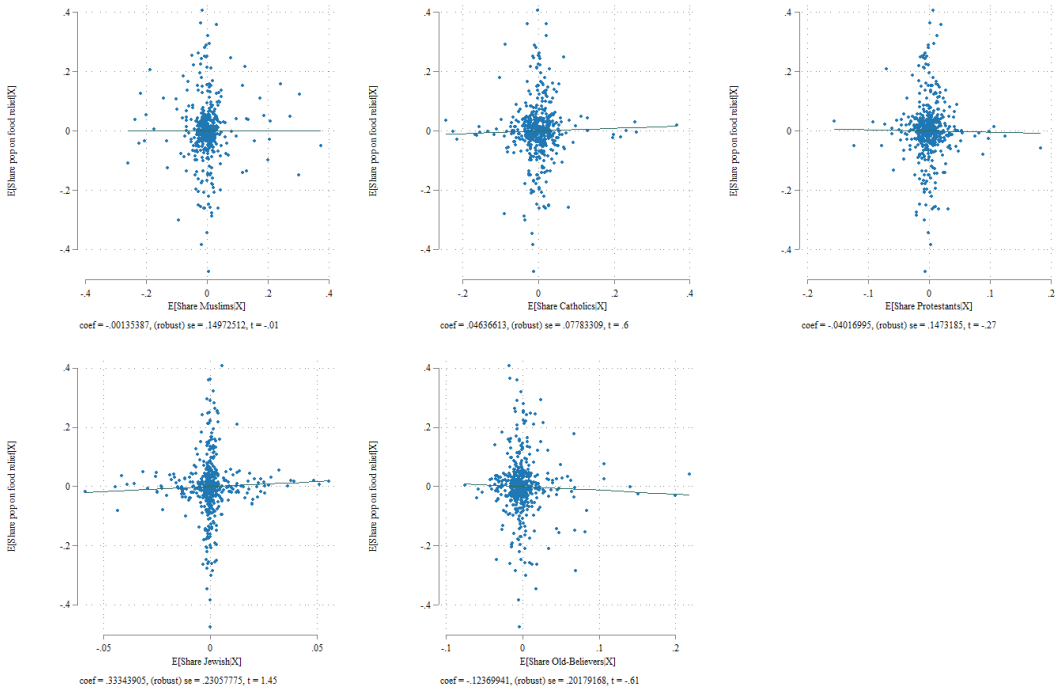
Notes: This table demonstrates the interaction effect between famine and different groups of motifs depicting women as 1) submissive and stupid, 2) involved in domestic affairs, 3) violent, 4) physically active, 5) sexual, and 6) other motifs. Controls+ include the term – Famine intensity × Year < 1892 interacted with average infant mortality (1887-1890, 1893-1897), latitude, longitude, exogenous soil productivity, average temperature and precipitation (GAEZ), ruggedness, distances to the railroad, to the nearest navigable river, coastline and capitals (Moscow and Saint Petersburg). *** p<0.01, ** p<0.05, * p<0.1.

Figure A13. Permutation test



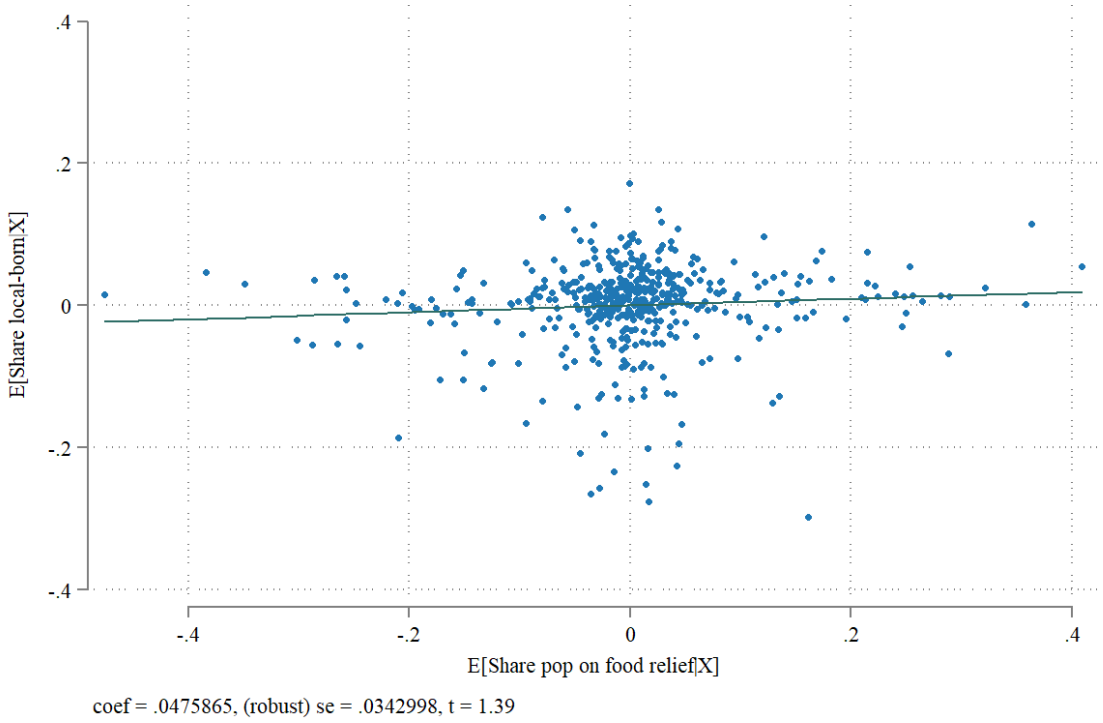
Notes: This figure shows the distribution of the estimated coefficients on interaction *Bias*Famine*Year < 1892* from the regressions with false treatment (1000 iterations). The distribution of gender motifs is randomly reshuffled across ethnic groups in the original Folklore dataset. We then link it to the 1897 Census to compute a population-weighted average of the bias measure at the district level. The test indicates that the randomly generated predictors does not deliver the same or stronger magnitude estimate as the true model. Regression specification corresponds to Table 2, column 4 (Panel A).

Figure A14. Famine intensity and shares of main religious minorities in the population



Notes: This figure plots conditional scatterplots of famine intensity vs. shares of religious minorities in the population. Other controls include 1891 infant mortality rate, birth ratio, literacy rate, latitude, longitude, exogenous soil productivity, average temperature and precipitation (GAEZ), ruggedness, distances to the railroad, to a nearest navigable river, coastline and capitals, and province dummies.

Figure A15. Famine intensity and Share of local residents in the population in 1897



Notes: This figure plots the share of population receiving a food assistance from the government in 1891-1892 vs. the share of local-born residents in the district population by 1897. The controls include 1891 infant mortality rate, birth ratio, literacy rate, latitude, longitude, exogenous soil productivity, average temperature and precipitation (GAEZ), ruggedness, distances to the railroad, to a nearest navigable river, coastline and capitals, and province dummies.

Appendix B. Binary Treatment & Matching

To provide additional support for our findings, we apply an alternative estimation technique based on binary treatment and a weighting scheme that allows us to balance the average values of observable characteristics between the treated and control groups, thereby mimicking the conditions of the experiment. We assign districts to the treated or control group based on whether the folklore-based gender bias measure is above or below the 75th percentile. Table B1 presents the mean values of control variables for these two groups (columns 1 and 2). To ensure covariate balance between the treated and control units, we apply the entropy balancing approach proposed by Hainmueller (2012) (column 3).

This approach involves solving the optimization problem to calculate the weights:

$$\sum_{T_i=0} w_i V_i = \sum_{T_i=1} V_i$$

Here, w_i represents the observational weight, and V_i denotes the value of a specific variable for observation i. The weights are determined in such a way as to eliminate the difference in the mean values of the variables between the treated and control groups.

Table B1. Covariate balance test before and after weighting of the observations

Variable	Average treated	Average Control	Average Control Weighted
Av infant mortality rate	18.713	21.568	18.715
Log railroad dist.	37.385	34.425	37.380
Soil productivity	81.405	83.045	81.406
Log coast dist.	57.139	61.063	57.142
Log dist. Moscow	65.087	62.104	65.084
Log dist. Saint-Petersburg	64.214	68.358	64.215
Longitude	34.235	38.198	34.236
Latitude	55.800	53.401	55.798
Av. Precipitation	50.914	47.720	50.913
Av. Temperature	48.674	54.592	48.677
Log river dist.	42.617	41.766	42.616
Ruggedness	22.905	31.368	22.909

We then apply the obtained weights in our regression framework with binary treatment indicator (see Table B2).

Table B2. Effect of famine and gender bias on sex ratios (binary treatment)

	(1)	(2)	(3)	(4)	(5)	(6)
	Dependent variable: M / F Ratio					
Famine intensity > 75p × Year < 1892	0.005	0.007**	0.003	-0.003	-0.006	-0.008
	(0.003)	(0.003)	(0.004)	(0.011)	(0.006)	(0.008)
			0.017**	0.022*	0.017*	0.020
Bias > 75p × Famine > 75p × Year < 1892						
			(0.008)	(0.013)	(0.009)	(0.016)
Bias > 75p × Year < 1892			-0.009**	-0.007	-0.006	0.001
			(0.003)	(0.008)	(0.006)	(0.008)
Sex ratio at birth		0.100***	0.102***	-0.004	0.117***	0.014
		(0.027)	(0.027)	(0.049)	(0.028)	(0.038)
Cohort size		0.022*	0.025*	0.033	0.046	0.057
		(0.013)	(0.013)	(0.025)	(0.032)	(0.060)
Cohort FE	✓	✓	✓	✓	✓	✓
District FE	✓	✓	✓	✓	✓	✓
Province × Cohort FE					✓	✓
Weights				✓		✓
Observations	5,006	5,006	5,006	5,006	5,006	5,006
R-squared	0.260	0.264	0.266	0.221	0.358	0.423

Notes: This table tests the robustness of the main findings to alternative definition of treatment. We substitute continuous measures of famine intensity and gender bias with binary indicators. Famine intensity > 75p defines districts with famine intensity above 75th percentile. Bias > 75p denotes districts with bias above 75th percentile. Columns (4) and (6) report estimates of weighted regressions. To maintain a balanced distribution of covariates across districts with low vs. high gender bias (see Figure 5) we reweight observations in the dataset following the approach by Hainmueller 2012 (STATA ebalance command). Then we apply these weights in our main regression analysis in columns (4) and (6). *** p<0.01, ** p<0.05, * p<0.1.

The estimation results of the model with binary treatment and weights generally support the baseline findings. The coefficient on triple interaction term appears with expected sign is statistically significant in most of the specifications.

Appendix C. Male-stopping rule

The male stopping rule implies that some families will continue fertility until they have a son. In the simplest example, consider two households, with and without a male stopping rule. Assume that the household with the male stopping rule has no other consideration or constraint other than having one son. The household without the male stopping rule aims for two children. The expected household size in places with and without the male stopping rule will be two[54], but note that girls will always live in bigger households. In addition, note that the variance of household size will always be larger among the households with the male stopping rule. In cases with higher "targeted" fertility, aiming for a certain number of sons will always result in higher variance.

Note that the male stopping rule can only exist in a context where parents can exert some control over their fertility. While far from perfect, there is some evidence to suggest women had a rudimentary knowledge of fertility control through the calendar method and prolonging breastfeeding (Gilyarovsky, 1866; also see Jayachandran and Kuziemko, 2011 on how women in India shorten breastfeeding for their daughters to achieve next pregnancy, resulting in adverse developmental outcomes for girls).

We hypothesize that places with more negative gender norms would be more likely to have higher variance in household size. Larger households would suffer higher mortality shocks from famine, and daughters would be more likely to live in larger households. This would then result in the main finding of our paper – sex ratios increase after a famine in places with more negative gender norms. We digitize data on household sizes by district in the European part of the Russian Empire and correlate the household size mean and variance with the negative gender bias.

Figure C1.A shows, first the relationship between negative female bias and mean household size. As expected, we find that there is no relationship between these, as male stopping rule would not necessarily induce higher household sizes. The mean household size of 6 (in other words, if we assume a household usually consists of two parents and children, would imply having, on average, 4 children) is consistent with an average target of 4 children and with the male stopping rule at 2 sons.[55] However, it is important to note here that the household sizes 6-10 were binned together in the data, which we coded as 8, so we are missing variation in the 6-10 family size group.

Figure C1.B shows that there is a strong correlation between negative female bias and the *variance* in household size, suggesting that the male-stopping rule could have existed in this society, particularly in places with more negative bias. That is, places with high negative gender bias would be more likely to stop early if the target number of sons was reached early, resulting in smaller households. They would also be more likely to continue fertility until this number was achieved, even if it meant larger families.

[54] This follows from calculating the mean of the geometric distribution, with a success rate of 0.5.
[55] This follows from calculating the mean of the negative binomial distribution, with the target of 2 successes, and success rate of 0.5.

Figure C1. Female negative bias and household size mean and variance

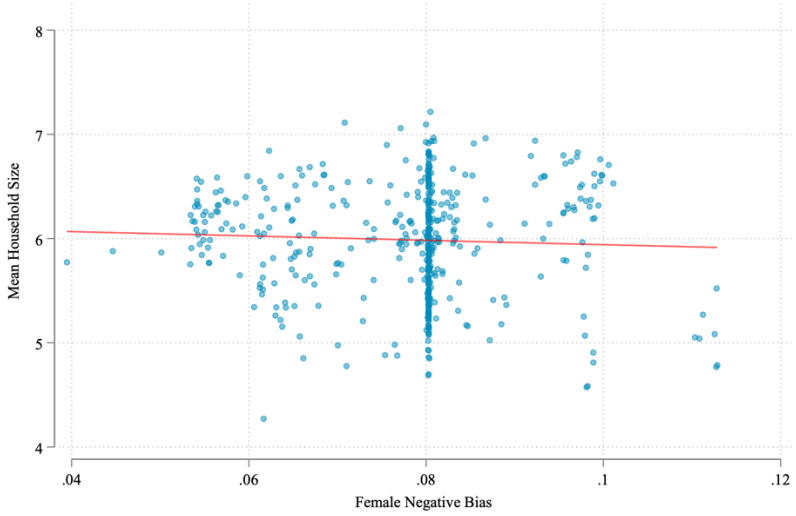


Figure C1.A. Female Negative Bias and Mean Household Size

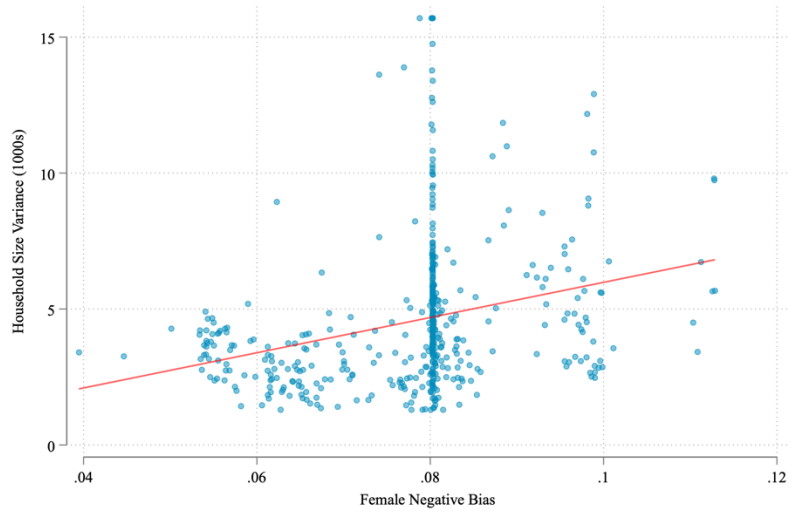


Figure C1.B. Female Negative Bias and Household Size Variance

Notes: Source: First Universal Census of the Population of the Russian Empire, 1897. Publications from various years. Tables II. Household size variance was calculated by weighting each household size by the number of households of a given size. Note that households larger than 5 were binned into sizes 6-10, and 11+. We assign size 8 to the former, and 12 to the latter.

Appendix D. Using Climate Shock for Prediction of Food Relief

The following tables replicates the main findings of the paper using predicted food supply as a measure of Famine intensity

Table D1. Effect of famine and gender bias on sex ratios: accounting for the impact of economic factors

	(1)	(2)	(3)	(4)	(5)	(6)	(7)
	Dependent variable: M / F Ratio						
Bias × Famine (pred) × Year < 1892	2.811***	2.877***	2.802***	2.831**	2.811***	2.833***	2.985***
	(1.069)	(1.100)	(1.067)	(1.119)	(1.064)	(1.067)	(1.140)
High-skill employment F/M ratio × Famine (pred) × Year < 1892		-0.081					-0.116
		(0.104)					(0.112)
High education F/M ratio × Famine (pred) × Year < 1892			0.042				0.060
			(0.056)				(0.058)
Industry employment × Famine (pred) × Year < 1892				-0.008			-0.034
				(0.131)			(0.133)
Urbanization × Famine (pred) × Year < 1892					0.030		0.038
					(0.082)		(0.082)
Share cattle producers × Famine (pred) × Year < 1892						0.090	0.111
						(0.126)	(0.145)
Cohort FE	✓	✓	✓	✓	✓	✓	✓
District FE	✓	✓	✓	✓	✓	✓	✓
Province × Cohort FE	✓	✓	✓	✓	✓	✓	✓
Controls+	✓	✓	✓	✓	✓	✓	✓
Observations	5,006	5,006	5,006	5,006	5,006	5,006	5,006
R-squared	0.365	0.365	0.365	0.365	0.365	0.365	0.366

Notes: This table shows estimates of the joint effect of the 1891/92 Famine and gender bias on sex ratios. Additional controls include the interaction of famine intensity (predicted values) and cohort dummy with the indicators of women's economic status: women employment rate in high skilled occupations relative to man, and women high school education relative to man. In addition, we include general economic indicators: industrial employment rate, urbanization rate, and share of workers involved in cattle breeding. Controls+ include the term – Famine intensity × Year < 1892 interacted with average infant mortality (1887-1890, 1893-1897), latitude, longitude, exogenous soil productivity, average temperature and precipitation (GAEZ), ruggedness, distances to the railroad, to the nearest navigable river, coastline and capitals (Moscow and Saint Petersburg). We do not report lower order interaction terms for better exposition. Standard errors are clustered at the district level. *** p<0.01, ** p<0.05, * p<0.

Table D2. Effect of famine and gender bias on sex ratios: accounting for religion

	(1)	(2)	(3)	(4)	(5)	(6)
	Dependent variable: M / F Ratio					
Folklore bias × Famine (pred) × Year < 1892	2.811***	2.028*	3.343***	3.443***	2.640**	2.721**
	(1.069)	(1.227)	(1.128)	(1.140)	(1.083)	(1.084)
Famine (pred) × Share Muslims × Year < 1892		0.113				
		(0.136)				
Famine (pred) × Share Catholics × Year < 1892			-0.984**			
			(0.467)			
Famine (pred) × Share Protestants × Year < 1892				-0.512**		
				(0.218)		
Famine (pred) × Share Jewish × Year < 1892					-1.136	
					(1.680)	
Famine (pred) × Share Old-Believers × Year < 1892						0.290
						(0.262)
Cohort FE		✓	✓	✓	✓	✓	✓	✓
District FE		✓	✓	✓	✓	✓	✓	✓
Province × Cohort FE		✓	✓	✓	✓	✓	✓	✓
Controls+		✓	✓	✓	✓	✓	✓	✓
Observations	5,006	5,006	5,006	5,006	5,006	5,006
R-squared	0.365	0.365	0.366	0.366	0.365	0.365

Notes: This table shows the sensitivity of the coefficient on Bias × Famine × Year < 1892 to additional interactions of famine intensity with shares of religious minorities in population: Muslims, Catholics, Protestants and Old-Believers (subdivision within Orthodox Church). Famine intensity is the predicted share of the district population receiving food relief from the government. Controls+ include the term – Famine intensity × Year < 1892 interacted with average infant mortality (1887-1890, 1893-1897), latitude, longitude, exogenous soil productivity, average temperature and precipitation (GAEZ), ruggedness, distances to the railroad, to the nearest navigable river, coastline and capitals (Moscow and Saint Petersburg). We do not report lower order interaction terms for better exposition. Standard errors are clustered at the district level. *** p<0.01, ** p<0.05, * p<0.1

Table D3. Effect of famine and gender bias on sex ratios in urban areas (placebo treatment)

	(1)	(2)	(3)	(4)	(5)	(6)
	Dependent variable: M / F Ratio					
Famine (pred) × Year < 1892	-0.038	-0.034	-0.085	-0.333	6.943	-5.331
	(0.032)	(0.032)	(0.295)	(0.433)	(10.612)	(15.580)
Bias × Famine (pred) × Year < 1892			0.652	1.233	1.233	-1.422
			(3.904)	(4.845)	(3.748)	(4.819)
Bias × Year < 1892			-0.468	-1.125	-1.328*	-0.840
			(0.441)	(1.104)	(0.708)	(1.224)
Cohort FE	✓	✓	✓	✓	✓	✓
District FE	✓	✓	✓	✓	✓	✓
Province × Cohort FE				✓		✓
Controls+					✓	✓
Observations	5,006	5,006	5,006	5,006	5,006	5,006
R-squared	0.358	0.357	0.357	0.358	0.358	0.357

Notes: This table replicates Table 2 for the sample of *urban areas* (cities and towns). Due to the small size of the cohorts in some towns and resulting extreme sex ratios, the values in the tails of the distribution (1th and 99th percentiles) are winsorized. Controls+ include the term – Famine intensity × Year < 1892 interacted with average infant mortality (1887-1890, 1893-1897), latitude, longitude, exogenous soil productivity, average temperature and precipitation (GAEZ), ruggedness, distances to the railroad, to the nearest navigable river, coastline and capitals (Moscow and Saint Petersburg). We do not report lower order interaction terms for better exposition. Standard errors are clustered at the district level. *** p<0.01, ** p<0.05, * p<0.1.

Appendix E. Food Relief and Harvest

In this section, we provide additional evidence justifying the choice of our baseline measure of famine intensity – the share of the district's population on food relief. We first demonstrate that this measure is mainly orthogonal to the share of the non-Russian population (Table E1), and second, we show the strong "first stage" - the positive link between food relief and mortality shock (Table E2-4).

Table E1 reports the estimates of the regression analysis of the link between the measures of famine intensity and the share of the non-Russian population.

Table E1. Food relief and 1891 Harvest as predictors of mortality in 1891

	(1)	(2)	(3)	(4)
	Dependent variable:			
	Share pop on food relief	Food loan pc	Months on relief	1891 Harvest pc
1891 Harvest pc.	-0.009***	-0.050***	-0.653***	
	(0.002)	(0.009)	(0.115)	
Share non-Russian	-0.045	-0.277	-3.894	-2.164**
	(0.033)	(0.218)	(2.835)	(1.099)
Share literate pop	-0.053	0.229	0.333	7.716
	(0.095)	(0.667)	(8.553)	(4.762)
Log railroad dist, 1870	-0.007	-0.010	-0.568	-0.356**
	(0.005)	(0.025)	(0.410)	(0.176)
Outcome mean	0.141	0.574	9.32	8.00
District controls	yes	yes	yes	yes
Province fe	yes	yes	yes	yes
Observations	492	491	491	492
R-squared	0.832	0.762	0.781	0.548

Notes: the alternative measures of famine intensity (food loan pc and months on relief) are taken from Chanrnysh 2022. Other control variables include literacy rate, exogenous soil productivity, distances to the capitals (Moscow and Saint-Petersburg), coast line, latitude and longitude, level of precipitation and temperature, ruggedness (GAEZ).

Table E2 reports the estimates of the statistical "horse race" between 1891 harvest and the share of population on food relief as predictors of 1891 infant mortality.

Table E2. Food relief and 1891 Harvest as predictors of mortality in 1891

	(1)	(2)	(3)	(4)	(5)	(6)
	Dep var: 1891 Infant Mortality					
Share pop on food relief	10.853***		11.449***	6.023***		5.855***
	(1.108)		(1.257)	(1.707)		(1.789)
Harvest 1891		-0.117**	0.026		-0.054	-0.004
		(0.052)	(0.050)		(0.048)	(0.049)
District controls	yes	yes	yes	yes	yes	yes
Province FE				yes	yes	yes
Observations	501	492	492	501	492	492
R-squared	0.620	0.534	0.616	0.793	0.785	0.792

Notes: Other control variables include literacy rate, exogenous soil productivity, distances to the capitals (Moscow and Saint-Petersburg), coast line, latitude and longitude, level of precipitation and temperature, ruggedness (GAEZ). 1891 Harvest measure amount of cohorts-specific controls including birth ratio at birth and cohort size. Controls+ include district-level controls that interacted with famine intensity (1891 harvest), defined as a difference between the total amount of grain collected and sowed (in pud per square of rural land). We take an average across three main crops rye, wheat spring, wheat summer, and potatoes weighted by their sown square.

In table E3 we demonstrate the robustness of the previous findings by substituting our own measure of 1891 harvest with the alternative measure of harvest drop based on the difference between predicted and actual harvest in 1891.

Table E3. Food relief and 1891 Harvest (alternative measure) as predictors of mortality in 1891

	(1)	(2)	(3)	(4)	(5)	(6)
	Dep var: 1891 Infant Mortality					
Share pop on food relief	10.853***		11.532***	6.023***		5.678***
	(1.108)		(1.267)	(1.707)		(1.770)
Harvest Drop		-0.923**	0.292		-0.534	-0.188
		(0.401)	(0.422)		(0.358)	(0.371)
District controls	yes	yes	yes	yes	yes	yes
Province FE				yes	yes	yes
Observations	501	492	492	501	492	492
R-squared	0.620	0.534	0.617	0.793	0.785	0.792

Notes: Other control variables include literacy rate, exogenous soil productivity, distances to the capitals (Moscow and Saint-Petersburg), coast line, latitude and longitude, level of precipitation and temperature, ruggedness (GAEZ). Harvest Drop Is the difference between actual and predicted harvest levels. The prediction are based on linear OLS model using the sample of districts not affected by the famine and exogenous climate and geographic variables as predictors.

In table E4, we perform our main analysis by substituting the measure of famine intensity based on food relief with 1891 harvest. In areas with larger harvests, we expect boys not to receive preferential treatment, as household resources would be less constrained. Using our preferred controls and fixed effects (columns 1–2), we find that the coefficient on the interaction term is insignificant. Then we restrict the sample by excluding districts with low harvests that did not receive aid and high harvests that did. We assume that areas without any government food support should typically have harvests above the 25th percentile. If such areas fall below this threshold, it likely reflects either measurement error or the ability to compensate for a poor harvest through grain reserves. Similarly, districts with relatively good harvests that nonetheless received government support might indicate that the food constrains were actually high. After applying this sample restriction, the coefficient increases in magnitude and approaches conventional levels of statistical significance (Table E4, column 3). These findings suggest that while the variation in the 1891 harvest generally aligns with food relief, some districts' harvest data fail to fully reflect actual food constraints. This discrepancy likely contributes to the insignificant estimates reported earlier (Table E4, columns 1-2). Excluding these outliers brings the estimates in line with our baseline results.

Table E4. 1891 Harvest, Folklore Bias and Sex Ratios

	(1)	(2)	(3)
	Dep var: M/F Ratio		
1891 Harvest × Year < 1892	-0.001	-0.018	-0.030
	(0.002)	(0.043)	(0.053)
Bias × 1891 Harvest × Year < 1892	**0.019**	**-0.041**	**-0.075**
	(0.030)	**(0.039)**	**(0.046)**
Bias × Year < 1892	-0.343	-0.026	0.066
	(0.296)	(0.305)	(0.340)
Cohort-specific controls	yes	yes	yes
Cohort FE	yes	yes	yes
District FE	yes	yes	yes
Province × Cohort FE	yes	yes	yes
Controls+		yes	yes
Restricted sample			yes
Observations	4,920	4,920	3,740
R-squared	0.374	0.379	0.395

Notes: Bias is the folklore-based measure of discrimination against women. 1891. Cohorts-specific controls including birth ratio at birth and cohort size. Controls+ include district-level controls that interacted with famine intensity (1891 harvest), defined as a difference between the total amount of grain collected and sowed (in pud per square of rural land). We take an average across three main crops rye, wheat spring, wheat summer, and potatoes weighted by their sown square. Year < 1892 is a binary indicator that equals one for the cohorts born before 1892 and thus affected by the 1891 Famine. Controls+ include the term – 1891 Harvest × Year < 1892 interacted with average infant mortality (1887-1890, 1893-1897), latitude, longitude, exogenous soil productivity, average temperature and precipitation (GAEZ), ruggedness, distances to the railroad, to the nearest navigable river, coastline and capitals (Moscow and Saint Petersburg). We do not report lower order interaction terms for better exposition.

In Table E5, we demonstrate robustness of our main findings to the alternative measures of famine intensity. *Food loan pc* denotes per capita size of bread loan provided by the government to the population at the district level. *On food relief months* refers to the duration of food relief program.

Table E5. Baseline and alterative measures of the 1891/92 Famine intensity

	(1)	(2)	(3)
	Dep var: M/F Ratio		
Share pop on food relief× Year < 1892	1.140		
	(1.764)		
Bias × Share pop on food relief × Year < 1892	2.962***		
	(1.047)		
1891 Food loan pc × Year < 1892		0.438	
		(0.499)	
Bias × Food loan pc × Year < 1892		0.519**	
		(0.257)	
On food relief months × Year < 1892			-0.019
			(0.029)
Bias × On food relief months × Year < 1892			0.039**
			(0.015)
Bias × Year < 1892	-0.541	-0.451	-0.514
	(0.419)	(0.389)	(0.391)
Cohort-specific controls	yes	yes	yes
Cohort FE	yes	yes	yes
District FE	yes	yes	yes
Province × Cohort FE	yes	yes	yes
Controls+	yes	yes	yes
Observations	5,006	4,980	4,980
R-squared	0.364	0.365	0.365

Notes: Bias is the folklore-based measure of discrimination against women. Year < 1892 is a binary indicator that equals one for the cohorts born before 1892 and thus affected by the 1891/92 Famine. The measures *Food loan pc* and *On food relief months are taken from Charnysh (2022).* Controls+ include the term – Famine intensity × Year < 1892 interacted with average infant mortality (1887-1890, 1893-1897), latitude, longitude, exogenous soil productivity, average temperature and precipitation (GAEZ), ruggedness, distances to the railroad, to the nearest navigable river, coastline and capitals (Moscow and Saint Petersburg). We do not report lower order interaction terms for better exposition.

Appendix F

This appendix details the rationale for excluding the 1892 birth cohort from the treated group in our main specification, and shows the robustness of the results to its inclusion. The decision is based on the specific conditions this cohort faced due to the timing of famine severity and relief.

1) Timing of Food Relief

Figure A2 demonstrates that the bulk of state-sponsored food relief arrived in the first half of 1892. This means that for the 1892 cohort food availability was improved, particularly for those born in the second half of the year. This alleviation of food constraints could have altered household rationing behavior in a way that confounds the pure famine effect.

This contrasts to pre-1892 cohorts who experienced the famine at its peak severity (autumn 1891), after harvest reserves were depleted and before significant government aid arrived.

2) Different Vulnerability

The 1892 cohort was in utero or breastfeeding during the famine's peak, making them less dependent on the solid food that was the subject of intra-household allocation decisions. Likely older siblings from the 1890/91 and earlier cohorts were directly competing for scarce solid food, placing them at greater risk from discriminatory rationing.

Robustness

As a robustness check, we assess the sensitivity of our results to the exclusion of the 1892 cohort by redefining the treatment group to include it. Table F1, Panels A and B, presents a comparison between the baseline estimates (from Table 2) and those from the expanded model. The results demonstrate that the key findings are robust to this change in sample definition. Specifically, in our preferred specification (Column 6), the coefficients remain statistically significant and are quantitatively close in magnitude, indicating that our conclusions are not driven by the exclusion of the 1892 cohort.

Table F1. Effect of famine and gender bias on sex ratios with and without 1892 cohort

	(1)	(2)	(3)	(4)	(5)	(6)	(7)
	Dependent variable: M / F Ratio						
	Panel A: 1892 cohort excluded from treatment						
Famine intensity × Year < 1892	0.013**	0.017***	-0.047	-0.127	2.075*	1.140	-0.202*
	(0.006)	(0.006)	(0.049)	(0.085)	(1.174)	(1.764)	(0.103)
Bias × Famine × Year < 1892			0.816	1.756*	1.983***	2.962***	2.461*
			(0.635)	(1.049)	(0.597)	(1.047)	(1.334)
Bias × Year < 1892			-0.247**	-0.390	-0.122	-0.541	0.131
			(0.123)	(0.361)	(0.153)	(0.419)	(0.275)
Sex ratio at birth		0.101***	0.101***	0.116***	0.100***	0.112***	0.079***
		(0.027)	(0.027)	(0.028)	(0.027)	(0.028)	(0.028)
Cohort size		0.022*	0.025*	0.047	0.035**	0.051	0.041**
		(0.013)	(0.013)	(0.032)	(0.015)	(0.035)	(0.018)
Observations	5,006	5,006	5,006	5,006	5,006	5,006	5,006
R-squared	0.260	0.265	0.266	0.358	0.276	0.364	0.355
	Panel B: 1892 cohort included in treatment						
Famine intensity × Year < 1892	0.006	0.011*	-0.049	-0.119	1.716	1.063	-0.168**
	(0.006)	(0.006)	(0.049)	(0.083)	(1.088)	(1.672)	(0.066)
Bias × Famine × Year < 1892			0.770	1.618	1.739***	2.870***	1.859**
			(0.635)	(1.022)	(0.598)	(1.084)	(0.884)
Bias × Year < 1892			-0.217*	-0.390	-0.096	-0.562	0.113
			(0.121)	(0.372)	(0.148)	(0.446)	(0.234)
Sex ratio at birth		0.093***	0.093***	0.110***	0.093***	0.106***	0.075***
		(0.026)	(0.026)	(0.026)	(0.026)	(0.026)	(0.027)
Cohort size		0.021*	0.023*	0.041	0.034**	0.044	0.036**
		(0.012)	(0.012)	(0.032)	(0.014)	(0.035)	(0.017)
Cohort FE		✓	✓	✓	✓	✓	✓
District FE		✓	✓	✓	✓	✓	✓
Province × Cohort FE				✓		✓	
Controls+					✓	✓	
District specific linear trends							✓
Observations	5,507	5,507	5,507	5,507	5,507	5,507	5,507
R-squared	0.246	0.250	0.251	0.347	0.259	0.353	0.333

Notes: Panel A replicates baseline estimates of the paper (Table 2) with exclusion of 1892 cohort from the sample. Panel B reports the estimates with 1892 included in treated cohorts. Controls+ include the term – Famine intensity × Year < 1892 interacted with average infant mortality (1887-1890, 1893-1897), latitude, longitude, exogenous soil productivity, average temperature and precipitation (GAEZ), ruggedness, distances to the railroad, to the nearest navigable river, coastline and capitals (Moscow and Saint Petersburg). Standard errors are clustered at the district level. *** p<0.01, ** p<0.05, * p<0.1

Appendix G

Below are example motifs from the original Berezkin (2015) catalogue, as reported in Michalopoulos and Xue (2021). We draw these examples from motifs in the top quartile of the Michalopoulos and Xue (2021) measure for each motif attribute. For instance, a value of 0 for female submissiveness indicates that the motif does not address this theme, while higher values indicate a stronger association. We further restrict the sample to motifs from ethnic groups that can be matched to the 1897 Census ethnicity categories.

Motif examples:[56]

High in submissiveness:

(1) A rival or adversary kidnaps man's wife or bride or takes her sending her man off. The man gets her back.

(2) An imposter pretends to be the hero to take his position and/or to marry or to violate a woman.

(3) Supernatural male person (often a snake, a dragon) hides clothes of a human girl or sits on it. To return her clothes she had to become his wife.

(4) An unrecognized hero comes to a place where his bride or wife has to marry another man or is tuned into a slave. Despite expectations, he gets to strain a tight bow killing his rivals.

(5) Disguised as a bird, small animal or insect, the hero lets himself be caught by a girl to copulate with her or to steal valuables possessed by her family or he lets himself be caught by the girl's father to receive opportunity to get to his daughter.

(6) In the beginning of time two or more men (animal people) have only one woman.

(7) A girl gets sick. A man comes to cure, and the cure is to copulate or to attempt to copulate with her.

High in stupidity:

(1) A man has accumulated a store of provisions (saved some money, etc.) and tells his wife that it is for the long winter (Christmas, emergencies, etc.). A trickster (beggar) comes to the woman and tells her that his name is Long Winter, etc. She gives him the provisions.

[56] While we mainly use the Google API English translations as used by Michalopolous and Xue (2021), we refine some of the translations from Russian, where the tool was not accurate.

(2) Stupid woman beliefs that frogs will spin or weave for her and throws her yarn (ball of threads) into the water.
(3) Two girls or young women meet a demon. One or both are not aware of danger. One is eaten up, another escapes.
(4) Because a girl or young woman offends a cat (rare: a dog), it plays a trick that brings the girl a misfortune (usually extinguishes the fire and in search of it the girl gets to a demon).