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Long-term health and socioeconomic consequences of early-life exposure to the 1959–1961 Chinese Famine

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ABSTRACT

This research investigates long-term consequences of early-life malnutrition by examining effects of the 1959–1961 Chinese Famine. Taking into account temporal and geographic variations in famine severity, we construct a difference-in-differences estimator to identify effects of early-life exposure to famine on perceived health and socioeconomic outcomes in midlife. Using a sample of 1716 adults born in 1955–1966 in rural China from a nationally representative survey—the 2005 Chinese General Social Survey—we find that the famine had adverse effects on mid-life health for males born into families where at least one parent was a Communist Party member and females regardless of parental party membership. Being born during the famine had no effects on years of education or income for either gender. Quantile regressions suggest intense mortality selection among males who had no party-affiliated parents. Our study highlights the importance of timing and contexts of life experiences in shaping health.

1. Introduction

In recent years, health studies have begun using a life-course approach more and more often, emphasizing the long-term impact of experiences during critical periods of growth on subsequent health and human capital (Kuh et al., 2003). In particular, exposure to adverse environmental influences during critical developmental stages, such as prenatal and early-postnatal periods, is found to lead to permanent changes in gene expression as well as brain and body functions as early as childhood, with increased disease risk in adulthood (Barker, 1995, 2007). From a policy perspective, it is important to understand the lasting effects of early-life malnutrition (Barker, 2003; Lumey et al., 2011; Victora et al., 2008), given that child under-nutrition is a problem not only in many developing countries (Black et al., 2008; de Onis and Blössner, 2000; Pelletier et al., 2012), but in some developed countries as well. In the United States, for example, an estimated 16.7 million children under the age of 18 lived in food insecure households in 2011 (Coleman-Jensen et al., 2012).

It is difficult, however, to measure nutritional intake during prenatal life, infancy, and early childhood. Thus, situations in which nutritional deprivation is created exogenously such as famines provide a unique opportunity to study the impact of early-life under-nutrition on subsequent health (Brown and Susser, 2008). While the 1944–1945 Dutch Famine is the most extensively examined famine in the literature, the 1959–1961 Chinese Famine lasted much longer and had more severe consequences resulting in an estimated 16.5–30 million excess deaths over the three-year period (Ashton et al., 1984; Banister, 1987; Coale, 1981; Peng, 1987; Yao, 1999). In this study we take advantage of this natural experiment to investigate

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the consequences of the Chinese Famine for perceived health and socioeconomic outcomes in midlife. Previous research on the Chinese Famine mostly relied on samples with limited regional coverage and yielded mixed findings (Almond et al., 2007; Chen and Zhou, 2007; Gørgens et al., 2012; Huang et al., 2010, 2012; Luo et al., 2006; Meng and Qian, 2009; Mu and Zhang, 2011; St Clair et al., 2005; Xu et al., 2009; Yang et al., 2008), and none of them examined perceived health, an important predictor of mortality (Idler and Benyamini, 1997).

Drawing on the life course perspective, we use nationally representative data from the 2005 Chinese General Social Survey (CGSS) to investigate the long-term health and socioeconomic consequences of early-life exposure to the 1959–1961 Chinese Famine. We use a difference-in-differences method taking advantage of the temporal and regional variations in famine intensity. Recognizing the importance of social contexts in shaping the linkage between early life experiences and later life chances (Elder et al., 2003), we examine famine effects separately among four subgroups at different risk levels of nutritional deprivation, classified by gender and family background. By doing so we extend prior studies that examined the powerful role of social forces in producing health inequalities (Link and Phelan, 1995). We also contribute to the literature by demonstrating that, contrary to conventional wisdom, when combined with biological mechanisms (such as mortality selection), advantageous family background does not necessarily serve as a buffer when mid-life health is the focal outcome. Our study sheds light on the relationship between early-life nutritional deficiency and later health in China, a large, poor country in the midst of famine (Ashton et al., 1984), and in recent decades as it has experienced dramatic increases in the overall health and economic standing of its citizenry (Chen et al., 2010). This study also serves as an important step in assessing the healthcare needs of individuals born during the famine, who are now moving through retirement.

The paper is organized as follows. We first review empirical evidence from the Dutch and Leningrad famine studies. Next, we provide historical accounts and empirical findings of the 1959–1961 Chinese Famine. This is followed by a description of the data, sample, measures, and analytical strategies. Adopting a difference-in-differences approach, we estimate famine effects on health and socioeconomic outcomes in midlife, first by gender and then by gender and family background. We conclude with a discussion of key findings.

2. Background

2.1. Early life as a critical period: evidence from the Dutch and Leningrad famines

Inspired by the life course framework—timing of lives, in particular—the "critical period" model suggests that the same life event may have different implications for individual development depending on when each experienced the event (Kuh et al., 2003). Under-nutrition *in utero*, for example, is hypothesized to be particularly devastating with lasting effects on health in later life (Barker, 1995, 1998, 2007).

The Dutch hunger winter (1944–1945) and the siege of Leningrad (1941–1944) are two extensively studied famines that provide key evidence for the "critical period" model, but with conflicting findings. Surviving adults prenatally exposed to the Dutch famine had higher mortality up to age 50, worse self-rated health, higher coronary heart disease risk, reduced glucose tolerance, higher BMI, and increased risk of psychological disorders (Brown et al., 2000; Neugebauer et al., 1999; Ravelli et al., 1999; Roseboom et al., 2000a, 2000b, 2001a, 2001b, 2003). In contrast, the Leningrad famine did not see significant long-term effects: intrauterine malnutrition was not associated with glucose intolerance, dyslipidaemia, hypertension, or cardiovascular disease in adulthood (Stanner et al., 1997; Stanner and Yudkin, 2001). The Leningrad famine not only lasted longer than the Dutch famine, but was preceded and followed by relative shortages of food. Under the influence of mortality selection, therefore, estimates of the impact of the Leningrad famine based on data of survivors may be biased toward zero (Roseboom et al., 2000a). These different findings indicate a complex interplay between historical events, social forces, and biological mechanisms, and highlight the importance of examining the famine effects in the unique Chinese context.

2.2. The 1959–1961 Chinese Famine and its short- and long-term health consequences

As one of the largest human catastrophes, the 1959–1961 Chinese Famine has received much scholarly attention. Lin and Yang (2000) argued that the Chinese Famine was jointly determined by urban-biased policies and the decline in food availability. To extract as much agricultural surplus as possible to facilitate the heavy industry-oriented development strategy, rural residents were forced to deliver quotas at prices set by the government, thus bearing most of the brunt of food supply reduction (Coale, 1981; also see Fig. 1). Following previous studies (Chen and Zhou, 2007; Huang et al., 2012; Li et al., 2010; Meng and Qian, 2009; Mu and Zhang, 2011), we focus on the rural sample in our analysis. Key to our research, exposure to the famine—determined by birth date and province of birth—is exogenous, enabling us to better identify the famine effect.

Earlier studies on the Chinese Famine revealed a sudden increase in mortality during the famine years (Fig. 1). The estimated total number of premature deaths during the famine ranged from 16.5 to 30 million (Ashton et al., 1984; Banister, 1987; Coale, 1981; Peng, 1987; Yao, 1999). Retrospective data also revealed a sharp increase in miscarriages and stillbirths during that time (Cai and Wang, 2005).

In terms of long-term effects of the Chinese Famine, an analysis of the 1988 Chinese national fertility survey showed no differences in mortality up until the late 20s and early 30s among the 1954–1958 pre-famine cohort, the 1959–1962 famine cohort, and the 1963–1967 post-famine cohort (Song, 2009), which might be partially explained by mortality selection at

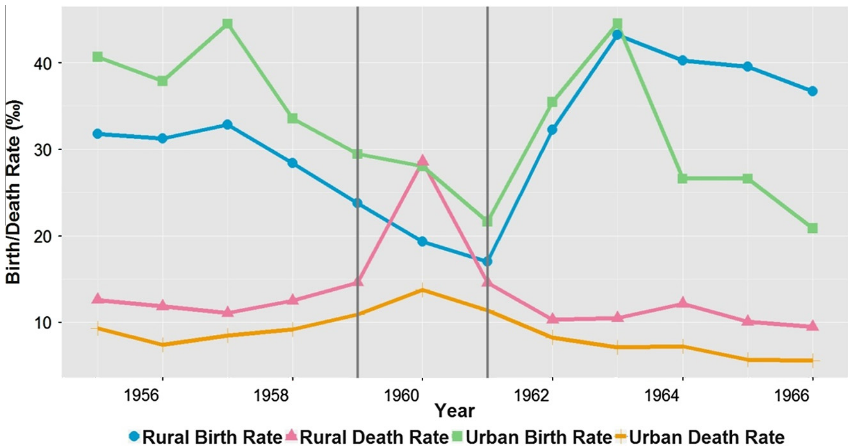

Fig. 1. Birth rates and death rates in rural and urban China; 1955–1966. (Source: Lin and Yang, 2000; Table 2, p. 145).

earlier ages. Some studies found negative effects for other health outcomes. For example, drawing on data from the China Health and Nutrition Survey (CHNS) which covers selected provinces, Chen and Zhou's (2007) difference-in-differences estimates showed that rural Chinese born in 1959 suffered most severely in terms of height reduction. Absent the famine, they would have grown 3.03 cm taller. This effect did not go away even when mortality selection was explicitly controlled for (Gørgens et al., 2012). Similarly, Huang et al. (2010) found that famine exposure reduced the height of rural females born in 1958 and 1959 by 1.7 and 1.3 cm, respectively, based on data collected from three provinces in China between 1993 and 1996. Also using the difference-in-differences approach, Luo et al. (2006) reported that rural females born during the famine years had a higher probability of being overweight. Meng and Qian (2009) adopted an instrumental variables method; their two-stage-least-squares estimates revealed that exposure to famine at ages 1–3 reduced height by 1.3% (2.1 cm), weight by 4.4% (2.4 kg), and upper arm circumference by 3.4% (0.8 cm), but no effect was found for hypertension, BMI, or weight-for-height. In an ingenious design, Almond et al. (2007) reported that mothers born in 1961 on mainland China who later migrated to Hong Kong were 8% more likely to give birth to a child of low birth weight (<2500 g), whereas no significant effects were detected for their Hong Kong-born counterparts. To the extent that the emigrants are healthier and/or from more advantaged families, they might underestimate the true famine effect.

Although previous studies have examined a wide range of health outcomes including disability, obesity, height, weight, weight-for-height, hypertension, and mental health, the majority of them drew on samples collected from a few provinces, limiting generalizability. Further, few studies have examined perceived health (Lumey et al., 2011), despite the fact that self-assessed health is a well-recognized predictor of subsequent mortality (Idler and Benyamini, 1997). To the best of our knowledge, only one paper (Roseboom et al., 2003) has studied the effects of prenatal exposure to the Dutch famine on adults' perceived health, finding those exposed to famine in early, but not mid or late, gestation more often rated their health as poor, but no such study has been conducted in the context of China.

2.3. Gender differences in the famine effects on subsequent health

That males and females are affected by the famine differently is well established in previous studies. The long-term famine effects appeared to be more pronounced among women than among men in China (Almond et al., 2007; Gørgens et al., 2012; Luo et al., 2006; Mu and Zhang, 2011; Ravelli et al., 1999; Yang et al., 2008). Significant famine effects were found among women but not among men on reduced height (Gørgens et al., 2012), higher incidence of disability (Mu and Zhang, 2011), and higher odds of being overweight (Luo et al., 2006; Yang et al., 2008). Additionally, Huang et al. (2012) reported increased risk of mental disorders among female famine survivors, but either decreased or no risk of mental illness among male famine survivors. Early-life mortality selection and differential treatment offer two explanations for such gender differences (Mu and Zhang, 2011).

First, males biologically exhibit greater vulnerability to prenatal insult than females (Kline et al., 1989; Kraemer, 2000). Evidence of male's higher excess mortality during natural disasters is well documented in China and elsewhere (see Macintyre, 2002; Sen, 1981; Song, 2009). As a result, male survivors are more likely to come from the upper end of health distribution than female survivors, leading to a more attenuated long-term famine effect on men than on women.

Second, the differential treatment hypothesis suggests that hard times force parents to allocate limited resources to the more valued children (Das Gupta and Li, 1999). In traditional Chinese society this means the sons (Das Gupta et al., 2003), as only sons can carry the family name, and sons instead of daughters are expected to provide support for elderly parents. Given the lack of a formal social security system in rural China, parents need to rear at least one son to care for them in their old age (Croll, 2000). Hence, boys may have received more food than their sisters, who then bore the brunt of malnutrition during the famine. In sum, despite pointing to different sources, both mortality selection and differential treatment lead us to hypothesize that rural women were more adversely affected by the Chinese Famine than rural men:

Hypothesis 1. Early-life exposure to the 1959–1961 Chinese Famine had more negative effects on perceived health in midlife for rural females than for rural males.

2.4. Family background as a moderating context

Social disruptions such as famines, wars, and ethnic cleansing, and how the subsequent effects are distributed across social groups are the focus of many empirical studies. Strauss and Thomas (1998) showed that the heights of Vietnamese adult men and women increased very rapidly for those born between 1925 and 1955, but remained unchanged for the Vietnam War cohort born from the mid 1950s to 1975. In particular, the stature of the poorest was hit hardest by the war. These findings indicate that the identity of those who will suffer most from catastrophic events can be traced to an unequal distribution of desirable resources, which follows a sociological logic (see also Link and Phelan, 1995). In the case of the Chinese Famine, the resource at stake is food. Who will suffer least from the famine can be determined by identifying who was advantaged in gaining access to food even when it was scarce.

Unlike Western society where education, income, and occupation loom large in access to a variety of resources (Elo, 2009), the stratification system in rural China during the Maoist era was political in nature (Bian, 2002; Walder, 1989). Shortly after it came into power, the Chinese Communist Party destroyed the economy-based system and replaced it with a system of access to goods, jobs, and power based on political power, especially in rural areas.[1] Political power took two forms: membership in the Communist Party and employment in a cadre position. Party members were a select group, and during the Maoist era, were usually recruited from those who successfully demonstrated their loyalty to the party (Bian et al., 2001), for example, through a "good" family background as designated by the state and based on father's or grandfather's occupation (e.g., poor/lower-mid peasant or revolutionary military). A cadre position was one of political or administrative leadership, and these appointments were highly controlled by local party organizations even as recently as the late 1980s (Oi, 1989). Thus, cadres—especially those with real power—are usually party members (Morduch and Sicular, 2000; Wu and Treiman, 2004).

Compared with ordinary peasants, those tightly connected with the Communist Party had preferential access to valuable resources and were in a much more strategic position to use the system for their own personal benefit. In other words, even some born during the famine years escaped severe deprivation because the household belonged to a more powerful class and had more access to food. For example, at the commune level, party cadres were awarded a higher ration of food than those they managed, and they were also permitted to exceed their shares (Thaxton, 2008). Further, anecdotal evidence suggests that during the famine, local party cadres often abused their power by illegally trading food with others in their political networks, inflating the amount of rations their district required and appropriating the surplus, distorting the transfer and distribution of food to peasants by stealing. In essence, they saved themselves and their families at the expense of ordinary peasants who had no political power (Dikötter, 2010; Thaxton, 2008).

Given this evidence, we hypothesize that the long-term famine effects are moderated by family background, especially among male famine survivors, for two reasons. First of all, recent studies suggest that males are more susceptible to adverse childhood circumstances such as food shortage. Salonen et al. (2009) found that lower SES during childhood measured through father's occupation was associated with higher BMI only in men, and their explanation is that the nutrition of boys *in utero* is more dependent on mothers' diet in pregnancy, whereas the nutrition of girls relies more on maternal protein metabolism (as reflected in mothers' height) (Eriksson et al., 2010). Second, given the preference for sons in rural China, parents were less likely to allocate food, if they had any, to girls than to boys. In other words, girls were not likely to be well-fed regardless of family background, particularly during a famine with increased resource constraints, whereas access to food might differ by family circumstances, to some extent, among boys. Thus, we hypothesize that the buffering role of family background is more pronounced for males:

Hypothesis 2. Early-life exposure to the 1959–1961 Chinese Famine had more negative effects on perceived health in midlife for those born into non-party-affiliated families than for those born into party-affiliated families, especially among males.

[1] One reviewer raised the issue that, even in the era before China's economic reform in 1978, family background was determined by multiple factors, such as work sectors (state vs. non-state owned), occupation, education, and military experience. It is true that occupation and work sectors played a prominent role in determining one's life chances in urban China (see Lin and Bian, 1991), but these two indicators were not relevant in rural China until the rise of the township and village enterprises in the 1980s, before which off-farm employment was not an option for rural residents (Bian, 2002). Indeed, in the CGSS (data used for our analysis), no further questions were asked regarding respondents' parents' work unit if their parents' occupation was "doing farm work." Education was a dubious resource during the Maoist era given the top leaders' inherent skepticism toward intellectuals (Bian, 2002; Walder, 1989), as evidenced by their last ranking among all nine "black" (or "bad") officially-designated family background categories. Thus, the better educated were unlikely to have advantages in food access during the famine. Indeed, in Appendix Table A, we used parental education to measure family background, but no significant difference in famine effects was found between those whose parents had more or less education. The Appendix shows junior high school as a cutoff point, but using primary school or senior high school yields the same pattern. (Results available upon request). This confirms that education was not the key dividing line in rural China's social stratification system during the Maoist era.

2.5. Long-term socioeconomic consequences of early-life exposure to the famine

A nutritional shortage may have adverse effects on educational and economic outcomes. However, existing literature is scarce and yields inconsistent findings regarding the famine effect on adult SES. No systematic influence of the Dutch famine was found on cognitive ability or income later in life (Stein, 1975; Scholte et al., 2012). Studies drawing on rural samples reported that the 1959–1961 Chinese Famine had no effect on working hours (Meng and Qian, 2009; Shi, 2011), but led to reduced educational attainment, labor force participation, and farming income (Almond et al., 2007; Chen and Zhou, 2007; Meng and Qian, 2009). Given the mixed evidence, we tentatively develop our last hypothesis:

Hypothesis 3. Early-life exposure to the 1959–1961 Chinese Famine led to lower SES in adulthood for the rural population.

3. Data and methods

3.1. Data

We use data from the 2005 Chinese General Social Survey (CGSS). Initiated in 2003, the CGSS is an annual or biannual survey that provides rich information on Chinese people's socioeconomic achievement and quality of life. The 2005 CGSS was carried out jointly by the Hong Kong University of Science and Technology and Renmin University of China. It used multistage stratified random sampling to generate a nationally representative sample of the adult population in both rural and urban China (Bian and Li, 2012). The sample consists of 10,372 individuals ages 18 to 69, with 6098 and 4274 individuals from urban and rural areas, respectively. The 2005 CGSS is ideal for our study because of its relatively large sample size, national coverage, and detailed information on socio-demographic characteristics and health measures.

3.2. Sample

The analytic sample for this study consists of 1716 individuals born in rural areas from 1955 to 1966. Because the 2005 CGSS did not collect information on place of birth, largely following Qian and Hodson (2011), we use both father's and mother's occupation to determine whether respondents were born in rural or urban areas. If both parents were "doing farm work" when the respondent was 14, we classify the respondent as born in a rural area, which is appropriate given the rare social and geographic mobility before the late 1970s and the Chinese tradition of children living with their fathers.[2]

We define those born in 1955–1958 as the *pre-famine cohort*, those born in 1959–1962 as the *famine cohort*, and those born in 1963–1966 as the *post-famine cohort*. People born in 1962 are included in the famine cohort because some of them were in gestation during the famine that lasted through late 1961. Due to the lack of information on birth month, we cannot distinguish who in the 1962 birth cohort had prenatal famine exposure and who did not. A similar problem exists for the 1959 birth cohort. To the extent that the misclassified persons (i.e., persons born in 1959 or 1962) are more akin to the pre-famine (or post-famine) than to the famine cohort, we underestimate the true famine effect. Another problem is selection at birth. The birth rate declined by nearly 50 percent during the famine compared with "normal" years (Ashton et al., 1984; also see Fig. 1). Were more advantaged or healthy couples more likely to delay childbearing, thereby confounding the relationship between early-life exposure to famine and later health? Almond et al. (2007) used census data reporting that the 1959–1961 cohort were born to better educated mothers than adjacent cohorts, and healthier parents were more likely to succeed in giving birth during the famine. The latter suggests an underestimate of the true famine effects. To test whether our results were sensitive to education-based fertility selection, we controlled for parents' education in Appendix Table B, but no substantive change was observed.

3.3. Measures

3.3.1. Outcomes

The 2005 CGSS surveyed eight health measures based on the SF-36 instrument. The SF-36 is a commonly used instrument with 36 items measuring health-related quality of life across eight dimensions of health (Ware, 2000), which are powerful predictors of hospitalization, morbidity, and mortality (Fan et al., 2002; Lowrie et al., 2003; Rumsfeld et al., 1999). Further, the Chinese SF-36 has been validated as a reliable instrument for evaluating both physical and mental health status (Wang et al., 2008).

The original wordings for the eight health measures in the 2005 CGSS are: During the past month, (1) How would you rate your health in general? (self-rated health) (2) Has your health limited your daily life, e.g., walking, or stair climbing?

[2] As one reviewer pointed out, the Up to the Mountains and Down to the Countryside Movement initiated in 1968 lasted until the late 1970s and sent many urban dwellers to the countryside. The majority were teenagers who finished senior high school (usually ages 15–18), but a small portion were intellectuals in their 30s and 40s. Thus, we may have misclassified a handful of their urban-born children in our sample. Using both parents' occupations alleviates this issue, but to the extent that urban-born individuals had briefer and less severe famine exposure than the rural-born, our estimates can be conceived as a lower bound of the true famine effect.

(physical functioning) (3) Has your health limited your work, within or outside household? (role-physical) (4) Do you feel any pain in your body? (bodily pain) (5) How would you rate your energy in general? (vitality) (6) Have you been bothered by emotional problems such as feeling anxious, depressed, or irritable? (mental health) (7) Has your physical and emotional health limited your social activities with family, friends, neighbors, or groups? (social functioning) and (8) Has your emotional health limited your usual activities related to work, study, or others? (role-emotional).

Factor analysis shows that the eight items reflect a single underlying concept, based on the Kaiser criterion of retaining those factors with eigenvalues of at least 1 (Rust and Golombok, 2009). Therefore, we sum up the eight items to create one composite measure—*general health*, the health outcome used for the following analysis. When constructing the general health measure, we reverse-coded the eight items so that higher scores indicate better health. We also rescaled the self-rated health and bodily pain items from 1–6 to 1–5, so that each of the eight items contributed an equal share to the composite measure. The Cronbach's alpha for this measure is 0.94, indicating high reliability.

Socioeconomic status in adulthood is measured through *years of schooling* and *personal income in year 2004*, two indicators widely used to capture individual SES (Adler et al., 1994; Hauser and Carr, 1995). We take the natural log of income to correct for right skewness.

3.3.2. Explanatory variables

As will be elaborated in Section 3.4. Analytical Strategy, we exploit both temporal and geographic variations in famine intensity. Temporal variation is captured by birth cohort as described in Section 3.2. Sample. Geographic variation is captured by the province-level excess death rate provided in Huang et al. (2012), and is calculated as the gap between the death rate in famine years (1959–1961) and the average death rate in the three years before the famine (1956–1958). The severity of the famine varied considerably and idiosyncratically across the 26 provinces in our analytic sample (out of mainland China's 29 provinces at the time),[3] ranging from 0.13 per thousand in Shannxi to 28.63 per thousand in Sichuan. We also experimented with three other measures of famine intensity: excess death rate (Yang, 1998, mortality difference between 1960 and non-famine years 1956–1958), cohort size shrinkage index (Huang et al., 2012), and fertility reduction index (Huang et al., 2012). These different indices were highly correlated at the provincial level (Huang et al., 2012); not surprisingly, our results did not change substantively with alternative measures of famine intensity (see Appendix Table C). To confirm that the excess death rate is not capturing variations across provinces in other key determinants of health such as economic development or health infrastructure, we tested for changes in physicians per 10,000 population and hospital beds per 10,000 population from 1956–1958 to 1959–1961.[4] Neither was associated with province-level famine intensity ($p = .75$ and .16 from Spearman's rank test, respectively). We did not find a good economic indicator due to data limitation, but Meng and Qian (2009) reported that the grain procurement system caused the famine to be severer in regions that typically produced more grain (usually the more economically developed areas).

3.3.3. Moderators

Women and men, as well as individuals from different family backgrounds, were affected by the Chinese Famine in different ways, given the strong preference for sons and the decisive role of political power in shaping life chances during the Maoist era (Dikötter, 2010; Thaxton, 2008; Walder, 1989). To reflect such contextual differences, we run separate models for the following four groups: males born into party-affiliated families, males born into unaffiliated families, females born into party-affiliated families, and females born into unaffiliated families. If respondents had at least one parent who was a Communist Party member when they were 14 years old, these respondents are coded as born into party-affiliated families. If respondents had no parent with Communist Party membership, they are coded as born into unaffiliated families (or non-party-affiliated families). Given our lack of knowledge regarding the exact year that parents became party members, it is possible that some parents joined the party after respondents were born. But note that the key here is the connection to the Communist Party apparatus and access to resources as a result of such ties. Thus our construction of party-affiliated vs. unaffiliated families should not pose an issue; even those parents who joined the party after having children should still be more connected with the party than unaffiliated residents (see Bian et al., 2001 for a discussion on the persistent political screening involved in becoming party members until the early 1990s). The 2005 CGSS does not allow us to distinguish cadres and party members as no information was collected on parents' cadre status. However, given that rural cadres were usually selected from party members (Morduch and Sicular, 2000; Wu and Treiman, 2004), we expect a similar pattern when using parental cadre status as the contextual factor. Lastly, given the small sample size for those born into party-affiliated families, we caution any over-interpretation of our findings.

[3] The 2005 CGSS did not survey Ningxia, Qinghai, and Tibet, three western provinces that had low to average levels of famine intensity.

[4] Both measures come from the China Data Center at the University of Michigan (http://chinadatacenter.org), an international center that integrates Chinese historical, social, and natural science data. In an email exchange, a staff member replied that these two measures were obtained from the National Bureau of Statistics of China. As one reviewer pointed out, the reliability of the data, especially of those collected during the famine period, is open to question. But note that the interest here is *changes* from pre-famine to famine periods across provinces; to the extent that bias resulted from data recording or reporting was stable either over time or across provinces, our general findings should not be hampered.

3.4. Analytical strategy

To identify the effects of early-life under-nutrition due to famine, one strategy is to take advantage of both the temporal and geographic variation in famine intensity (see Chen and Zhou, 2007; Huang et al., 2012; Luo et al., 2006). Specifically, one could track health outcomes for individuals residing in less severely affected provinces born during/before and after the famine, and then compare these differences with the corresponding differences for individuals residing in more severely affected provinces. This comparison produces the difference-in-differences estimator (DID) (Angrist and Pischke, 2009):

$$\Delta^2 = (Y^{FAMINE}_{Severe} - Y^{POST}_{Severe}) - (Y^{FAMINE}_{Less\ Severe} - Y^{POST}_{Less\ Severe}), \tag{1}$$

where the subscript "Severe" denotes residing in a province that was severely affected by the Chinese Famine (as measured by excess death rate), and the superscript denotes born in the famine period or after the famine period. Y indicates the outcomes (i.e., general health, years of schooling, and logged income). The estimator in Eq. (1) assumes that, were it not for the Chinese Famine, differences in adult health/SES between famine- and post-famine-cohort would have been similar across provinces.

The DID estimates can be conveniently computed in the regression framework using a continuous measure of famine severity level:

$$Y_{ij} = \beta_0 + \beta_1 pre_{ij} + \beta_2 fam_{ij} + \beta_3 edr_j + \gamma_1(edr_j \times pre_{ij}) + \gamma_2(edr_j \times fam_{ij}) + \sum_{k=1}^{26}\delta_k p_j + \varepsilon_{ij} \tag{2}$$

where Y_{ij} denotes the health (Hypotheses 1 and 2) or SES (Hypothesis 3) outcomes for individual i born in province j. *pre* and *fam* are two dichotomous variables indexing pre-famine cohort and famine cohort, respectively (post-famine cohort as the reference group). We use excess death rate, edr_j, to proximate province-level famine severity. To further ensure that our results are not driven by unobservable differences across provinces, we include province fixed effects (p_j) whenever the sample size is large enough (i.e., except when the samples are men or women from party-affiliated families).[5] Key to the current study is γ_2, the coefficient for the interaction term, which, in the DID setting, captures the effect of famine exposure *in utero* or infancy on health/SES in adulthood.

We use ordinary least squares (OLS) regressions, and adjust for age in all models given the life course patterns of both health and SES outcomes. The standard errors are clustered at the provincial level to account for correlation between observations within the same province. Missing data (mostly income) do not exceed 6% for any variable, so we report models where listwise deletion is used.

4. Results

4.1. Descriptive statistics

Table 1 reports descriptive statistics by cohorts and gender. Slightly more than half (54%) of the respondents are women. Respectively, 589 (34%), 413 (24%), and 714 (42%) individuals are from pre-famine, famine, and post-famine cohorts. Overall, men report significantly higher scores than women for all health indicators. When disaggregated by birth cohorts, pre- and post-famine cohort men have significantly better health than women from the same cohort for all health measures, whereas in the famine cohort, self-rated health and mental health do not significantly differ by gender, suggesting attenuated gender differentials in health for this cohort. Additionally, men have higher adult SES than women, as measured by education and income, while family background measured through parental party membership shows no significant gender difference.

4.2. Long-term famine effects on mid-life health

Models 1 and 2 in Table 2 present DID estimates for men and women, respectively. The 1959–1961 Chinese Famine did not appear to affect men's mid-life health (Model 1), but as expected in Hypothesis 1, famine had a significantly adverse effect on mid-life health for women: each additional 1‰ increase in excess death rate due to the famine resulted in a lowered general health by .193 point ($p < .01$, Model 2) among females born during the famine, compared with females born after the famine. These estimates take into account time-invariant differences between provinces and any inherent cohort differences. To put this into perspective, consider two provinces—Sichuan (excess death rate 28.63‰), a severely affected province, and Heilongjiang, a less severely affected province (excess death rate 1.75‰). For no other reason than being born in Sichuan, the

[5] We thank a reviewer for this suggestion. Note that, when the province dummies are not added, the main effect of excess death rate can be interpreted as the famine effects for those who were born after the famine, which should not be significant given that members from the post-famine cohort were not exposed to the famine, and this is indeed what we found (results available upon request). However, this interpretation no longer holds if we additionally control for province fixed effects because it requires one to think of a situation when excess death rate can change from one value to another while at the same time the province dummies are fixed at a given value, which does not make empirical sense given that excess death rate is a time-invariant province-level variable. We therefore do not present the main effects of excess death rate so as not to confuse readers. Our main findings, however, are substantively the same no matter whether we include province fixed effects or not (results available upon request).

Table 1
Descriptive statistics, by cohort and gender.

	Total (*N* = 1716)		Pre-famine cohort (*N* = 589)		Famine cohort (*N* = 413)		Post-famine cohort (*N* = 714)	
	Men (*N* = 791)	Women (*N* = 925)	Men (*N* = 308)	Women (*N* = 281)	Men (*N* = 184)	Women (*N* = 229)	Men (*N* = 299)	Women (*N* = 415)
Outcomes								
Health								
General Health Scale (8–40)	33.16	31.28***	32.65	29.96***	32.78	31.23*	33.93	32.21***
	(6.35)	(6.75)	(6.27)	(7.00)	(6.71)	(6.59)	(6.15)	(6.53)
Self-rated health (1–5)	3.64	3.39***	3.61	3.21***	3.54	3.43	3.73	3.49**
	(1.09)	(1.07)	(1.07)	(1.06)	(1.09)	(1.05)	(1.10)	(1.08)
Physical functioning (1–5)	4.37	4.15***	4.31	3.98***	4.34	4.11*	4.45	4.29*
	(0.90)	(0.96)	(0.90)	(1.01)	(0.96)	(0.94)	(0.86)	(0.91)
Role-physical (1–5)	4.35	4.14***	4.26	3.94***	4.32	4.11*	4.46	4.29*
	(0.93)	(0.99)	(0.92)	(1.06)	(0.99)	(0.94)	(0.89)	(0.93)
Bodily pain (1–5)	4.21	4.00***	4.14	3.86**	4.17	3.94*	4.32	4.12*
	(1.05)	(1.06)	(1.07)	(1.09)	(1.10)	(1.06)	(0.99)	(1.03)
Vitality (1–5)	3.78	3.53***	3.72	3.36***	3.70	3.48*	3.90	3.66**
	(1.05)	(1.07)	(1.04)	(1.09)	(1.11)	(1.05)	(1.01)	(1.06)
Mental health (1–5)	4.08	3.85***	4.04	3.77**	4.00	3.86	4.18	3.89***
	(0.97)	(1.03)	(0.96)	(1.07)	(0.99)	(0.99)	(0.97)	(1.02)
Social functioning (1–5)	4.37	4.12***	4.31	3.93***	4.35	4.16*	4.44	4.23**
	(0.86)	(0.94)	(0.85)	(0.98)	(0.92)	(0.93)	(0.84)	(0.91)
Role-emotional (1–5)	4.35	4.11***	4.27	3.92***	4.35	4.14*	4.44	4.22**
	(0.88)	(0.96)	(0.88)	(1.01)	(0.89)	(0.90)	(0.87)	(0.94)
SES								
Years of schooling	8.35	6.11***	7.95	5.22***	9.00	6.03***	8.35	6.77***
	(3.64)	(4.27)	(3.66)	(4.56)	(3.45)	(4.27)	(3.70)	(3.94)
Logged personal income in 2004	8.46	7.69***	8.32	7.54***	8.55	7.61***	8.55	7.83***
	(1.34)	(1.88)	(1.33)	(2.09)	(1.36)	(1.80)	(1.35)	(1.78)
Predictors								
Excess death rate (‰)	8.06	7.97	–	–	–	–	–	–
	(7.50)	(7.68)	–	–	–	–	–	–
Parent's party membership	0.09	0.09	0.09	0.10	0.11	0.07	0.08	0.10
	(0.29)	(0.29)	(0.29)	(0.30)	(0.32)	(0.25)	(0.28)	(0.30)
Age	44.62	43.93***	48.71	48.67	44.28	44.07	40.62	40.64
	(3.73)	(3.60)	(1.15)	(1.12)	(1.18)	(1.13)	(1.10)	(1.12)

Notes: 1. For each variable, the first row shows means and the second row shows standard deviations.
2. We test whether there is gender difference within each of the following samples: total sample, pre-famine-cohort, famine-cohort, and post-famine-cohort. Results are shown in columns denoting coefficients for "Women".
*** $p < 0.001$.
** $p < 0.01$.
* $p < 0.05$.

reduction in general health score was on average 5.19 points (=.193 * [28.63 − 1.75]) higher for Sichuan-born females than for females born in Heilongjiang.

In Models 3–6 in Table 2, we disaggregate the sample into four groups based on gender and parental party membership. A postestimation test (achieved via the "suest" command in Stata) indicates that the coefficients of the DID interaction terms for the four groups are significantly different from one another ($p = .0047$). In particular, we find that the Chinese Famine had significantly adverse effects on mid-life health for males born into party-affiliated families. Recall Model 1 showed that males as a whole were not affected by early-life exposure to the famine, but this null finding obscures heterogeneity along the dimension of family background. Among men whose parents were party members, each additional 1‰ difference in excess death rate led to a lower general health of males born during the famine by .513 point ($p < .001$, Model 4), compared with males born after the famine, while no significance was found for males whose parents were not party members (Model 3). Although insignificant, the famine effect on mid-life health appears to be positive among males born into non-party-affiliated families (.007, not significant, Model 3), probably due to high mortality selection within this group (Huang et al., 2012). A postestimation test shows that the difference of the interaction terms in Models 3 and 4 is significant ($p = .0013$), suggesting that for males, the effects of early-life under-nutrition are contingent on family background. Mortality selection is one potential explanation and we provide supportive evidence below.

A different pattern emerges for females. We find the famine effect is significant and negative only for those born into non-party-affiliated families; each additional 1‰ increase in excess death rate brought about by the famine widens the mid-life health disadvantage of the famine cohort by .211 point ($p < .05$, Model 5) vis-à-vis those born after the famine. Unlike males, however, for females the difference in the interaction terms between Models 5 and 6 does not reach significance based on a postestimation test ($p = .364$). This indicates that early-life under-nutrition adversely affected women's mid-life health,

Table 2
Difference-in-differences estimates of famine effects on mid-life perceived health, based on OLS regressions.

	Model 1: men	Model 2: women	Model 3: men, non-party-affiliated families	Model 4: men, party-affiliated families	Model 5: women, non-party-affiliated families	Model 6: women, party-affiliated families
Cohort (Ref. = Post-famine cohort)						
Pre-famine cohort	−2.084	−2.161	−1.627	−8.506*	−2.586	−0.465
	(1.571)	(1.499)	(1.490)	(3.902)	(1.655)	(6.003)
Famine cohort	−1.327	0.562	−1.199	−4.159	0.649	−0.894
	(1.113)	(0.811)	(1.180)	(2.573)	(0.912)	(4.383)
EDR *						
Pre-famine cohort	−0.023	0.088	−0.037	−0.072	0.094	0.145
	(0.058)	(0.068)	(0.050)	(0.145)	(0.085)	(0.146)
Famine cohort	−0.031	−0.193**	0.007	−0.513***	−0.211*	0.031
	(0.062)	(0.066)	(0.087)	(0.116)	(0.079)	(0.221)
Age	0.093	−0.079	0.059	0.838	−0.049	−0.233
	(0.159)	(0.181)	(0.158)	(0.542)	(0.193)	(0.675)
Constant	23.383**	22.480**	24.622***	2.114	21.616*	41.033
	(6.372)	(7.696)	(6.394)	(22.741)	(8.131)	(26.980)
Observations	791	925	716	75	840	85

Notes: 1. We do not show estimates for the main effect of excess death rate (see footnote 5 in the text). Estimates for province fixed effects are not shown for Models 1–3 or Model 5 for simplicity. Robust standard errors clustered at the province level are shown in parentheses.
2. Postestimation tests of the EDR * Famine-Cohort interaction terms across models: M3 = M4 = M5 = M6 ($p = .0047$); M3 − M4 = M5 − M6 ($p = .0433$); M3 = M4 ($p = .0013$); M5 = M6 ($p = .3640$).
*** $p < 0.001$,
** $p < 0.01$.
* $p < 0.05$.

regardless of their family background, a pattern significantly different from what is observed in men ($p = .043$). Taken as a whole, we find both supportive and contradictory evidence of Hypothesis 2: there are gender differentials in the moderating role of parental party membership, but contrary to our hypothesis, advantageous family backgrounds either intensify the negative famine effects (for males) or have no modifying role (for females).

In contrast to what has been found for the famine cohort, we observe neither statistically significant nor substantively important famine effects for the pre-famine cohort (the *pre* by *edr* interaction), despite their longer exposure to the famine. This lends support to the idea that it is the *timing* of exposure to adverse environments—*in utero* and/or early infancy—that largely determines health in later life.

Table 2 presents the adverse famine effects using a cohort categorization to capture the temporal variation in famine intensity. Is there a particular year of birth that made a single-year birth cohort particularly vulnerable? Replacing the two cohort indicators with birth year dummies (1955, 1956, ..., 1962), in Appendix Table D we show that males born into party-affiliated families in 1959 (−1.943, $p < .001$) and 1960 (−.550, $p < .01$), and females born into non-party-affiliated families in 1961 (−.401, $p < .1$) and 1962 (−.349, $p < .01$) were significantly and negatively affected by the famine. Further, as evidence suggests that the Chinese Famine began as early as 1958 (Ashton et al., 1984; Chang and Wen, 1997), we conducted a sensitivity analysis wherein the 1958 birth cohort is coded as in the famine cohort. Results do not change for males, but the famine effect for females born into non-party-affiliated families is no longer significant (see Appendix Table E). We consider the implication of these findings in Section 5.

4.3. Long-term famine effects on mid-life health: quantile regression

Famine effects may be different for survivors located in different areas in the health distribution (Chen and Zhou, 2007; Meng and Qian, 2009). To capture these differences, we estimate quantile regressions (from the 20th to the 80th percentile) via the "qreg" command in Stata. The estimation method is similar to OLS, but instead of minimizing sum of squared residuals it minimizes sum of absolute residuals; we use asymptotical standard errors. Note that in quantile regressions, "quantile" refers to those of the dependent variable, general health.

In general, famine effects on mid-life health as estimated from survivors are a function of two components: scarring effect and mortality selection. The former captures the long arm of early-life insults, while the latter is difficult to quantify without detailed information on the famine-induced deaths. Quantile regressions can be helpful in at least assessing the extent of mortality selection (Meng and Qian, 2009). Assuming that the strongest survive, mortality selection should be largest in lower quantiles, those who are the least healthy. As such, with heightened levels of mortality selection as typically occurred during the famine, the estimated effects of the famine on mid-life health should be attenuated toward zero or even become positive in the lower quantiles because, in this area, we are comparing individuals in the "control" group (post-famine cohort) with those in the "treatment" group (famine cohort), who would be healthier on the distribution absent the famine-induced selection. In contrast, the mortality selection bias is smaller for individuals in the higher quantiles because those individuals have more comparable "control" groups. Empirically, this means that higher mortality selection can be inferred if more attenuated (than the average) or even positive famine effects are observed in lower quantiles but not in higher quantiles.

For social groups constructed by gender and parental party membership, Fig. 2 plots estimates of their respective famine effects (i.e., interaction terms between famine cohort and excess death rate) from quantile regressions, denoted by circles and connected by a solid line, with the 95% point-wise confidence bands represented by the gray area around the solid line. The dashed horizontal lines represent the average famine effects derived from Table 2. We combine women from party-affiliated and the unaffiliated families, given that the famine effects do not differ between the two groups.

The contrast between men from the unaffiliated families and men from party-affiliated families is striking as shown in Fig. 2. The quantile-specific famine effects track closely with the dotted average level for the former group. More importantly, we observe a decline in famine effects from positive to negative in the lower quantiles (from 20th to 40th), lending some support to mortality selection. Males born into party-affiliated families, however, are adversely affected by early-life exposure to the famine across the board, with such effects particularly strong in the lower quantiles (below the dotted average level) and gradually attenuated in the higher quantiles. This suggests that among males born into party-affiliated families, the least strongest suffered greatly from the famine, but unlike their counterparts whose parents had no political power, they managed to survive, possibly facilitated by their parents' advantage in access to food. The pattern is less clear for women, among whom the famine effects do not seem to vary much across quantiles, probably because of females' biological survival advantage. Taken as a whole, these quantile regression results provide suggestive evidence of mortality selection for males from non-party-affiliated families, although the results should be interpreted with caution given small sample sizes and wide confidence intervals at some quantiles.

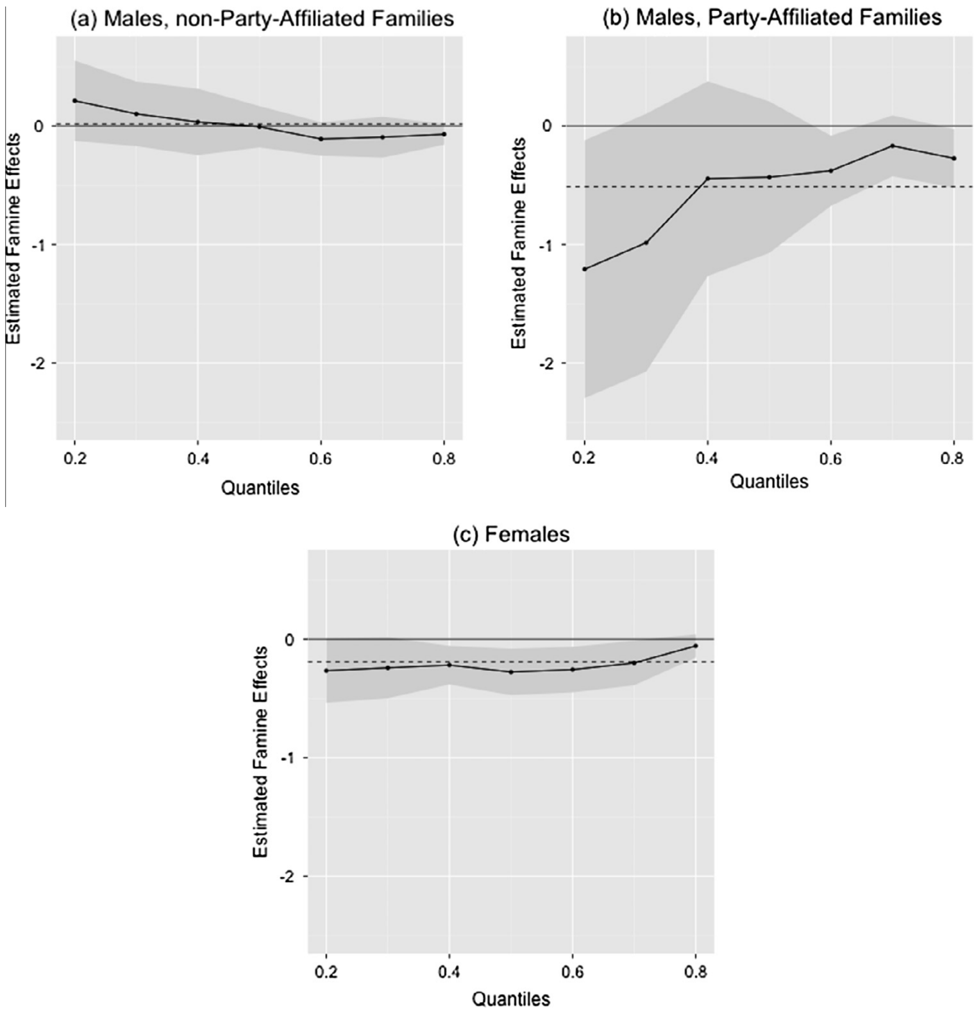

Fig. 2. Quantile regression estimates of famine effects: by gender and family background.

4.4. Long-term famine effects on schooling years and income

For both genders and regardless of parental party membership, being born during the 1959–1961 Chinese Famine had no effects on schooling years or income (Table 3), except marginal evidence that women from non-party-affiliated families were negatively affected in terms of their income (−.040, $p < .1$, Model 3 in Panel B). Therefore, Hypothesis 3 is rejected.

5. Discussion

This paper exploits temporal and geographic variations in the severity of the 1959–1961 Chinese Famine to estimate the long-term health and socioeconomic consequences of early-life under-nutrition. We use a sample of 1716 individuals born in rural China from 1955 through 1966 from a nationally representative dataset, and examine a composite self-perceived general health measure composed of eight dimensions (self-reported health, physical functioning, role-physical, bodily pain, vitality, mental health, social functioning, and role-emotional). Employing the difference-in-differences technique, our findings indicate that exposure to acute under-nutrition early in life had negative effects on health in midlife, especially for males born into party-affiliated families and females regardless of their parents' party membership. Our findings echo the key principles of the life course perspective that emphasize the role of timing and contexts of life experiences in shaping subsequent life chances (Elder et al., 2003; Umberson et al., 2010).

First of all, the timing of famine exposure matters for mid-life health. The Chinese Famine lasted three years, rendering it challenging to conduct a sharp contrast distinguishing the *in utero* effect from the infancy effect (Almond and Currie, 2011).

Table 3
Difference-in-differences estimates of famine effects on schooling years and income, based on OLS regressions.

	Model 1: men, non-party-affiliated families	Model 2: men, party-affiliated families	Model 3: women, non-party-affiliated families	Model 4: women, party-affiliated families
Panel A: schooling years				
Cohort (Ref. = Post-famine cohort)				
Pre-famine cohort	−0.148	4.927	−0.930	0.225
	(1.098)	(3.306)	(0.700)	(4.033)
Famine cohort	0.393	5.428*	−0.307	1.462
	(0.716)	(2.061)	(0.674)	(2.660)
EDR *				
Pre-famine cohort	−0.051	−0.038	−0.060+	−0.073
	(0.053)	(0.185)	(0.030)	(0.098)
Famine cohort	0.004	−0.096	−0.059	−0.064
	(0.069)	(0.170)	(0.035)	(0.106)
Age	−0.010	−0.611	−0.036	−0.205
	(0.110)	(0.456)	(0.112)	(0.459)
Constant	21.240***	33.877+	30.306***	15.979
	(4.378)	(18.667)	(4.902)	(18.521)
Observations	716	75	840	85
R-squared	0.098	0.135	0.154	0.052
Panel B: income (logged)				
Cohort (Ref. = Post-famine cohort)				
Pre-famine Cohort	0.259	−1.590+	−0.216	1.961
	(0.292)	(0.913)	(0.489)	(1.746)
Famine cohort	0.169	−0.555	0.134	−0.754
	(0.217)	(0.418)	(0.245)	(1.563)
EDR *				
Pre-famine cohort	0.010	0.049	−0.033	−0.045
	(0.029)	(0.045)	(0.022)	(0.028)
Famine cohort	0.007	0.053	−0.040+	0.062
	(0.029)	(0.035)	(0.021)	(0.065)
Age	−0.067+	0.003	0.005	−0.195
	(0.037)	(0.085)	(0.045)	(0.176)
Constant	13.640***	9.518*	19.046***	16.125*
	(1.575)	(3.418)	(1.866)	(6.913)
Observations	685	74	776	78
R-squared	0.141	0.187	0.173	0.099

Note: We do not show estimates for the main effect of excess death rate (see footnote 5 in the text). Estimates for province fixed effects are not shown for Models 1 and 3 for simplicity. Robust standard errors clustered at the province level are shown in parentheses.
*** $p < 0.001$.
** $p < 0.01$.
* $p < 0.05$.
+ $p < 0.1$.

Table A
Difference-in-differences estimates of famine effects on midlife perceived health across parental educational groups, based on OLS regressions, Junior High School (JH) as the cutpoint.

	Model 1: men, parents < JH	Model 2: men, parents ⩾ JH	Model 3: women, parents < JH	Model 4: women, parents ⩾ JH
Cohort (Ref. = Post-famine cohort)				
Pre-famine cohort	−1.008	−12.005*	−2.564	3.928
	(1.576)	(4.384)	(1.827)	(5.016)
Famine cohort	−0.557	−6.778**	0.387	1.963
	(1.166)	(2.118)	(0.951)	(3.632)
EDR *				
Pre−famine cohort	−0.019	−0.074	0.113+	−0.628**
	(0.058)	(0.170)	(0.056)	(0.203)
Famine cohort	−0.032	−0.008	−0.198*	−0.259
	(0.067)	(0.107)	(0.074)	(0.276)
Age	−0.015	1.229*	−0.036	−0.281
	(0.154)	(0.483)	(0.213)	(0.645)
Constant	27.639***	−14.496	20.288*	42.689
	(6.152)	(20.031)	(8.945)	(25.919)
Observations	691	86	806	100

Notes: 1. We do not show estimates for the main effect of excess death rate (see footnote 5 in the text). Estimates for province fixed effects are not shown for Models 1 and 3 for simplicity. Robust standard errors clustered at the province level are shown in parentheses.
2. Postestimation tests of the EDR * Famine-Cohort interaction terms across models: M3 = M4 = M5 = M6 (p = .4026); M3 – M4 = M5 – M6 (p = .7737); M3 = M4 (p = .8517); M5 = M6 (p = .8316).
*** $p < 0.001$.
** $p < 0.01$.
* $p < 0.05$.
+ $p < 0.1$.

Table B
Difference-in-differences estimates of famine effects on mid-life perceived health, based on OLS regressions, parents' schooling years as an additional covariate.

	Model 1: men	Model 2: women	Model 3: men, non-party-affiliated families	Model 4: men, party-affiliated families	Model 5: women, non-party-affiliated families	Model 6: women, party-affiliated families
Cohort (Ref. = Post-famine cohort)						
Pre-famine cohort	−2.080	−2.339	−1.609	−8.370*	−2.745	−1.147
	(1.617)	(1.507)	(1.574)	(3.949)	(1.648)	(6.127)
Famine cohort	−1.297	0.581	−1.135	−3.994	0.646	−1.159
	(1.122)	(0.788)	(1.200)	(2.687)	(0.884)	(4.524)
EDR *						
Pre-famine Cohort	−0.023	0.083	−0.037	−0.069	0.090	0.112
	(0.060)	(0.068)	(0.053)	(0.148)	(0.085)	(0.133)
Famine cohort	−0.021	−0.208**	0.016	−0.522***	−0.223**	0.015
	(0.061)	(0.067)	(0.089)	(0.128)	(0.079)	(0.218)
Age	0.113	−0.023	0.078	0.835	−0.001	−0.069
	(0.162)	(0.183)	(0.162)	(0.557)	(0.198)	(0.696)
Schooling years of parents	0.139*	0.154*	0.132	0.123	0.133*	0.198
	(0.065)	(0.063)	(0.082)	(0.159)	(0.056)	(0.141)
Constant	22.034**	18.577*	23.332**	1.708	18.203*	33.194
	(6.486)	(7.780)	(6.508)	(23.451)	(8.381)	(28.039)
Observations	777	906	702	75	822	84

Notes: 1. We do not show estimates for the main effect of excess death rate (see footnote 5 in the text). Estimates for province fixed effects are not shown for Models 1–3 or Model 5 for simplicity. Robust standard errors clustered at the province level are shown in parentheses.
2. Postestimation tests of the EDR * Famine-Cohort interaction terms across models: M3 = M4 = M5 = M6 (p = .0107); M3 – M4 = M5 – M6 (p = .0435); M3 = M4 (p = .0021); M5 = M6 (p = .3601).
+ $p < 0.1$.
*** $p < 0.001$.
** $p < 0.01$.
* $p < 0.05$.

Nevertheless, we show that those born before the famine were not affected by the famine in any significant way in terms of their health, and among women, the lasting famine effects on health were concentrated in the 1961 and 1962 birth cohorts. The Chinese Famine peaked in 1960 and remained severe in 1961 (Peng, 1987), meaning that the 1961 and 1962 birth cohorts were mostly *in utero* during the harshest part of the famine, leaving markers on their health several decades later. Indeed, this finding is particularly remarkable if we take into account (1) mortality selection, that these two birth cohorts (especially the 1961 birth cohort) should be composed of the fittest babies, and (2) that some of the 1962 birth cohort were

Table C
Difference-in-differences estimates of famine effects on mid-life perceived health, based on OLS regressions, using different famine severity measures.

	Model 1: men	Model 2: women	Model 3: men, non-party-affiliated families	Model 4: men, party-affiliated families	Model 5: women, non-party-affiliated families	Model 6: women, party-affiliated families
Panel A: excess death rate (mortality difference between 1960 and non-famine years 1956–1958)						
Cohort (Ref. = Post-famine cohort)						
Pre-famine cohort	−1.819	−1.693	−1.543	−8.259^{+}	−2.122	0.237
	(1.472)	(1.470)	(1.425)	(3.972)	(1.623)	(5.912)
Famine cohort	−1.439	0.726	−1.497	−4.973*	0.816	−1.579
	(1.139)	(0.837)	(1.277)	(2.281)	(0.968)	(4.722)
EDR *						
Pre-famine cohort	−0.033	0.028	−0.031	−0.081	0.032	0.049
	(0.027)	(0.040)	(0.032)	(0.124)	(0.047)	(0.101)
Famine cohort	−0.011	−0.106*	0.020	−0.192***	−0.116*	0.064
	(0.045)	(0.048)	(0.063)	(0.038)	(0.049)	(0.133)
Age	0.101	−0.102	0.069	0.900	−0.071	−0.286
	(0.157)	(0.183)	(0.156)	(0.531)	(0.194)	(0.691)
Constant	24.476***	26.199**	25.758***	−0.384	25.272**	43.922
	(6.344)	(7.747)	(6.373)	(22.299)	(8.147)	(27.765)
Observations	791	925	716	75	840	85
Panel B: cohort size shrinkage index						
Cohort (Ref. = Post-famine cohort)						
Pre-famine cohort	−1.625	−1.822	−0.830	−7.671	−2.045	−3.653
	(1.913)	(2.198)	(1.619)	(7.171)	(2.637)	(7.507)
Famine cohort	−1.189	4.119*	−1.925	1.583	4.875*	−3.529
	(2.237)	(1.971)	(2.744)	(5.211)	(2.193)	(8.432)
Cohort size shrinkage index *						
Pre-famine cohort	−0.157	0.136	−0.268	−0.128	0.104	1.103
	(0.300)	(0.588)	(0.296)	(1.486)	(0.695)	(0.853)
Famine cohort	−0.095	−1.216*	0.184	−2.180^{+}	−1.407*	0.729
	(0.488)	(0.532)	(0.648)	(1.127)	(0.556)	(1.656)
Age	0.095	−0.096	0.062	0.712	−0.067	−0.256
	(0.161)	(0.183)	(0.160)	(0.562)	(0.193)	(0.683)
Constant	33.274***	36.779***	34.319***	6.428	35.630***	42.716
	(6.646)	(8.055)	(6.888)	(23.750)	(8.359)	(27.033)
Observations	791	925	716	75	840	85
Panel C: fertility reduction index						
Cohort (Ref. = Post-famine cohort)						
Pre-famine Cohort	−1.775	−2.447	−1.319	−5.220	−3.116	−0.748
	(1.736)	(2.114)	(1.513)	(7.772)	(2.634)	(6.661)
Famine cohort	−0.589	4.278^{+}	−1.603	3.248	4.887^{+}	−3.150
	(2.222)	(2.288)	(2.858)	(4.037)	(2.419)	(9.719)
Fertility reduction index *						
Pre-famine cohort	−0.122	0.327	−0.158	−0.882	0.406	0.563
	(0.270)	(0.611)	(0.292)	(1.911)	(0.720)	(1.067)
Famine cohort	−0.260	−1.373^{+}	0.125	−2.843**	−1.551*	0.734
	(0.582)	(0.708)	(0.789)	(0.802)	(0.712)	(2.051)
Age	0.089	−0.100	0.058	0.753	−0.068	−0.331
	(0.161)	(0.185)	(0.160)	(0.558)	(0.194)	(0.689)
Constant	54.890***	78.273***	56.144***	3.826	76.142***	44.961
	(6.615)	(7.417)	(6.710)	(23.168)	(7.809)	(27.125)
Observations	791	925	716	75	840	85

Note: We do not show estimates for the main effect of excess death rate (see footnote 5 in the text). Estimates for province fixed effects are not shown for Models 1–3 or Model 5 for simplicity. Robust standard errors clustered at the province level are shown in parentheses.
*** $p < 0.001$.
** $p < 0.01$.
* $p < 0.05$.
$^{+}$ $p < 0.1$.

not exposed to the famine *in utero*. This finding lends support to the critical period notion that early life conditions have long-lasting effects on or even "program" later health (Barker, 1995, 1998). That famine effects for females become non-significant once the 1958 cohort is added also fits well with the critical period model, given that the majority who were born

Table D
Difference-in-differences estimates of famine effects on mid-life perceived health, based on OLS regressions, using individual birth years.

	Model 1: men, non-party-affiliated families	Model 2: men, party-affiliated families	Model 3: women, non-party-affiliated families	Model 4: women, party-affiliated families
Birthyear 1955	2.787	−0.981	1.101	7.611
	(3.177)	(5.886)	(3.507)	(10.203)
Birthyear 1956	−0.528	−6.056	−1.429	−3.893
	(2.872)	(5.992)	(3.021)	(6.797)
Birthyear 1957	0.833	−9.841	−0.155	8.400
	(2.767)	(6.461)	(2.563)	(8.050)
Birthyear 1958	1.420	−4.998	−2.574	2.358
	(2.521)	(5.961)	(2.374)	(6.835)
Birthyear 1959	0.454	8.059*	1.527	
	(2.253)	(3.765)	(2.089)	
Birthyear 1960	−0.243	−2.712	0.328	8.040
	(2.285)	(4.086)	(2.374)	(5.374)
Birthyear 1961	−1.684	−2.345	1.886	5.785
	(2.717)	(3.235)	(1.725)	(3.487)
Birthyear 1962	0.161	−11.327+	2.614+	−5.068
	(1.338)	(5.988)	(1.293)	(5.135)
EDR * Birthyear 1955	−0.053	−0.150+	0.038	0.302
	(0.056)	(0.080)	(0.089)	(0.824)
EDR * Birthyear 1956	0.099	−0.001	0.148+	0.660
	(0.103)	(0.133)	(0.083)	(0.445)
EDR * Birthyear 1957	−0.034	−0.048	0.051	−0.477
	(0.128)	(0.262)	(0.102)	(0.912)
EDR * Birthyear 1958	−0.219*	0.192	0.182	0.264
	(0.088)	(0.521)	(0.171)	(0.195)
EDR * Birthyear 1959	−0.088	−1.943***	−0.131	−0.305
	(0.092)	(0.296)	(0.082)	(0.464)
EDR * Birthyear 1960	0.201+	−0.550**	−0.028	−0.256
	(0.113)	(0.154)	(0.090)	(0.247)
EDR * Birthyear 1961	0.254	−0.908	−0.401+	
	(0.263)	(0.542)	(0.211)	
EDR * Birthyear 1962	−0.030	1.230	−0.349**	0.435
	(0.098)	(0.858)	(0.096)	(0.322)
Age	−0.292	0.504	−0.304	−0.945
	(0.311)	(0.653)	(0.344)	(0.816)
Constant	39.020**	15.514	30.596*	69.943*
	(12.982)	(26.950)	(14.744)	(32.726)
Observations	716	75	840	85

Note: We do not show estimates for the main effect of excess death rate (see footnote 5 in the text). Estimates for province fixed effects are not shown for Models 1 and 3 for simplicity. Robust standard errors clustered at the province level are shown in parentheses.
*** $p < 0.001$.
** $p < 0.01$.
* $p < 0.05$.
+ $p < 0.1$.

in 1958 should have spent most of their *in utero* period in 1957, thus diluting any long-term consequences of early-life insults. A recent study by De Rooij et al. (2010) found that the most critical period of exposure for a fetus was during the first three months of a pregnancy, which explains why even a few months' exposure (or non-exposure) to the famine had made such a large difference for later health for females in the 1962 (or 1958) birth cohort.

In addition, long-term famine effects are contingent on social contexts. In this study we uncover two important contexts: gender and family background. Consistent with previous studies, females seem to bear the brunt of negative famine consequences. Mortality selection and differential treatment are two potential explanations. Our data, a truncated sample only consisting of individuals who were still alive in 2005, are not suitable to quantify the extent of selectivity, although the finding that famine effects vary little across quantiles among women but not among men hints the sex-specific mortality selection pattern. Other studies also provide some insights. Mu and Zhang (2011) exploited heterogeneity in the culture of son preference across nineteen large ethnic groups and showed that mortality selection was the driving force for the observed gender difference in the impact of the Chinese Famine on disability rates. Differential treatment is also possible, especially in rural China, where sons were treated more favorably than daughters. Unfortunately, we cannot test this hypothesis directly given that the 2005 CGSS did not collect information that could proxy for family resource allocation (e.g., sibling compositions). Nevertheless, we want to mention that prenatal sex identification through ultrasound B machines was not possible before the 1980s in China (Zeng et al., 1993), so females born during the Chinese Famine were protected from unfavorable treatment at least *in utero*, and any gender differences resulting from differential treatment are most likely due to parental neglect after birth rather than *in utero*. In practice, both mortality selection and differential treatment were probably at work, as evidenced by our finding that females but not males are sensitive to alternative cohort construction. Specifically, females

Table E
Difference-in-differences estimates of famine effects on mid-life perceived health, based on OLS regressions, 1958 in the famine cohort.

	Model 1: men	Model 2: women	Model 3: men, non-party-affiliated families	Model 4: men, party-affiliated families	Model 5: women, non-party-affiliated families	Model 6: women, party-affiliated families
Cohort (Ref. = Post-famine cohort)						
Pre-famine cohort	−1.043	−0.407	0.101	−13.829*	−0.180	1.282
	(1.711)	(1.487)	(1.759)	(6.623)	(1.717)	(5.847)
Famine cohort	−0.536	0.292	−0.140	−6.422+	0.661	−2.083
	(1.090)	(0.953)	(1.210)	(3.432)	(1.081)	(3.807)
EDR *						
Pre-famine cohort	0.027	0.059	0.013	−0.006	0.075	−0.322
	(0.067)	(0.055)	(0.052)	(0.166)	(0.071)	(0.434)
Famine cohort	−0.074	−0.099	−0.042	−0.450**	−0.126	0.141
	(0.059)	(0.078)	(0.079)	(0.139)	(0.093)	(0.187)
Age	−0.066	−0.241	−0.172	1.277+	−0.290	−0.089
	(0.195)	(0.163)	(0.210)	(0.714)	(0.182)	(0.554)
Constant	29.818**	28.816***	34.482**	−15.506	31.264***	35.185
	(8.755)	(6.776)	(9.418)	(29.366)	(7.456)	(22.395)
Observations	791	925	716	75	840	85

Notes: 1. We do not show estimates for the main effect of excess death rate (see footnote 5 in the text). Estimates for province fixed effects are not shown for Models 1–3 or Model 5 for simplicity. Robust standard errors clustered at the province level are shown in parentheses.
2. In this table, famine cohort includes respondents born between 1958–1962, pre-famine cohort includes people born 1955–1957, and post-famine cohort includes those born between 1963 and 1966.
3. Postestimation tests of the EDR*Famine-Cohort interaction terms across models: M3 = M4 = M5 = M6 ($p = .0046$); M3 − M4 = M5 − M6 ($p = .0411$); M3 = M4 ($p = .0277$); M5 = M6 ($p = .2940$).
*** $p < 0.001$.
** $p < 0.01$.
* $p < 0.05$.
+ $p < 0.1$.

born in 1958 were at a much higher risk of being neglected during the famine period than males of the same cohort, so adding females of the 1958 birth cohort who survived the famine should exacerbate the extent of mortality selection, attenuating the estimated famine effect toward zero.

Family background is another contextual factor shaping the long-term consequences of famine. A common theme emerging from disaster studies is that social forces operate even in seemingly neutral natural disasters, with less advantaged groups affected most severely (e.g., Strauss and Thomas, 1998). Thus, it is striking that virtually no previous famine research incorporated family background into their analytic framework.

What have we gained by explicitly taking into account family background as an important contextual factor? Along with prior studies we found a null famine effect for men, but when disaggregated by parents' political power, a noticeable adverse famine effect on mid-life health was identified among men from party-affiliated families, significantly greater than that for men from the unaffiliated families. Is it against the general pattern detected in previous literature? Not necessarily. We resort to the complex interplay between biological and social forces to explain this seemingly counterintuitive finding: differential mortality selection by family background. During the famine period, families with tighter ties to the Communist Party were more likely to mobilize resources to obtain more food. Given male's greater biological vulnerability to prenatal insult (Kline et al., 1989; Kraemer, 2000; Salonen et al., 2009), rural males from ordinary peasant families were subject to especially severe mortality selection. Quantile regressions provide suggestive evidence supporting this argument: we observe a declining trend of estimated famine effects from lower to higher quantiles for males born into non-party-affiliated families, suggesting high levels of mortality selection that yielded a very robust group of survivors in the lower quantiles (see the first plot of Fig. 2). Males born into party-affiliated families were adversely affected across the board, but the effect was particularly severe in the lower quantiles (see the second plot of Fig. 2). This seems to indicate that many of the weakest were able to survive, but their health was nevertheless permanently harmed. In contrast to males, we find no significant difference in the negative famine effects between females born into party-affiliated and the unaffiliated families, possibly due to less prenatal mortality selection among females and a widespread preference for sons regardless of family background.

Unlike several other studies, we did not find that the 1959–1961 Chinese Famine had any negative consequences for education or income. We propose possible reasons. Meng and Qian (2009) reported that *in utero* exposure to famine caused an 8.6% (0.6 years) reduction in educational attainment, but this actually refers to the effects estimated on the 90th percentile that they argue as a way to correct for mortality selection. We do not follow this strategy in our analysis. Chen and Zhou (2007) found that the 1959 and 1960 cohorts had lower annual working hours and lower farming income because of the famine. Our results differ from theirs probably because we rely on a nationally representative sample instead of the eight provinces they examined. In still another study, Almond et al. (2007) used census data, but the large sample size almost guaranteed significant findings. The effect sizes reported in their study were small.

This study has several limitations. First, our samples for males and females born into party-affiliated families are small, which might yield the problem of overfitting. Future studies relying on larger samples are essential to addressing this issue. Second, we are unable to distinguish two types of parents' political credential given data limitation: Communist Party membership and cadre status. Although prior studies suggested that these two are highly correlated, to what extent using alternative measures affects our conclusion is a task future studies could attend to. Third, because the 2005 CGSS did not collect birthplace information, we have to rely on parents' occupation when the respondent was 14 to obtain the rural sample. Although the tight administrative control virtually eliminated unauthorized rural-to-urban migration in the pre-reform era (Wu and Treiman, 2004), we are not sure to what extent this measure accurately captures places of birth. Fourth, the long duration of the Chinese Famine makes it impossible to conduct a sharp contrast distinguishing the *in utero* effect from the infancy effect. Studies based on other famines that are short and abrupt might shed more light on such distinctions.

Despite these limitations, our study extends previous research on the long-term consequences of famine in general and of the 1959–1961 Chinese Famine in particular. To our knowledge, no previous studies have used a nationally representative sample to examine the effects of the Chinese Famine on perceived health, an important predictor of subsequent mortality. Moreover, our findings add nuances to previous famine research by revealing who in the population is affected most harshly through the confluences of biological and social forces. Having experienced one of the largest famines in human history, these individuals are now moving through their retirement years. Our study suggests the necessity of paying special attention to this cohort, with programs designed specifically to meet their special needs.

Appendix A

See Tables A–E.

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