LONG-TERM EFFECTS OF THE 1959-1961 CHINA FAMINE: MAINLAND CHINA AND HONG KONG*

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Abstract

This paper estimates the effects of maternal malnutrition exploiting the 1959-1961 Chinese famine as a natural experiment. Using the 1% sample of the 2000 Chinese Census, we find that around age 40, fetal exposure to acute maternal malnutrition has compromised a range of socioeconomic outcomes, including: literacy, labor market status, wealth and marriage market outcomes. Women married men with less education, and men married later, if at all. In addition, maternal malnutrition reduced the sex ratio (males to females) in two generations – those prenatally exposed and their children – presumably through heightened male mortality. This tendency toward female offspring is interpretable in light of the Trivers-Willard (1973) hypothesis, according to which parents in poor condition should skew the offspring sex ratio toward daughters. Hong Kong Natality micro data from 1984-2004 further confirm this pattern of female offspring among mainland-born residents likely exposed *in utero*.

1 Introduction

The monumental policy failure of Mao's Great Leap Forward resulted in mass starvation and the premature death of between 18-30 million people(Ashton *et al.* , 1984; Peng, 1987; Lin, 1990; Yao, 1999; Li & Yang, 2005). This unprecedent Famine, which peaked in 1960, provides a natural experiment in acute nutritional deprivation.

This paper evaluates the long term effects of maternal malnutrition on adult outcomes. Our focus on the prenatal period is motivated by a number of factors. There is growing evidence that fetal health may be particularly important for adult health outcomes, a linkage popularized by Barker (1992). The original formulation of the "fetal origins hypothesis" linked maternal nutritional status to chronic diseases of offspring in adulthood, but fetal health may also impact socio-economic outcomes. The prevalence of maternal malnutrition in many poor countries (e.g., DeRose *et al.* (2000)) makes this relationship one of continuing relevance. Moreover, evidence that poor maternal nutrition impairs health *and* socioeconomic outcomes may help explain the strong relationship between health and socioeconomic outcomes, apparent at both the individual and national levels (Case *et al.* , 2002; Cutler *et al.* , 2006).

Although measures of prenatal nutritional status correlate negatively with adult socioeconomic outcomes, evidence of a causal relationship is surprisingly scarce (see e.g., Rasmussen (2001); Walker *et al.* (2007)). The employment of the Chinese Famine (henceforth the Famine) as a natural experiment reduces the scope for omitted factors that might generate positive associations between measures of fetal health and adult socioeconomic outcomes in the absence of a causal pathway.

Employing the 1% sample of the 2000 Chinese Population Census, we find that education, labor force attachment, as well as outcomes on the marriage market, were compromised for the cohort exposed prenatally to greater famine intensity as measured by the death rate. We also find that greater fetal exposure to the Famine reduced the the cohort's sex ratio, suggesting greater male vulnerability to maternal malnutrition. Perhaps even more interestingly, we find an "echo effect" of the Famine on the next generation: children whose mothers were exposed prenatally also register Famine impacts. In particular, Famine-exposed mothers are more likely to give birth to daughters. To our knowledge, ours is the first study to trace the offspring sex ratio to the *in utero* environment of the parent.

While a narrow age window reduces the problem of confounders, our "control" groups were not untouched by the Famine: the older cohorts experienced it directly as infants and toddlers and the younger cohorts were the children of famine survivors. The difficulty in finding an appropriate control group motivates two additional approaches. One, we utilize geographic variation in Famine severity to generate comparisons *within* birth cohorts. Two, while the Famine was endemic in mainland China, Hong Kong, then a British colony, was untouched. The Famine resulted in a large inflow of mainland Chinese into Hong Kong. Since Hong Kong Natality data include information on the birth place of the parents, we can compare mainland born and Hong Kong born mothers in these data. Results from both approaches are consistent with the Famine having negatively impacted a broad spectrum of outcomes.

The remainder of the paper is organized as follows. Sections 1.1-1.3 describe the background of the Famine and reviews the previous Famine literature. Section 2 describes the 2000 Chinese Population Census and the 1984-2004 Hong Kong Natality files. Section 3 reports descriptive results, and Section 4 reports regression results examining the long-term effect of the Famine. Section 4.3 presents results from the Hong Kong natality data. Section 5 concludes.

1.1 Famine background

The Famine began in the fall of 1959 and impacted all regions of China. Grain output dropped 15% in 1959 and another 15 percent in 1960 (Li & Yang, 2005, page 846). By 1962, birth and death rates had returned to normal levels.

While weather conditions contributed to the Famine, the radical agricultural policies of the Great Leap Forward were chiefly to blame (Lin, 1990; Li & Yang, 2005). In a breakneck attempt to overtake Britain and eventually the U.S., labor was diverted from agriculture to industry while grain procurement from rural areas was increased. At the same time, collectivization of agricultural production resulted in shirking and falling productivity (Lin, 1990). The political climate encouraged provincial leaders to overstate grain production. Despite widespread starvation, China was a net exporter of grain throughout 1960 (Lin & Yang, 2000).

Famine intensity varied by region (Peng, 1987). Rural death rates rose to 2.5 times pre-Famine levels. Urban residents fared better but were not spared, death rates in the peak year 1960 were 80% above pre-Famine levels (China Statistical Press, 2000). Central provinces such as Anhui, Henan and Sichuan were the worst hit, while northeastern provinces such as Heilongjiang and Jilin were relatively spared. By 1961, death rates had returned to normal in more than half of the provinces, but remained high in, for instance, the southern provinces Guangxi and Guizhou (close to Hong Kong).

1.2 Famine studies: Epidemiology

The best epidemiological evidence to date linking maternal nutritional deprivation to subsequent adult outcomes derives from the cohort *in utero* during the 1944-45 Dutch famine. While the seminal study found limited effects at age 18 (Stein *et al.* , 1975), at middle age, this cohort exhibited a broad spectrum of effects including: self-reported health (Roseboom *et al.* , 2001b), coronary heart disease morbidity (Roseboom *et al.* , 2001b; Bleker *et al.* , 2005), and adult antisocial personality disorders (Neugebauer *et al.* , 1999). These, and studies of the 1866-1868 Finnish Famine and the Nazi Seige of Leningrad, have focused exclusively on health outcomes. Epidemiological findings from the Chinese Famine include heightened risk of schizophrenia from prenatal exposure (Clair *et al.* , 2005) and overweight of women (Luo *et al.* , 2006).

1.3 Famine studies: Economics

A number of recent studies evaluating the Famine's impact on the socioeconomic outcomes of survivors have used the the China Health and Nutrition Surveys (CHNS) (Chen & Zhou, 2007; Meng & Qian, 2006; Gorgens *et al.* , 2005). The CHNS is a panel dataset that begun in 1989 of health and economic outcomes of approximately 4,000 Chinese households from nine provinces (out of 31 provinces or province level administrative regions). The small sample size combined with the collapse of fertility during the Famine necessitates the inclusion of ages well after birth as "treated."[1] However, broad "early childhood" hypotheses makes it difficult to reject alternative explanations. The possibility then that events at other ages – for instance the Cultural Revolu-

[1] 171 rural CHNS respondents and 107 urban respondents were born 1960-61 (Chen & Zhou, 2007) table 2.

tion and the forced "rustification" of students in outlying areas – confounds results is a concern.

Chen & Zhou (2007) considered those up to age 6 as treated, and found that those born in 1955, 1957, 1959, 1960 and 1962 were stunted, with the largest height reductions for the 1959, 1960, and 1962 birth cohorts. In an analogous specification, they found reduced labor supply of those born in 1959 and 1960, and lower wealth as measured by the size of residence for birth cohorts 1958 and 1959.

Meng & Qian (2006) considered the following birth cohorts as potentially affected: 1952-54, 1955-58, 1959-60, with cohorts born 1961-64 as the reference group. Using reductions in cohort size as a proxy for Famine severity (assumed to occur through Famine mortality), their OLS estimation returned mixed results, and little evidence for a particularly strong effect for the 1959-60 cohort. Instrumenting for cohort size, using per capita grain production in 1997, they found a small negative effects on education, but a substantial (25%) reduction in hours worked for the 1959-60 cohort.

Gorgens *et al.* (2005) studied adult heights of cohorts exposed to the Famine in childhood using the CHNS data. They argued both that children who survived the Famine did not show any stunting and that stunting did occur. They reconcile these two arguments by a third: Famine mortality was concentrated among shorter people. The net effect of stunting and selection, the authors argued, made the height of survivors appear unchanged. However, the claim that no stunting is observed among survivors is controversial (Yan, 1999; Chen & Zhou, 2007; Morgan, 2006).

2 Data

Our main data set is the 1% sample of the 2000 Chinese Population Census.[2] The 1% sample covers more than 11 million records, and, to our knowledge, has not been used to evaluate long-term effects of the Famine.[3] These data include information on education, labor market status and residence of respondents. Demographic information includes sex, birth year and month, marriage and fertility information. Unlike preceding Census surveys and the CHNS data, the 2000 Census records the province of birth, eliminating the potential for confounding due to internal migration. The 2000 census captures Famine cohorts near age 40, and therefore near the flat portion of their occupation and earnings profile. Moreover, for women, fertility was close to complete. Death rates peaked in 1960 but were elevated in 1959 and 1961 as well, see table 1. We use those born 1956-1964, a sub-sample which covers three pre-Famine years and three post-Famine years. Our relatively narrow birth interval is intended to increase the similarity of the unobserved, later-life factors and their effects on Census outcomes.

Our second data set is the The Hong Kong Natality microdata (1984-2004) files, derived from the universe of birth certificates. These data include information on maternal country of birth. Restricting the sample to mothers either born in mainland China or Hong Kong in the years 1957-1965 yields some 600,000 records, approximately one-third of whom emigrated from the Mainland. The Hong Kong data provide an important control group since all of Mainland China (rural and urban) was afflicted by the Famine.

[2] Conducted by the Chinese National Bureau of Statistics for Mainland China.

[3] Shi (2006) used a 0.1% sub-sample.

2.1 Measuring the Famine

We use two measures of famine intensity: death rates and average month of birth.

Death rates We use the (all-age) death rate (China Statistical Press, 2000) by year and province to calculate two proxies for famine intensity. We have data for 29 out of the 31 provinces (or province level divisions).

- First, for every person, we calculate the weighted average of the death rate in the province of birth for the duration of the fetal period, henceforth "weighted death rate" or **wdr**. For example, a person born in January 1960 in Beijing is assigned 1/9th of Beijing's 1960's mortality rate and 8/9th of Beijing's 1959's mortality rate. This weighted death rate ranged from 0.0054 to 0.0686 (per person). The lowest within year range was for 1956, when the **wdr** ranged from 0.0066 to 0.0197, a difference of 0.013. The greatest within year range was for 1960, when the **wdr** ranged from 0.0069 to 0.0686, a difference of 0.0617. The max difference-in-differences was thus 0.049.

- Second, we collapse this weighted death rate by month and year of birth, thus calculating a population weighted national average for each month and year, henceforth "aggregate weighted death rate" or **awdr**. During the study period, this measure ranged from 0.010 to 0.022, a difference of 0.012. Thus measured, those born towards the end of 1960, early 1961, were exposed to the highest mortality, Figure 1.

Average month of birth For Hong Kong, we cannot rely on mortality data since immigrants were a highly selected group, both geographically (the Famine hit bordering provinces later) and due to the particular migration policies in place (further described in Section 4.3). To obtain a proxy for when the Famine peaked for this group of immigrants, we use average month of birth. The motivation for this proxy is that for the northern hemisphere, Famines tend to be most severe during the winter months. This reduces fertility disproportionately in the later half of the year, thereby lowering the average month of birth (Stein *et al.* , 1975).[4] This proxy indicates 1961 as the worst Famine year for Hong Kong mothers born on the mainland, Figure 3. There was no corresponding change for Hong Kong born mothers, Figure 4. For comparison, this method indicates 1960 as the worst year for mainland China, Figure 2 (consistent with the mortality data).

3 Descriptive Results

We start by presenting outcomes by date of birth in a series of figures. These figures suggest that those born in 1960 had worse socio-economic outcomes. Recall that this cohort was *in utero* during the period with the highest death rate as measured by the weighted death rate (Figure 1).

In 2000, the 1960 and flanking birth cohorts were more likely to not be working, be supported by other household members, live in smaller homes and less likely to head their households, (Figure 5).

A final descriptive pattern of note is that women born around 1960 bore more female children and were more likely to have lost a child, Figures 6 and

[4]Authors' tabulation of appendix table 4 data in (Stein *et al.* , 1975).

7.

4 Regression Results

To investigate whether higher prenatal famine intensity compromised adult outcomes, we focus on the birth cohorts 1956-1964, and estimate:

$$y_{it} = \beta_0 + \theta \cdot \mathbf{awdr}_t + \beta_1 \cdot \texttt{YOB} + \beta_2 \cdot \texttt{YOB}^2 + \beta_3 \cdot \texttt{YOB}^3 + \lambda_{province} + \varepsilon_{it}, \quad (1)$$

where y_i denotes the outcome for individual i. $\mathbf{awdr}_t$ denotes the aggregate weighted death rate by birth year and month of birth t,[5] YOB denotes birth year, $\lambda_{province}$ denotes a vector of province dummies. θ measures the departure of outcomes for birth cohorts exposed prenatally to greater famine intensity as proxied by the death rate. Thus (1) allows for a flexible cohort profile within a narrowly-defined birth interval, and assesses whether the prenatal death rate contributes additional explanatory power, as reflected by $\hat{\theta}$. We estimate equation (1) separately for men and women.

Results from estimating (1) for 2000 census outcomes are reported in Tables 1-3. Table 1 shows a consistent, negative effect of prenatal Famine exposure on labor market outcomes. Greater famine intensity is associated with higher likelihood of being illiterate and not working. During the Famine, **awdr** increased by 1.2 percentage points, implying, e.g., that the most Famine exposed cohorts were 7.5% ($0.5052 \times 0.012/0.081$) [women] and 9% ($0.1585 \times 0.012/0.021$) [men] more likely to be illiterate; 3% ($0.4714 \times 0.012/0.189$) [women] and 5.9% ($0.4017 \times 0.012/0.082$) [men] more likely to not work; and, for women, 13% ($0.0448 \times 0.012/0.004$) higher risk of disability.

[5] While this variable has no regional variation, it was constructed using information on the province (and month) of Census respondents' birth in weighting the mortality data , see Section 2.1.

The census does not have any direct measure of earnings, but there is information on housing, which may serve as a proxy for income and wealth. Thus measured, greater fetal Famine exposure also reduced adult economic status.

We also estimate equation (1) using the marriage market variables as outcomes (Table 2). While marriage was nearly universal for women, inspection of who they married reveals that Famine exposed women married men with less education. For men, both the extensive and intensive margins were affected. Men were 6.5% ($0.4902 \times 0.012/0.09$) more likely to be unmarried and were 8.2% ($0.2676 \times 0.012/0.039$) more likely to have never married. Moreover, they married at older ages (1.5 months) and were 0.7 % ($0.5145 \times 0.012/0.87$) less likely to head their household.

Prenatal famine exposure also raised male (relative to female) mortality as evidenced by survival around age 40. The most exposed cohort was 3% ($1.3147 \times 0.012/0.51$) more female (Table 3, column 1). The most striking finding, however, is that prenatally exposed women bore more girls, the most Famine exposed a a 0.4%-age point (0.3194×0.012) lower sex ratio (column 2).[6] To anticipate results, the Hong Kong data (derived from birth certificates) will corroborate this pattern.

4.1 Geographic variation in Famine intensity

A second test of our hypothesis is to isolate the geographic variation in the Famine and make comparisons exclusively within birth cohorts. This approach reduces the potential for confounding from later-life events with age-specific

[6]Similar results are obtained when the logit transform of the proportion of male children is the dependent variable.

effects (e.g. if the Cultural Revolution delayed school entry among seven-year olds). To that end, we evaluate whether regional variation in the Famine predicts the magnitude of damage. We estimate:

$$y_{itj} = \beta_0 + \theta \cdot \mathbf{wdr}_{tj} + \gamma_{yob} + \lambda_{province} + \varepsilon_{itj}, \tag{2}$$

where θ is the parameter of interest, t denotes year of birth and j the province of birth. The mortality rate is the weighted death rate (**wdr**) previously described for the individual's birth date (year and month) and province of birth. As in equation (1), we include vectors of province of birth dummies ($\lambda_{province}$), and, as the goal is to isolate the geographic variation in health induced by the famine, we absorb the average differences for each birth year by including a vector of year of birth dummies (γ_{yob}).

Results from estimating (2) provide qualitatively distinct evidence that regional differences in malnutrition impact outcomes in adulthood (Tables 4-6). For post-Famine events to bias estimates from (2), they would have to mimic the geographic variation of the Famine (and disproportionately impact the same *in utero* cohorts to constitute an alternative explanation for estimates from both (1) and (2)).

To start with Table 4, for women, the coefficient on the weighted death rate (**wdr**) is positive and significant for "Disabled" and and negative and significant for residential area. For men, all outcome variables were of the expected signs and statistically significant.

The lowest within year mortality range was for 1956, when the **wdr** ranged from 0.0066 to 0.019667, a difference of 0.013. The greatest within year range was for 1960, when the **wdr** ranged from 0.0069 to 0.0686, a difference of 0.0617. The max difference-in-differences was thus 0.049. Attributing this increase in the intra-year difference (across provinces and birth months) to the Famine, our estimates imply that the most Famine-exposed women were about 50% ($0.049 \times 0.0418/0.004$) more likely to be disabled, and the corresponding figure for men was 47% ($0.049 \times 0.0585/0.006$). As for housing, the worst exposed women lived in houses that were 2.8 (0.049×58.95) square-meters smaller than the least affected, and the corresponding figure for men was 2.6 square meters. Highest famine exposure was also associate with a 30% increase in the risk of being illiterate and a 10% increase in the risk of not working (for whatever reason).

Men from high-Famine areas were less likely to be married (15%), more likely to never have married (20%), married older (3.3 months), and were less likely to head their household (3%) (Table 5). For women, the point estimates are of the same signs as those for men, but are insignificant.

By contrast, the sex ratio results are stronger in the cross-section and indicate that areas with greater famine intensity had a more female population both measured by survivors in 2000 and their offspring (Table 6). The estimated effect sizes are rather large, where prenatal exposure to the greatest Famine intensity is associated with a 0.8%-age point reduction in the sex ratio (sons to daughters).

Rural versus urban We also estimate the above models separately for those born in rural versus urban regions. We find a Famine effect on the labor and marriage market outcomes for both, although the effects for the rural sample were larger (presumably reflecting the greater severity of the Famine in rural areas). For both rural and urban areas, we find that the Famine reduced the sex ratio of the *in utero* cohort and again in the next generation (results available from authors).

Province of residence Finally, we note that estimates reported in Tables 1-6 are essentially unchanged when fixed effects for the 2000 province of residence are included along with the province of birth dummies.

4.2 Potential Biases

As the Famine both raised mortality and reduced fertility, Famine cohorts were approximately 25-50% smaller than neighboring cohorts in the 2000 Census. To the extent that Famine-induced mortality was negatively selective, as would seem most plausible (especially insofar as health is concerned), estimates of damage to survivors are biased toward zero.

Negative selection into fertility is a greater potential concern, since this could generate the appearance of effects absent any true damage. However, historical evidence suggests that the Famine, unlike the subsequent Cultural Revolution, hit poorer individuals the hardest, e.g., Cai & Wang (2005). The Dutch Famine provides further evidence: fathers of children conceived in the winter of 1944-45 were more likely to have non-manual occupations (Stein *et al.* , 1975).

Direct evidence on selection into fertility is available from the China Fertility surveys (conducted in 1985 and 1988), which include information on the respondent's mother's educational attainment (further information in the Appendix). Plotting the share of women whose mothers had no education, primary or less, secondary or more, or who did not know their mother's education, the 1959-61 birth cohorts do not appear any worse than adjacent cohorts (Figure 8). If anything, the 1959-61 birth cohorts had mothers with more education than adjacent cohorts.

Cohorts born during 1962 to 1964 did not directly experience the Famine, and in this respect may constitute a better control group for the Famine exposed cohorts than those born in the 1950s (who were exposed to higher mortality rates and malnutrition in childhood). Re-estimating equations (1) and (2) on the sample restricted to birth cohorts 1959-1964, we obtain similar, if not slightly stronger, results (available on request).

Another possible source of bias is that those born during famines may be born to inherently more fecund women or parents who have a stronger desire for many offspring, and possibly, less interest in the quality of each individual child. Whereas we cannot control for parental preferences (other than note as above, that the maternal education of the Famine cohorts was if anything better than that of adjacent cohorts), we can investigate sibship size using a recent survey: The 2005 Urban Chinese Education and Labor Survey conducted by the Ministry of Education in 12 cities in China, covering some 10,000 households.[7] The 1959-61 cohorts do not appear to have more siblings, as shown by Figure 9. Rather, these birth cohorts are on trend (linear and decreasing in year of birth).[8]

[7]The 2000 census does not have information on sibship size. Neither can it be inferred from the relationship variable for a household, since most adult siblings live in different households. Finally, the earliest publicly available Chinese census was conducted in 1982, when the 1959-1961 cohorts were in their early 20s.

[8]This is confirmed by a regression of sibsize on a dummy for birth cohorts 1959-61, controlling for a linear trend in birth year. The coefficient on this dummy is about zero, with a very large standard error (not reported).

4.3 Hong Kong: An Unaffected City

A shortcoming of the analysis using the 2000 (mainland) census is the want of a truly unexposed control group. The Famine struck all parts of China, rural and urban. Moreover, while adjacent cohorts offer the closest comparison, these cohorts were also likely exposed: either directly as infants or toddlers or indirectly in that they were born into families with recent Famine exposure.

Hong Kong Natality data offer a potential solution to this problem. Communist China severely restricted out-migration. This policy was, however, temporarily and dramatically suspended during a six-week period in the spring of 1962 when a large number of mainlanders entered Hong Kong (Burns, 1987). Among the refugees were mainland born children, who themselves show up as parents in the 1984-2004 Hong Kong Natality files.

The migration of mainland residents to Hong Kong, during and in the years after the Famine, provides a common environment for those affected by the Famine (mainland immigrants) and those who were not (Hong Kong born).

The Hong Kong Natality microdata allow us to focus on second generation birth outcomes, specifically low birth weight and sex of offspring. Low birth weight may be a negative outcome because it is a correlate of poor adult health and economic performance. As for sex of offspring, a daughter may not be a poor outcome. Still, it may signal poor parental condition, see Section 5.

We estimate a modified version of equation (1) separately on the sub samples of mainland born and Hong Kong born mothers giving birth in Hong Kong 1984-2004. Based on the dramatically lower average month of birth among mainland born mothers born in 1961 (Figures 3 and 4), we substitute the dummy variable `I(1961)`, which takes on the value 1/100 for those born in 1961, for the death rate (**awdr**).

A dummy for the sex of the child is also included when the dependent variable is birth weight since males are on average heavier than females. The birth interval is shifted forward one year from the mainland Census regressions, that is, we focus on births to parents themselves born 1957-1965. Furthermore, we restrict the sample to singleton births. We find that mothers born in 1961 were 8% (0.247/0.030) more likely to give birth to a child of low birth weight (less than 2,500 grams) and 1.2 percent (0.00629/0.52) less likely to give birth to a son than mothers born in adjacent years (Table 7). No significant effects were detected for the Hong Kong born mothers, despite their greater numbers.

5 Summary and Discussion

Observing Famine survivors near age 40, we find that maternal malnutrition compromised a range of census outcomes. Higher famine intensity – by virtue of either time or place of birth – was associated with greater risk of being illiterate, out of the labor force, marrying later (men), and marrying spouses with less education (women). Furthermore, wealth also appeared compromised (as proxied by size of residence). This broad spectrum of damage is observed in the absence of negative selection into childbearing during the Famine, as measured by maternal education. To the extent that selection into childbearing during famine was in fact positive (as maternal education and previous research would suggest), our estimates of long-term damage are conservative. As the nutritional status of pregnant women continues to lag other demographic groups in many developing countries, our findings have implications for the targeting of policies designed to improve health and economic outcomes.

Perhaps the most intriguing result is that Famine exposure lowered the sex ratio of not only the first, but also the second generation. Prenatally exposed women were themselves more likely to bear daughters. This pro-female effect is all the more noteworthy given the well documented prevalence of son preference in mainland China. Famine-induced reductions in the sex ratio are consistent with empirical work finding lower sex ratios for unmarried or poorly educated mothers (Almond & Edlund, n.d.). While the magnitude of the Famine's effect on the sex ratio may appear small, it was several times the found effect size of marital status in U.S. natality data (Almond & Edlund, n.d.) and similar to differences between mothers reporting living with a partner around the time of conception in survey data and those who did not (Norberg, 2004).

Trivers & Willard (1973) proposed that evolution would favor parental ability to vary the sex ratio of offspring according to condition: parents in poor condition would favor daughters and parents in good condition would favor sons. Their argument was based on the observation that while the average number of offspring to males and females equalize, the reproductive success of a male offspring tends to be more resource-sensitive. Maternal malnutrition has been observed to correlate with more female births, e.g., Andersson & Bergström (1998). Pathways include heightened rates of male fetal deaths, as was found to be the case during the Dutch famine (Roseboom *et al.* , 2001a). Another possibility is that starvation affects early cell division of male and female embryos differentially (Cameron, 2004).

To our knowledge, ours is the first large scale quasi-experimental evidence of a Trivers-Willard effect in human populations. It is also the first evidence (quasi-experimental or otherwise) of an intergenerational "echo-effect" of maternal nutritional status on the sex ratio (to our knowledge). The support we find for the Trivers-Willard hypothesis suggests that modest movements in the sex ratio may reflect profound changes in maternal conditions.

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Appendix

Variable Definitions:

Census 2000

wdr Weighted death rate for the gestation period, assuming 9 month gestation, and province of birth. For example, a person born in January 1960 in Beijing is assigned 1/9th of Beijing's 1960's mortality rate and 8/9th of Beijing's 1959's mortality rate.

awdr Aggregate weighted average death rate, the **wdr** collapsed by month and birth year. Thus, it is the population weighted mean of **wdr** by month and year of birth.

mean Mean of dependent variable.

Province The province of birth. Our results are robust to inclusion of dummies for province of residence.

Illiterate Dummy indicating that the respondent was either illiterate or semi-literate.

Don't work Dummy indicating that the person did not work for more than 1 hour between October 25 and October 31 (in 2000).

On leave from job Not working because on leave, training, or seasonal lay-off.

Supported by other HH members Main income source was support by other household members.

Disabled Dummy indicating that the person does not work because he/she has "lost ability to work."

House area Area of home, in square meters.

Unmarried Dummy indicating that the respondent was unmarried at the time of the census.

Never married Dummy indicating that the respondent had never married.

Spousal education Includes head-spouse couples only. Education is in years.

Marriage age Age in months at time of first marriage.

Household head Dummy indicating that the respondent was household head. Includes only respondents living in "family type" households (as opposed to "collectives").

Male Dummy indicating that the respondent is male.

Sons/Kids Fraction sons among ever borne children. Excludes women who had not borne any children.

No kid Dummy indicating that the woman had borne no children.

Child mortality Number of children ever borne minus number of surviving children (at the time of the census) divided by the number of children ever borne, by year and quarter of birth of mother.

Hong Kong Natality data

I(1961) Dummy indicating that the mother was born in 1961, scaled by 1/100.

Low BWT Low birth weight. Dummy indicating that child weighed less than 2,500 grams at birth.

China Dummy for whether born in mainland China.

China Fertility Surveys

The China Fertility surveys were carried out in 1985 and 1987 in the following provinces: Hebei, Shaangxi, Liaoning, Guangdong, Guizhou, Gansu; and the municipalities of Beijing and Shanghai. (We have not been able to access data for Shandong.) In total, some 46,000 ever-married women between 15-49 years of age were interviewed, providing detailed information on pregnancy history. These data are available from the Office of Population Research, Princeton University, http://opr.princeton.edu/Archive/cidfs/.

Table 1: **2000 Census: Labor and Housing Outcomes for 1956-1964 Birth Cohorts**

	Illiterate	Don't work	Disabled	House Area
		Women		
mean	0.081	0.189	0.004	87.162
awdr	0.5052**	0.4714***	0.0448*	-220.1528***
	[0.2169]	[0.1530]	[0.0250]	[48.4753]
N	786156	786156	786156	772260
R^2	0.09	0.06	0	0.1
		Men		
mean	0.021	0.082	0.006	83.933
awdr	0.1585*	0.4017***	0.0657	-104.7566**
	[0.0784]	[0.1131]	[0.0426]	[38.3963]
N	818103	818103	818103	790342
R^2	0.04	0.04	0	0.08

awdr – aggregate weighted death rate by birth year and month.

mean – mean of dependent variable.

Standard errors clustered at province of birth in square brackets. * significant at 10%; ** significant at 5%; *** significant at 1%.

Table 2: **2000 Census, Marriage Market Outcomes, 1956-1964 Birth Cohorts**

	Unmarried	Never married	Spousal ed.[a]	Marriage age [b]	Household head [c]
			Women		
mean	0.061	0.004	9.057	269.237	0.118
awdr	0.2608	-0.0013	-6.3342**	67.4994**	-0.0998
	[0.1632]	[0.0249]	[2.4652]	[28.5417]	[0.1633]
N	786156	786156	685989	783015	786156
R^2	0.01	0.01	0.07	0.07	0.01
			Men		
mean	0.090	0.039	8.060	290.898	0.870
awdr	0.4902***	0.2676**	-0.1692	125.1309***	-0.5145**
	[0.1285]	[0.1035]	[2.5349]	[28.5395]	[0.2302]
N	818103	818103	683041	785927	818103
R^2	0.01	0	0.1	0.06	0.02

awdr – aggregate weighted death rate by birth year and month.

mean – mean of dependent variable.

[a] Includes head-spouse couples only. Education is in years.

[b] Marriage age is in months.

[c] Includes those residing in family units (i.e., excludes those residing in collectives).

Standard errors clustered at province of birth in square brackets. * significant at 10%; ** significant at 5%; *** significant at 1%.

Table 3: **2000 Census: Sex Ratio outcomes, 1956-1964 Birth Cohorts**

		Women[b]	
	Male[a]	Sons/Kids	No child
mean	0.51	0.548	0.007
awdr	-1.3147***	-0.3194**	0.0712
	[0.2651]	[0.1368]	[0.0503]
N	1604259	773291	786156
R^2	0	0	0

awdr – aggregate weighted death rate by birth year and month.

mean – mean of dependent variable.

[a] dummy – equals 1 if respondent male.

[b] Pertains to children borne.

Standard errors clustered at province of birth in square brackets. * significant at 10%; ** significant at 5%; *** significant at 1%.

Table 4: **2000 Census: Cross-sectional Variation in Famine Severity, Labor Market and Housing Outcomes, 1956-1964 Birth Cohorts**

	Illiterate	Don't work	Disabled	House Area
		Women		
wdr	0.1659	0.0953	0.0418***	-58.9501**
	[0.1269]	[0.1657]	[0.0116]	[22.0095]
N	764786	764786	764786	751352
R^2	0.08	0.06	0	0.1
		Men		
wdr	0.1231*	0.1628**	0.0585***	-52.1040*
	[0.0688]	[0.0666]	[0.0170]	[28.5949]
N	795408	795408	795408	768522
R^2	0.02	0.04	0	0.08

Standard errors clustered at province of birth in square brackets. * significant at 10%; ** significant at 5%; *** significant at 1%.

Table 5: **2000 Census: Cross-sectional Variation in Famine Severity, Marriage Market Outcomes, 1956-1964 Birth Cohorts**

	Unmarried	Never married	Spousal ed.[a]	Marriage age [b]	Household head [c]
			Women		
wdr	0.0505	0.0217	0.0794	14.7224	-0.0701
	[0.0623]	[0.0130]	[1.5906]	[19.9378]	[0.1297]
N	764786	764786	668672	761879	760726
R^2	0.01	0	0.06	0.07	0.01
			Men		
wdr	0.2666***	0.1555**	1.5938	67.6296***	-0.5089***
	[0.0696]	[0.0634]	[1.3770]	[22.8696]	[0.1183]
N	795408	795408	665857	764670	779087
R^2	0.01	0	0.1	0.06	0.02

wdr Weighted average death rate for the gestation period, assuming 9 month gestation. Varies by province and month and year of birth.

[a] Includes head-spouse couples only. Education is in years.

[b] Marriage age is in months. [c] Includes those residing in family units (i.e., excludes those residing in collectives).

Standard errors clustered at province of birth in square brackets. * significant at 10%; ** significant at 5%; *** significant at 1%.

Table 6: **2000 Census: Cross-sectional Variation in Famine Severity, Sex ratio outcomes, 1956-1964 Birth Cohorts**

		Women[b]	
	Male[a]	Sons/Kids	No child
wdr	-0.3264**	-0.1693**	0.0325
	[0.1390]	[0.0797]	[0.0251]
N	1560194	752418	764786
R^2	0	0	0

[a] dummy – equals 1 if respondent male.

[b] Pertains to children borne.

Standard errors clustered at province of birth in square brackets. * significant at 10%; ** significant at 5%; *** significant at 1%.

Table 7: **1984-2004 Natality outcomes in Hong Kong: mainland vs Hong Kong born mothers**

	Mother born:			
	Mainland		Hong Kong	
	Low BWT[a]	Son[b]	Low BWT[a]	Son[b]
mean	0.031	0.52	0.039	0.517
`I(1961)`	0.247**	-0.629***	0.014	-0.009
	[0.099]	[0.121]	[0.037]	[0.074]
N	198452	198452	393419	393419
R^2	0	0	0	0

`I(1961)` – dummy, equals 1/100 if mother born in 1961.

mean – mean of dependent variable.

[a] – dummy, equals 1 if birth weight was less than 2,500 grams.

[b] – dummy, equals 1 if child male.

Regression results from estimating equation 1 (without the province dummies), where the birth weight regressions also include a dummy for the sex of the child.

Standard errors clustered by year of birth in square brackets. * significant at 10%; ** significant at 5%; *** significant at 1%.

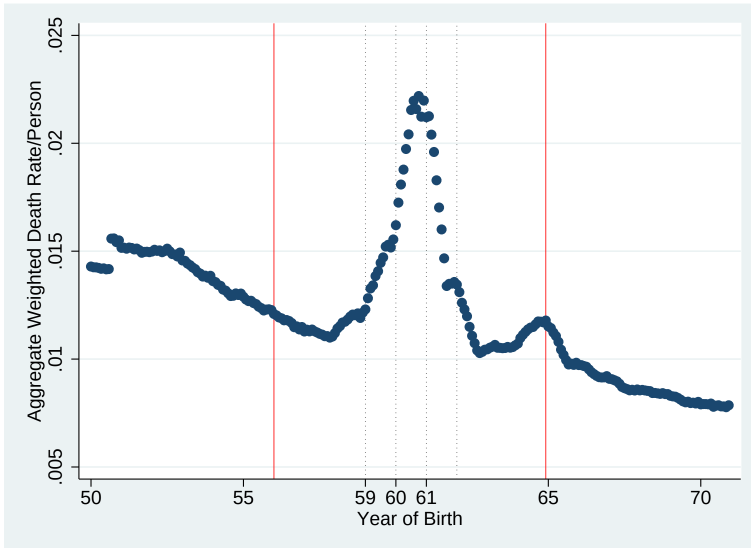

Figure 1: Aggregate Weighted Death Rate by Year and Month of Birth, Mainland China

Note: Authors' calculations based on all age death rates by year and province as reported by China Statistical Press (2000).

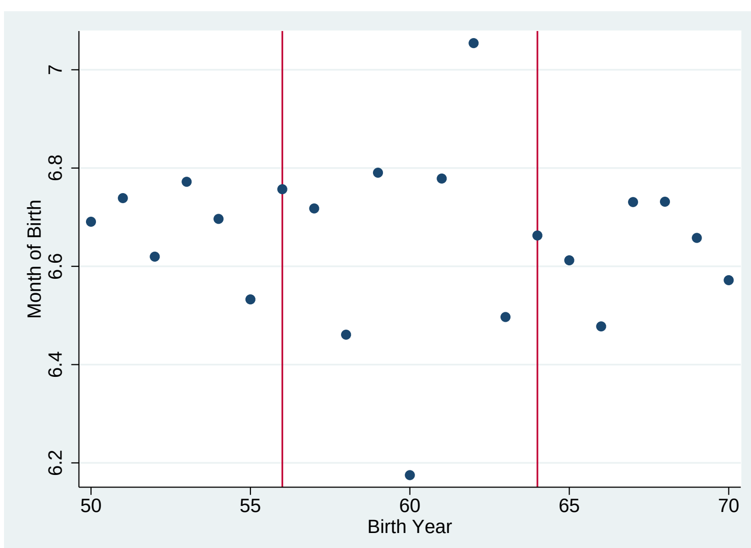

Figure 2: Average Month of Birth, Mainland China

Source: 2000 census.

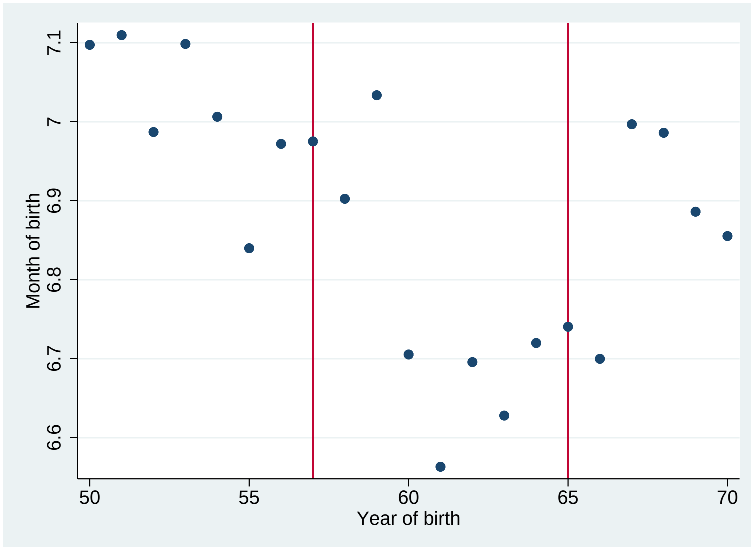

Figure 3: Average month of birth, Hong Kong Mothers born in Mainland

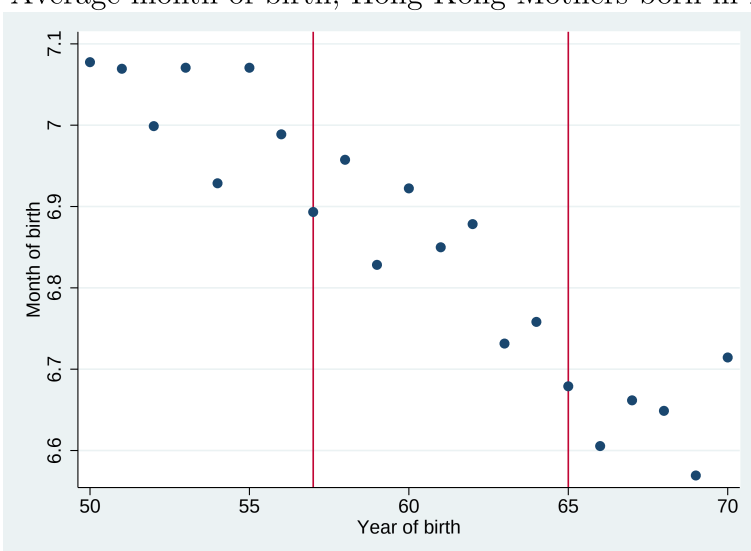

Figure 4: Average month of birth, Hong Kong Mothers born in Hong Kong

Source: Hong Kong Natality microdata.

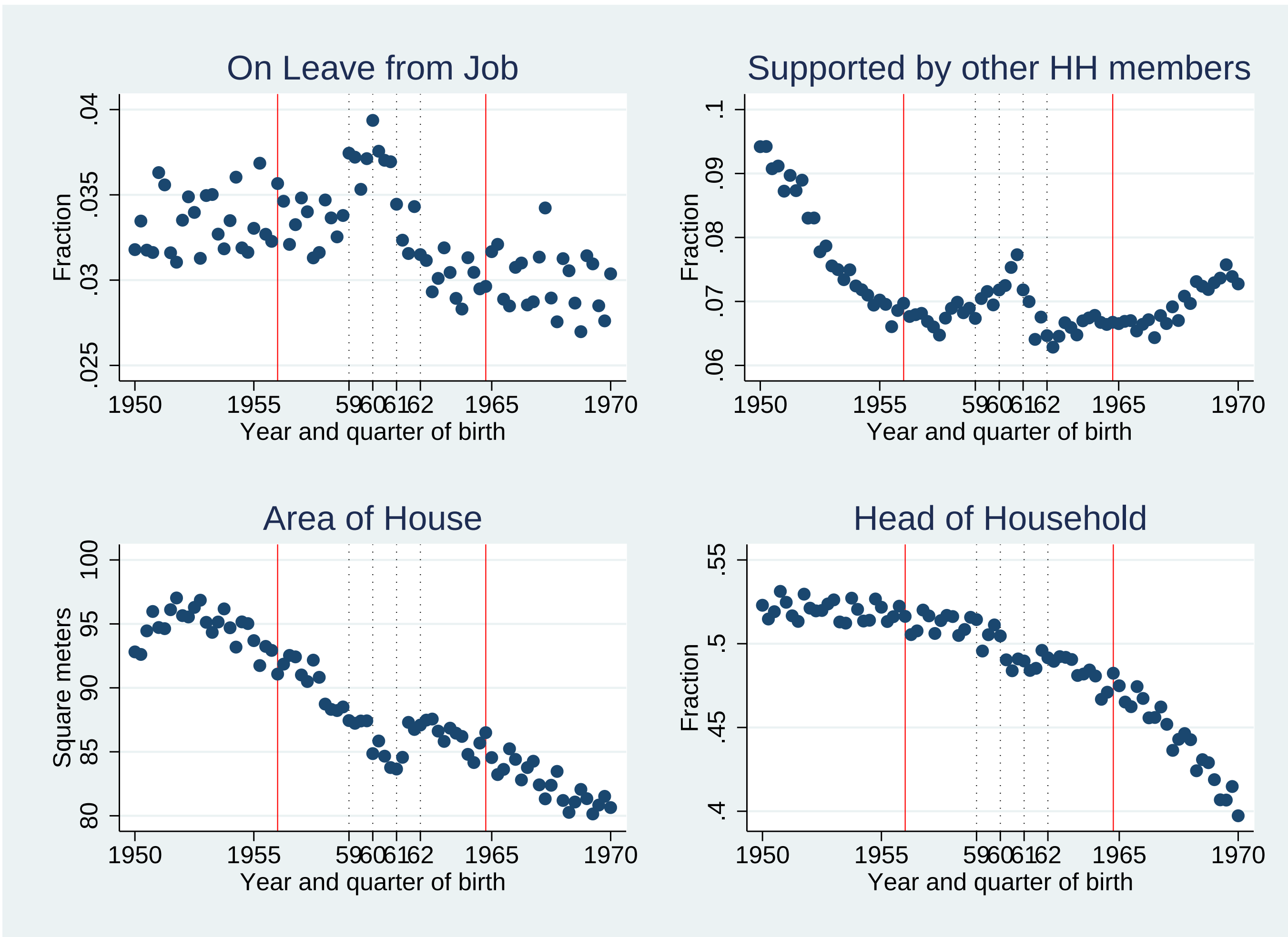

Figure 5: Socio-economic outcomes by year and quarter of birth

Source: 2000 census.

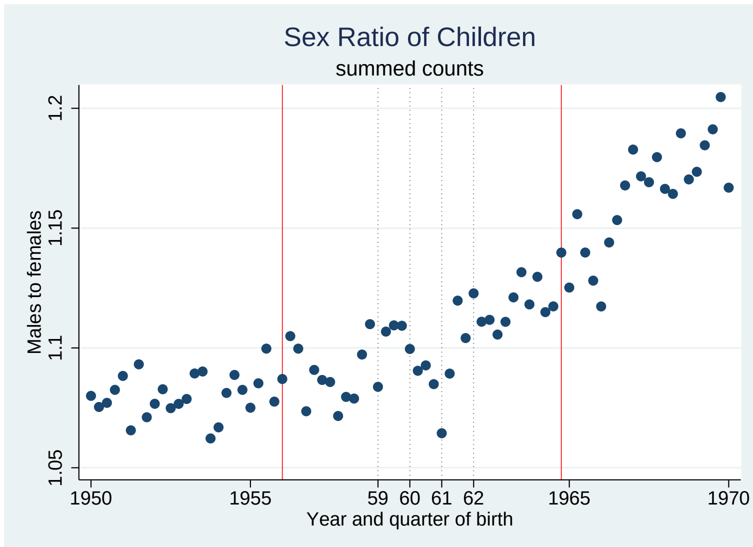

Figure 6: Offspring sex ratio by mother's quarter of birth

Source: 2000 census.

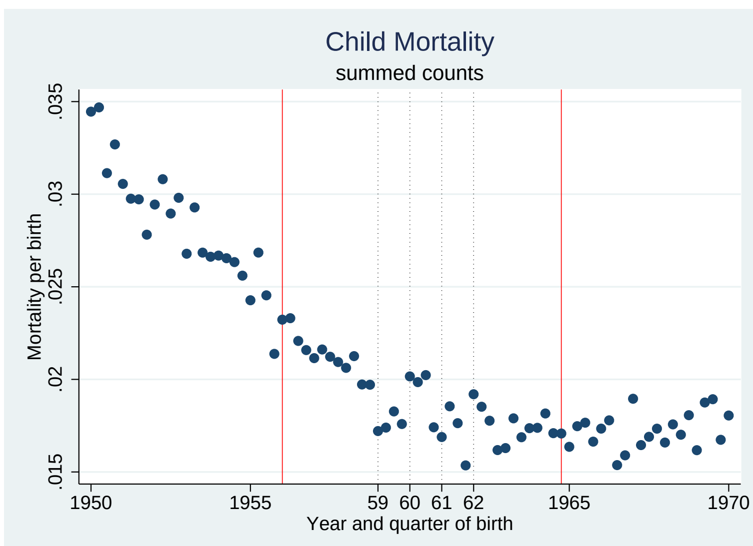

Figure 7: Child mortality by mother's quarter of birth

Source: 2000 census.

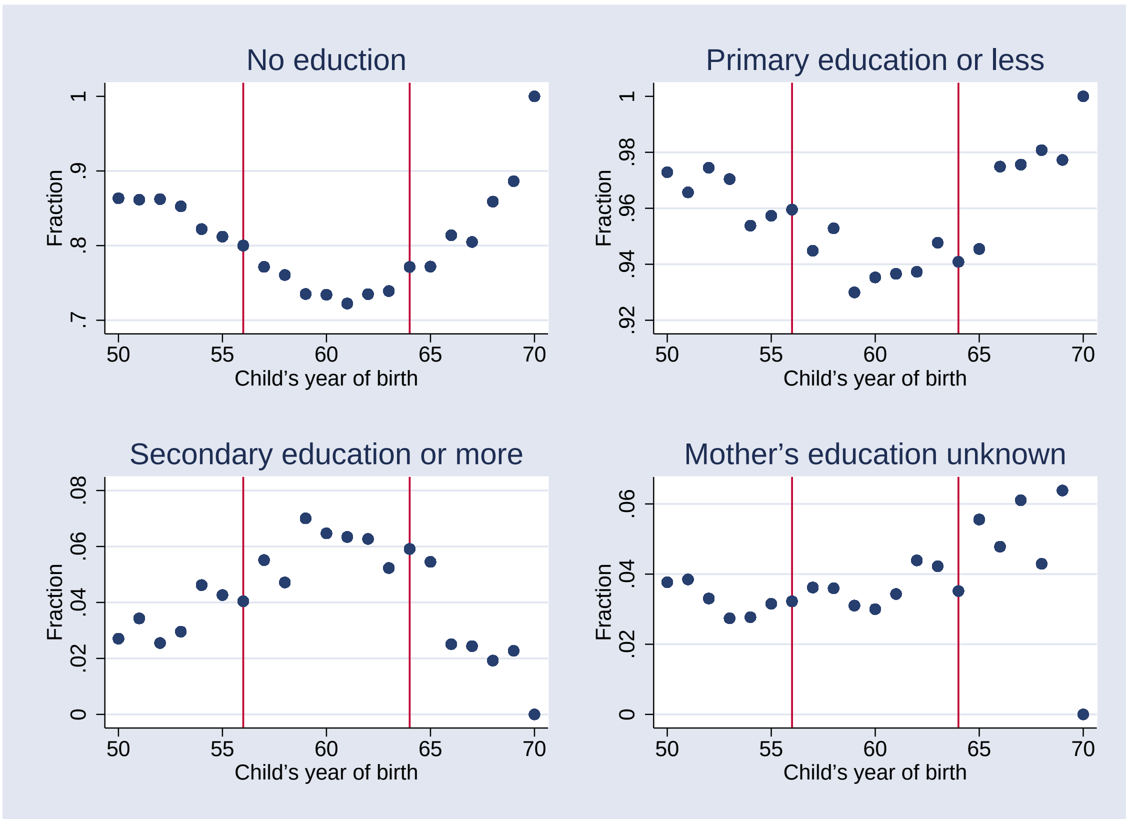

Figure 8: Mother's education by respondent's year of birth

Source: China Fertility surveys 1985/87.

Note: For Mother's education unknown, the universe is all respondents. For the remainder, the universe is those who knew their mother's education.

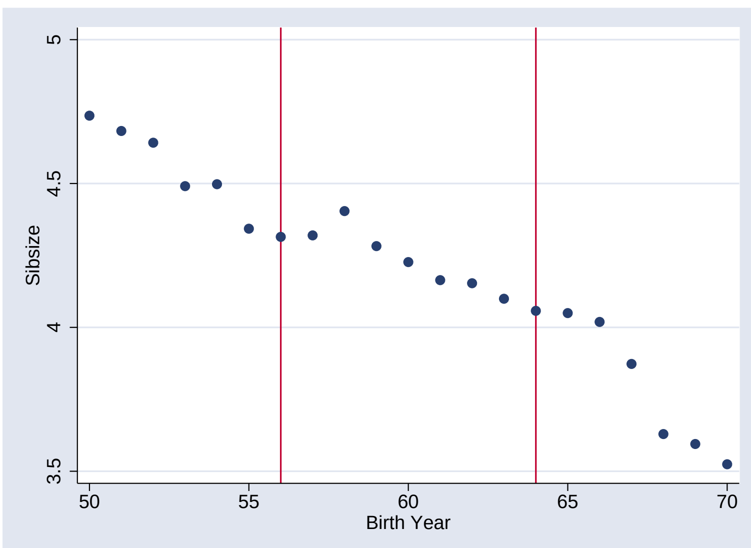

Figure 9: Number of siblings by respondent's year of birth

Source: 2005 Urban Chinese Education and Labor Survey.