

LARGE DEMOGRAPHIC SHOCKS AND SMALL CHANGES IN THE MARRIAGE MARKET

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Abstract

Between 1958 and 1961, China experienced a drastic famine. The famine substantially reduced birth rates and also adversely affected the health of these famine-born cohorts. This paper provides non-parametric estimates of the total effects of the famine on the marital behavior of famine-affected cohorts in rural Sichuan and Anhui. These reduced form estimates incorporate general equilibrium and heterogenous treatment effects. The paper uses the Choo-Siow model to decompose observed marital outcomes into quantity and quality effects. A decline in marital attractiveness of famine-affected cohorts, which is correlated with an increase in marital childlessness, provides support for the external validity of the Choo Siow decomposition. The small observed changes in marriage rates of the famine-born cohorts are due to a substantial decline in their marital attractiveness. (JEL: J1, O53)

1. Introduction

The “Great Leap Forward” was a national-level political and economic experiment carried out in China between 1958–1961.¹ Collectivization of farming, which began in the mid-fifties, increased in speed and scope. Rural labor was reallocated from agriculture towards inefficient industry. In many localities, strong political incentives contributed to official exaggeration of grain yields and output, leading to a reduction in sown area, and excessive state procurement and export.

The Greap Leap Forward resulted in one of the most severe famines in Chinese history. Estimates of famine-related mortality range from 15 to 30 million deaths. Peng (1987) estimates that births lost or postponed resulted in about 25 million fewer

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1. For an overview of the Great Leap Forward, see Lardy (1987).

births.² In general, the countryside was struck much harder than cities.³ The economic experiment was abandoned by early 1962. The mortality rate rapidly fell and the birth rate also quickly recovered.

While the drop in birth rates is widely recognized, much less is known about the effects on those who were born during the famine. The medical literature reports that individuals suffering nutritional deprivation either *in utero* or in their infancy face severe deleterious long-run health effects (See Barker, 1992). Recent research by Gorgens et. al. (2005), St. Clair et. al. (2005), and Luo, Mu and Zhang (2006) provide some direct confirmation for this in the case of China, focusing on such health-related outcomes as stunting, obesity, and schizophrenia.

Not all the effects of the famine on the famine-born cohorts were negative. Due to the drop in the birth rates, the famine-born cohorts were small in number relative to adjacent birth cohorts. This scarcity should increase their relative values in both the marriage and labor markets. Their increased value in the labor market should also further add to their desirability in the marriage market.

Ceteris paribus, the net effect of the famine on marital outcomes of famine-born cohorts is an aggregation of three effects: (1) a negative attractiveness effect due to adverse health outcomes that reduces demand for famine-born spouses; (2) a positive attractiveness (wage) effect due to relative scarcity of famine-born cohorts in the labor market that increases their demand as spouses; and (3) due to customary gender differences in ages of marriage, there is an increase in spousal demand for famine-born cohorts because of their relative scarcity in the marriage market.

The famine also affected both pre and post-famine-born cohorts. Pre-famine-born cohorts were young children at the time of the famine. The famine would have adversely affected their health and human capital. The least fortunate among them would have died. So again there are quantity and quality effects on those who survived into marriageable age. Meng and Qian (2009) provide direct evidence on the long run health and labor market outcomes of these cohorts. How the marital attractiveness of these individuals compares with that of the famine-born cohort is difficult to assess a priori.

The upshot of the above is that there are large quantity and complicated quality changes to the famine-affected cohorts. These individuals also match with each other in the marriage market. Thus, the observed marital outcomes of the famine-affected cohorts are equilibrium responses to these quantity and quality changes.

The objective of this paper is to examine the effects of the famine on the marital outcomes of the famine-born and adjacent cohorts in Sichuan, an agricultural province

2. Ashton et. al. (1984) provide estimates at the national level of the magnitude of the demographic crisis. See Peng (1987) for a detailed analysis of the impact of the famine at the provincial level. Lin and Yang (2000) and Li and Yang (2006) examine causes of the famine. Becker (1997) provides a narrative account.

3. The famine extended beyond China's traditional famine belt region. For example, Sichuan province, where mass famines were rare, was one of the hardest struck.

which was severely affected by the famine.^{4,5} We study the rural population because the famine disproportionately affected rural rather than urban communities.⁶

Figure 1(a) shows the distribution of individuals by age in the 1990 Census for rural Sichuan. Due to a high long-run birth rate and mortality rates that increase with age, population by age should be declining with age. The famine-born cohort was 29-31 years old men and women. It is clear from Figure 1(a) that the population of the pre-famine-born cohort (32-34) was also adversely affected by the famine. That is, young children during the famine were less likely to survive to adulthood. On the other hand, the panel also shows the quick recovery in fertility (and subsequent survival to adulthood) after the famine.

Our paper provides two kinds of estimates of the effects of famine on marital behavior. First, using a first difference methodology, we provide non-parametric estimates of the net effect of the famine on the marital outcomes of the famine-affected cohorts. To this end, we estimate a set of summary statistics, the total gains to marriages, which are a complete description of marital behavior by all participants at a point in time. We estimate total gains for two time periods. In the control period, we assume that marital behavior was unaffected by the famine; in the treatment period, marital behavior was affected by the famine. The estimated differences in total gains between the two periods provide an estimate of the net causal effect of the famine on marital behavior. We implement our first difference empirical strategy by comparing for Sichuan the marital behavior of the famine-affected cohorts in 1990 to their same age counterparts in 1982.⁷

Our net causal-effect estimates incorporate both general equilibrium and heterogeneous treatment effects, and are consistent as long as our first-difference methodology is able to control for other factors that may have affected marital behavior between the two periods. We observe marital behavior of same-age cohorts that is eight years apart. Absent the famine, we argue that other cultural, social or policy changes did not differentially affect the marital behavior of these cohorts.

Unlike the pre-famine born cohorts, marital behavior of the famine-born cohort was significantly different than their same aged counterparts 8 years earlier. Unless policies targeted the marital behavior of the famine-born cohort alone, the non-linear effects captured by our first-difference estimator for the famine-born cohorts suggest

4. Similar results for Anhui province, which was also hard hit by the famine, are presented in our working paper, Brandt, Siow and Vogel 2009.

5. In 1957, 86% of the labor force were in the primary (agriculture, fisheries and forestry) sector in Sichuan. Between 65%-70% of GDP originated in the primary sector. These data are taken from Xin Zhongguo wushi nian tongji ziliao huibian, 2005). Further details are in Peng (1987).

6. Peng(1987) documents that excess mortality was more severe in rural than in urban areas. At the national level, the excess crude death rate for the urban population between 1958–1962 was 13.84 compared to 7.94 for the two preceding years. By comparison, the excess crude death rate for the rural population rose from 11.45 to 24.45 over the same period. See Peng (1987), p. 646.

7. Initially, we used a difference in differences strategy. These estimates were difficult to interpret because the famine was national in scope and the control provinces that we used were also affected by the famine.

that the differences in marital behavior between the two censuses may be attributed to the effects of the famine alone.

Second, in order to decompose the estimated net effects into quantity and quality effects, we first estimate a structural marriage-matching model that will exactly fit the 1982 marriage distribution. Our non-parametric structural model is the CS marriage-matching model (Choo and Siow 2006).⁸ Due to its flexibility and ease of use, recent researchers have used CS and its extensions to study different aspects of empirical marriage matching.⁹ It fits the 1982 data perfectly. The parameters of the CS model are the total gains statistics described above. For a particular observed marital match, the total gains measures the average marital output which two randomly-chosen individuals with those characteristics would enjoy if they were matched in a marriage relative to remaining single. We next use the estimated model to predict the response of the marriage market in 1990 to changes in observed population supplies in 1990. Finally, the estimated differences between the actual and predicted 1990 marriage distributions are estimates of changes in marital quality due to the famine.¹⁰

For this paper, there are two threats to the interpretation of our decompositions. First, the CS model may not be a good approximation of marriage matching behavior, in which case the estimated total gains do not capture true changes in marital qualities. Second, even if the model is conceptually correct, there were many social and economic changes in China between 1982 to 1990.¹¹ Insofar as these changes have significant effects on observed marital behavior, they will contaminate our estimates of the changes in total gains. As we have already argued above, there were unlikely to be significant policy changes which differentially affected marital behavior of the famine born cohort and not the two adjacent aged cohorts. We also deal with the two threats to identification by providing a test of external validity of CS using data on childlessness.

Many studies have shown that the famine-affected cohorts suffered adverse health outcomes. Ill health will reduce the output of marriages drawn from this cohort. To the extent that cohort-specific lower marital outputs due to the famine were observed

8. A marriage matching function is a production function for marriages (Pollard 1997, Pollack 1990). Inputs are population vectors of types of individuals. Output is a matrix of who marries whom, and who remains unmarried.

9. They include the effect of the legalization of abortion on marriage formation (CS), famine effects on marriage in China (Porter, forthcoming), marriage and divorce in Denmark (Bruze, Svarer and Weiss, forthcoming), evolution of marital preferences in the US (Chiappori, et. al., 2010), positive assortative matching in spousal education attainment (Siow, forthcoming), marriage matching and spousal labor supply (Choo and Seitz 2013). Chiappori and Salanie (2014) has an excellent survey of CS and its extensions. Also see Mourife and Siow (2014).

10. Our decomposition is based on residual analysis, a common tool in economics (E.g. interpretation of Solow residuals in macroeconomics, estimates of discrimination in log earnings regressions.).

11. For example, Wang and Zhou (2010) show that divorce rates were rapidly rising in that period. Rural to urban and interprovincial migration were also increasing. Since the absolute number of divorces was small, we do not believe that the increases in the divorce rates were large enough to contaminate our estimates based on the stocks of marriages. We discuss the problem of migration on our estimates in the data appendix and again argue that they are not quantitatively significant for our purpose.

and/or anticipated by individuals before they marry, measures of marital output should predict post-marital ill health outcomes. An extreme negative marital outcome is infertility in the marriage. In the 1982 and 1990 census, each woman was asked how many children she had given birth to. For each census year, we construct a measure of childlessness for each type of marital match, which is the proportion of those marital matches that report having no children. This construction of childlessness is completely unrelated to the marriage matching data used to estimate total gains.

Using each type of marriage match as an observation, we regress the change in childlessness between the two census on the change in estimated total gains. Our results show that, for each type of marital match, an increase in total gains is associated with a fall in childlessness between the two census. Since childlessness occurred after marriage, this result shows that our estimate of changes in total gains predicts future changes in childlessness. We interpret the increase in childlessness of the famine born at least as partially due to direct famine effects on their health as children. We cannot rule out the alternative, but to us far fetched, hypothesis that the increase in childlessness of the famine born was only due to negative socioeconomic effects which affected the famine born's health after the famine was over.

Since our test of the usefulness of total gains is not rejected, we use the CS model to decompose the estimated marital changes into quantity and quality effects. In Sichuan, there were small observed changes in the marriage rates of the famine-born cohorts relative to their adjacent-aged peers. There were large changes in choices of marital partners and total gains for famine-born cohorts relative to their adjacent-aged peers. To a first approximation, our decomposition shows that the benefit that the famine-born cohort derived from their relative scarcity was offset by their decline in marital attractiveness. The main substantive conclusion of this paper is that the small observed changes in marriage rates of the famine-born cohorts are due to a substantial decline in their marital attractiveness. Our main methodological conclusion is that allowing for general equilibrium effects and estimating heterogeneous treatment effects are important because the pre-famine-born, famine-born and post-famine-born cohorts experienced very different but linked outcomes.

Our 2008 working paper contained two other extensions. First, we repeated the above analysis with data from rural Anhui. Our results were qualitatively and quantitatively similar to what we have here. Second, because the CS model allows for individuals to have multidimensional characteristics, we also estimated the CS model for both Sichuan and Anhui where individuals match by both age and educational attainment. The main findings were: (i) To a first order, adding educational attainment as an additional attribute for individuals to match on did not significantly change the estimates of losses in average total gains across birth cohorts by age. (ii) There were first-order effects within an age cohort. Famine-born individuals with low education suffered a larger drop in marital attractiveness than their more educated peers.

Finally, it will be useful to state some limitations of our study. In general, the lack of data in the censuses limits the degrees of observed heterogeneity we can admit in our empirical analysis. Although several recent papers provide dynamic extensions of

CS (Siow 2009; Bruze et. al. forthcoming), our data are not suited for these dynamic investigations because we do not know the durations of marriages.

2. Literature Review

Our paper is substantively related to three literatures. First, it is related to the literature that studies the effects of the “Great Leap Forward” on social and economic outcomes (E.g. Almond et. al., 2007; Chen and Zhou, 2007; Geogens et. al. 2007; Meng and Qian, 2010; Mu and Zhang, 2010; Porter, forthcoming). Of these papers, Almond et. al. and Porter study the effects on the marriage market. We build on their work.

Both of these two papers use a regression framework to study the causal effects of famine-related changes on marital behavior of the affected cohorts. Both papers focus on homogenous treatment effects. Utilizing data on cohorts born between 1956-64, Almond et. al. analyze the effect of the severity of pre-natal exposure on marriage rates, age at marriage, spousal age differences and other spousal characteristics. They conclude that the famine had modest effects on marriage rates, but caused the affected cohorts to marry later, increased spousal age gaps, and decreased spousal education gaps. They attribute this to the decline in marital attractiveness of the famine-born cohort. Porter, on the other hand, uses marital-share weighted adult sex ratios to capture the effects of relative scarcity on marital outcomes, controlling for marital attractiveness along the same lines as Almond et. al. Using data for cohorts born between the 1930s and 1970s, she finds that as women became more scarce, men married later and adapted their behavior to be more competitive in the marriage market. Women also exercised greater bargaining power post-marriage.

A famine effect on marital behavior in a marriage market is a general equilibrium phenomenon. A regression using individual-level data and across marriage market variation in sex ratios to estimate a famine effect on marital outcomes ignores these general equilibrium effects and also imposes a single homogeneous treatment famine-effect parameter.¹² Our paper incorporates all general equilibrium effects and estimates heterogenous treatment effects. We are more restrictive in one important respect: The regression framework controls for a wider variety of individual characteristics than we are able to do.¹³ Thus we view our framework as complementary to the above papers.

Second, our paper is also related to a literature looking at the effect of early childhood nutrition, including that experienced in utero, on life outcomes such as health, education and fertility. Maternal nutritional deprivation has been linked to a

12. The inability of the regression framework using individual level data to deal with general equilibrium effects is well known (E.g. Imbens and Woolridge 2008; Heckman, Lochner and Taber 1999). Using marriage rate regressions, Angrist 2002 found that the causal effect of changes in the sex ratio on the male marriage rate is inconsistent with that found for the female marriage rate. Choo and Siow 2006 also provide another example.

13. Conceptually, our framework allows for many individual characteristics. Practically, as we increase the number of characteristics, we will run into a thin cell problem.

broad spectrum of adult outcomes including obesity, schizophrenia, and coronary heart disease morbidity. A study in Guatemala (Ramakrishnan et. al. 1999) identifies positive effects of nutritional supplements on reproduction outcomes for those in utero and up to three years of age. Studies based on the 1944-45 Dutch Famine (Kyle, et. al., 2006 and Elia et. al., 2005) only find a link between post-natal exposure to famine and female reproduction in terms of childlessness, age at first delivery, and family size. In the Chinese context, Almond et. al. (2007) find a positive but statistically weak correlation between the severity of pre-natal exposure to famine and the likelihood of females having children. They find no effect for males. We build on their work.

Finally, our paper is related to the literature that studies the effect of exogenous variations in the sex ratios on marital outcomes (Akabayashi, 2006; Bhrolchain 2001; Brainard 2006; Esteve i and Cabré 2004; Francis 2007). Many of the exogenous variations in sex ratios have both a quality and quantity dimension due to the effects of war (e.g. Brainard; Esteve i and Cabré; and Francis). Even the variations in sex ratios due to superstition about being born in “unlucky” years have a quality dimension (Akabayashi). Individuals born in these “unlucky” years suffer a social stigma that makes them less desirable in the marriage market. But they benefit from being relatively scarce in the labor and marriage market. Our framework can be applied to disentangle these two effects in these environments.

Previous researchers who have studied marriage rates and large exogenous sex ratio changes in other contexts have also often found small effects of these changes on marriage rates, e.g. Bhrolchain (2001). As we do in this paper, these researchers attribute these small effects to flexible spousal choices in the face of large changes in sex ratios. But small changes in marriage rates due to a large exogenous change in sex ratio may be deceiving. In the case of this paper, the small changes mask large offsetting quantity and quality effects.

Methodologically, we build on Bergstrom and Lam’s (1994; hereafter BL) and Foster and Khan’s (2000) empirical studies of the Swedish and Bangladeshi marriage markets using frictionless marriage matching models. CS extends BL, and Foster and Khan in two ways. First, CS is non-parametric and second, CS endogenizes the marriage rates of different birth cohorts.¹⁴

CS has been extended and generalized in several directions. Chiappori and Salanie (2014) provide an excellent survey on recent developments. See Mourifie and Siow (2014) for some particularly tractable empirical extensions.

3. Methodology

Let t denote the year of the census, $t = \{1982, 1990\}$. At each year, individuals are differentiated by their age and or education. Let j denote type j women and i denote

14. While not a conceptual advance, relaxing the tight parameterization of BL is important for fitting the spousal matching distribution. Endogenizing the behavior of the unmarried individuals is new to CS and it allows us to fit age-varying marriage rates.

type i men. $j = 1, \dots, J$ and $i = 1, \dots, I$. F^t is the population vector of women at time t with typical element f_j^t , the number of women of type j in year t . M^t is the population vector of men at time t with typical element m_i^t , the number of men of type i in year t .

μ_j^t is the number of married women of type j year t . μ_i^t is the number of married men of type i year t . Let the number of unmarried type i men be $\mu_{i0}^t = m_i^t - \mu_i^t$, and the number of unmarried type j women be $\mu_{0j}^t = f_j^t - \mu_j^t$. Let μ_{ij}^t be the number of type i men married to type j women. There are $I \times J$ types of marriages at time t . Let μ^t , the marriage matching distribution, be an $I \times J$ matrix whose typical element is μ_{ij}^t .

In order to predict changes in the marriage distribution due to changes in population vectors, M^t and F^t , and other parameters which affect the marriage market, π^t , demographers and economists estimate marriage matching functions (MMF):

$$\mu^t = G(M^t, F^t, \pi^t)$$

The CS MMF is defined by the $I \times J$ equations:

$$\ln \frac{\mu_{ij}^t}{\sqrt{\mu_{i0}^t \mu_{0j}^t}} = \pi_{ij}^t; \quad i = 1, \dots, I; j = 1, \dots, J \quad (1)$$

The total gains to marriage, π_{ij}^t , has the following interpretation. At date t , pick a i type man and a j type woman at random. π_{ij}^t is the difference between the expected utilities that this randomly chosen pair will derive from marriage and the sum of the expected utilities they will derive from remaining unmarried. The left hand side of equation (1) is the log ratio of the number of (i, j, t) marriages versus the geometric average of those types of individuals remaining unmarried. Intuitively, we expect this ratio to increase with π_{ij}^t . CS says that the log of this ratio is equal to π_{ij}^t .

A derivation of CS is given in the appendix. Decker et. al. 2013 and Graham 2013 provide the comparative statics for this model. An advantage of CS is that it allows a type of individual to be defined by multiple characteristics. How many characteristics should a researcher use depends on the study of interest.

From a descriptive point of view, CS fits any marriage matching distribution from a single marriage market exactly. In other words, any generalization of CS will not be identified with data from a single marriage market which is what we have from the 1982 census. CS is easy to estimate. π_{ij}^t is estimated by the left hand side of equation (1).

Given M^t and F^t , Decker, et. al. show that μ^t is uniquely determined by the matrix of parameters π^t with typical element π_{ij}^t . So from a purely descriptive point of view, π^t is an alternative complete description of the marriage distribution μ^t . In other words, we can use π_{ij}^t to describe the marriage distribution at time t independent of whether CS is the true model of the marriage market or not.

Let $\Delta X \equiv X^{t'} - X^t$, for any X . Then $\Delta \pi$ is a complete description of the changes in marital behavior between the two periods. It imposes no a priori structure on behavior between the two periods. Thus:

$\Delta\pi$ is a consistent estimate of the net causal effects of the famine on marital behavior for the famine-affected cohorts as long as t and t' are valid control and treatment periods.

$\Delta\pi$ will not be a consistent estimate of the net causal effects of the famine on marital behavior if there are other factors which also affected the marriage markets between those two periods. This is a standard problem of using a time difference estimator to estimate causal effects. We argue that this is not a significant concern in our context. First, the two censuses are 8 years apart, which is a short time for significant changes to social and cultural values regarding marriage. We are also not aware of such policy changes. Empirically, the marital behavior of women in the 1990 census, born immediately before the famine, was by and large similar to the same aged women in the 1982 census, born 8 years earlier. The small difference in marital behavior was that these pre-famine born women tended to avoid marrying slightly younger (famine born) men relative to their same age peers 8 years earlier. Thus the marital behavior of these pre-famine born women adds to the evidence that there were not significant interim policy changes which affected the marital behavior of women 8 years apart in the censuses.

Second, marital behavior of the famine born cohort were significantly different than their same aged counterparts 8 years earlier. Unless policies targeted the marital behavior of the famine born cohort alone, the non-linear effects captured by our first difference estimator for the famine born cohorts may be attributed to the effects of the famine alone.

Even if we accept that $\Delta\pi$ is a consistent estimate of the net causal effects of the famine on marital behavior, we cannot decompose the net causal effect into quality versus quantity adjustments. In order to disentangle observed changes in marital behavior between effects due to changes in marital surpluses and effects due to population supplies, we need to predict what the counterfactual marriage distribution in t' , $\hat{\mu}^{t'}$, would have been with changes only in population supplies, $M^{t'}$ and $F^{t'}$, holding π^t unchanged. Based on the CS model, we can use equation (1) to produce $\hat{\mu}^{t'}$.

Although the CS model provides a way to disentangle observed changes in marital behavior between effects due to changes in marital surpluses and effects due to population supplies, there is no way to test the internal validity of the CS model using marriage matching data alone because the CS model is just identified from such data. We provide a test of the CS model in our context using data on childlessness.

The main implication of CS is that total gains is a measure of the gains to marriage relative to not marrying. Having children is perhaps the most important component of marital output for most marriages. The inability to have children will severely reduce marital output. So we should expect an indicator of infertility to be negatively related to the CS measure of total gains. This is the relationship which we wish to test.

In the 1982 and 1990 census, each woman was asked how many children she gave birth to. For each census year, we construct a measure of childlessness for each type of marital match that is equal to the proportion of those marital matches which reported having no children. This construction of childlessness is completely unrelated to the

marriage matching data used to estimate total gains. Using each type of marriage match as an observation, we investigate the relationship between changes in childlessness and the changes in total gains, with and without controlling for changes in sex ratio and changes in population sizes between the two census. If the CS model is valid, we expect a negative relationship between changes in childlessness and changes in total gains. The reason for looking at differences is so that we can control for age effects in having children.

We study childlessness rather than fertility because the number of children is more within the control of the married couple and also policy experiments operative at that time. Although childlessness is also endogenous, we assume that most married couples wanted at least one child and therefore the absence of children is primarily due to exogenous factors such as infertility of the couple.¹⁵

4. Summary Data and Sex Ratios

All the data presented here come from the 1% household sample of the 1982 Census of China and the 1% clustered sample of the 1990 Census of China. Wang (2000) and Mason and Lavelly (2001) are useful resources on the details of the census and data samples. In our analysis, we only use those data pertaining to individuals who reside in rural counties. This can be rationalized on two grounds: first, the countryside was more affected by the famine than the cities,¹⁶ and second, the rural marriage market was largely self-contained, and highly local in nature.

In the Data Appendix, we discuss how rural is defined, and examine several other data-related issues, including migration. Migration may matter in a number of ways for our analysis. First, migration of rural born out of Sichuan, and in-migration into rural Sichuan may bias our estimates of the sex ratios that we use in the construction of our marriage matching functions. Second, some of the migration may have been for marriage.

Unfortunately, the 1982 Census contains no information on migration, and for the 1990 Census the information is limited to migration that occurred between 1985 and 1990. Using the more complete data provided by the 2000 Census, we are able to construct alternative estimates of the sex ratios based on the rural-born population, including those currently living outside the province. As detailed in the appendix, the differences with the estimates we use are small. Moreover, the bias is fairly similar across the age cohorts that make up our analysis.

The Census does not directly link who is married to whom in a household. It provides a marriage indicator for each individual. The data appendix discusses how we match up married couples in a household based on age differences and marital statuses. There may be a small number of imputation error when an older sibling

15. We have experimented with using fertility as a dependent variable with qualitatively similar results.

16. See footnote 10.

married a younger spouse than a same gender younger sibling. Since we use the same imputation method for both censuses, there is no reason to expect such imputation errors to systematically bias our first difference results.

Table 1 provides some summary statistics for rural counties in Sichuan from the 1982 and 1990 census. The average spousal age differences in Sichuan ranges between two to three years. Any observed spousal age difference is an equilibrium outcome determined by marriage market conditions. Under average marriage market conditions existing in China at the times of the 1982 and 1990 census, the average spousal age difference was about three years. Since the age of marriage is recorded in the census by years, it will be pedagogically convenient for us to assume that the customary equilibrium spousal age difference is three years. The marriage rates for women are higher than those for men in both provinces and both census. High education is defined as completing elementary school. The table shows that less than 50% of men or women in rural Sichuan completed elementary school.

The first-order impact of the famine on the marital behavior of individuals would have been on the famine-born cohort and their customary spouses. For men who usually marry women three years younger, the customary spouses for the famine-born men were the post-famine-born women. For famine-born women, their customary spouses were the pre-famine-born men. Thus, we consider individuals born between 1956 to 1964 to be the famine-affected cohorts. We observe the marital behavior of individuals in 1982 and 1990. For convenience, the ages of these individuals in 1990 are given in Table 2.

Our main interest is to examine the behavior of the famine on marital behavior in the 1990 census. The reason for focusing on the 1990 census is that by 1990, the post-famine cohort was 26-28 years old. Most women of that age category and older would have acquired their permanent marital status. Except for 26 and 27 years olds, most men of that age category and older would also have acquired their permanent marital status.

We will use individuals of the same age and characteristics in 1982 as controls for their counterparts in 1990. That is, the control group for post-famine individuals are those who were 26-28 in 1982, the control group for famine-born cohort are those who were 29-31 in 1982, and the control group for the pre-famine cohort are those who were 32-34 in 1982. In general, as shown in Table 2, individuals in the control groups, of age 26 and older in 1982, were not affected at birth by the famine. There is one year of overlap. Individuals of age 26 in 1982, used as a control group for pre-famine 26 years olds in 1990, are also in the post-famine group in 1990. Whether this year of overlap will affect the results is an empirical issue which will be resolved shortly.

The eighties were a period of active social and economic changes in China, including a marriage reform act of 1981. Most of the social and economic changes in the eighties, including migration, have level and or trend effects.¹⁷ Thus, we need

17. China's New Marriage Law in 1981 increased the age of marriage to 20 and 22 for females and males, respectively. In our analysis, the youngest individual in our control group would have been 26 in

to make a case that it is reasonable to use a first-difference strategy to study the impact of the famine on marital outcomes. We make our case in two steps. First, we will show that in 1982, within a province, marital behavior of the three control groups differed from each other by at most a smooth age trend due to lifecycle effects, i.e., absent smooth lifecycle effects, the marriage behavior of 26-28, 29-31, and 32-34 years olds in 1982 was similar to each other.¹⁸ The shape of total gains was similar for all three cohorts. That is, conditional on marriage, there was no difference in the choice of spouses between the three cohorts. Since we are comparing marital behavior by age groups, life-cycle effects in the timing of marriage are controlled for by the levels of the total gains distributions.

Second, the observed changes in marital behavior in 1990 for the famine-affected cohorts follow a distinct non-linear age pattern that coincides with the population changes for those cohorts in 1990. We do not know any other social or economic change that followed this distinct age pattern and also only impacted these cohorts. Alternatively, we are essentially assuming that total gains in 1990 would have been the same as in 1982 absent the famine.

5. Sichuan

Figure 1(a) shows the number of individuals by age in rural counties in Sichuan in 1990. The pre-famine cohort, 32-34, was affected by the famine. There were less of them than 35 or 36 years olds. Absent the famine, due to population growth and mortality risk, there should be less older individuals rather than more in a given census year. Thus, the 32-34 years olds were adversely affected by the famine.

The famine-born cohort, 29-31, is substantially smaller than the adjacent cohorts, reflecting primarily the fall in the birth rates of that cohort. Recovery of the birth rates after the famine was very rapid. There is no visible impact of the famine on cohort sizes after 1964, ages 25 or younger in 1990.

Figure 1(a) also shows that there were less 35 and 36 years olds than 37 olds, which implies that these cohorts were also affected by the famine. We do not directly study their marital behavior because of our focus on the marital behavior of the famine-born cohort with their adjacent aged peers. The analysis of the famine-affected cohorts takes into account that they could and did marry individuals 35 years old and older in 1990.

1982, suggesting that any effect of the new marriage law on their behavior through whom they could marry would have been marginal.

18. The presence of lifecycle effects in marital behavior necessitates using time variation as a control group, i.e. we cannot use the marital behavior of 40 year olds in 1990 as a control for the behavior of 30 year olds in 1990 due to lifecycle differences between the two groups.

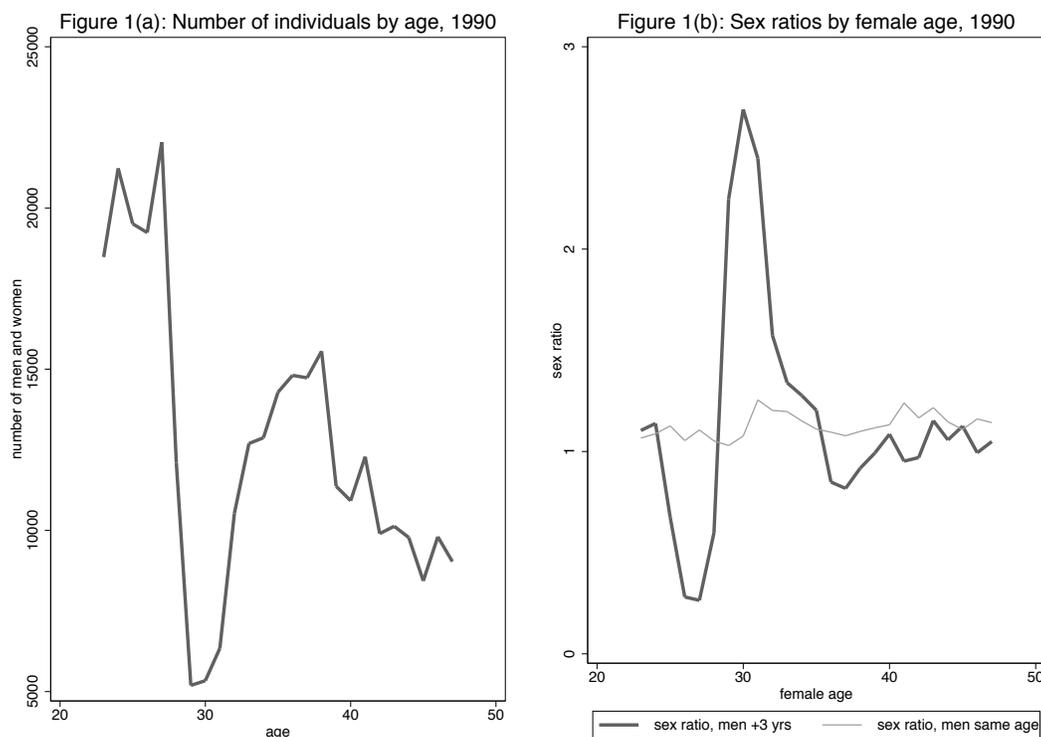


FIGURE 1

5.1. Marriage Rates

Figure 1(b) shows two sex ratios by age. The dashed line is the sex ratio of men to women for same age men and women. In general, the sex ratio is slightly larger than one, which will have an effect on the male versus female marriage rates.¹⁹ The famine had little to no impact on the sex ratio. There is little evidence that male children were significantly favored over female children among the famine-affected cohorts. The solid line is the sex ratio by women's age where the men were three years older than the women. Here, the effects of the famine are very clear. The sex ratio was above 2.5 for famine-born women because there were relatively more pre-famine-born men. Also the sex ratio fell to 0.25 for post-famine-born women because there was a relative scarcity of famine-born men. If individuals valued the customary age of marriage, there should have been large marriage market effects on the famine-affected cohorts. Figure 2 plots the marriage rates for men and women by age in 1990 and 1982. First, the marriage rates for women in 1982 were similar for all three control age groups (26-28, 29-31, 32-34). There is no evidence that the age pattern of customary sex ratios in Figure 1(b) had any impact on the marriage rates of these women in 1982. The 26

19. The appendix shows that the sex ratios are influenced by differences in migration between males and females in the two provinces. In Sichuan, the higher out migration of females compared to men raises the sex ratio. The opposite is true in the case of Anhui.

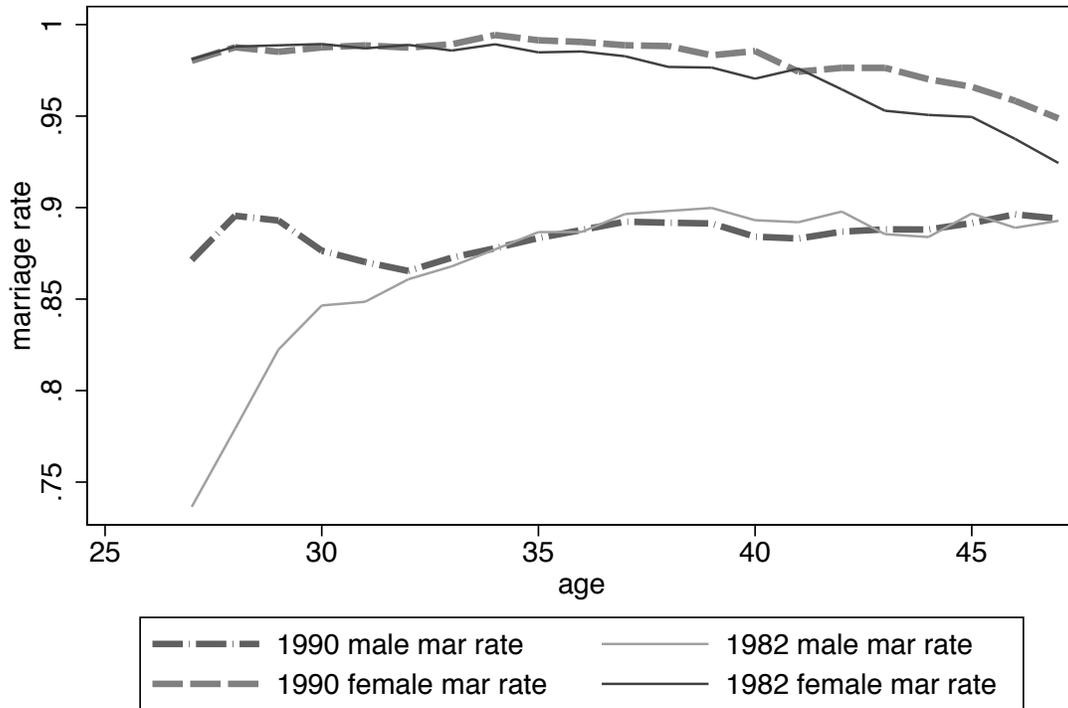


FIGURE 2. Sichuan marriage rates, 1990, 1982.

years olds in 1982, which overlapped with the famine-affected cohorts in 1990, do not display any unusual behavior in 1982. Thus, the female marriage rates in 1982 provide no evidence against using 1982 as a control group for 1990 behavior.

In both census years, 1982 and 1990, and at all ages, female marriage rates exceed 0.95. For women younger than age 40, marriage rates for women of the same age were essentially the same in 1990 and 1982. In other words, the famine-affected women in 1990 had the same marriage rates as their same age peers in 1982. Figure 1(b) showed that the famine-born women were in relative scarcity and the post-famine women were in relative surplus when compared to their customary spouses. This strongly suggests that the famine-affected women also married non-customary spouses and that these substitutions to a first order left the marriage rates of famine-affected women unchanged.

In general, the male marriage rates in both 1982 and 1990 were lower than the female marriage rates, consistent with the sex ratio being larger than one in rural Sichuan.

In 1982, the marriage rates for men followed a relatively smooth concave upward trend with age, with a small flattening out at age 30. There is no unusual movement at age 26, the year of overlap. There is no evidence that the age pattern of customary sex ratios in Figure 1(b) had any impact on the marriage rates of these men in 1982. Thus, the male marriage rates in 1982 provide no evidence against using 1982 as a control group for 1990 behavior.

In 1990, the marriage rates for famine-affected men were different from unaffected cohorts. The marriage rates of post-famine and famine-born men were higher than their older peers. Compared with 1982 men of the same ages, the marriage rates of pre-famine-born men in 1990 were not significantly different. Compared with 1982 men of the same ages, the marriage rates of famine and post-famine-born men in 1990 were significantly higher. Thus, both across-age comparisons in 1990, and across-years comparisons suggest that the marriage rates of famine and post-famine-born men were positively affected by the famine.

Based on marriage rates between 1990 and 1982, a tentative conclusion is that the marriage rates of famine-affected women in 1990 were unchanged. The marriage rates of pre-famine-born men were unaffected whereas the marriage rates of famine-born and post-famine-born men increased in 1990. These conclusions are consistent with the findings in Almond et. al. The marriage rates of famine-born men increased by less than 5% compared to their 1982 peers. The marriage rates of post-famine-born men increased by substantially more, 5% to 15% more than their 1982 peers. But famine-born men are scarce. It is therefore surprising that the increase in their marriage rates was so modest.

A caveat is necessary. While it is tempting to interpret the difference in male marriage rates for the post and famine-born cohorts as due to the famine as we do above, the case for such an interpretation is weak. The comparison of the female marriage rates showed no difference between treatment and control cohorts. There was no difference between post-famine treatment and control for men. The difference between treatment and control for both post and famine-born men was in the same direction. Thus, what we observed is a shift in marital behavior for young men in 1990. Other social and economic changes could have also affected the marital behavior of these young men in 1990, most notably, rising family incomes with the implementation of rural reforms in the late 1970s. The divergence in marriage rates for post-famine males grew as individuals were born further away from the famine years, contrary to the expectation that the effects of the famine were less for individuals born further away from the famine years. Thus, in addition to famine effects, there were potentially other shocks which also affected the marital behavior between 1982 and 1990. Whether the quantitative effects of these other shocks are large enough to bias our marriage matching results will be investigated in Section 5.4.

5.2. Total Gains

This section studies the changes in the marriage market between 1982 and 1990 through the lens of the CS model. To investigate the change in total gains for famine-affected cohorts, we first consider total gains of individuals who were born before the famine. Figure 3(a) shows total gains for 27, 30 and 33 years old women and their spouses from -3 years to +6 years older in 1982. Total gains for 30 and 33 years old women and their spouses were similar in 1982. Total gains for 27 years old women, while similar in shape, were lower than the other two age cohorts. This is expected because the marriage rate for 27 years old women was lower than the other two age

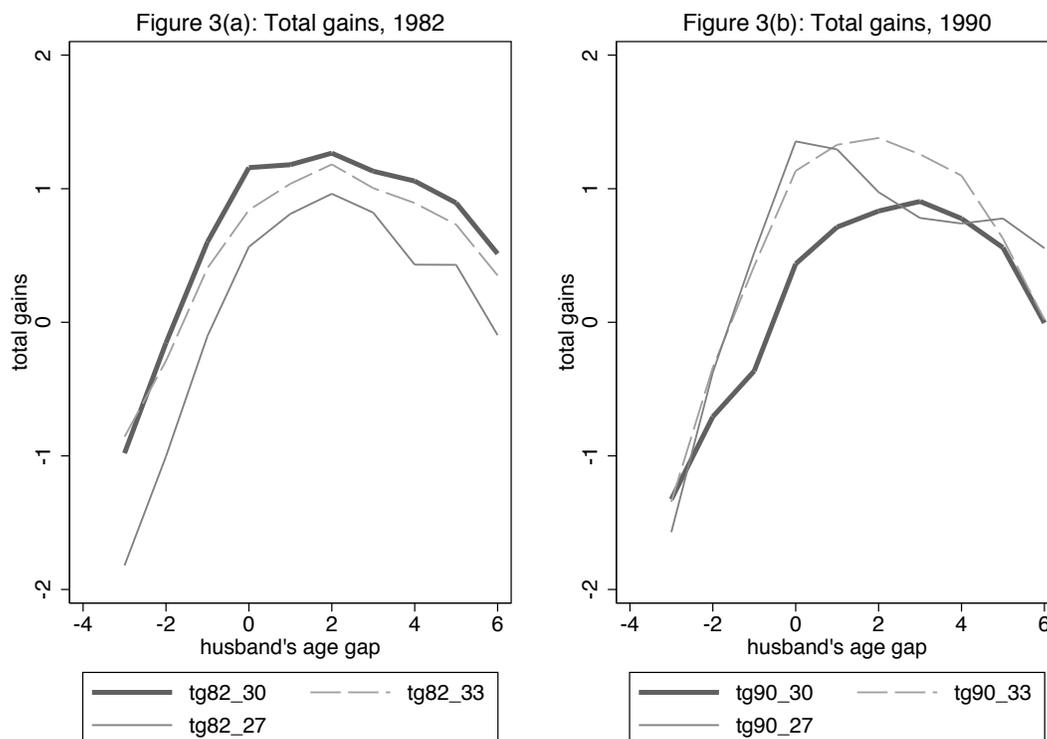


FIGURE 3

cohorts. Thus, it is reasonable to use the 1982 individuals as control groups for their same age 1990 peers.

Figure 3(b) plots total gains of three age cohorts of women in 1990, 27 (post-famine-born), 30 (famine-born) and 33 (pre-famine-born) and their husbands. Starting with pre-famine-born 33 years old women and their spouses, total gains is a smooth concave function in husband's age gap. Total gains from $[0,4]$ were relatively similar. Total gains of post-famine-born 27 years old women and their spouses were in general similar to the pre-famine women. Where they differ, and total gains were lower for post-famine women, were with husbands between 1 to 3 years older. For post-famine-born women, these husbands were famine-born men. Thus, marrying famine-born men resulted in lower total gains to marriage relative to the pre-famine women with spouses 1 to 3 years older.

Total gains for famine-born women, age 30, and their spouses were significantly lower than that of pre and post-famine-born women. Figure 2 shows that the marriage rates of famine-born women were similar to pre and post-famine women. There are some small differences in the marriage rates of the husbands (measured by age gaps) of famine-born and other famine-affected women. But there were large differences in the customary marriage sex ratios as shown in Figure 2. These large differences in the customary marriage sex ratios should have resulted in significantly different marriage rates for famine-born and other famine-affected women. But because the marriage rates for all the famine-affected women were roughly the same in spite of large differences

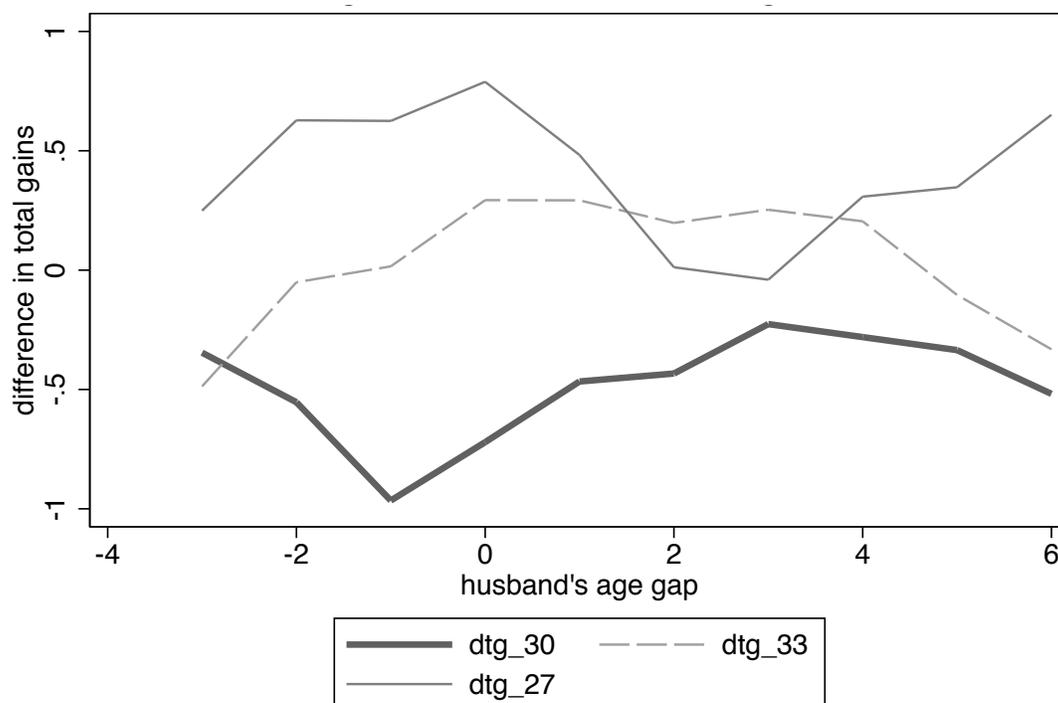


FIGURE 4. 1990-1982 total gains.

in customary marital sex ratios, total gains for the famine-born women had to be lower than the other famine-affected women.

Figure 4 presents the difference in total gains between 1990 and 1982 for the same age women and their spouses. The differences in total gains were negative for the famine-born women, age 30, for all spousal ages. The difference in total gains for 33 years old women between the two census was mostly a little larger than zero. It was negative for marriages with famine-born husbands. The difference in total gains for 27 years old women was largely above zero. It dipped to zero for famine-born husbands. Both pre and post-famine-born women primarily had larger total gains from marriage than their same age 1982 counterparts unless they married famine-born husbands. The famine-born women and their spouses had total gains that were lower than their 1982 counterparts. This shows that the famine had a significant, concentrated negative effect on famine-born cohorts, both men and women.

Finally, the non-linear heterogeneous changes in total gains for famine-affected cohorts are hard to explain by other factors which may have also affected the marriage market between the two periods.

5.3. A Structural Interpretation of the Changes in Total Gains

So far, our discussion of the changes in total gains has been descriptive. The change in marital behavior for the pre, post and famine-born cohorts are sufficiently different from each other to suggest that the famine had significant effects on the marital

behavior of the famine affected cohorts. How important are quantity versus quality effects? We will first argue that it is implausible that the observed total effects of the famine are due exclusively to one or the other effect.

Figure 2 showed that the marriage rate of famine born women was similar to their same aged counterparts eight years earlier, and the marriage rate of famine born men was higher than those of their same aged counterparts. So if famine born individuals were less desirable in the marriage market, their marriage rates suggest that both genders were able to take advantage of their relative scarcity to attain their relatively high marriage rates. On the other hand, it is difficult to explain the differences in total gains in Figure 4 with quantity effects alone. If famine-born women were relatively scarce, and there was no change in the marital attractiveness of these women, it is hard to explain why the difference in total gains was negative for famine-born women, age 30, for all spousal ages. Thus, the changes in marriage rates and total gains suggest that both quantity and quality effects are important in explaining the change in marital behavior of the famine-affected cohorts. The next step is to quantify the importance of the two factors.

We will use the CS model to do that decomposition. According to CS, equation (1) says that the estimated total gains are estimates of structural parameters of the CS model. In particular, the total gains for an $\{i, j\}$ marriage is the average payoff for a spouse in that marriage relative to them remaining unmarried. The difference in total gains measures the change in this relative payoff between t' and t .

It is easy to use CS to interpret Figure 4. The differences in total gains were negative for the famine-born women, age 30, for all spousal ages. They are particularly negative for same age or slightly younger husbands. In other words, the marital output of marriages with famine-born women was lower compared to their same age counterpart in 1982. And if their husband is also famine-born, the payoff was even lower. On the other hand, the marital output of marriages with pre-famine-born women was marginally higher than their same aged counterpart in 1982, validating the Meng and Qian (2010) hypothesis that children who survived the famine were positively selected. Finally, the difference in total gains for 27 years old women was largely positive, but dipped below zero for famine-born husbands. What this means is that the marital outputs of both famine-born men and famine-born women suffered a substantial drop relative to their 1982 same aged peers. Thus, the CS model unambiguously shows that the famine-born cohort suffered a significant reduction in marital attractiveness.

5.4. Testing the Structural Interpretation of Changes in Total Gains

Marriage matching data from the two census cannot be used to test the internal validity of the CS model. Equivalently, CS has no over-identifying restriction. If CS is false, our first difference methodology will generate changes in total gains independent of any actual changes in marital output. We argued that the observed changes in total gains for the famine-affected cohorts are consistent with a fall in marital output for the famine-affected cohorts. This argument is based on our interpretation of the data and not a test of the model's internal validity. With no change in marital output, non-CS

models can generate changes in total gains as a result of changes in population vectors between two time periods.²⁰

There are other reasons why our estimation strategy may also not be valid even if CS is valid. There were many social and economic changes in China between 1980 to 1990 that might affect our estimates of total gains. For example, Wang and Zhou (2010) show that divorce rates were rapidly rising in that period. Rural to urban and inter-provincial migration was also increasing in that period. To the extent that these social and economic changes have significant intercept and trend changes on marital behavior, they may contaminate our estimates of the changes in marital gains. The extent of the potential biases due to these concerns is empirical which is what we will turn to next..

The objective of this section is to test the external validity of the CS model in our context. Since marriage-matching data alone cannot be used to test the model, we will use other data to test CS. Many studies have shown that the famine-affected cohorts suffered adverse health outcomes. Ill health will reduce marital output of that couple. To the extent that cohort-specific lower marital outputs due to the famine were observed and/or anticipated before marriage, our CS measure of total gains should predict post-marital ill health outcomes.

An extreme negative marital outcome is infertility in the marriage. In the 1982 and 1990 census, each woman was asked how many children did they give birth to. We define childlessness of $\{i, j\}$ matches in year t , C_{ij}^t , as the logarithmic fraction of those married couples who reported not having given birth as of the time of the census. Childlessness is not the same as infertility because a childless couple may have children in the future. Thus C_{ij}^t depends on behavioral considerations as well as infertility.

We use C_{ij}^t as a proxy for low marital output in $\{i, j\}$ marriages at time t . Infertility is an extreme health outcome which affects a small fraction of married couples. For both census years, less than 2% of married women at each age between 30 and 50 reported having no birth.²¹ Childlessness is only observed after the marriage. Thus while marriage market participants cared about infertility, it is unlikely that they could have accurately predicted infertility among potential spouses. What they were more likely to be able to observe at the time of the marriage decision was the general health and socioeconomic well-being of their potential spouses. Consistent with the empirical evidence on childhood or fetal malnutrition and later reproductive success, we assume that infertility was negatively related to adult good health and high socioeconomic well-being.

There are several determinants of the level of childlessness and changes in childlessness between the two census. (i) From a level perspective (within a census), childlessness is affected by life-cycle considerations. (ii) Between the two census, changes in childlessness may also be affected by other social and economic changes that occurred between 1982 and 1990. (iii) Holding marital output constant, changes in

20. We thank one of the referees for providing such an example.

21. For lifecycle reasons, the childlessness rates higher for younger married women.

childlessness may be affected by changes in marriage market conditions which affect intra-household resource allocations. Rasul 2006, Ashraf et. al. 2010, and others have shown that husbands prefer to have more children than their wives. So an increase in the sex ratio (ratio of men to women) will weakly increase childlessness. And (iv), childlessness is affected by infertility, ill health and/or poor socioeconomic outcome for the married couple.

We are interested in studying how changes in total gains, as a measure of changes in marital output, predict future childlessness. So we are interested in the effect of the fourth factor discussed in the previous paragraph. In order to do that, we have to control for the first three factors in the previous paragraph. Our empirical work proceeds in two steps. First, holding factors (i), (ii) and (iii) constant, following Almond et. al. (2007), we show that a more severe famine effect, as measured by lower population sizes for the $\{i, j\}$ marital matches, resulted in more childlessness for couples in $\{i, j\}$ matches. Second, holding factors (i), (ii) and (iii) constant, we investigate whether an increase in $\{i, j\}$ total gains reduces childlessness in $\{i, j\}$ marriages. Finally, we compare, with and without holding factors (i), (ii) and (iii) constant, whether changes in total gains have more predictive power than changes in population sizes in predicting changes in childlessness.

Using samples from two different age ranges, we regress the difference in childlessness between the two census years, $\Delta C_{ij}^t = C_{ij}^{90} - C_{ij}^{82}$, on differences in total gains, $\Delta \pi_{ij}^t$, differences in sex ratio, $\Delta SR_{ij}^t = \Delta(\ln m_i^t - \ln f_j^t)$, differences in population, $\Delta P_{ij}^t = \Delta \ln(m_i^t + f_j^t)$, and unrestricted quadratic functions in spousal ages.²² The first difference estimation strategy controls for variation in childlessness due to lifecycle behavior, factor (i) above. We include an intercept in all the regressions to control for a common time effect which affects all $\{i, j\}$ pairs across the two census. The unrestricted quadratic in ages controls for smooth changes by spousal ages in childlessness which may be due to other social and economic changes between 1982 and 1990, factor (ii) above.²³ Changes in the sex ratio, ΔSR_{ij}^t , control for changes in marriage market conditions which may affect the intra-household allocation of resources, factor (iii) above. To proxy for changes in marital output, we use two different measures, differences in total gains, $\Delta \pi_{ij}^t$, and differences in population, ΔP_{ij}^t . Differences in population have been used in this literature to proxy for changes in the health status of the famine-affected cohorts (E.g. Almond, et. al.). An increase in ΔP_{ij}^t is an increase in health in 1990 relative to 1982.

A qualification is necessary. CS is a static model. If the above concerns lead to delay in marriage for the famine-affected cohorts, and delay leads to increases in childlessness, then measured childlessness may overestimate true childlessness. How this concern will affect our estimates is hard to evaluate.

22. Age controls are i, j, i^2, j^2 and $i \times j$.

23. There are two main changes of potential concern: (i) The marital reform act of 1980 which led to an increase in the legal age of marriage. (ii) Market liberalization of the rural economy in the late seventies and eighties. Neither change nor other changes specifically targeted the famine-affected cohorts.

To be included in our samples, an $\{i, j\}$ cell must have non-missing ΔC_{ij}^t and $\Delta \pi_{ij}^t$ values. It is important to note that if every $\{i, j\}$ couple in the census in year t reports at least one birth, then C_{ij}^t is undefined because we are using the log of the fraction of couples without children. So ΔC_{ij}^t is defined only if there is at least one incidence of childlessness in each census year. This censoring is more prevalent in sparse $\{i, j\}$ marriage cells.²⁴

Our first sample consists of spouses between the ages of 18 to 50. To focus more closely on the famine-affected cohorts in 1990, our second sample consists of spouses between the ages of 25 to 40. Table 3 provides the summary statistics for the two samples. There were 419 observations in the first sample. The mean of ΔC_{ij}^t is -0.403 implying that childlessness fell between 1982 and 1990. Over the same period, total gains, π_{ij}^t , increased by an average of 0.126 while the proportion of males fell, and the total population increased.

For the sample with the narrower age range, there are 149 observations. The mean of ΔC_{ij}^t was more negative than with the larger sample. This means that the changes in childlessness with the famine-affected cohorts in 1990 were larger than the excluded cohorts, which is to be expected if the famine negatively affected fertility. The standard deviation of ΔC_{ij}^t is also greater than in the larger sample. The larger dispersion in childlessness—even though the age range is narrower—suggests that something other than life-cycle factors affected childlessness for the smaller age range sample. Consistent with the famine, the mean of ΔP_{ij}^t is smaller than in the larger sample.

Figure 5 shows the plot of ΔC_{ij}^t against $\Delta \pi_{ij}^t$. The line is a simple linear regression through the points. In general, there is a negative relationship between the change in childlessness and the change in total gains. As discussed above, there are two other factors, (ii) and (iii), which are not controlled for in Figure 5.

Table 4 presents estimates of OLS regression of ΔC_{ij}^t on ΔSR_{ij}^t , ΔP_{ij}^t , and age effects. Our first sample consists of marriages in which the spouses were between 18 and 50 years of age. There are 434 observations out of a potential 1089 observations.²⁵ Column (1) includes only two covariates, ΔSR_{ij}^t and ΔP_{ij}^t . The point estimate for ΔSR_{ij}^t is 0.139 with a standard error of 0.044, providing evidence that marriage market tightness matters, with an increase in the sex ratio of 1% increasing childlessness by 0.14%. The sign of the point estimate is consistent with the finding in other papers that wives generally prefer less children than their husbands. The point estimate for ΔP_{ij}^t is -0.190 with a standard error of 0.093, implying that an increase in the population by 1% decreases childlessness by 0.19%, which supports the hypothesis that the famine had negative health and fertility effects for the famine-affected cohorts. Since there is population growth in general across the two census, as well as other social and

24. For example if $\mu_{ij}^t < 10$, it is common for every marriage which is observed to have at least one birth.

25. 330 missing observations were due to no observed marriage in the $\{i, j\}$ cell in 1982 and/or 1990. The rest of the missing observation were due to no childless marriage in the $\{i, j\}$ cell in 1982 and/or 1990.

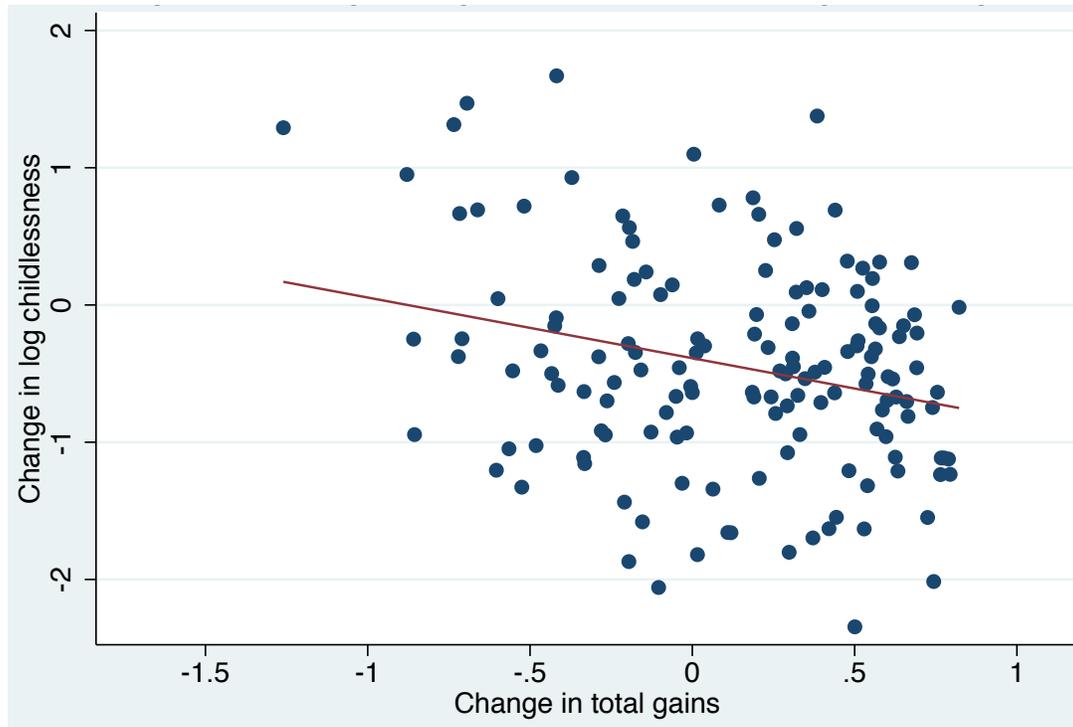


FIGURE 5. Change in log childlessness on change in total gains.

economic changes between the two census which are not age neutral, Column (2) adds an unrestricted quadratic function of spousal ages to the regression. Adding age effects does not change either the quantitative or qualitative importance of ΔSR_{ij}^t and ΔP_{ij}^t in determining changes in childlessness.

To focus more narrowly on the famine-affected cohorts in the 1990 census, our second sample restricts the spousal ages to lie between 25 and 40 years of age, giving us 149 observations. Column (3) includes only two covariates, ΔSR_{ij}^t and ΔP_{ij}^t . The point estimate for ΔSR_{ij}^t is 0.224 with a standard error of 0.109. Now a 1% increase in the sex ratio increases childlessness by 0.224%, which is larger than the estimate from the smaller sample. The standard error doubles which is not surprising given the smaller sample. The point estimate for ΔP_{ij}^t remains negative and roughly of the same magnitude as that in the larger sample, but is estimated less precisely. Column (4) adds an unrestricted quadratic function in spousal ages. While the point estimates are qualitatively similar to those in the earlier columns, the precision of the estimates is reversed from that in column (3).

In summary, the regressions in Table 4 show that changes in marriage market tightness, and changes in health of the married cohorts (as proxied by changes in population size) affect changes in childlessness. Adding age controls to the regressions does not change the point estimates of ΔSR_{ij}^t and ΔP_{ij}^t significantly but increases their standard errors.

Table 5 presents estimates of OLS regression of ΔC_{ij}^t on $\Delta \pi_{ij}^t$ and other covariates. There are no age effects in Table 5. Columns (1) to (4) use the larger age range

sample. Column (1) includes only $\Delta\pi_{ij}^t$. The point estimate is -0.362 and the standard error is 0.054. A 1% increase in the change in total gains (more than two standard deviations) lowers the change in childlessness by 0.36%, which is slightly less than half the standard deviation of ΔC_{ij}^t . The negative point estimate is consistent with our interpretation that ΔC_{ij}^t is a good proxy for the change in marital output between the two census. By comparing R^2 , $\Delta\pi_{ij}^t$ by itself explains more than twice the variation in ΔC_{ij}^t than ΔSR_{ij}^t and ΔP_{ij}^t did in column (1) of Table 4.

Column (2) adds ΔSR_{ij}^t to the regression. The point estimate on $\Delta\pi_{ij}^t$ and its standard error barely changes from that in column (1). The point estimate on ΔSR_{ij}^t is 0.144 and the standard error is 0.043, which are comparable to that in column (1) in Table 4. So whether we control for “health changes” via $\Delta\pi_{ij}^t$ or ΔP_{ij}^t , the effect of changes in ΔSR_{ij}^t on ΔC_{ij}^t is similar. This evidence supports the hypothesis that when wives have more bargaining power in marriage, they are likely to demand less children. Column (3) includes $\Delta\pi_{ij}^t$ and ΔP_{ij}^t as covariates. The point estimate and standard error on $\Delta\pi_{ij}^t$ are similar to that from the first two columns. The point estimate on ΔP_{ij}^t has the wrong sign and the estimated standard error is large. Column (4) includes $\Delta\pi_{ij}^t$, ΔSR_{ij}^t and ΔP_{ij}^t as covariates. The point estimate on $\Delta\pi_{ij}^t$ is more negative, while the point estimate on ΔP_{ij}^t is positive but statistically insignificant at the 5% significance level. As a health or marital output indicator, the point estimate on ΔP_{ij}^t has the incorrect sign. Since both $\Delta\pi_{ij}^t$ or ΔP_{ij}^t are health indicators, the estimated incorrect sign on ΔP_{ij}^t is probably due to collinearity. Still, the results in columns (3) and (4) suggest that $\Delta\pi_{ij}^t$ is a better indicator of changes in marital output caused by the famine than ΔP_{ij}^t .

Columns (5) to (8) use the smaller sample to repeat the same regressions as columns (1) to (4). In general, the estimated coefficients on $\Delta\pi_{ij}^t$ are more negative. Although the estimated standard errors are larger, the point estimates are all statistically significant at the 5% significance level if not lower. The point estimates on ΔSR_{ij}^t remains positive but the standard errors are larger. The point estimates on ΔP_{ij}^t continue to have the incorrect signs and the standard errors remain large.

Table 6 adds unrestricted spousal age quadratics as covariates to the regressions in Table 5. Comparing the R^2 's between Table 5 and 6, age effects explain a significant amount of variation in ΔC_{ij}^t . With the larger sample, the point estimates and standard errors on $\Delta\pi_{ij}^t$ and ΔSR_{ij}^t remain similar. The point estimate on ΔP_{ij}^t becomes positive (incorrect sign) and statistically significant. With the smaller sample in columns (5) to (8), the point estimates for $\Delta\pi_{ij}^t$ remain negative but is only statistically different from zero at the 5% significance level in columns (5) and (6). The estimated coefficients on ΔP_{ij}^t and ΔSR_{ij}^t are imprecise.

Summarizing the results in Tables 5 and 6, age effects explain a significant amount of variation in ΔC_{ij}^t . The variation that is explained is largely orthogonal to that explained by the other covariates. With or without age effects, the point estimates on $\Delta\pi_{ij}^t$ are consistently between -0.3 to -0.5 when ΔP_{ij}^t (the other health proxy) is excluded, and are usually statistically different from zero at the 5% significance level or lower. Together with the results in Table 4, the results strongly suggest that changes

in total gains are good predictors of changes in childlessness due to the famine, even after controlling for age-varying time effects, changes in marriage market tightness, and even when changes in total population are included in the regressions.

We investigated the robustness of our results by:

- (i) using an alternative definition of $\Delta P_{ij}^t = \Delta(\ln(m_{ij}^t) - \ln(f_{ij}^t))$.
- (ii) using an alternative definition of ΔC_{ij}^t where C_{ij}^t is measured in level rather than in log. In this case, we include cells in which $C_{ij}^t = 0$. This alternative definition increased the sample sizes by about 50%.
- (iii) weighting the observations in the regressions by the number of marriages in each cell. This weighting scheme gives more weight to more common types of marriages.
- (iv) using an alternative dependent variable, $\Delta \ln n_{ij}^t$, which is the difference in the log of the number of surviving children as of the time of the census.

In the first three robustness investigations, the point estimates on $\Delta \pi_{ij}^t$ are consistently negative. The t statistics are approximately the same in (i) and (ii). The standard errors are larger in (iii) because there is less variation in $\Delta \pi_{ij}^t$ when μ_{ij}^t is large. The point estimates and standard errors for ΔP_{ij}^t and ΔSR_{ij}^t were also similar to those reported here.

In robustness check (iv), our use of $\Delta \ln n_{ij}^t$ as an indicator for total gains is based on the assumption that total gains is non-decreasing in the number of surviving children, which may not be true. But because the observed fraction of childless couples for most marital matches was low, there is merit to checking whether a different dependent variable will generate similar results. We ran the equivalent specifications for columns (1) to (4) in Table 6, now with 680 observations per regression rather than 419 in Table 6. For all four specifications, the estimated elasticity of $\Delta \ln n_{ij}^t$ with respect to $\Delta \pi_{ij}^t$ was approximately 5% with a standard error around 2.5%. An increase in total gains over the two time periods for an (i, j) marital match was associated with an increase in the average number of surviving children for those matches between the two time periods. This finding is consistent with that using childlessness as a dependent variable.

The evidence in this section provides external validity to the CS measure of $\Delta \pi_{ij}^t$ as a change in marital output between the two period. We have shown that variations in population vectors as proxied by ΔSR_{ij}^t affected ΔC_{ij}^t . If CS is largely incorrect, then $\Delta \pi_{ij}^t$ will contain both changes in marital output as well as changes in population vectors. Our empirical finding that our point estimates of the effect of $\Delta \pi_{ij}^t$ on ΔC_{ij}^t are largely independent of variation in ΔSR_{ij}^t suggests that CS is largely correct.

6. Quantity versus Quality

Given the drop in the relative marital output of famine-born spouses, why did their marriage rates not drop? An explanation is that they were in relatively short supply in the marriage market. Since we do not reject the test of external validity of the CS model in our context, to quantify the effect of the changes in quantities on marital behavior,



FIGURE 6

we use the 1982 CS structural estimates to predict what the marriage distribution in 1990 would have been with 1990 population vectors and 1982 estimated parameters (i.e. $\hat{\mu}_{ij}^t$ using equation (1)).

The predicted marriage rates from the CS MMF behave sensibly. In Figures 6(a) and 6(b), the predicted marriage rates are above average for the famine-born cohorts and below average for the adjacent aged birth cohorts. No accounting constraint is violated. Note that actual female marriage rates were over 0.95 for most ages. Even though there were large changes in sex ratios of the customary spousal age differences for the famine-born cohorts, their predicted marriage rates remained below 1.²⁶ The predicted female marriage rates for famine-born cohorts were very similar to those predicted for adjacent aged cohorts. In other words, the CS MMF respects both the general equilibrium accounting constraints of MMFs and also captures the flexibility of individuals in their marital choices.

In Figure 6(a), famine-born and post-famine males had lower marriage rates than predicted by CS. Pre famine-born males had higher marriage rates than predicted by CS. Changes in relative scarcities of the different types of individuals caused by the

26. As shown in our working paper, this feature is violated by the standard demography MMF (Scheon 1981)..

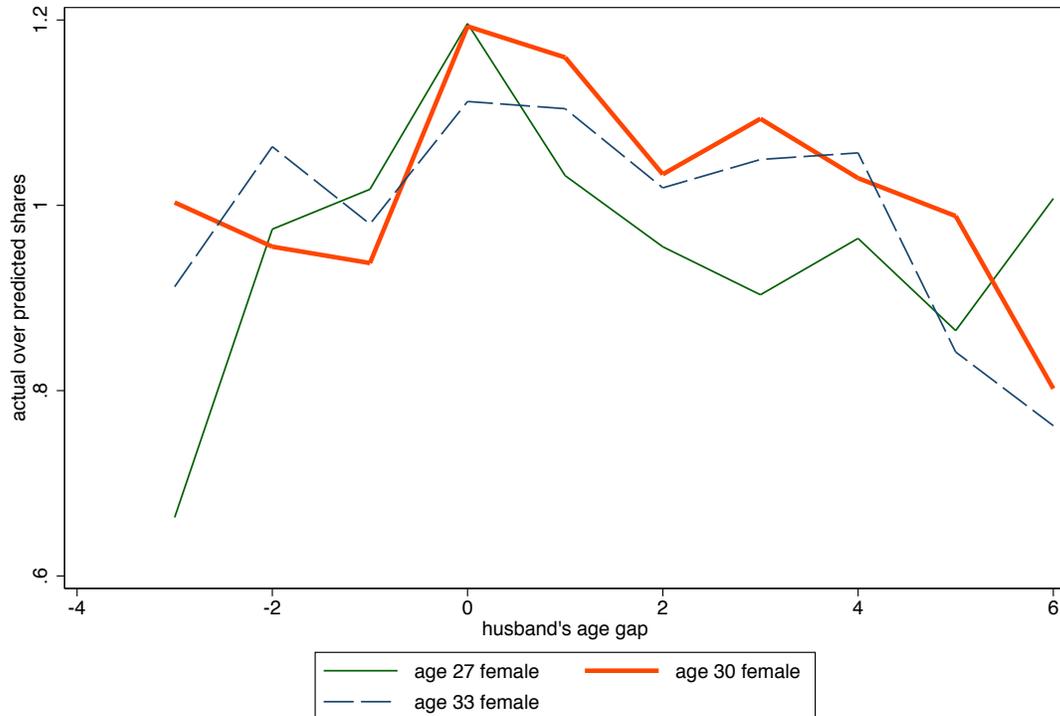


FIGURE 7. 1990 ratio of actual to predicted shares.

famine cannot explain these discrepancies. The famine must have also changed marital attractiveness to marriage for these cohorts.

Figure 6(b) shows that the discrepancies between predicted and actual female marriage rates were small. CS is able to generate predicted male marriage rates that were highly responsive to changes in population supplies and female marriage rates that were marginally responsive. It is clear that changes in population supplies alone cannot explain the observed changes in marriage rates. We also need to account for changes in marital attractiveness of the famine-affected cohorts.

Figure 7 shows the ratio of actual to predicted marriage shares for women in 1990 by the age gaps of their husbands. Consider first the marriage shares of 39 years old famine-born women. Their actual shares of own age and pre-famine-born husbands exceed the predicted shares. The actual shares are more concentrated among these husbands than is predicted by changes in population vectors alone. In other words, drop in marital attractiveness of the famine-born cohort led them to marry even more of their own type. Due to adding up constraints in the marriage market, when actual shares are different from predicted shares for one cohort, they also must be different for other cohorts. This shows up for pre-famine-born 33 years old women. Their actual shares of own age and older husbands also exceed the predicted shares. Finally for post-famine-born 27 years old women, the actual shares of own age husbands exceed the predicted shares and the reverse occurs for famine-born husbands. This is evidence that post-famine-born women “avoided” famine-born husbands.

Figures 4, 6(a), 6(b) and 7 provide a unified summary of the effects of the famine on the marital behavior of famine-affected cohorts. Figure 4 shows that the relative marital output of famine-born cohorts fell substantially. By itself, the drop in relative marital output would have substantially reduced the marriage rates of the famine-born cohorts. The famine also substantially reduced the relative supply of the famine-born cohorts. Figures 6(a) and 6(b) show that these reductions would have substantially changed the marriage rates of the famine-affected cohorts. To a first order, the simultaneous changes in quantities and qualities cancel each other out and result in small changes in marriage rates for the famine-affected cohorts. Figure 7 showed that famine born women disproportionately married more husbands of their own age than could be explained by population changes alone. Pre-famine born women also disproportionately married more husbands of their own age than could be explained by population changes alone. In other words, pre-famine born men, rather than marrying their customary brides three years younger, preferred to marry their same-aged brides.

6.1. Robustness Checks

Our 2008 working paper repeated the analysis with data from rural Anhui. Our results on the changes in total gains and decomposition of changes in marriage rates behavior were qualitatively and quantitatively similar to what we obtained here.

The famine had a negative impact on the health and human capital endowments of famine-born individuals (see Almond et. al. (2007), Gorgens et. al. (2005), St. Clair et. al. (2005), and Luo, Mu and Zhang (2006)). The CS model admits marriage matching along multiple dimensions. So we also estimated the CS model for both Sichuan and Anhui where individuals match by both age and educational attainment. To a first order, adding educational attainment as an additional attribute for individuals to match on did not significantly change the estimates of losses in average total gains across birth cohorts. There were larger effects within an age cohort. Famine born individuals with low education suffered a larger drop in marital attractiveness than their more educated peers. As a result, their predicted marriage rates were slightly lower than their higher educated peers. However the average predicted marriage rates across both two groups did not differ significantly from the predicted marriage rates using the model with only age that is discussed in this paper.

7. Conclusion

This paper has three conclusions. First, there were little changes in the marriage rates of the famine-born cohorts relative to their adjacent aged peers in 1990 or same age peers in 1982. To a first approximation, our decomposition shows that the benefit that the famine-born cohort derived from their relative scarcity is offset by their decline in marital attractiveness. The main substantive conclusion of this paper is that the small observed changes in marriage rates of the famine-born cohorts are due to a substantial decline in their marital attractiveness.

Second, we provide evidence of external validity of the CS measure of total gains as a measure of marital output. Variations in total gains between the two census predicted variations in future childlessness of those marital matches. Substantively, we add to the evidence that early nutritional deprivation due to a famine affects adult childlessness.

Finally, our main methodological conclusion is that allowing for general equilibrium effects and estimating heterogeneous treatment effects are important because the pre-famine-born, famine-born and post-famine-born cohorts experienced very different but linked outcomes. Ignoring these effects in estimation will obscure the losses in the marriage market experienced by the famine-born.

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Table 1 Sample Statistics on Sichuan Rural Counties, 1982 and 1990 Census

Census Year	1982	1990
Province	Sichuan	Sichuan
Numbers of men aged 18-50	188,081	225,261
Numbers of women aged 18-50	179,727	221,750
Share married men	0.737	0.719
Share married women	0.846	0.802
Mean (woman's - spouse's age)	-3.216	-2.664
Share of high edu (male)	0.302	0.440
Share of high edu (female)	0.195	0.295

Notes: high edu: graduate from elementary school.

Table 2 1990 Famine Affected Cohorts and Their 1982 Comparison Groups

	Pre Famine	Famine	Post Famine
Birth Years	1956-1958	1959-1961	1962-1964
1982 age	24-26	21-23	18-20
1990 age	32-34	29-31	26-28

Table 3 Summary Statistics of childlessness analysis samples

Variables	Age range 18-50			Min	Max
	Obs	Mean	Std. Dev.		
ΔC_{ij}^t	419	-0.40338	0.731232	-2.34515	1.750211
$\Delta \pi_{ij}^t$	419	0.126467	0.569819	-2.52071	2.082992
ΔSR_{ij}^t	419	-0.16255	0.685185	-2.65066	1.503534
ΔP_{ij}^t	419	0.205898	0.346508	-1.15733	1.439799
i	419	35.10024	8.553209	18	50
j	419	30.61814	8.797143	18	50
Variables	Age range 25-40			Min	Max
	Obs	Mean	Std. Dev.		
ΔC_{ij}^t	149	-0.44695	0.768268	-2.34515	1.66987
$\Delta \pi_{ij}^t$	149	0.134326	0.455414	-1.25981	0.821832
ΔSR_{ij}^t	149	0.095165	0.628872	-1.55033	1.503534
ΔP_{ij}^t	149	-0.00229	0.335564	-1.15733	0.403261
i	149	33.89262	4.242865	25	40
j	149	30.73826	4.180709	25	40

Table 4 Changes in Log Childlessness by Changes in SR and POP

	(1)	(2)	(3)	(4)
	$\Delta C'_{ij}$	$\Delta C'_{ij}$	$\Delta C'_{ij}$	$\Delta C'_{ij}$
$\Delta SR'_{ij}$	0.139 (0.044)**	0.139 (0.050)**	0.224 (0.109)*	0.042 (0.126)
$\Delta P'_{ij}$	-0.190 (0.093)*	-0.226 (0.107)*	-0.141 (0.199)	-0.287 (0.212)
Age quadratic		y		y
Age range	18-50	18-50	25-40	25-40
Observations	434	434	149	149
R-squared	0.03	0.07	0.04	0.16

Notes: Robust standard errors in parentheses

* significant at 5%; ** significant at 1%

Table 5 Changes in Log Childlessness by Changes in Total Gains

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
	ΔC_{ij}^t	ΔC_{ij}^t	ΔC_{ij}^t	ΔC_{ij}^t				
$\Delta \pi_{ij}^t$	-0.362 (0.054)**	-0.349 (0.052)**	-0.382 (0.059)**	-0.392 (0.058)**	-0.442 (0.141)**	-0.372 (0.148)*	-0.748 (0.235)**	-0.667 (0.236)*
ΔSR_{ij}^t		0.144 (0.043)**		0.161 (0.043)**		0.143 (0.108)		0.121 (0.107)
ΔP_{ij}^t			0.067 (0.094)	0.154 (0.093)			0.548 (0.295)	0.509 (0.291)
Age range	18-50	18-50	18-50	18-50	25-40	25-40	25-40	25-40
Obs.	419	419	419	419	149	149	149	149
R-sq.	0.08	0.10	0.08	0.10	0.07	0.08	0.09	0.10

Notes:

Robust standard errors in parentheses

* significant at 5%; ** significant at 1%

Table 6: Changes in Log Childlessness by Changes in Total Gains and Age Effects

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
	ΔC_{ij}^t	ΔC_{ij}^t	ΔC_{ij}^t	ΔC_{ij}^t	ΔC_{ij}^t	ΔC_{ij}^t	ΔC_{ij}^t	ΔC_{ij}^t
$\Delta \pi_{ij}^t$	-0.381 (0.052)**	-0.364 (0.052)**	-0.473 (0.066)**	-0.463 (0.065)**	-0.342 (0.159)*	-0.337 (0.160)*	-0.526 (0.298)	-0.526 (0.298)
ΔSR_{ij}^t		0.107 (0.049)*		0.120 (0.050)*		0.037 (0.123)		0.043 (0.124)
ΔP_{ij}^t			0.291 (0.128)*	0.321 (0.128)*			0.300 (0.386)	0.309 (0.389)
Age- quadratic	yes	yes	yes	yes	yes	yes	yes	Yes
Age range	18-50	18-50	18-50	18-50	25-40	25-40	25-40	25-40
Obs.	419	419	419	419	149	149	149	149
R-sq.	0.13	0.14	0.14	0.15	0.18	0.18	0.18	0.18

Notes:

Robust standard errors in parentheses

* significant at 5%; ** significant at 1%