DISCUSSION PAPER SERIES

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IZA DP No. 4243

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June 2009

Forschungsinstitut
zur Zukunft der Arbeit
Institute for the Study
of Labor

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June 2009

ABSTRACT

Large Demographic Shocks and Small Changes in the Marriage Market*

This paper provides non-parametric estimates of the total effects of famine in China on marital behavior of famine affected cohorts in rural areas of Sichuan and Anhui. The reduced form estimates incorporate general equilibrium and heterogeneous treatment effects, two important components of equilibrium marital behavior. Next the paper uses a structural model of the marriage market to decomposed observed marital outcomes into quantity and quality effects. The structural estimates show that the famine reduced the marital attractiveness of the famine-born cohort. The conclusion is that the small observed changes in marriage rates of the famine born cohort are due to a significant decline in marital attractiveness.

JEL Classification: J12

Keywords: marriage market, famine

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* We thank SSHRC for financial support. We also thank various seminars' participants for useful comments.

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 industry: People were moved from villages to work in urban factories, while agricultural land and labor were directed towards steel production in "backyard furnaces". 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. This was in addition to a state rationing system that already favored urban industrial workers to the detriment of the villages.

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 births.[2] In general, the countryside was struck much harder than cities.[3] The economic experiment was abandoned by early 1962. The mortality rate quickly 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 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.

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

[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.

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 marriageble age.[4] How the marital attractiveness of these individuals compares with that of the famine born cohort is difficult to assess apriori.

The upshot of the above is that there are large quantity and complicated quality changes to the famine affected cohorts. The affected 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. Due to the complex changes and responses, there were heterogenous responses by different affected cohorts.

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 and Anhui, two agricultural provinces which were severely affected by the famine.[5] We study the rural populations because the famine disproportionately affected rural rather than urban communities.[6]

Figures 1 and 1a show the distributions of individuals by age in the 1990 Census for rural Sichuan and Anhui, respectively. Due to a high long-run birth rate and mortality rates that increase with age, population by age

[4]Meng and Qian (2006) provide direct evidence on the health changes of these cohorts.

[5]In 1957, 86 and 92 percent of the labor force were in the primary (agriculture, fisheries and forestry) sector in Sichuan and Anhui, respectively. Between 65-70 percent 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.

should be declining with age. The famine born cohort was 29-31 years old men and women. As we can see in Figure 1 for Sichuan, 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. Figure 1a plots population by age for Anhui, which was also heavily affected by the famine. Again, the drop in population began with the pre-famine cohort and accelerated for the famine born cohort, and then recovered quickly after the famine.

The paper provides two kinds of estimates of the effects of famine on marital behavior. First, using a first-difference methodology, we non-parametrically estimate the total effect of the famine on the marital outcomes of the famine affected cohorts. To this end, we estimate a set of 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 different 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 total effect of the famine on marital behavior. Our total effect estimates incorporate both general equilibrium and heterogeneous treatement effects. The estimates 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. Due to the unusual non-linear effects of the famine on marital outcomes, we will argue that our choices of the treatment and control groups are credible. We implement our first difference empirical strategy by comparing, within a province, the marital behavior of the famine affected cohorts in 1990 to their same age counterparts in 1982.[7]

Second, using a non-parametric structural marriage matching model, we decompose the estimated total effects into quantity and quality effects. Our structural model will be the CS marriage matching model (Choo and Siow 2006).[8] The non-parametric total gains estimates provide one to one non-

[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.

[8]A marriage matching function is a production function for marriages (Pollard 1997,

parametric estimates of the structural parameters of the CS model.

For both Sichuan and Anhui, 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. The main methodological conclusion is that allowing for general equilibrium effects and estimating heterogenous treatment effects are important because the pre-famine born, famine born and post-famine born cohorts experienced very different but linked outcomes.

Can the decline in marital attractiveness of the famine born cohort be captured by differences in educational attainment of the various cohorts? To answer this question, we re-estimate the CS model, allowing now for matching by both age and educational attainment. We show that the changes in educational attainment for the famine born cohort are insufficient to explain changes in marital behavior of that cohort. The change in marital attractiveness of the famine born cohort is not well captured by the changes in their educational attainment.

1.1 Literature review

Our paper is related to two 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.; Chen 2007; Geogens, et. al. 2004; Meng and Qian; Porter 2007). 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 the famine on marital behavior of the affected cohorts. Both papers focus on homogenous treatment effects. Outcomes include marriage rates, age at marriage, spousal age differences and other spousal characteristics. They conclude that the famine had modest effects on marriage rates

Pollack 1990). Inputs are population vectors of types of individuals. Output is a matrix of who marries whom, and who remains unmarried.

caused the affected cohorts to marry later, increased spousal age gaps, and decreased spousal education gaps. The two papers attribute the outcomes to different causal mechanisms. Almond et. al. emphasize the decline in marital attractiveness of the famine born cohort whereas Porter emphasizes the relative scarcity of famine born cohort in the marriage market. Almond et. al. use variations in provincial mortality rates of the famine affected cohorts as a regressor and interpret its estimated effect as due to changes in marital attractiveness. Porter uses marital-share weighted adult sex ratios of the famine affected cohorts within and across provinces as a regressor and interprets its estimated effect as due to changes in sex ratios. Since the famine increased mortality and decreased birth rates (and therefore eventual population supplies to the marriage market), it is difficult to disentangle empirically the two effects in a regression framework.

A famine effect on marital behavior in a marriage market is a general equilibrium phenomenon. A regression using individual-level data of spousal characteristics on individual attributes ignores these general equilibrium effects.[9]

Our paper relaxes the "either or" hypothesis, 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.[10] Thus we view our framework as complementary to the above papers.

Our paper is also related to the literature which 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

[9]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.

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

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. Bergstrom and Lam (1994) and 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 offsetting large quantity and quality effects.

2 Methodology

This paper uses three statistics to study marital behavior: marriage rates, marriage shares and a marital accounting scheme, total gains to marriage.

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 type i men. $j = 1, .., J$ and $i = 1, ..I$. 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 $\mu_{i0}^t = m_i - \mu_i^t$, and the number of unmarried type j women be $\mu_{0j}^t = f_j - \mu_j^t$. 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 be an $I \times J$ matrix whose typical element is μ_{ij}^t.

The following accounting identities, which are general equilibrium constraints, have to be satisfied:

$$\begin{aligned}
\mu_{0j} + \sum_{i=1}^{I} \mu_{ij} &= f_j \ \forall\ j \qquad (1)\\
\mu_{i0} + \sum_{j=1}^{J} \mu_{ij} &= m_i \ \forall\ i\\
\mu_{i0}, \mu_{0j}, \mu_{ij} &\geq 0 \ \forall\ i, j
\end{aligned}$$

The first equation above says that the sum of unmarried type j women plus all the type j women in different types of marriages must equal the number of type j women. The second equation above says that the sum of unmarried type i men plus all the type i men in different types of marriages must equal the number of type i men. Finally, the number of individuals in any marital choice must be non-negative.

Rather than working with μ^t, M^t and F^t, most researchers use simpler statistics to describe the marriage market at time t. The most common statistics are the marriage rates:

$$r_i^t = \frac{\mu_i^t}{m_t^i};\ r_j^t = \frac{\mu_j^t}{f_j^i};\ i = 1, .., I;\ j = 1, ..J$$

Marriage rates are equivalent to the marriage odds ratio for different types of individuals:

$$o_i^t = \frac{\mu_i^t}{m_t^i - \mu_i^t};\ o_j^t = \frac{\mu_j^t}{f_j^i - \mu_j^t};\ i = 1, .., I;\ j = 1, ..J$$

Marriage rates or odds ratios are not summary statistics for marital behavior. To study who marries whom, we first study spousal shares by types of husbands and wives:

$$s_{j|i}^t = \frac{\mu_{ij}^t}{\mu_i^t};\ i = 1, .., I;\ j = 1, ..J \tag{2}$$

$$s_{i|j}^t = \frac{\mu_{ij}^t}{\mu_j^t};\ i = 1, .., I;\ j = 1, ..J \tag{3}$$

$s_{j|i}^t$ is the share of type j women among type i's wives. $s_{i|j}^t$ is the share of type i men among type j's husbands. Shares are informative about spousal subsitution patterns, i.e. who to marry. By definition, the sum of the shares across different types of spouses for the same type of individual is one. Thus the shares are not informative about the choice of whether to marry or not.

To investigate how substitution affects the decision to marry and vice versa, we need a statistic that will link the two effects. To that end, let the total gains to an $\{i, j\}$ marriage, π_{ij}^t, be:

$$\pi_{ij}^t = \ln \frac{\mu_{ij}^t}{\sqrt{\mu_{i0}^t \mu_{0j}^t}};\ i = 1, .., I; j = 1, ..J \tag{4}$$

The numerator is the number of $\{i, j\}$ marriages at time t. The denominator is the geometric average of the number of unmarrieds, μ_{i0}^t and μ_{0j}^t at time t. Let Π^t be the $I \times J$ matrix with typical element π_{ij}^t.

We can rewrite total gains as:

$$\pi_{ij}^t = \ln \sqrt{o_i^t o_j^t} \sqrt{s_{i|j}^t s_{j|i}^t};\ i = 1, .., I; j = 1, ..J \tag{5}$$

π_{ij}^t, the total gains to $\{i, j\}$ marriages, is the average of the log odds of marriages for i type men and j type women plus the average of the log shares. Thus the total gains to marriage combines substitution patterns with marriage rates.

We will now show that Π^t is an alternative accounting scheme for marital behavior, μ^t.[11] Π^t has $I \times J$ elements and μ^t also has $I \times J$ elements. Given Π^t, $M^t$and F^t, we can use the system of $I \times J$ equations (4) to recover μ^t. By definition, there an admissible solution to this system of equations, namely μ^t. CS shows that the solution is locally unique.[12]

Consider a new time t' with different marital matches and population supplies, $\mu^{t'}$, $M^{t'}$ and $F^{t'}$. We can estimate new total gains, $\Pi^{t'}$:

$$\pi_{ij}^{t'} = \ln \frac{\mu_{ij}^{t'}}{\sqrt{\mu_{i0}^{t'} \mu_{0j}^{t'}}};\ i = 1, .., I; j = 1, ..J \tag{6}$$

$\Pi^{t'}$ is a complete description of the marriage distribution in time t'. Let $\Delta X \equiv X^{t'} - X^t$. Then $\Delta\Pi$ is a complete description of the changes in marital behavior between the two periods.

Let t denote a period in which the famine did not affect marital behavior. We use Π^t to characterize marital behavior at time t. Let t' denote a period in which the famine affected marital behavior. We will use $\Delta\Pi$ as our estimates of the causal effect of the famine on the marital behavior of famine affected individuals. Because $\Pi^{t'}$ is a complete accounting scheme for marital behavior at time t', it imposes no apriori restriction on marital behavior. $\Delta\Pi$ can be estimated nonparametrically. Thus $\Delta\Pi$ is a consistent estimate of the causal

[11] It cannot deal with marital choices with zero observation. I.e. π_{ij}^t must be finite.

[12] Global uniquness may also apply because the solution has to satisfy all the accounting identities in (1). In widely different applications (Choo and Siow 2006 and 2006a) including this paper, we have not found multiple solutions.

effect of the famine on marital behavior for the famine affected cohorts as long as t and t' are valid control and treatment periods.

If there are changes other than the famine which affect marital behavior between t and t', then $\Delta\Pi$ will be an inconsistent estimate of the causal effects of the famine. There were other social and economic changes between 1982 and 1990, including migration behavior, that may have affected marital behavior between the two periods. We will assume that these other changes are not correlated with ΔM and ΔF, the changes in population supplies due to the famine. As we will show later, the behavior of $\Delta\Pi$ is so different for the pre-famine, famine born and post-famine cohorts, and so closely tied to changes in population supplies caused by the famine, that it will be difficult to imagine other factors causing these changes in marital behavior. We cannot identify level and trend effects of the famine to the extent that these other factors have level and trend effects on marital behavior.

The above is as far as we can go in reduced-form estimation. Tautologically, total gains at time t is a function of population vectors and exogenous parameters at time t:

$$\Pi^t = \kappa(M^t, F^t, \Lambda^t)$$

The famine simultaneously affected population vectors, $M^{t'}$ and $F^{t'}$, and exogenous parameters, $\Lambda^{t'}$, at time t'. Thus:

$$\Delta\Pi = \kappa(M^{t'}, F^{t'}, \Lambda^{t'}) - \kappa(M^t, F^t, \Lambda^t) \tag{7}$$

In order to disentangle observed changes in marital behavior between effects due to changes in marital preference and effects due to population supplies, we need to posit a model for $\kappa(.)$ and to estimate Λ^t and $\Lambda^{t'}$. CS is such a model. In their static model of the marriage market, total gains measures the expected marital gain to a random $\{i, j\}$ pair marrying relative to them not marrying. The thought experiment is as follows. Consider a randomly chosen type i male marrying a randomly chosen type j female. We compare the marital output of this randomly chosen couple to the geometric average of what they would have obtained if they remained unmarried. So once the individuals are chosen, they can only compare whether to marry or forever remain unmarried. This measure of relative expected marital gain is unaffected by marriage market conditions because we are not choosing the couple based on relative scarcity, nor do we allow them to marry other individuals.

Thus in CS, Π^t are exogenous, independent of population vectors, M^t and F^t, or other determinants of marriage market conditions at time t. In other words, CS implies:

$$\Pi^t = \kappa(\Lambda^t) = \Lambda^t \tag{8}$$
$$\Delta\Pi = \Delta\Lambda \tag{9}$$

Not only do $\kappa(.)$ not depend on M^t or F^t, it is also the identify function.

$\Delta\Pi$ is always a reduced form estimate of the total effect of the famine on marital behavior. Equation (9) says that the difference in total gains measures exactly the changes marital attractiveness between the two time periods, independent of the changes in population vectors. By using CS or equivalently equation (9) to interpret the estimates of $\Delta\Pi$, we have moved from a reduced form to a structural interpretation.

If we have estimates of Π^t, we can predict what the new marriage distribution $\widehat{\mu}^{t'}$ will be with new population vectors $M^{t'}$ and $F^{t'}$, and $\Pi^{t'} = \Pi^t$ for all i and j by solving:

$$\pi_{ij}^t = \ln \frac{\widehat{\mu}_{ij}^{t'}}{\sqrt{\widehat{\mu}_{i0}^{t'}\widehat{\mu}_{0j}^{t'}}};\ i = 1,..,I; j = 1,..J \tag{10}$$

If the only effect of the famine was to change population supplies between t and t', then $\pi_{ij}^{t'} = \pi_{ij}^t$ and the marital distribution at time t' should be $\widehat{\mu}^{t'}$ as in the system of equations (10). The predicted marriage distribution, $\widehat{\mu}^{t'}$, satisfies the non-trivial accounting constraints described in equations (1) (CSS and Siow (2008)). So the CS model implies that $\widehat{\mu}^{t'}$ is an estimate of the effects of quantity changes caused by the famine on marital behavior at time t'.

If the new actual marriage distribution, $\mu^{t'}$, differs from the predicted marriage distribution, $\widehat{\mu}^{t'}$, $\Delta\Pi \neq 0$. Because total gains completely describe the marriage distribution, the previous statement is always true. Without a model of marital behavior, we do not know how much of $\Delta\Pi$ is due to changes in marital attractiveness, $\Delta\Lambda$, and how much is due to changes in population supplies, ΔM and ΔF. The bite of CS is that it implies that $\Delta\Pi$, the changes in total gains, only measure changes in marital attractiveness, $\Delta\Lambda$, due to the famine.

If the CS model is incorrect, then our decomposition of the causal effects of the famine into quantity and quality changes will be wrong. Because CS is a fully saturated (just identified) model, the validity of CS in this particular application is untestable. CS is unlikely to be exactly right. Siow (2008) shows that generalized versions of CS will in general not reduce to (8). Thus CS should be regarded as an approximate model of actual marital behavior.

We provide an indirect test to show that CS is not vacuous. In demography, CS is known as a marriage matching function (MMF), an empirical model of the marriage market. Although the objective has been known for some time, it has been difficult to come up with empirically tractable and yet behaviorally attractive MMFs.[13] The main difficulty is how to deal with alternative spousal choices while minimizing a priori identifying restrictions. The current standard MMF in demography is the harmonic mean MMF (E.g. Qian and Preston (1993); Schoen (1981)):

$$\frac{\mu_{ij}^t}{m_i^t} + \frac{\mu_{ij}^t}{f_i^t} = \lambda_{ij}^t;\ i = 1, .., I; j = 1, .., J$$

Like CS, this model will fit any cross section data perfectly. As a structural model, it has two deficiencies. First, it ignores the problem of alternative spousal choices. Changes in $m_{i'}^t$ or $f_{j'}^t$ do not affect μ_{ij}^t. Second, the accounting constraints, (1), are not imposed.

We will show that using the harmonic mean MMF in our context leads to inadmissible predictions, predicted marriage rates which exceed one.

3 Summary Data and Sex Ratios

All the data presented here come from the one percent household sample of the 1982 Census of China and the one percent clustered sample of the 1990 Census of China. Wang (2000) and Mason and Lavely (2001) are useful resources on the details of the censuses 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,[14] 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

[13].... the frustrating search for a mathematically tractable and sociologically realistic “marriage function”. (Preston and Richards (1975))

[14]See footnote 10.

in number of ways for our analysis. First, migration of rural born out of Anhui (Sichuan), and in-migration into rural Anhui (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.

Table 1 provides some summary statistics for rural counties in Anhui and Sichuan from the 1982 and 1990 censuses. The average spousal age differences in the two provinces ranged 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 censuses, the average spousal age difference was about three years. Because the censuses collect ages by years, we assume that the customary equilibrium spousal age difference is three years.

The marriage rates for women are higher than that for men in both provinces and both censuses.

High education is defined as completing elementary school. The table shows that less than fifty percent of men or women in rural Sichuan and Anhui 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 the immediate table above, 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.[15] Thus, we need 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 lifecyle effects, the marriage behavior of 26-28, 29-31, and 32-34 years olds in 1982 was similar to each other.[16] Conditional on marriage, the distribution of spousal shares were similar for all three cohorts. So there was no difference in the choice of spouses between the three cohorts. Since we are comparing marital behavior by age groups, we are controlling for lifecycle effects in the timing of marriage.

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. So although there were

[15] 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 1982, suggesting that any effect of the new marriage law on their behavior through whom they could marry would have been marginal.

[16] The presence of lifecyle effects in marital behavior neccesitates 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.

likely other social and economic changes which affected the marital behavior of these famine affected cohorts, including new off-farm opportunities, migration opportunies and rising incomes, these other changes were primarily level and trend changes which were orthogonal to the changes in population supplies due to the famine which followed a very distinctive non-linear age pattern. We do not know any other social or economic change that followed this distinct age pattern and also only impacted these cohorts.

4 Sichuan

Figure 1 shows the number of individuals by age in rural counties in Sichuan in 1990. The pre-famine cohort, 32-34, were 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 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.

4.1 Marriage rates

Figure 2 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.[17] The famine had little to no impact on the sex ratio. There is little evidence that male children were significantly favored over female

[17] 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.

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 3 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 2 had any impact on the marriage rates of these women in 1982. The 26 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 2 earlier 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 2 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 summarized in Figure 4. They are also consistent with the findings in Almond et. al. The marriage rates of famine born men increased by less than 5 percent compared to their 1982 peers. The marriage rates of post-famine born men increased by substantially more, 5 to 15 percent 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 were in the same direction. Thus what we observed is a shift in marital behavior for young men in 1990. Other social 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 surely were other shocks which also affected the marital behavior between 1982 and 1990.

4.2 Marriage shares

To set the stage, it is convenient to have an idea of what customary marital shares were. Figure 5 plots the distributions of husbands by spousal age differences for women who were 33, 30 and 27 in 1982. Because they were born substantially before the famine, the marital behavior of 1982 women of those ages should have been unaffected by the famine. We differentiate husbands by their age gaps over their wives, from -3 years to +6 years. Husbands within these 10 years ages interval account for 83-96% of all husbands. The figure shows that there is essentially no difference in the distributions of husbands by spousal age differences for these different aged women in 1982. The marital behavior of 1982 women of those ages was unaffected by the famine; and their spousal choices as represented by spousal age differences did not change with their age in 1982.

The 27 years olds in 1982, which are one-year removed from the overlap cohort of age 26, did not display any unusual behavior in 1982. Figure 5 is the strong case for using the 1982 cohorts as the control cohorts.

Turning to the effects of the famine, Figure 6 shows the marital partners of three age cohorts of women in 1990: 27 (post-famine born), 30 (famine born) and 33 (pre-famine born). First consider 33 years old women who were born before the famine. Since women generally marry older men, Figure 2 tells us that the customary husbands of these women are not scarce. The largest share of husbands was two years older. For 33 years old women in 1990, the age distribution of their spouses looks the same as their same age peers in 1982 in Figure 5.

30 years old women were famine born women. They are scarce relative to their customary husbands. Compared with the shares of 33 years old women, their marriage shares distribution shifted to the right. Although they could replicate the share distribution of the pre-famine women because they were scarce relative to older men, more of them married older husbands.

27 years old women were born after the famine. They suffer a relative scarcity of customary husbands. As shown in the figure, their marriage share distribution is almost symmetric around age gap $[-2, 2]$. It flattens out between age gap 2 to 4 and then increases. Thus, post-famine women married a much larger share of own age or younger men, and also significantly older men. The share of husbands in the age gap $[-2, 2]$ was 0.83. Figure 6 shows that in 1990, the distributions of husbands by spousal age differences were significantly different between pre-famine born, famine born and post-famine born women.

Using the behavior of women in 1982 as control groups, Figure 7 plots the ratio of 1990 husbands' shares to 1982 shares for 27, 30 and 33 years old women. If there is no difference in shares between 1982 and 1990 for the same age women, then the ratio should be 1. Consider the case of 33 years old women. The ratio of shares are slightly above 1 for age gap between [0,4]. In both 1982 and 1990, most of the husbands of these 33 years old women fell in the age gap between [0,4] years. The ratio of shares are lower than 1 below [0,4]. This says that in the [0,4] range, pre-famine born women in 1990 had the same relative distribution of husbands by spousal age differences as their 1982 counterparts. But pre-famine women in 1990 had substantially less younger husbands outside the range. Less younger husbands can be explained by the relative scarcity of famine born men.

Famine born women in 1990, age 30, behaved very differently from their 1982 counterparts. They were far more likely to marry older men and far less likely to marry men of the same age or younger.

Post-famine women in 1990, age 27, also behaved very differently from their 1982 counterparts. They were far more likely to marry same age or younger men, also pre-famine born men, and far less likely to marry famine born men. So here, post-famine women avoided the scarce famine born men. What is interesting is their increased demand for substantially older, pre-famine born men. Both 27 and 30 years old women in 1990 had relatively more demand for significantly older men.

Taken as a whole, Figure 7 shows that different cohorts of famine affected women responded differently in their spousal choices. Consistent with Almond et. al. and Porter, there is a small increase in the average spousal age gap for the famine affected cohorts relative to their 1982 peers. However this average effect masks the heterogeneous responses by the different famine affected cohorts. It is also important to note that Figure 7 generates a pattern of responses for the famine affected cohorts that would be hard to rationalize based on other social and economic changes which occured in China.

4.3 Total gains

To preview what we will find, recall that marriages rates of famine born women were the same as their pre and post-famine born peers. The marriage rates of famine born men were lower than their post-famine born peers. But famine born men and women are relatively scarce in the marriage market. Let j denote the cohort of famine born women and i denote their customary spouses. To a first order, o_j^t did not change and j type women are scarce relative to type i men. $s_{i|j}^t$ must fall and by equation (5), π_{ij}^t must fall.

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 8 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 were lower than the other two age cohorts. Thus it is reasonable to use the 1982 individuals as control groups for their same age 1990 peers.

Figure 9 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 3 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 in customary marital sex ratios, total gains for the famine born women had to be lower than the other famine affected women.

Figure 10 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 censuses 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 were 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 heterogenous 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.

4.4 Quantity and quality

So far, our discussion was descriptive. Taken together, the first differences in the three statistics strongly suggest that the famine had substantial effects on the marital behavior of the famine affected cohorts. The change in marital behavior for the pre, post and famine born cohorts are sufficiently different from each other that it is difficult to explain these first difference changes by other social economic factors.

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 7, which plots the ratio of 1990/1982 marriage shares of husbands by his age gap, provides strong evidence for the importance of quantity effects. Figure 7 shows that famine born women were able to marry a larger share of husbands of customary age than the 1982 same aged wives. The famine born wives also avoided their share of same-aged husbands compared to 1982 same-aged peers. In addition, Figure 7 shows that post-famine wives also married a lower share of famine born husbands compared with their same aged peers. So both famine born and post-famine wives relatively avoided famine born husbands. Yet the marriage rates of famine born men were higher than their 1982 peers. The relative scarcity of famine born men and women reconciles this set of disparate observations.

On the other hand, it is difficult to explain the differences in total gains in Figure 10 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 marital shares 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 now use the CS model to do that decomposition. Equation (8) 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, as per equation (9).

It is easy to use CS to interpret Figure 10. 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 were lower compared to their same age counterpart in 1982. And if their husband is also famine born, the payoff was even worse. On the other hand, the marital output of marriages with pre-famine born women were marginally higher than their same aged counterpart in 1982, validating the Meng and Qian hypothesis that children who survived the famine were positively selected. Finally, the difference in total gains for 27 years old women were largely above zero. It 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 substantial drop relative to their 1982 same aged peers. Thus the CS model unambiguously shows that the famine born cohort suffered a substantial drop in marital attractivenss.

Given this 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. To quantify the effect of the changes in quantities on marital behavior, 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. $\widehat{\mu}_{ij}^{t'}$ using equation (10)). We also make a similar 1990 prediction using the harmonic mean MMF with 1982 estimated parameters.

Figures 11a and 11b show for Sichuan, the predicted male and female marriage rates from the two models respectively. For both genders, the predicted marriage rates from the harmonic mean MMF often exceed 1, an inadmissible prediction. These violations occur because the harmonic mean MMF does not impose required general equilibrium accounting identities, ignores substitution effects, and the changes in sex ratios of customary spousal age differences were large. Thus as previous researchers have observed, the standard MMF used by demographers is a poor empirical MMF.

On the other hand, the predicted marriage rates from the CS MMF behave sensibly. In Figures 11a and 11b, 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 with large changes in sex ratios of the customary spousal age differences for the famine born cohorts, their predicted marriage rates remained below 1. 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. These two attributes show the advantage of the CS MMF over the harmonic mean MMF.

In Figure 11a, 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 famine cannot explain these discrepancies. The famine must have also changed marital attractiveness to marriage for these cohorts.

Figure 11b 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 11e 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 were more concentrated among these husbands than were 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, (1), 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 10, 11a, 11b and 11e provide a unified summary of the effects of the famine on the marital behavior of famine affected cohorts. Figure 10 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 11a and 11b 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 cancelled each other out and resulted in small changes in marriage rates for the famine affected cohorts. If marital output did not change, Figure 11c shows that famine born women would have been less likely to marry, and post-famine born women would have been more likely to marry famine born men than what they did.

4.5 Education effects

Since 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), Luo, Mu and Zhang (2006)), their educational attainment may have been affected. This may enable us to use the change in educational attainment of that cohort to proxy for their drop in marital attractiveness, thereby explaining their drop in total gains to marriage.

Figure 12 shows the fraction of women who had less than a primary school education by age. Not surprising, in both the 1990 and 1982 censuses, the fraction grew with the age of the women. It is difficult to see the change in educational attainment at the levels for the famine born cohort. This implies that it is unlikely that the change in educational attainment of the famine born cohort will be able to explain the changes in total gains to marriage of that cohort.

Figure 13 shows the log difference (growth rate) of the fraction of women who had less than a primary school education by age. In the 1982 census, the growth rate fell rapidly by age for women below age 35. Such a decline in the growth rate should be expected if women increased their educational attainment over time. In the 1990 census, the growth rate also fell by age for women above age 32. For women between ages 28 and 32, the growth rate of the fraction of women with less than a primary school eduction formed a valley with a bottom at age 30. Based on deviations from trend growth, the famine affected cohorts had less education than they would otherwise have. Although we do not present the results here, the results for male educational attainment are similar. Figure 13 raises the possibility that changes in the growth rate of education of the famine affected cohorts may be able to explain some of the fall in total gains to marriage of those cohorts.

Denote individuals with more than a primary education as high education and those with a primary education or less as low education. Figure 14 shows the ratio of 1990 to 1982 female marriage rates by age and education. High education women were less likely to marry in 1990 than in 1982 compared with their low education peers. For high educated famine born women (age 30), the ratio is slightly higher than their adjacent aged peers. For low educated famine born women, the ratio is slightly lower than their adjacent aged peers. These small differences in marriage rates suggest that high educated famine born women may have fared better than their low education counterparts.

To evaluate the overall effects of the changes in educational attainment on marital behavior, we estimate total gains π_{ij}^{1982} for every $\{i, j\}$ match where a type is defined by the individual's age and education (high versus low). We then use the estimated total gains, π_{ij}^{1982}, to predict the marriage distribution in 1990, $\widehat{\mu}_{ij}^{1990}$, due to changes in population vectors alone. Figure 15 plots the ratio of the own predicted gender marriage rates in 1990 to the actual marriage rates in 1990 by age, $\widehat{r}_g^t(r_g^t)^{-1}$, $g = i, j$.

Figure 15, where predictions depend on age and education, looks remarkably similar to Figure 8 where the predictions only depend on age. In other words, the change in educational attainment of the famine born cohort did not have a first-order impact on predicted marriage rates.

To examine this more closely, Figure 16 plots 1990-1982 total gains for marriages with two high educated spouses. Total gains for post-famine women were higher than their 1982 peers. But note that relative total gains for post-famine married women fell as the age of their husbands increased from own age to famine born husbands. So famine born men also had lower total gains. Total gains for famine born and pre-famine women were slightly lower than their 1982 peers. Figure 17 plots total gains for marriages with two low educated spouses. Total gains for pre and post-famine women were higher than their 1982 peers. On the other hand, total gains for famine born women were substantially lower than their 1982 peers. Figures 16 and 17 suggest that among famine born women, low educated women suffered a larger drop in total gains than high educated women.

5 Anhui

In general, the effects of the famine on the marital behavior of the famine affected cohorts in Anhui were about the same as for Sichuan.

Figure 1a and 2a present the number of individuals by age in 1990 in Anhui, and the same age sex ratio as well as the sex ratio for men three years older by female age. These two figures are similar to what we saw for Sichuan. In particular, the timing of the effects of the famine on birth rates in both provinces were the same.

Figure 3a shows the male and female marriage rates in 1982 and 1990. As the figure shows, female marriage rates between 1982 and 1990 were very similar. Across most ages, male marriage rates were lower in 1982. One potential explanation for the higher male marriage rates in 1990 is that relatively more unmarried men left the province in 1990 than in 1982.[18]

To see the effects of the famine more clearly, Figure 4a shows the ratio of 1990 to 1982 marriage rates by gender and age. There is a small dip in the marriage rate of famine born females relative to pre and post-famine born cohorts. This dip is more visible for males. Quantitatively, the effects are not large, within 2-3% difference. These dips are different from their Sichuan counterparts.

Note however that because famine born individuals are relatively scarce, these dips in marriage rates should be reflected in significant drops in total gains for famine born individuals.

Figure 5a shows that in 1982, Anhui women at ages 27, 30 and 33 had very similar shares of husbands by spousal age differences. Figure 6a shows that in 1990, Anhui women at ages 27, 30 and 33 had different shares of husbands by spousal age differences.

Figure 7a shows the 1990 shares divided by the 1982 shares. For 1990 33 years old wives (pre-famine born females), their shares of same age or older husbands were roughly the same as for the 1982 wives. On the other hand,

[18] As discussed earlier, our identification of the famine effect is not based on levels or trend changes.

the 1990 share of younger husbands (famine born males) were lower than their 1982 peers.

For 1990 30 years old (famine born) wives, their shares of own age husbands were lower than older or younger husbands compared with their 1982 peers. Finally, 1990 27 years old (post-famine) wives, their shares of own age spouses were significantly higher than famine born spouses, but comparable to pre-famine born spouses when compared with their 1982 peers. That is, post-famine born wives were more likely to have own age or pre-famine born spouses.

The behavior of 33 years old (pre-famine) and 27 years old (post-famine) wives with respect to their choices of husbands were similar to that of Sichuan women. However famine born Anhui wives were more likely to marry own aged spouses compared with their Sichuan counterparts.

Figure 8a shows the total gains for 27, 30 and 33 years old women in 1982. Although the shape of the three curves are similar, total gains for 27 years old women were lower than for 30 and 33 years old women. Figure 9a shows the total gains for 27, 30 and 33 years old women in 1990. Total gains for 30 years old women were significantly lower than for 33 years old and roughly comparable to 27 years olds. Total gains for own age spouses for 30 years old women were lower than those for pre and post-famine women.

Figure 10a shows 1990 total gains minus 1982 total gains. The most stark feature is that the difference in total gains for 30 years old women were lower than that for 27 and 33 years old women at all spousal age differences. In other words, famine born women were less desirable as spouses compared to their same age counterparts in 1982.

For 33 years old women (i.e the pre-famine cohort and their same age counterparts), the difference in total gains were above zero with own age and older husbands. It was lower than zero for younger (famine born) husbands. Finally, total gains for 27 years old (post-famine born) women were higher than their 1982 peers for own age and younger men, but lower than their 1982 peers for older (famine born) men. Thus Figure 10a unambiguously show that famine born individuals had lower total gains to marriage compared with 1982 same age peers or with pre and post-famine born individuals.

Figures 11c and 11d show the ratios of 1990 predicted marriage rates by age to actual marriage rates using π_{ij}^{82} where the type of an individual is defined by their age. The model underpredicts the number of marriages for post-famine born women. It substantially over predicts the number of married famine born men, and under predicts the marriage rates of pre- and post- famine born men. Thus we know that total gains must have changed between 1982 and 1990 for the same age groups.

Figure 11f shows the ratio of 1990 actual to predicted marriage shares using π_{ij}^{82} where the type of an individual is defined by their age. 30 years old famine born women married much more own age men and less pre-famine born men than were predicted. Post famine born 27 years old women were less likely to marry famine born men than were predicted. Thus post-famine women "avoided" famine born men.

Figures 12a and 13a show the fraction of women with primary or lower education and the growth in that fraction by age in 1990 and 1982 respectively. As in Sichuan, educational attainment declined with age. In Figure 13a, there is a sharp jump in the growth rate of low education famine born, age 30, women which shows that the famine born generation had less educational attainment than expected. But as shown in Figure 12a, the level effect is small.

Figure 14a shows the ratio of 1990 to 1982 female marriage rates. For women with low education, the marriage rates of the famine born cohort were slightly lower than pre and post-famine born women. For women with high education, famine born women suffer a larger dip in their marriage rates relative to pre and post-famine born women.

Figure 15a shows the ratio of predicted marriage rates by age to actual marriage rates using π_{ij}^{82} where the type of an individual is defined by their age and education. The Figure is almost identical to Figure 8a where the type of an individual is defined by age alone. What this means is that accounting for the changes in educational attainment between same age 1982 and 1990 indviduals did not account for the changes in marital quality of the famine born individuals.

Figure 16a shows 1990 minus 1982 total gains for high educated women with high educated men. famine born high educated women (age 30) had weakly lower total gains compared with their 1982 peers. Total gains for pre-famine born high educated women were roughly the same as their 1982 peers. Post famine born educated women had higher total gains than their 1982 peers, although there was a small relative dip with older (famine born) men.

Figure 17a shows 1990 minus 1982 total gains for low educated women with low educated men. famine born low educated women (age 30) had significantly lower total gains compared with their 1982 peers. Total gains were relatively higher with same age males. Total gains for pre-famine born high educated women were roughly the same as their 1982 peers for own age and older men. They were lower with younger (famine born) men. Post famine born educated women had the same total gains than their 1982 peers with own age or pre-famine born men. They had lower gains with famine born men.

Comparing figures 16a and 17a, high educated famine born individuals suffered a smaller loss in total gains to marriage than their low educated counterparts. This finding is comparable to that for Sichuan.

6 Conclusion

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. Thus the main 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.

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Data Appendix

All data are either from the 1% household sample of the 1982 Census of China, or the 1% clustered sample of the 1990 Census of China. Our samples are individuals in rural counties of Sichuan and Anhui, over the age of 18. There are issues common to both censuses, as well as differences between them, that required attention for the analysis in this paper. This appendix describes the most important of those issues.

Sampling. The microdata sample for 1982 was sampled at the household level, while the 1990 "clustered" sample was sampled at the level of the *administrative unit*. Lavely and Mason (2006) discusses the geographic coverage of the 1990 sample and compares summary statistics from the sample to tabulations from the complete census. They concludes that, aside from a few discrepencies, the sample "reproduces the geographic distribution of population and major population components quite well." We would not expect the different sampling methods to bias our results. The marriage markets we investigate are *de facto* defined as the subset of rural counties in each province. While the two census samples will contain exactly the same counties, they ought to both have a representative sample of rural households in each province.

Defining rural areas. Our analysis covers rural counties in Sichuan and Anhui. There has been much research and debate about the best definitions of rural and urban populations for various purposes (see Chan (1994), Martin (1992) and Shen (1995). The "official" definitions in fact changed between the 1982 and 1990 censuses. Our intention, in analyzing only rural areas, is to look at those areas most severely affected by the famine, and avoid conflating the very different effects suffered by cities and villages. Also, since the 1982 and 1990 use different classifications for the smaller geographic units, we wanted to select geographic areas on criteria which could be consistently applied in both censuses.

Both censuses define the largest three levels of geographic units according to a six digit *goubiao* (GB) code. The first two digits define the province (or municipalities like Beijing and Tianjin, which report directly to the central government). The second two digits define a prefecture within the province, and the last two label the counties within the prefecture. See Chan (1994) for a diagram of these levels in the two censuses. Within Sichuan and Anhui, we select those counties whose last two digits indicate they are *xian* (as opposed to *shi* or *shixiaqu*, which refer to urban areas). This definition will generally not capture the same households as those included in the official definition.[19]

Migration. Migration introduces a number of potential biases into our analysis. In the construction of sex ratios for the rural-born population at the provincial level, our population of interest is those males and females born in rural Anhui and Sichuan that survive to marriageable age. In the use of the 1982 and 1990 census data, this implies that we should include all individuals born in rural areas in Anhui and Sichuan that migrated either to the cities or to other provinces. Analogously, we should exclude those individuals born outside of Anhui and Sichuan that migrated into rural areas in these provinces. For the first difference methodology we use, we need information on migration in both census years in order to use alternative estimates of the sex ratios in our analysis.

In the 1982 census, no information is provided on place of birth, or migration. In the 1990 census, place of birth is again not provided, but information on migration between 1985 and 1990 is. In 1985, the famine affected cohort, i.e. those born between 1956 and 1964, would have been between the ages of 21-29, and by 1990, 26-34. Thus, data on migration between 1985 and 1990 will pick up some, but not all, of the movement in and out of rural areas in these provinces, especially that related to marriage.

In the case of Anhui, 1.02 percent of all women with rural registration between the ages of 26-34 in 1990 report migrating out of the province within the last five years. The percentage for marriage however, was only a third of this. This is slightly misleading because a significant number of these women would have already been married by 1985. For the youngest women, i.e. those born between 1962 and 1964, and in 1985 were between 21 and 23, 0.5 percent migrated for marriage. For males, 1.74 percent migrated, but only 0.14 percent was marriage-related. Migration into rural Anhui was slightly smaller, but in absolute terms, the numbers of males and females migrating in and out of rural Anhui were nearly offsetting.

In the case of Sichuan, 1.41 percent of all women between the ages of 26-34 migrated, with 0.88 percent marriage-related. As in the case of Anhui, the percentage was higher for younger women. For men, the percentage that migrated out was 1.51, but only one-tenth was because of marriage. The primary reason was for work. Flows into rural Sichuan were smaller than the outflows, with the net outflow of women for marriage equal to 0.63 percent of

[19] In earlier versions of the analyses, we also used a definition based on the share of individuals working in rural sectors. This definition resulted in a similar sample of counties.

the cohort. Overall, there was a small net outflow of females from Sichuan.

The 2000 census provides slightly richer information on migration that can be used to get a better sense of how serious these biases are. Of course, by 2000, attrition is also a much more serious problem. From the 2000 census, we know where an individual was born; where they were living at the time of the census; and if they had migrated in the last five years, the reason. The latter information is slightly less useful by 2000 because the cohort of interest would have been between the ages of 36- 44. We utilize this information along with that on an individual's registration status or hukou to identify the following: 1. Individuals with rural hukou that were born in Anhui (Sichuan) that are currently living in urban areas in Anhui or in other provinces; 2.Individuals with rural hukou that were born in Anhui (Sichuan) that are currently living in rural areas in other provinces; 3.Individuals with rural hukou that were born outside Anhui (Sichuan), but are currently living in rural areas in these two provinces. We construct an alternative estimate of the sex-ratio for rural-born Anhui (Sichuan) by adding individuals in groups 1 and 2 to those individuals who the census identifies as currently living in rural Anhui (Sichuan), and then subtracting individuals in group 3. There are two remaining omissions relating to the rural born between 1956 and 1964 that we cannot deal with. First, individuals that made it to marriageable age, but died before 2000. And second, individuals that were rural-born, but who at the time of the 2000 census had urban registration.

We graph old and new sex ratios in Figures A1 and A2. In the case of Anhui, larger out-migration flows by males and larger in-migration flows by women results in a slightly higher M/F ratio on average across all age cohorts. The bias is slightly higher for the famine born cohort, followed by the pre-famine cohort, and then the post-famine cohort. Overall, however, the bias (as measured by the ratio of the new sex ratio to the old sex ratio) is in the ballpark of 5% for the cohort born between 1956 and 1964. In the case of Sichuan, women are more likely than men to have migrated out of the province, with this only marginally offset by slightly higher in-migration of women than men into rural Sichuan. As a result, the new sex-ratio is slightly lower than the original estimate. This bias is fairly similar over all age cohorts, and again is in the vicinity of 5 percent. Overall, the small size of the bias, combined with the fact that the differences across cohorts are relatively small increases our confidence in our identification strategy.

Defining married couples within households. Except for household heads and their spouses, the census does not provide any definite way to determine who is married to whom within the household. Also, it provides no way to identify a person's spouse if that spouse does not live in the same household. For example, a male and female household member whose marital statuses are both "Married" and are identified "Children of the household head" may be married to each other (and one is a son-in-law or daughter-in-law). Or it is possible that they are both biological children of the household head who are married to spouses living outside the household.

Assuming that the first possibility is the more likely, we determine the married couples within households according to the following rules. First, we identify all "potential" married couples within each households as those who are of the opposite sex and have consistent relationships to the household head (both children of the head, or parents of the head, etc.). For children of household heads, we also required that potential couples be within five years of age.[20] If each person in the household has only one potential spouse, we define them to be married. If a person has multiple potential spouses, we assign married couples through positive assortive matching by age, e.g., the oldest married male child is married to the oldest married female child within five years his age.

Determining spouses was by far most problematic for children of household Heads, simply because there were more of these than parents, grandparents, grandchildren, etc. Fortunately, in the majority of households, there was only one potential married couple amongst the children.[21] In Sichuan, amongst households where there was at least one potential married child couple, 93.3% had only one potential couple, while 99.6% had two potential combinations or less. In Anhui, household sizes were slightly larger, as only 83% of households with at least one potential child couple had only one potential couple, while 93.6% had no more than two potential combinations.

Imputing missing spouses. For some individuals, there were no potential spouses in their household; for example, there were households where

[20] We performed tests of these rules by relaxing the age assumption, and comparing marriage age distributions of the full sample with the restricted sample of household heads and their spouses.

[21] We calculate the potential married couple combinations amongst children in a household to be $P(I, J)$ if $J < I$, or $P(J, I)$ otherwise, where I is the number of married male children in the household, J is the number of married female children in the household and $P(\cdot, \cdot)$ is the permutation operator. For example, a household with one married female child and two married male children has two potential child couple combinations. One with three female children and two male children (or vice versa) has six potential combinations.

the household head was married, but no spouse was present; or there were households with an odd number of married children. We imputed the age and education of these individuals' spouses by assigning values randomly from the distribution of spouse age and eduction for those of the same sex and age in that province with non-missing spouses.

Table 1: Sample Descriptive Statistics					
Census year	1982			1990	
Province	Sichuan	Anhui		Sichuan	Anhui
No. men 18-50	188,081	91,898		225,261	127,600
No. women 18-50	179,727	86,245		221,750	124,236
Share ever married men	0.737	0.719		0.719	0.700
Share ever married women	0.846	0.814		0.802	0.769
Mean (woman's - spouse's age)	-3.216	-2.776		-2.664	-2.436
Share of high edu (male)	0.302	0.349		0.440	0.474
Share of high edu(female)	0.195	0.149		0.295	0.207

Table 2: Famine affected cohorts and their comparison groups			
	Pre famine	Famine	Post famine
Birth years	1956-1958	1959-1961	1962-1964
1982 ages	24-26	21-23	18-20
1990 ages	32-34	29-31	26-28

Figure 1: Sichuan number of individuals by age, 1990

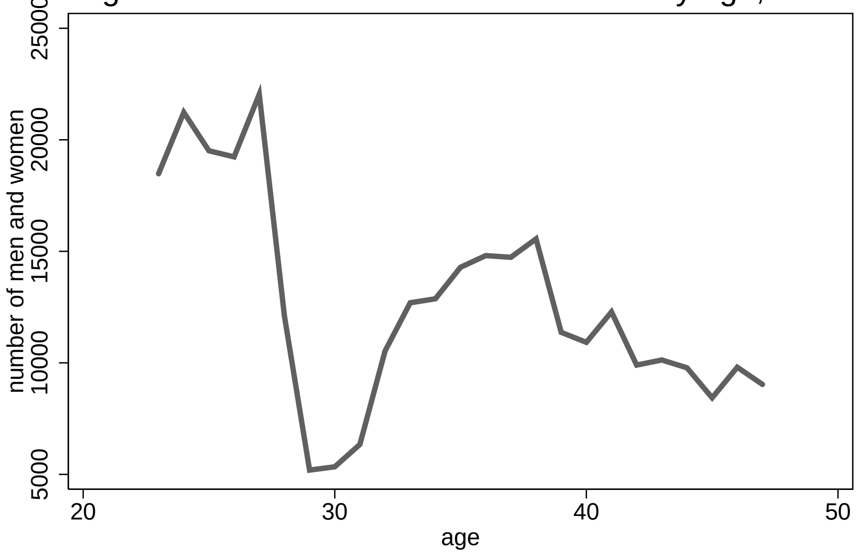


Figure 1a: Anhui number of individuals by age, 1990

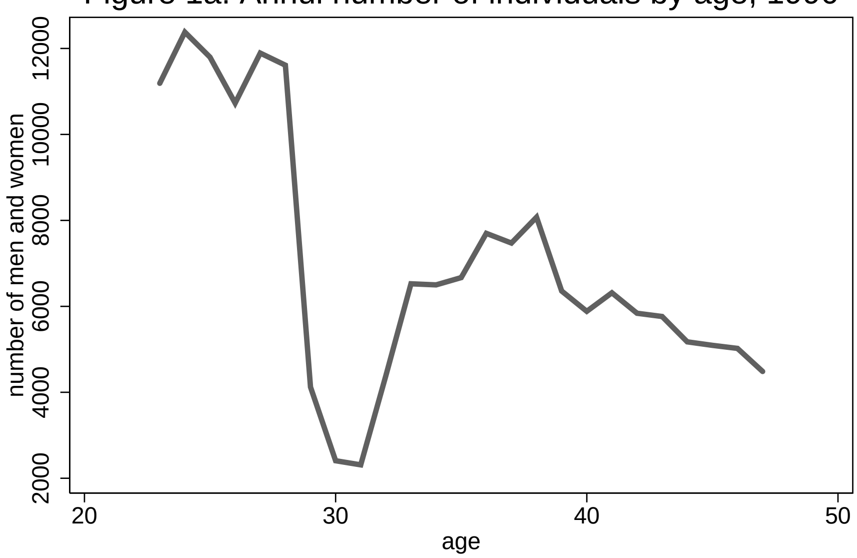


Figure 2: Sichuan sex ratios by female age, 1990

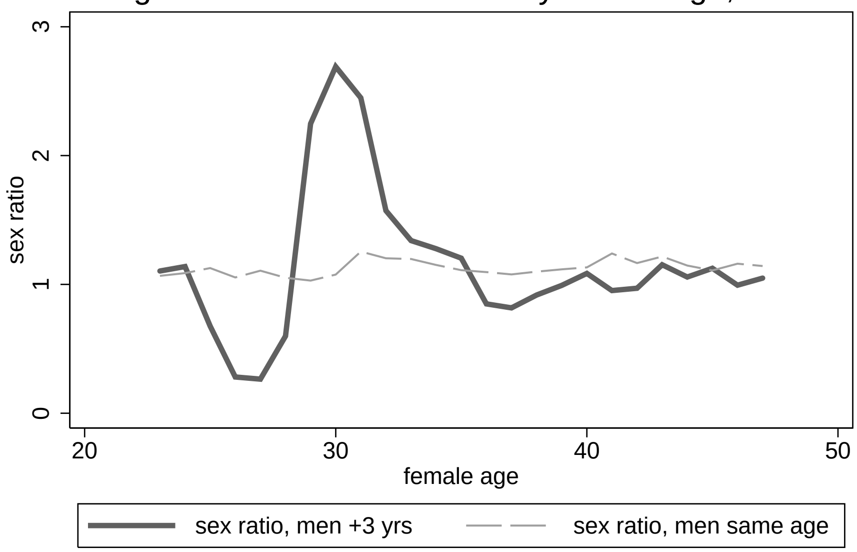


Figure 2a: Anhui sex ratios by female age, 1990

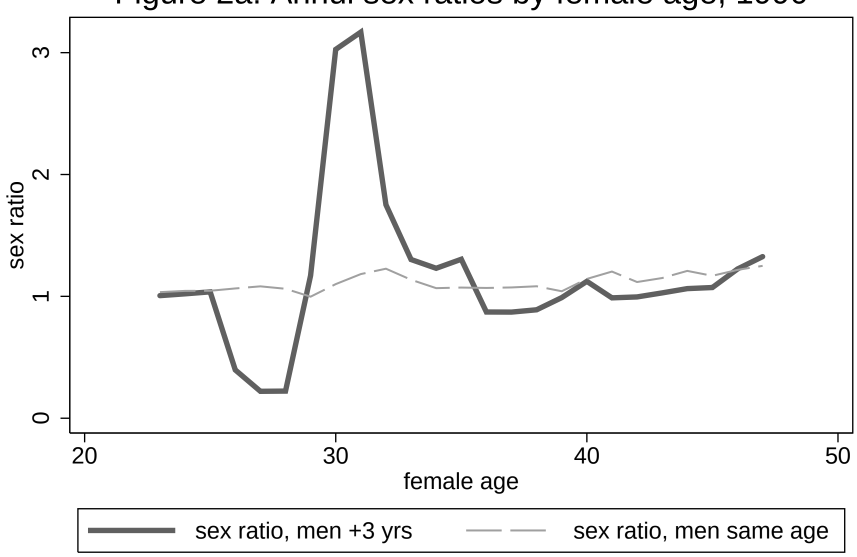


Figure 3: Sichuan marriage rates, 1990, 1982

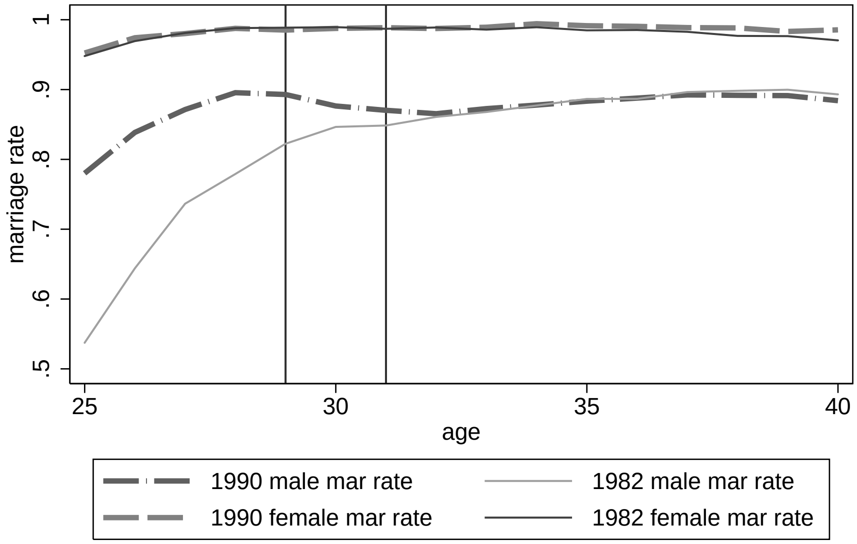


Figure 3a: Anhui marriage rates, 1990, 1982

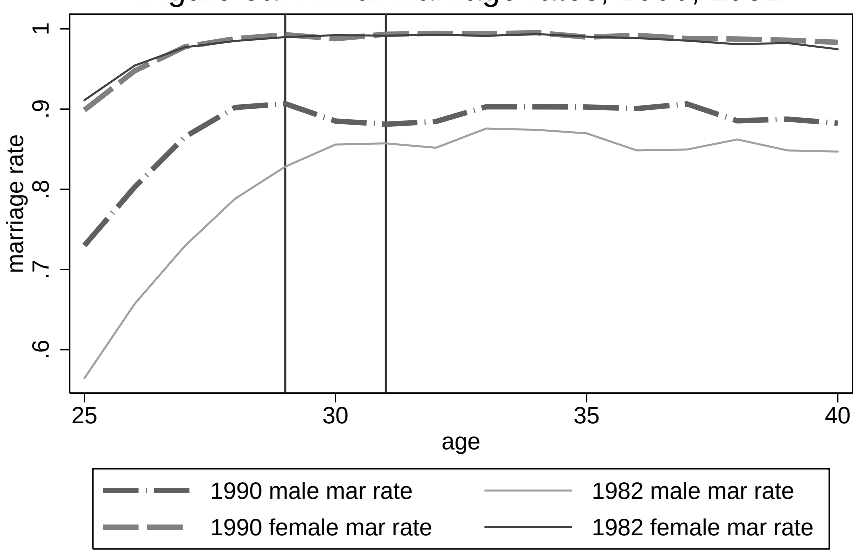


Figure 4: Sichuan 1990/1982 marriage rates

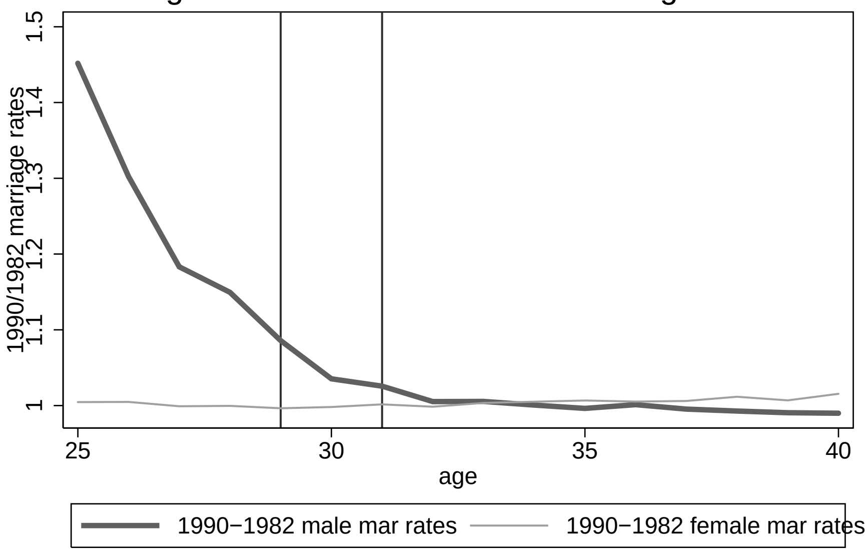


Figure 4a: Anhui 1990/1982 marriage rates

Figure 5: Sichuan share of husband, 1982

Figure 5a: Anhui share of husband, 1982

Figure 6: Sichuan share of husband, 1990

Figure 6a: Anhui share of husband, 1990

Figure 7: Sichuan 1990/1982 share of husband

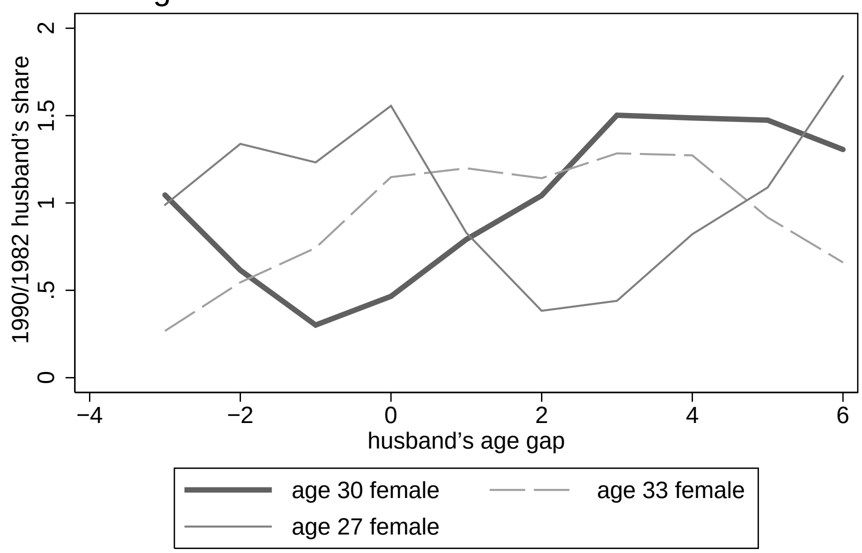


Figure 7a: Anhui 1990/1982 share of husband

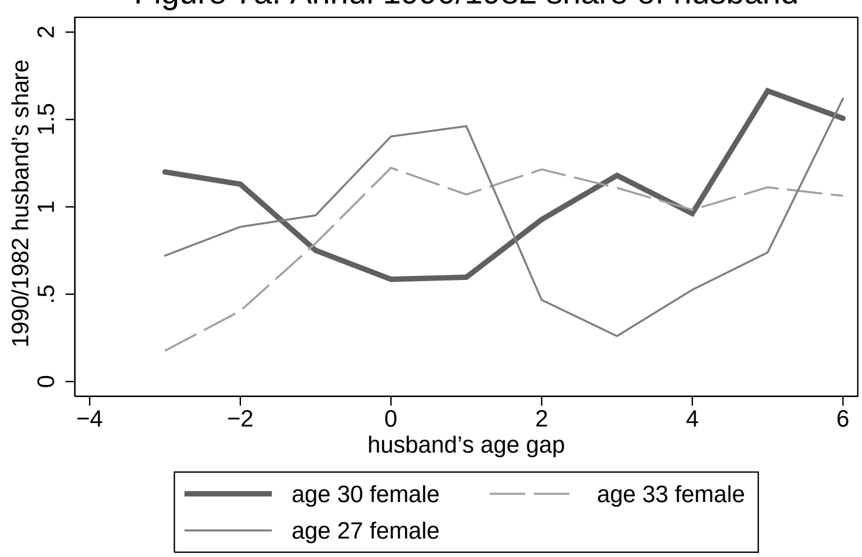


Figure 8: Sichuan total gains, 1982

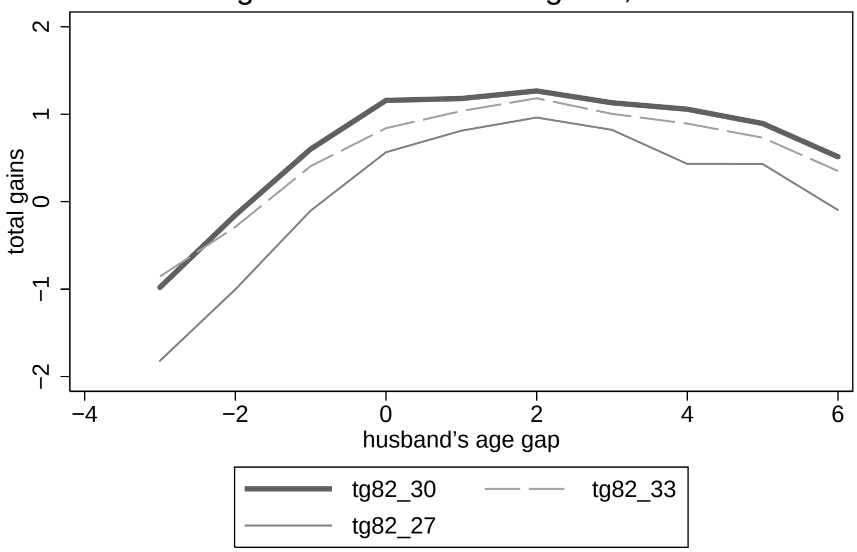


Figure 8a: Anhui total gains, 1982

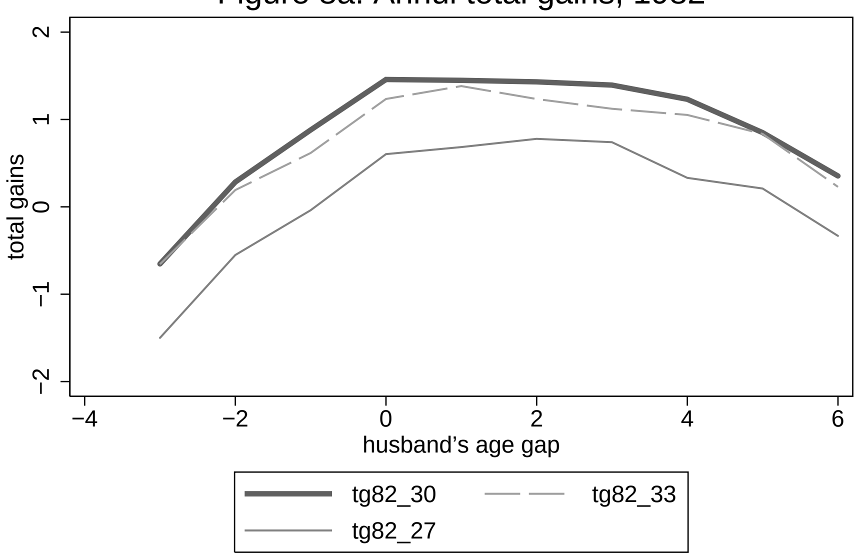


Figure 9: Sichuan total gains, 1990

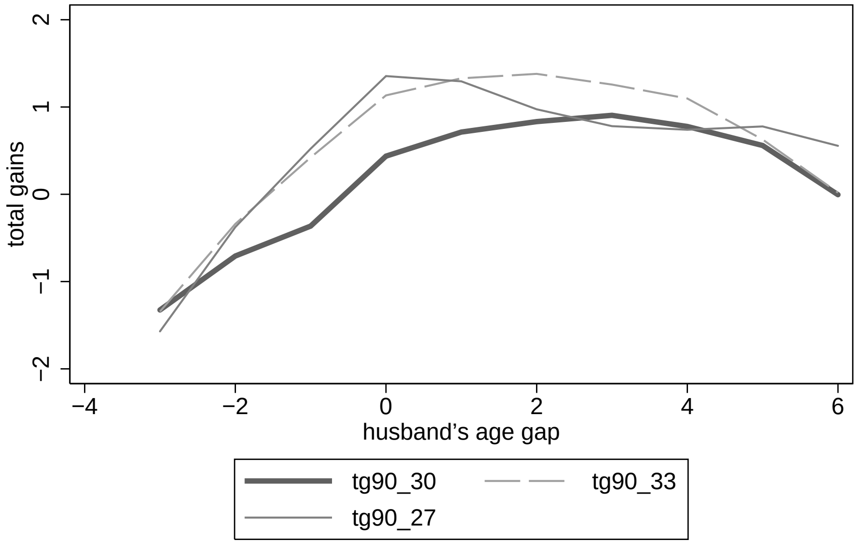


Figure 9a: Anhui total gains, 1990

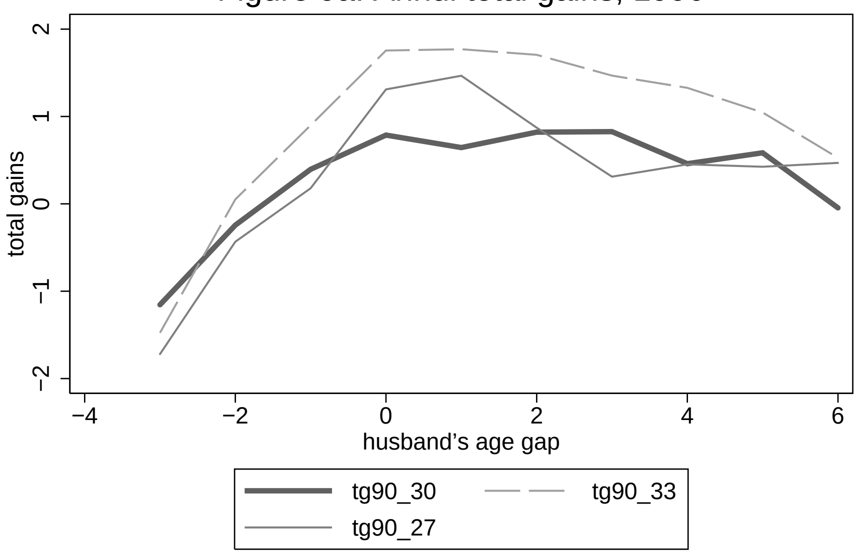


Figure 10: Sichuan 1990−1982 total gains

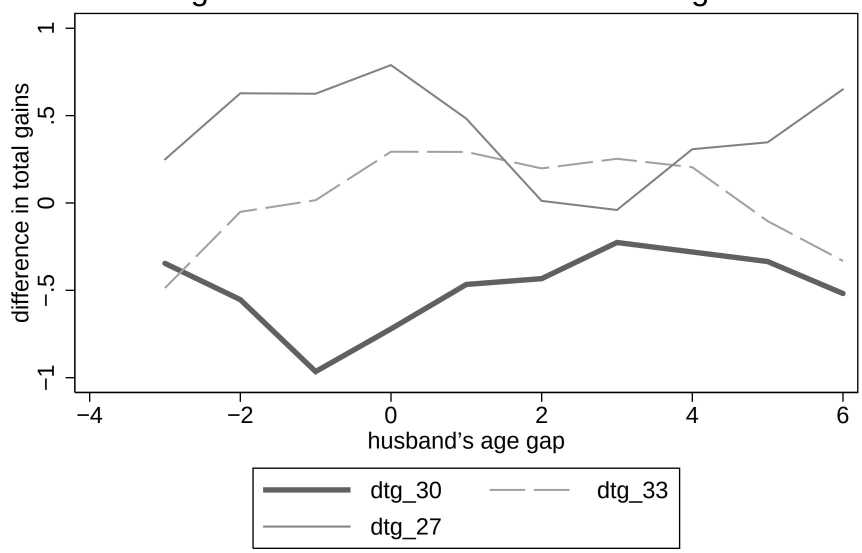


Figure 10a: Anhui 1990−1982 total gains

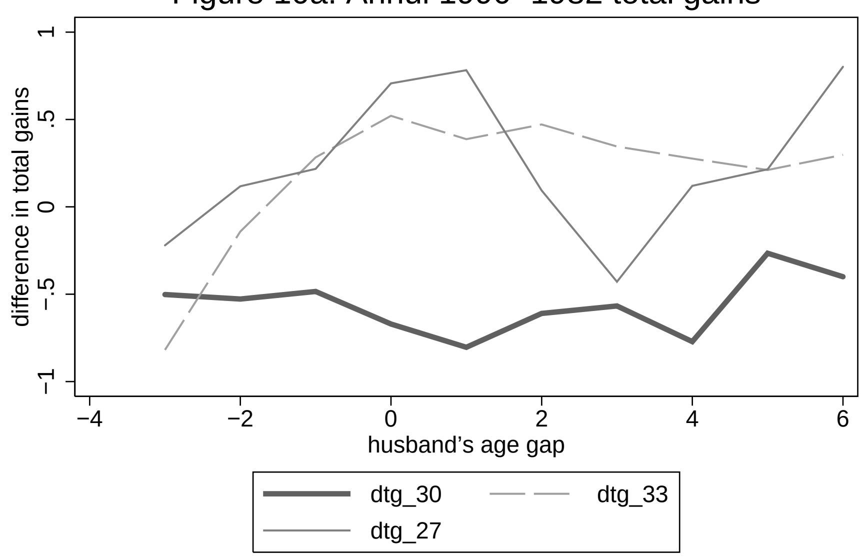


Figure 11a: 1990 Sichuan male marriage rates

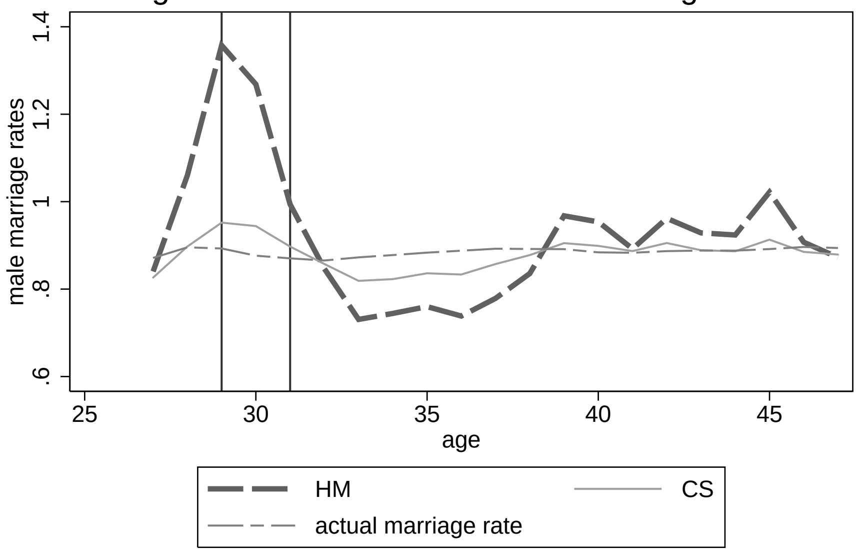


Figure 11b: 1990 Sichuan female marriage rates

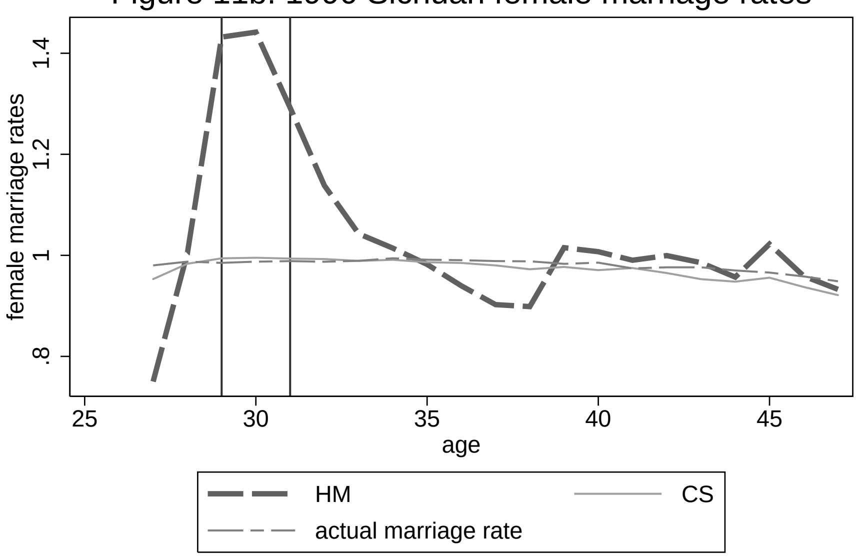


Figure 11c: 1990 Anhui male marriage rates

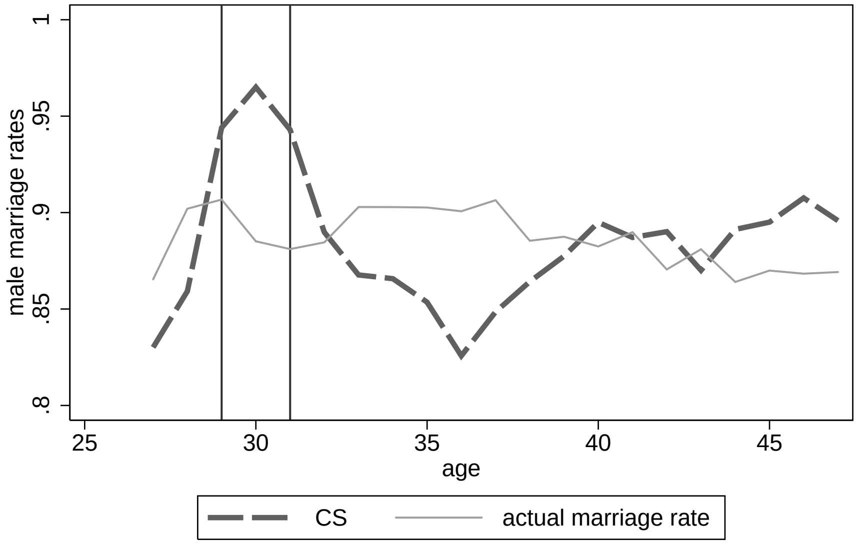


Figure 11d: 1990 Anhui female marriage rates

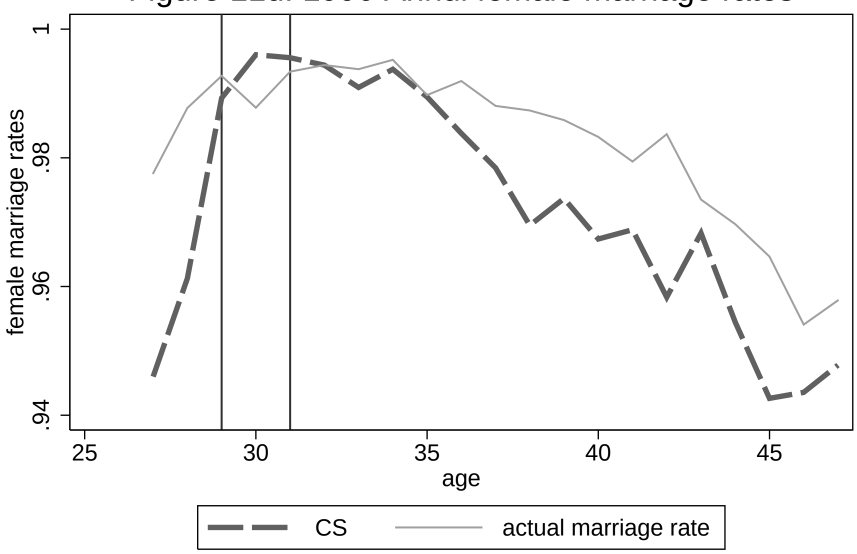


Figure 11e: Sichuan 1990 ratio of actual to predicted shares

Figure 11f: Anhui 1990 ratio of actual to predicted shares

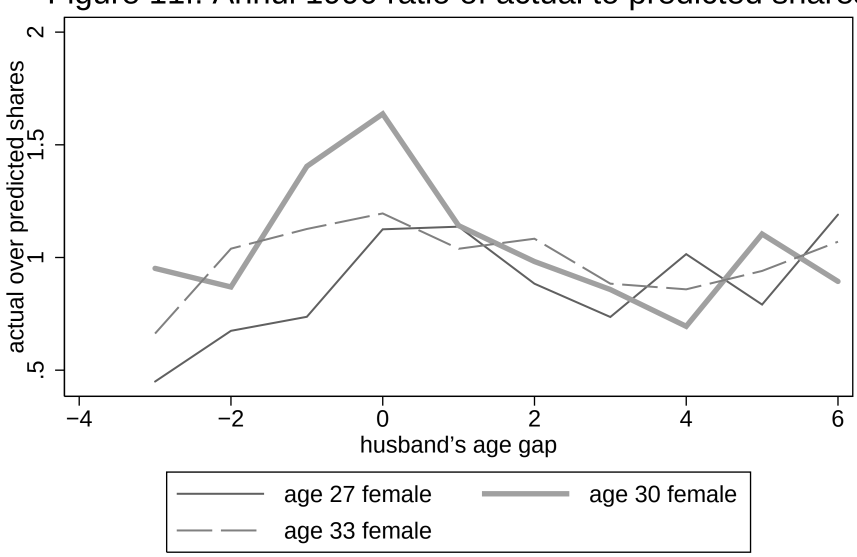


Figure 12: Sichuan fraction low education female

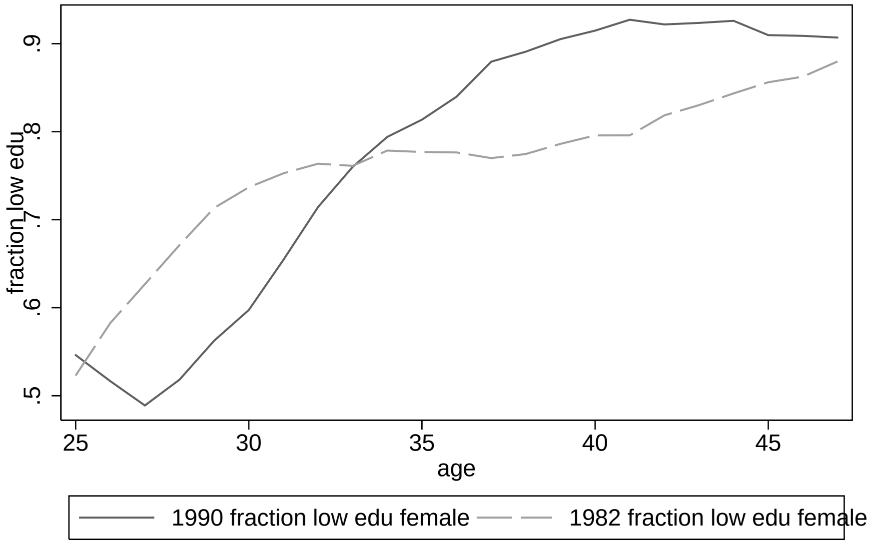


Figure 12a: Anhui fraction low education female

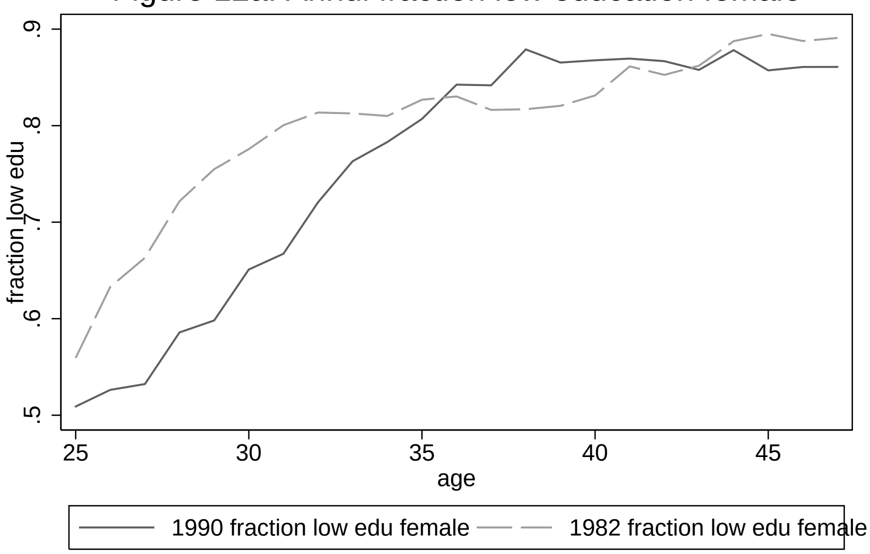


Figure 13: Sichuan growth of fraction low education female

Figure 13a: Anhui growth of fraction low education female

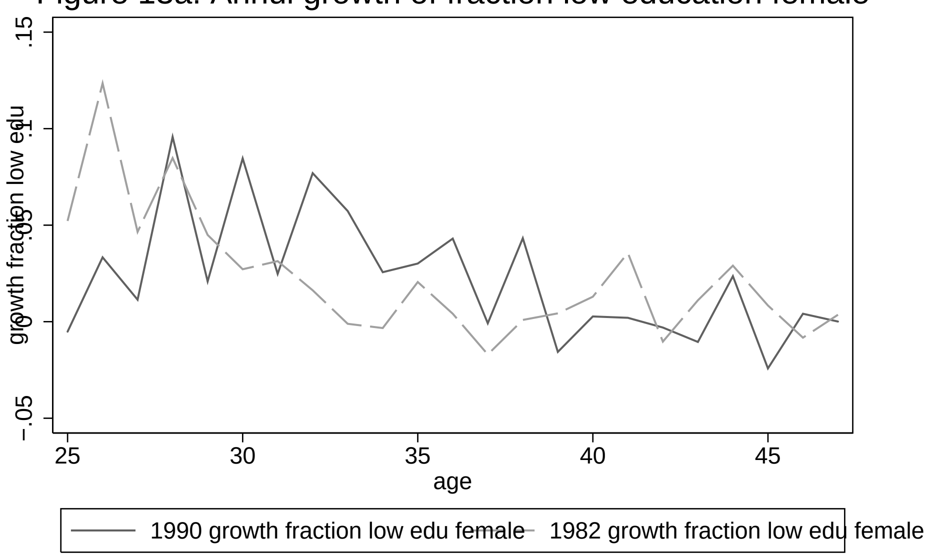


Fig 14: Sichuan 1990/1982 female mar. rates by edu

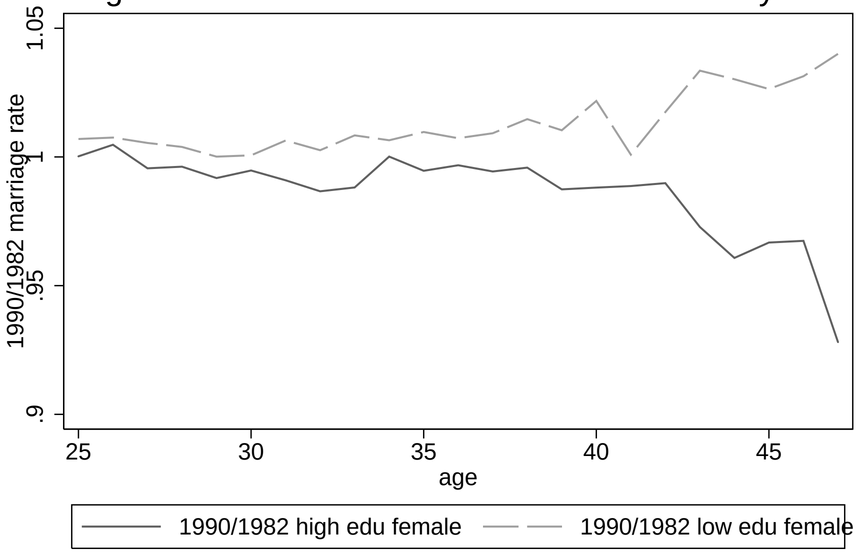


Fig 14a: Anhui 1990/1982 female mar. rates by edu

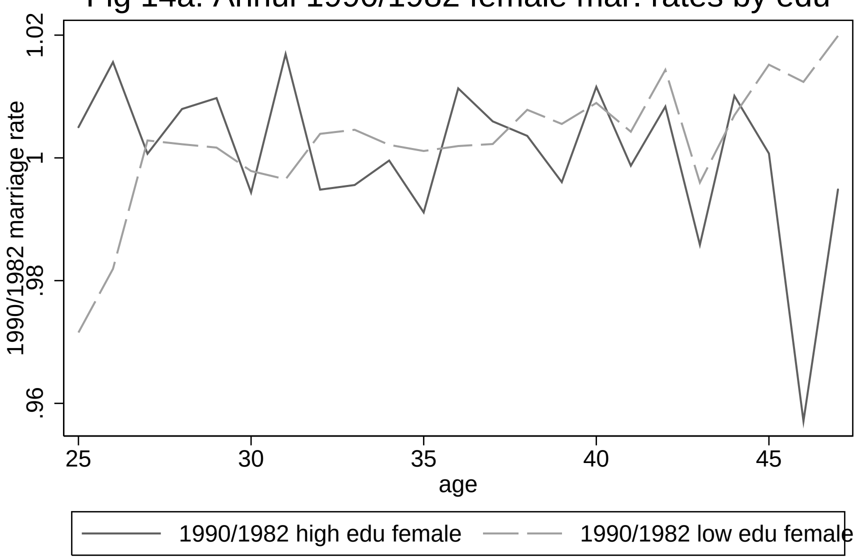


Figure 15: Sichuan predicted over actual mar. rates by age and edu

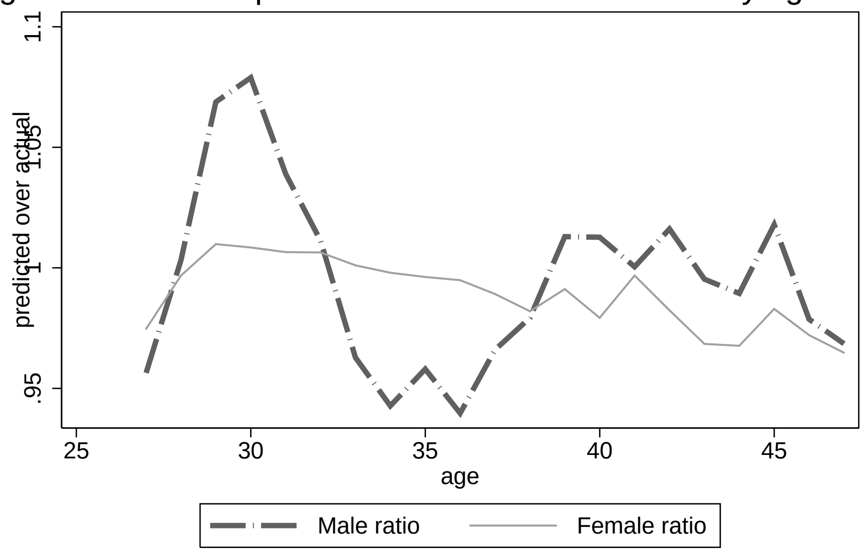


Figure 15a: Anhui predicted over actual mar. rates by age & edu

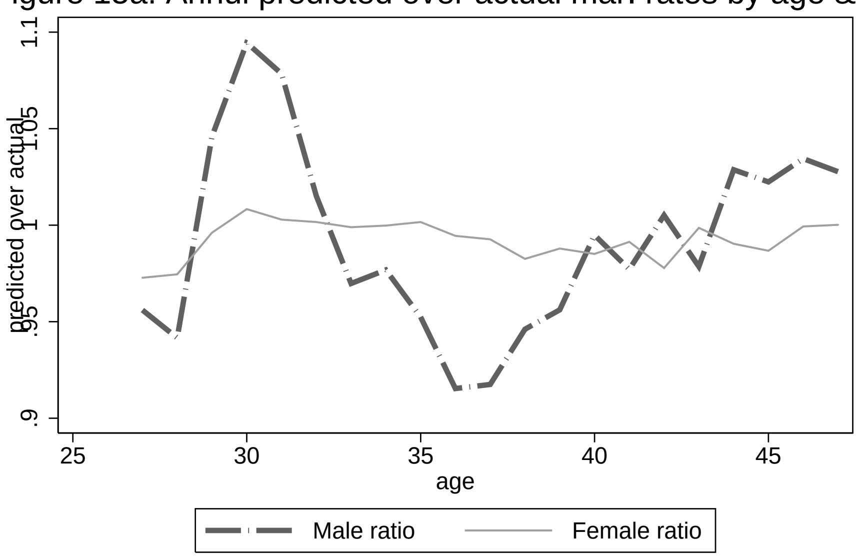


Figure 16: Sichuan HH 1990−1982 total gains

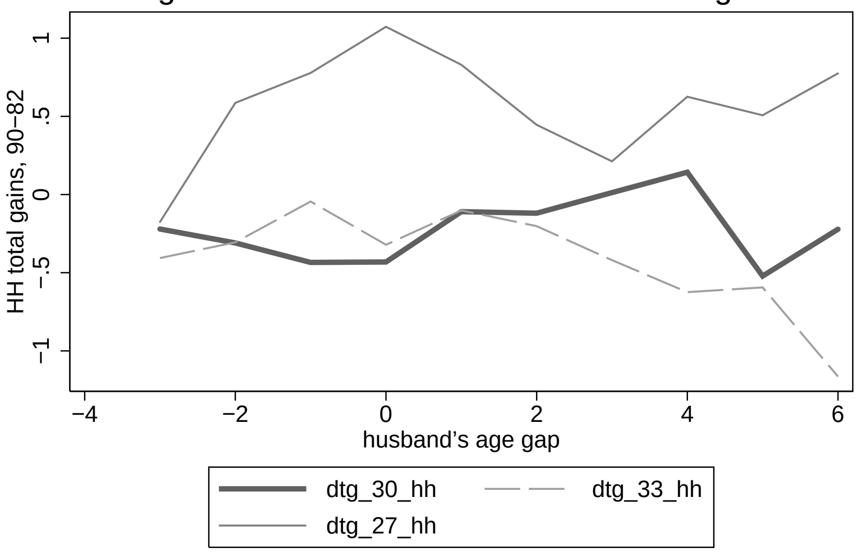


Figure 16a: Anhui HH 1990−1982 total gains

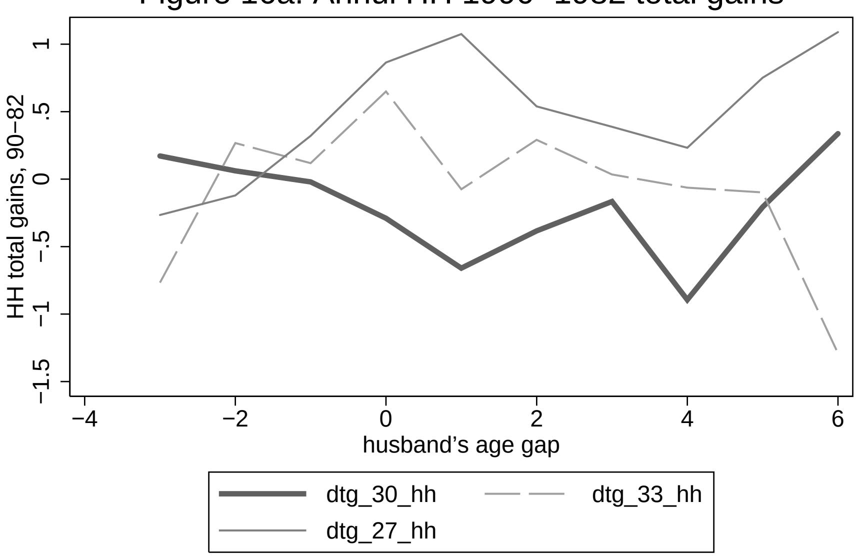
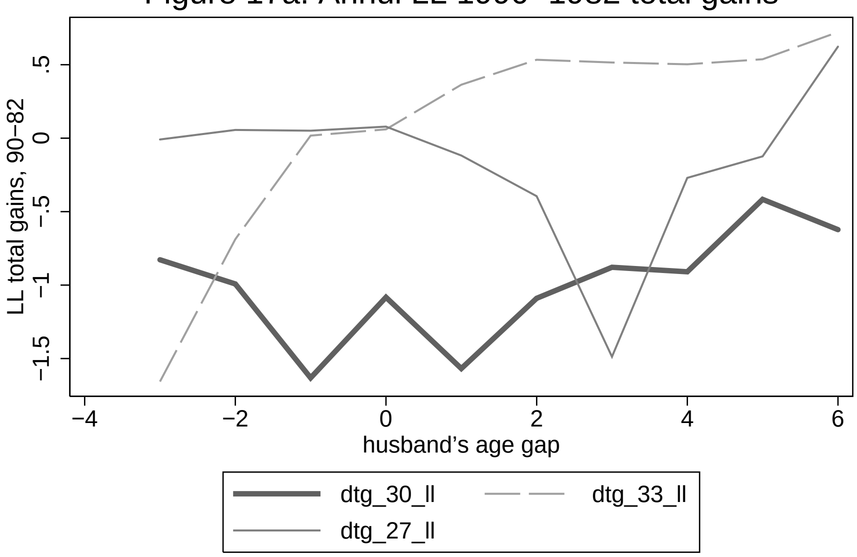


Figure 17: Sichuan LL 1990−1982 total gains

Figure 17a: Anhui LL 1990−1982 total gains



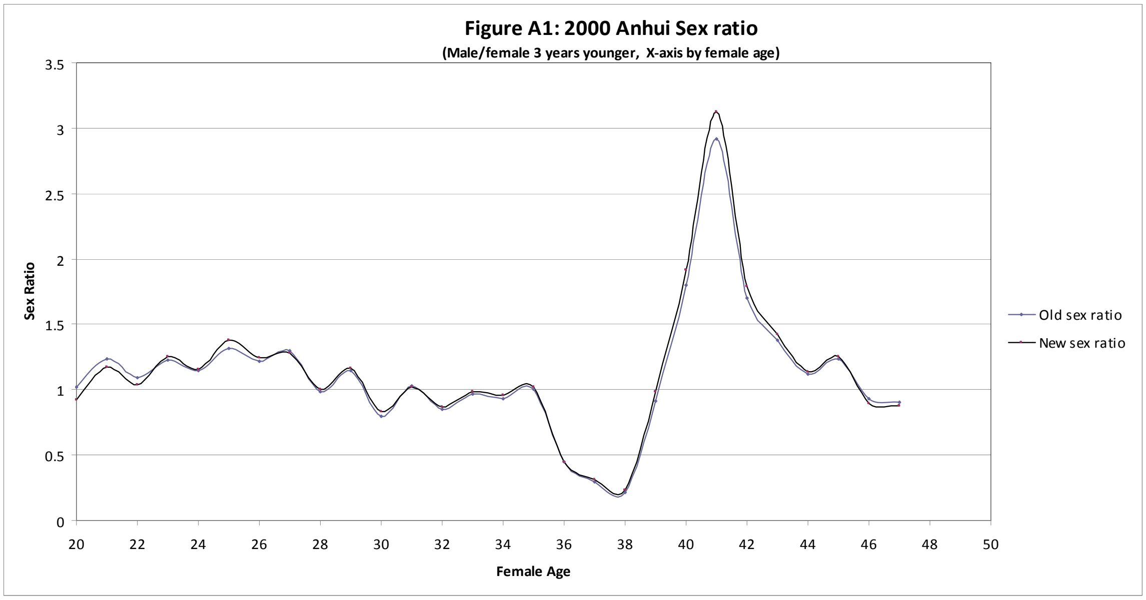


Figure A1: 2000 Anhui Sex ratio
(Male/female 3 years younger, X-axis by female age)
Sex Ratio
Female Age
Old sex ratio
New sex ratio
0
0.5
1
1.5
2
2.5
3
3.5
20
22
24
26
28
30
32
34
36
38
40
42
44
46
48
50

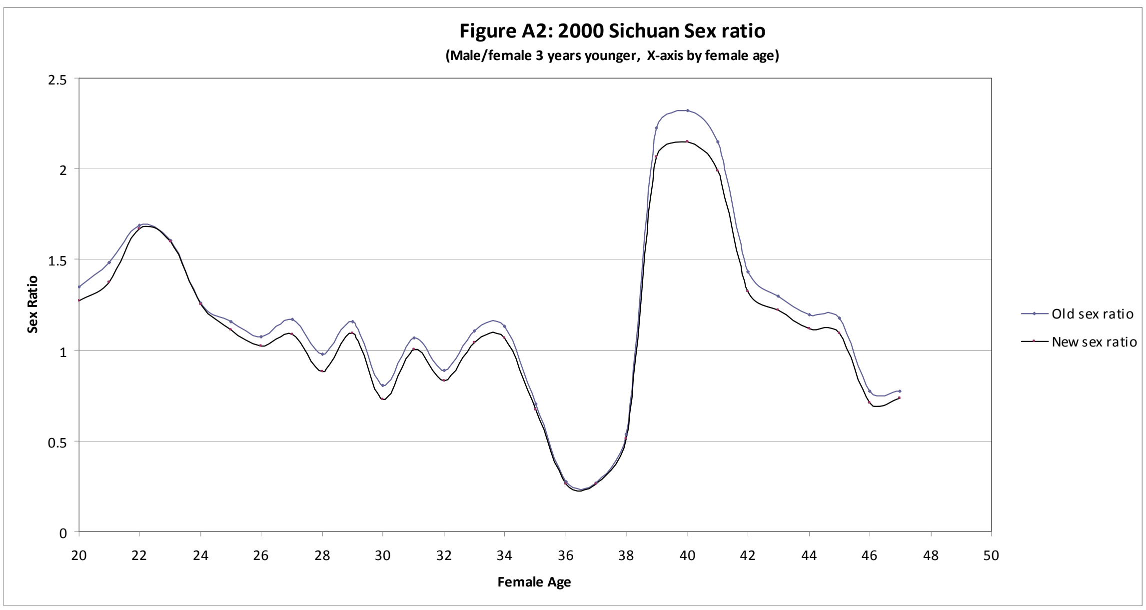


Figure A2: 2000 Sichuan Sex ratio
(Male/female 3 years younger, X-axis by female age)
Sex Ratio
Female Age
Old sex ratio
New sex ratio
0
0.5
1
1.5
2
2.5
20
22
24
26
28
30
32
34
36
38
40
42
44
46
48
50