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How do sex ratios in China influence marriage decisions and intra-household resource allocation?

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Abstract

This article examines how imbalanced sex ratios influence marriage decisions and household bargaining. Using data from the 1982 Chinese census, the traditional “availability ratio” is modified to reflect the degree to which men tend to marry women from different cohorts. This ratio reflects the average tendency of men to prefer women who are close in age to women who are several years younger than them by weighting cohort sizes using the proportion of people in the population who marry someone born in a different cohort. Given that men generally marry younger women, this ratio varies independently of the size of one's own birth cohort. Yet, the ratio fluctuates considerably across individuals, as the sizes of birth cohorts in China vary across time and regions. This enables us to examine how variability in such ratios may influence marriage decisions and household bargaining. The findings suggest that women exercise greater bargaining power once married. Results indicate that as women become scarcer in the marriage market, they have healthier sons. Men also delay marriage, and consume less tobacco and alcohol. This paper also highlights how sensitive findings may be to using this modified weighted availability ratio rather than a traditional unweighted availability ratio.

Keywords

household bargaining; marriage; China; availability ratio

1. Introduction

This paper examines how high sex ratios in China impact individuals' decision-making behavior once married. While a number of studies have provided explanations for such imbalanced ratios,³ their potential effects remain relatively unexplored. Becker's marriage market model and related theories suggest that as women become scarcer in the marriage market relative to men, they receive larger transfer payments from their husbands. That is, women's share of the “marital surplus” increases (Becker 1973; Choo and Siow 2006a, 2006b). Such allocation of household resources is also determined by a household

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³See the following studies: Coale 1984, 1991; Coale and Banister 1994; Oster 2005; Das Gupta 2005; Ebenstein 2007; Lin and Luoh 2008; Oster et al. 2008; Qian 2008; Almond et al. 2010; and Hamoudi 2010. As of 2005, China's sex ratio at birth was 119 boys for every 100 girls. Among children ages one to four this figure rose to 125 in the overall population, reaching as high as 142 in some provinces (Zhu et al. 2009).

bargaining process that results in what is termed the household “sharing rule” in the collective model of household bargaining. In this model, the sex ratio is a distribution factor that shifts the household sharing rule in favor of the scarcer sex (Browning et al. 1994; Chiappori et al. 2002).

In this paper, I test whether the sex ratio at marriage is such a relevant distribution factor in determining intra-household resource allocation in China.⁴ I find that high sex ratios improve the health of sons, supporting the hypothesis that women who face relatively favorable marriage market conditions have greater bargaining power once married.⁵ This is further supported by the fact that sex ratio effects remain unchanged when controlling for spouse characteristics. High sex ratios also delay the age at which men marry, and reduce the amount of tobacco and alcohol a husband consumes. In comparing the effects of an imbalanced sex ratio of 120, relative to the usual average ratio of 107, this difference of 13 men per 100 women would improve boys' short-term health by roughly 31%, would raise the age at which men marry by nearly six months, and would reduce men's daily tobacco consumption by 56%.

These findings suggest that at the time of marriage, spouses are able to commit future resource allocation, supporting prior research showing that a marriage market with assortative matching results in a unique sharing rule that lasts throughout the marriage and is, in part, a function of the sex ratio in the marriage market (Iyigun and Walsh 2007). Empirical evidence of such commitment was also found in South Korea (Lee 2007). Such commitment may be at least partly due to path dependence because of relatively high transition costs to altering initial arrangements (Arthur 1994). Though, such commitment may be limited and renegotiated, particularly when circumstances change. For example, social programs allocating money specifically to women also influence how resources are distributed within the household (Mazzocco 2007).

The findings in this paper also complement prior research on famine-affected cohorts and their relatively low marriage rates at particular ages (Brandt et al. 2008; Almond et al. 2010). However, related studies assume a fixed age difference between spouses.⁶ For example, the number of men aged 22 to 28 would be divided by the number of women aged 20 to 26, an estimation that ignores the degree to which a man aged 22 prefers to marry a woman aged 20 to one aged 26.

In contrast, the sex ratio developed in this article weights cohort sizes by the proportion of people in the population who marry someone born in a different cohort. In this particular case, where fertility has fluctuated considerably, cohort sizes vary significantly over very brief intervals. Without such weights, relatively large cohorts are equally weighted to smaller ones, which would overestimate the number of available mates, should smaller cohorts be preferred.

⁴In related work in the U.S. and Taiwan, a high sex ratio leads to improved bargaining power for women (Grossbard-Shechtman 1993; Angrist 2002; Grossbard and Amuedo-Dorantes 2007; Francis 2011).

⁵In prior research, increases in female income or greater control of household assets leads to greater investment in children (e.g., Thomas 1990; Duflo 2000, 2003; Rangel 2006; Qian 2008; Brown 2009; Rubalcava et al. 2009).

⁶For example, see: Grossbard-Shechtman (1993), Bergstrom and Lam (1994), Angrist (2002) Chiappori et al. (2002), Brandt et al. (2008), and Francis (2011).

The sex ratio developed in this article reflects the average tendency of men to prefer women who are close in age to women who are several years younger than them. We can account for such preferences across different ages by using the observed population distribution in age differences between spouses to proxy for the net gains from marrying someone of a given age. For example, the proportion of men in a population marrying women who are two years younger than them approximates an average likelihood of a particular man marrying such a woman. This is because such average rates have been shown to reflect the average benefit for a man to such a marriage, in comparison to him marrying a woman of a different age, or remaining single (Choo and Siow 2006a, 2006b).

Observed differences in ages for individual cohorts reflect a combination of preference parameters and the amount of net transfer payments made to the scarcer sex. In this case, the share of household resources given to a woman can be considered the “price” for marrying a particular woman. So I examine overall average age differences across spouses from cohorts not under study here, who face different prices. I also demonstrate that such population trends in average age differences between spouses are not influenced by particular cohorts that face especially stringent prices or marriage market conditions.

Finally, it is important to note that we can never be assured that population sex ratios are completely exogenous to the individual. To the extent possible, I control for some potential bias due to endogeneity or selection issues resulting from the circumstances unique to one's birth cohort. First, I do so by controlling for the size of one's birth cohort, as many sex ratio effects are determined by one's birth province and year. An additional control variable reflects the degree of competition from neighboring cohorts outside the marriage market such as competition for employment opportunities or education-related resources like teacher availability and class size. Another potential concern is that cohorts that differ in systematic abilities and life-forming early work experiences are averaged together in the ratio, which could potentially introduce biases related to these time-varying changes in unobservables. To address such potential biases, three additional controls are included in the regressions: the number of women in the youngest and oldest cohorts of the sex ratio, and the average completed level of education as of 2000 of all women who are included in the sex ratio. As these population-level variables may themselves be influenced by individual characteristics, I also estimate regressions excluding them. In addition to these robustness checks, I find that sex ratios do not influence possible placebo outcomes: one's height or the likelihood of completing schooling beyond the primary level. Finally, I discuss limitations to these methods.

In the context under study here, sex ratios resemble the degree of scarcity of women relative to men in other developing countries also experiencing sharp demographic changes that are somewhat similar to historical fertility fluctuations experienced in China.⁷ In introducing a new way to measure sex ratios characterizing marriage markets, this paper improves our

⁷One consequence of declines in fertility is a “marriage squeeze” (Schoen 1983), as men generally marry women who are several years younger. Women born during periods of low fertility marry men born earlier in larger cohorts. As a result, these men are faced with a relative scarcity of women, particularly if the women they most want to marry (those generally two to three years younger) cannot be easily substituted for slightly younger cohorts, also born during low fertility periods. The largest difference in sex ratios in this article is roughly 23%, where sex ratios range between 102 and 125. One related study has examined how a very large influx of

understanding of how shifts in the population affect marriage markets, and how imbalanced population sex ratios relate to individual household decisions. Overall, the findings suggest that if a man is born during a sustained period of low fertility or between periods of high and low fertility, he is faced with a great deal of competition in the market for a spouse. As a result, he must negotiate a smaller share of the marriage surplus.

The remainder of this paper is organized as follows. Section 2 explains how marriage market conditions are estimated. Section 3 outlines the estimation strategy and describes the data used in the analysis. Section 4 discusses the empirical results, and Section 5 concludes.

2. Estimating marriage market conditions in China

2.1. Background literature on sex ratios and availability ratios

Widely used sex ratios stagger male and female cohorts to reflect the fact that women generally marry older men.⁸ For example, the ratio would be the number of men aged 22 to 28 to the number of women aged 20 to 26. This ratio is equivalent to an un-weighted “availability ratio” (Goldman et al. 1984, Fossett and Kiecolt 1991).⁹

This ratio is very appropriate for the circumstances in which it has been used. However, in this particular case in China, where cohort sizes vary significantly over very brief intervals (often annually), there is one potential issue with such an estimation; the ratio does not take into account the degree to which a man aged 22 prefers to marry a woman aged 20 to one aged 26. If such preferences are not taken into account and relatively large cohorts are equally weighted to smaller ones, should smaller cohorts be preferred, the number of available mates would be overestimated. Indeed, a weighted availability ratio has been developed by demographers to describe a ratio in which the number of unmarried men and women of each age are weighted by the marriage rates of each age combination, thus incorporating the number of unmarried individuals remaining in each age group (Goldman et al. 1984). Critics of this weighted ratio claim that the assumptions behind it are too restrictive with respect to preference structures and that it is highly correlated with more general ratios (Fossett and Kiecolt 1991).

But there is another potential issue with this availability ratio: as the ratio incorporates the number of unmarried individuals, it reflects the number of eligible marriages at any given point in time once marriages have already begun to take place, rather than indicating the number of potential marriages before any marriages of the population take place. The reason this is important is that the availability ratio is influenced by the age at marriage. For example, supposing that all men marry younger women, since nearly all other women would be married, only younger women would be included in the ratio. But since both young and old men are single, they are both included in the ratio, resulting in nearly twice as many men to women in the ratio. At the extreme, the availability ratio would potentially be equal to

mainland Chinese soldiers to Taiwan influenced household bargaining (Francis 2011), where the influx of Chinese soldiers raised the sex ratio for those ages 20 to 24 by about 57% (from 97 men per 100 women to 152 men). While this is an important and unique study, it is unclear how generalizable the findings are beyond this particular case.

⁸For example, see: Grossbard-Shechtman (1993), Bergstrom and Lam (1994), Angrist (2002), Chiappori et al. (2002), and Grossbard and Amuedo-Dorantes (2007).

⁹For a comprehensive survey of research studies implementing this ratio, see Chiswick and Houseworth (2011).

two.¹⁰ This traditional availability ratio includes only non-married men and women, who have remained unmarried, in part because of the very same marriage market conditions the ratio is intended to estimate.

In addition, since a significant proportion of men marry someone two years younger than themselves,¹¹ it is evident that men would prefer this age gap to marrying someone two years older. The marriage rate observed in the population has been shown to be a good proxy for marriage preferences over different types or ages because it reflects the net gains in random utility from marrying a particular type (or age) over remaining single (Choo and Siow 2006a, 2006b).¹² In fact, social scientists have often taken the marriage rate to be a reasonable proxy for the gains to marriage (e.g., Wilson 1987).

2.2. A new “Generalized Weighted Availability Ratio”

In this sub-section, we construct a sex ratio which reflects the preferences men have for women of different ages without incorporating decisions resulting from undersupplies of women of different age groups. This ratio differs from prior weighted “availability ratios” (Goldman et al. 1984) in that it reflects the number of potential marriages in a population rather than the number of eligible marriages at any given time once marriages have begun to take place.

In constructing the ratio, we combine marriage rates between men age i and women age j with the particular number of men age i and women age j that are available for marriage in a given cohort. The distribution in marriage rates for different (i, j) pairings approximates the average price for an (i, j) match. This distribution in age differences does not reflect changing marriage market conditions because it is fixed across all observations and is estimated from the population census. As women become scarcer, men must adjust their behavior by marrying women with higher transfer costs, but such rising costs limit the extent to which men can adjust to such a rising scarcity of available women. Weighting specific cohort sizes by an overall population distribution in age differences approximates the relative costs (or gains) on average from marrying someone i or j years older.

This is similar to the notion of an average price for a market good, which cannot be adjusted by any individual consumer in a competitive market, regardless of the specific price negotiated by that consumer. Thus, the average population distribution of (i, j) pairings is similar to the average price for a good in a market of atomistic consumers and producers, which is unaffected by the individual price exchanged between a particular k producer and g consumer of any given (i, j) pairing. While the bargaining between one buyer and seller results in a given price, this negotiated price does not influence the average price for the good in the overall market. Similarly, the particular transfer payments from a husband k in population i to his wife g in population j does not influence the average transfer payment

¹⁰To illustrate this point further, while 78% of women over 15 in China were ever married, only 70% of men over 15 were ever married. While the total sex ratio was 1.05, the single sex ratio was 1.16. The single sex ratio by age also shows a strongly increasing pattern by age, until ages 36-40, when the single sex ratio declines.

¹¹In the 1982 Chinese census, only 10% of women were married to men born in their same year.

¹²Choo and Siow (2006a, 2006b) propose a static transferable utility marriage market model to demonstrate that the probability of a woman of type j marrying a husband of type i is proportional to her net gains from doing so. While their model refers to different types more generally, I define type by age here, since my focus is on people marrying across different ages.

amount across all i husbands to all j wives. This notion of atomistic buyers and sellers in a marriage market is particularly applicable because small changes in the ratio are examined here.

Sex ratios relevant to one's marriage market are estimated according to whether one is from a rural or urban area, and one's birth year and province. Interprovincial marriages were relatively rare through the late 1980s (Fan and Huang 1998), and fertility fluctuated considerably by province, particularly since some areas were more severely affected by the famine than others (Peng 1987). While this marriage market area might be overestimated for some; for others, a narrower regional definition such as one's county of birth, might underestimate the relevant marriage market. As the survey data used here only provides information on one's birth province, defining a marriage market based on any smaller geographic unit would not be possible without exacerbating issues of endogeneity. For example, one's county of residence is endogenously chosen and may have been influenced by marriage market conditions in one's county of birth. The possibility of such endogeneity bias is of much greater concern than overestimating the geographic area of the marriage market.

Urban and rural marriage markets are assumed to be separate, an assumption supported by prior research (Fan and Huang 1998). In China, every person is given either an urban or rural *hukou* (registration classification for residence). Through the 1980s, a person could change one's rural *hukou* to an urban one by graduating from a post-secondary institution in a designated urban area. However, individuals with rural *hukous* generally had difficulty in being accepted to these institutions. Rural areas are far poorer than urban areas, and for the cohorts in this sample, it has generally been highly undesirable for those with an urban *hukou* to marry rural *hukou* holders.

Considering urban and rural populations separately, availability ratios are derived using the 1982 Chinese population census available by province of residence.¹³ Cross-province migration did not become common until the mid-1980s, when agrarian reforms and relaxed restrictions on migration encouraged migrants to move to areas with greater economic opportunities.¹⁴ Around 99% of the enumerated population was residing in the same location during the 1982 census collection period where they had been registered at birth, with only 1% of the population living in a place different from where they had been registered (which in most but not all instances would be the place where they were born). Since many of these individuals did not move across provinces, less than 1% of the population lived in a province different from their place of registration (or birth).

Cohorts used in estimating ratios include men and women born between 1932 and 1981. Those born prior to 1932 were over 50 years old at the time of the census, and since mortality rates between men and women begin to diverge at this age, using data for these older cohorts would underestimate the number of men available in the marriage market. To prevent an overestimation of the number of men available for marriage, the 1982 Chinese

¹³Data was provided by the Minnesota Population Center.

¹⁴See the following studies: Chan (2001), Chan et al. (1999), Fan and Huang (1998), and Huang and Zhan (2005).

census is used because the timing is appropriate for the sample under study. The data reflects the number of men and women available for marriage in each province around the time at which those born during the high mortality period of the famine were marrying.¹⁵ By this time, they had either reached or were close to reaching marriageable age in 1982 (i.e., they were age 23 or younger).¹⁶

Now consider, for example, how one might construct a sex ratio reflecting the marriage market conditions for a man born in 1960. For computational simplicity, suppose he is interested in women who are between nine years younger and three years older than his age, that is, women born between 1957 and 1969. Supposing all these women are interested in marrying a man between three years younger and nine years older than them, women born in 1957 would be interested in marrying a man born between 1948 and 1960. Similarly, women born in 1969 would be interested in marrying a man born between 1960 and 1972. Thus, while a man born in 1960 is interested in marrying women born between 1957 and 1969, he is competing for these women with men born between 1948 and 1972.

Let F_{a-i}^{rp} be the number of women of age $(a-i)$ born in province p and urban or rural area r . Let us suppose a man is considering marrying women of different i types where i now indexes the age difference between him and his possible match. For example, for a man interested in marrying a woman between nine years younger and three years older, $i \in [-3, 9]$. Further, let us suppose that urban and rural men have different preferences for substituting across different types. Let ρ_{-i}^r be the fraction of men of all ages marrying women i years younger, where r indexes whether the sum is across all rural or urban men. Thus, let

$\sum_i \rho_{-i}^r F_{a-i}^{rp}$ be the number of women available to a man of age a born in province p . To estimate the number of men with whom he is competing for these women, let ρ_j^r be the fraction of women in urban or rural area r who have married a man j years older, where $j \in [-3, 9]$. Let M_{a-i+j}^{rp} be the number of men of age $(a-i+j)$ born in province p . Thus, $\sum_j \rho_j^r M_{a-i+j}^{rp}$ is the number of men available to a woman of age $(a-i)$. Now we take each of these weighted averages and weight them by the fraction of all married couples in urban or rural China in which the age difference between spouses is i . Summing across all i results in the number of men with whom a man of age a is competing. Taking the ratio of these two weighted averages results in the following Generalized Weighted Availability Ratio (GWAR) relevant to men:¹⁷

$$R_{m,a}^{rp} = \frac{\sum_i \rho_i^r \cdot \sum_j \rho_j^r M_{a-i+j}^{rp}}{\sum_i \rho_{-i}^r F_{a-i}^{rp}} \quad (1)$$

¹⁵According to the 1982 census, the sex ratio by year of birth declines significantly for the years of the Great Famine and those immediately following. Boys born during this time period were less likely to survive to marriageable ages.

¹⁶The mean age at marriage is 25 for men and 23 for women.

¹⁷We would similarly estimate availability ratios from a woman's perspective.

The denominator indicates the number of women “demanding” a man of age a , and the numerator indicates the number of men with whom a man of age a is competing (the “supply” side).

To estimate relevant weights in this particular weighted availability ratio, we need to determine the relevant population distribution in spousal age differences which would reflect preferences for marrying different age groups in the overall population, but would also minimize the extent to which such weights would be endogenous to a particular individual, cohort, or group, particularly one that has faced a shortage of potential spouses. To begin, we estimate the distribution in spouse age differences in the sample of married urban and rural men in the 1982 Chinese census born between 1932 and 1964.¹⁸ The third column in Table 1 provides the distribution in spousal age differences ranging from -3 years (husband is three years younger than his wife) to +9 years (husband is nine years older than his wife). This range accounts for 92% of couples.

In order to minimize the extent to which the Generalized Weighted Availability Ratio reflects biases due to changing marriage market conditions, rather than use this distribution in spouse age differences, we restrict the sample to cohorts born prior to those under study here. As availability ratios are not calculated for cohorts born between 1932 and 1937, ratios are constructed using the distribution in spouse age differences for this subsample.¹⁹

Three different variants of the distribution for this smaller subsample of the 1982 Chinese census are shown in Table 1.²⁰ Estimating GVAR1, where $i \in [-3, 9]$, requires fertility data for the thirteen cohorts above and below an individual. Using a distribution closer to the observed one would require fertility data from additional cohorts, which would further restrict the number of available observations. Results are also compared across ratios estimated using two other distributions (see Table 9). To widen the available data, the age difference is further restricted to $i \in [-2, 6]$ (GVAR2), requiring fertility data on nine cohorts above and below that of any given individual; as well as $i \in [0, 5]$ (GVAR3). Figures 1 and 3 depict the male GVARs for urban and rural China.

These ratios are compared to the Generalized Un-weighted Availability Ratio (GUAR) in Figures 2 and 4. To construct this ratio, a given individual's birth year is the median birth year of the cohorts included in the numerator. The denominator is a similar sum across cohorts of women, where the median cohort is two years younger than the median cohort of the men included in the numerator. GUAR1 is estimated using 15 adjacent cohorts; GUAR2 involves using nine cohorts; and GUAR3 uses five cohorts. For example, GUAR3 for a man born in 1952 would be the number of men born between 1950 and 1954 divided by the

¹⁸The sample is restricted to these years because mortality rates for men tend to increase after age 50 and men cannot legally marry below the age of 18.

¹⁹There are some limitations to this approach. First, weights determined by earlier cohorts introduce some measurement error in proxying for the weights relevant to the population under study, to the extent that these population averages differ across the two sets of cohorts. Second, these weights may be determined by some omitted socioeconomic factors which may also influence outcomes examined in the analysis here. I am grateful to a reviewer for pointing this out.

²⁰In general, distributions estimated from the two samples are quite similar. In comparison to the wider sample of men born between 1932 and 1964, men born between 1932 and 1937 are less likely to marry women closer in age, and are more likely to marry women who are increasingly younger (four to nine years younger in rural areas and three to nine years younger in urban areas).

number of women born between 1952 and 1956. This ratio places equal weight on the probability that a man aged 25 will marry a woman aged 22, or aged 27.

As the figures above show, differences in ratio calculations become important when fertility fluctuates sharply. Fertility grew steadily through the 1950s, and it was higher in those regions that were more economically developed (Peng 1989). During this initial period of fertility growth, all sex ratios declined similarly, though sex ratio estimates diverge during brief periods of fertility fluctuation. During the Great Famine, which affected most provinces between 1959 and 1961, declines in fertility began in the more developed provinces in the east (Peng 1989). In the years leading up to and during the famine, sex ratios differ considerably from one another. In the recovery period after the famine, fertility rose sharply. Fertility rates in rural and urban areas began to diverge, with fertility in urban areas returning to famine levels. The latter decline was also primarily due to government interventions. For example, urban youth were sent to temporarily live in rural areas during the Cultural Revolution and the government campaigned to reduce urban fertility. Thus, urban sex ratios differ from one another during this period of sharp fluctuation. In rural areas, fertility remained high until the 1970s, particularly in less developed areas (Birdsall and Jamison 1983). Rural sex ratios rise quite similarly because of the persistent decline in rural fertility in the 1970s.

3. Empirical model and data

Having described the different variants of availability ratios, I now outline the estimating equation used in determining how this ratio relates to marriage decisions and household bargaining.

3.1 Estimation strategy

The following regression is estimated:

$$\begin{aligned}
 y_{im,a}^{kp} = & \beta_0 + \beta_1(\text{sexRatio})_{m,a}^{kp} \\
 & + \beta_2(\text{OwnCohort})_{m,a}^{kp} \\
 & + \beta_3(\text{CompetingCohorts})_{m,a}^{kp} \\
 & + \beta_4(\text{AvgFemaleEduc})_{m,a}^{kp} \\
 & + \beta_5(\text{NumYoungerWomen})_{m,a}^{kp} \\
 & + \beta_6(\text{NumOlderWomen})_{m,a}^{kp} \\
 & + \beta_7 X_{im} + \mu_a + v_k + \lambda_p + (v_k + \lambda_p) + \varepsilon_{im,a}^{kp} \quad (2)
 \end{aligned}$$

Standard errors are clustered by the husband's hukou type and birth province. To address potential biases resulting from clusters numbering less than thirty, I implement a wild cluster bootstrap-*t* procedure using the method outlined by Cameron et al. (2008).²¹ I report *p*-values for coefficient estimates on sex ratios using this wild clustering method below the *p*-values based on the standard clustering correction approach.

The outcome measure for couple i of cohort a is denoted by $y_{im,a}^{kp}$, where the subscript m indicates that the variable applies to the male partner; k indicates whether the respondent is from a rural or urban area; and p indicates the province of birth. Outcomes include age at marriage; age gap between spouses; male daily tobacco and alcohol consumption; and boys' and girls' weight for age z scores and height for age z scores. The latter variables on children's health are constructed using reference data from the U.S. National Center for Health Statistics, where children's z scores are calculated by subtracting each measure from the median measure of the same age and sex group of children in the U.S. and dividing by the standard deviation for this group. In regressions on this outcome, the unit of analysis is the child, whose observation is taken from the most recent survey conducted while the child was age ten or younger. For other outcomes, the unit of analysis is the couple.

To control for circumstances unrelated to the marriage market that are specific to the male partner's birth cohort, the size of the male partner's birth cohort is included in all regressions. This variable controls for some selection issues related to the famine, school class size, and competition when entering the labor market.

Such competing pressures may also come from adjacent birth cohorts. For example, male partners may compete for jobs with men born a number of years later. Competition for limited schooling resources while growing up might also impact labor market opportunities into adulthood, which may influence one's attractiveness in the marriage market. To control for some selection issues related to such competitive pressures, the number of competing men in adjacent cohorts is included in regressions. This is simply the numerator of the sex ratio used in the regression.

Another potential concern is that averaging across years to estimate the sex ratio may introduce additional time-varying unobservables that are systematically related to both outcome measures and regressors. For example, during the 1980s and 1990s, educational curricula changed considerably and education capacities were expanded. Young women at the beginning of the 13 year window used in estimating sex ratios may have very different initial labor market experiences from young women at the end of the 13 year window. For example, women who completed schooling and entered the job market in the early 1980s may have had very different experiences from those who entered the market in the mid-1990s. One potential concern is that cohorts that differ in terms of systematic abilities and life-forming early work experiences are averaged together in the availability ratio. Such differences could potentially introduce biases that are related to these time-varying changes in unobservables.

To address such potential biases, three additional controls are included in the regressions: the number of women in the youngest and oldest cohorts of the sex ratio, and the average

²¹For each bootstrap iteration, I impose the null hypothesis that the sex ratio has no effect on outcome measures. I resample residuals at the cluster level to preserve any correlation between individuals within a province and hukou type. These residuals and the covariates are then used to obtain a predicted outcome y that does not include the effect of the sex ratio. This predicted y is then regressed on the complete set of covariates that includes the sex ratio. For the k th iteration, estimates of $\beta_{1,i}^*$ and its standard error yield corresponding Wald statistics. The distribution of these Wald statistics (under the null hypotheses) is collected over 999 iterations in order to compute critical values.

completed level of education as of 2000 of women who are included in the sex ratio. Education levels of cohorts are weighted in the same way they are weighted in the GVAR. This variable is constructed using data from the 2000 census because this census year is closer to the outcome variables from the CHNS survey than is the 1982 census. Note that regressions on this province-level average education of women in the ratio indicate that this measure is not correlated with availability ratios²²

Remaining unobservables that differ systematically across male cohorts and rural or urban provincial areas are largely controlled for by province and year of birth fixed effects (X_p and μ_a). A dummy for whether the respondent holds a rural or urban hukou is also included (v_k), along with interaction terms between the latter and birth province fixed effects. Thus, many remaining unobservable differences in factors such as education, employment, or other opportunities that differ systematically across birth cohorts and provinces are controlled for by these fixed effects.

Additional individual-specific controls in the regressions are indicated by X_m and include the height of the male partner of couple i (a common measure of an adult's long-term health). For regressions on child health—estimated separately for boys and girls—the age of the child is also included as a regressor.

An additional individual-specific regressor is included to control for the husband's exposure level to the famine. Following Almond et al. (2010), this is the weighted average death rate during the time the individual was in utero, and it is based on his month, year, and province of birth. Averaged death rates are weighted by the number of months one was exposed to higher than average death rates while in utero. For example, if one was born in February 1960, the weighted average death rate = $(2/9) * (1960 \text{ death rate}) + (7/9) * (1959 \text{ death rate})$. This data is available only for years of birth during and near the time of the famine. Given this measure captures exposure to the famine while in utero, those who were not in utero during the famine are given a zero value for this measure.

The correlation between individual famine exposure and GVAR1 is -0.1295. Similarly, GUAR1 is not highly correlated with the degree of famine exposure (-0.1115).²³ With these weakly negative correlations, those with higher ratios may have had slightly less exposure to the famine, and may be slightly healthier than those who had more exposure (Meng and Qian 2006; Chen and Zhou 2007; Almond et al. 2010). If this is the case, then men with a higher ratio may have healthier children not because of their wives' increased bargaining power due to a scarcity of women available to them in the marriage market, but because these men were less exposed to the famine. On the other hand, survival selection may bias marriage market effects in the opposite direction, in which case, men with a higher sex ratio, who were less exposed to the famine, would be less well-off than their counterparts (Gørgens et al. 2007; Mu and Zhang 2008). The effect of the GVAR on children's health would be underestimated if men with a higher sex ratio were less exposed to the famine and

²²These regressions were estimated at the province-residence type-birth year level. Results are available upon request from the author.

²³In regressions on famine exposure averaged to the province-year-residence type level, availability ratios do not predict this measure. Results are available upon request from the author.

were less healthy than those who survived the famine. To the extent possible, we control for such potential selection issues with the control variables outlined here.

Of course, as with other work in this area, there are limitations to this method in eliminating all selection bias and identifying a causal impact of sex ratios on outcomes. The weights used in estimating the GVAR are themselves general equilibrium outcomes of the marriage market. It is also possible that the number of males and females born in a given year are endogenously related to those born in the same year or in adjacent years.

Similarly, other regressors used here may also be endogenously related to the individual. For example, the province-level average education of women born in a certain time period may be related to individual characteristics. Further, these covariates may be related to one another. For these reasons, I also estimate regressions excluding these variables. First, I estimate regressions of outcomes on ratios alone. Second, I add the following controls: male hukou, male birth year and province fixed effects, interaction terms between hukou status and birth province fixed effects, and CHNS survey year dummies. I then estimate the full regressions outlined above.

To address the possible endogeneity bias of the weights in the GVAR, I estimated two-stage least squares regressions in which the GVAR was instrumented by the GUAR, and other variables generated with the weights were similarly instrumented by their unweighted counterparts. Durbin-Wu-Hausman tests indicated no statistically significant difference between IV and OLS estimates.²⁴ As the null hypothesis of exogeneity of the GVAR cannot be rejected, OLS is preferred to IV, since the OLS estimates are consistent and unbiased, and more efficient than the IV estimates. Had we rejected the null hypothesis, then OLS estimates would be inconsistent and biased, and IV estimates would be preferred. But the results of this test do not necessarily imply that the null hypothesis is true. It is possible that the unweighted variables are weak instruments for the weighted variables, in which case differences between IV and OLS would not be statistically significant. In addition, this method only addresses the possibility of endogeneity of the weights, but not the potential endogeneity of cohort size (numbers of males or females).

Finally, the effect of the sex ratio was also examined on two “placebo” outcomes: the likelihood that a husband completed schooling past primary and the husband's height.²⁵ One would expect height in particular to have been affected by negative health shocks such as the Great Famine rather than the marriage market. Regressions on these outcomes indicate that they are not significantly influenced by sex ratios. Coefficient estimates on sex ratios are not statistically significant, and magnitudes are relatively low. For example, while a unit increase in the GVAR reduces the likelihood of completing schooling past primary levels by 1.2%, an increase of one centimeter in an individual's height raises the likelihood by nearly 6%. Similarly, OLS regressions in which male height is the dependent variable indicate there is no impact of the availability ratio. Coefficient estimates on the ratios amount to an average increase in height of 0.017% per unit increase in the GVAR and an increase of 0.019% per

²⁴Results are available upon request from the author.

²⁵Results are available upon request from the author.

unit increase in the GUAR. These estimates point towards evidence that the ratios capture marriage market conditions rather than exposure to negative shocks such as the famine. Though, it is also possible that as sex ratios and height are both influenced by nutrition in utero and in childhood, then they may also be correlated with one another. In addition, education could be partly explained by sex ratios (see for example, Heer and Grossbard-Shechtman 1981).

3.2 Survey Data

The empirical analysis uses household data from the China Health and Nutrition Survey (CHNS), collected by the Carolina Population Center.²⁶ Individuals in the CHNS are matched to the province-level data corresponding to their province of birth, so as not to confound identification with potentially endogenous individual decisions on migration. Since the CHNS only asked household heads—most of whom are male—about their province of birth, all regressions are from the husband's perspective.

The survey includes approximately 4,400 households in nine provinces. Among these households, there are 3,669 married men who were born after 1933. As census data is only available for those born after 1931, the earliest cohort here was born in 1934. Note that these early cohorts are only included in the sample for sex ratios that require data on fewer cohorts. Depending on the number of age types in the ratio, sex ratios are estimated for slightly different samples. Among these couples, 2,370 children under age 10 are identified and matched to their parents.

While this is a panel data set, it has been used here as a cross-sectional data set, taking the most recent observation available for each individual. Dummy variables for each survey year are included in all regressions. For married couples, nearly 70% of the observations are from the 2006 survey; 8% from each of the 2004 and 2000 waves; 7% from the 1997 wave; and the remainder from the 1989 and 1991 waves. For the child sample, 20% of the observations are from the 1991 survey; 35% from 1993; 20% from 1997; 13% from 2000; and around 6% from the 2004 and 2006 surveys respectively.

Summary statistics of regressors and dependent variables are presented in Table 2.

4. Results

4.1. How marriage market conditions affect children's health

I first examine how sex ratios influence children's health. Short-term nutritional health of young children is generally measured by weight for age z scores. In Table 3, coefficient estimates on the weighted availability ratio are statistically significant for the sample of boys alone. For boys, the coefficient estimate is statistically significant at the 5% level under both clustering methods, and across all specifications.

The magnitude of the effect is considerable. To estimate the effect of marriage market conditions on the population, and to take into consideration the differences in the variation

²⁶A detailed description of this data can be found at <http://www.cpc.unc.edu/projects/china>.

of these two ratios, coefficient estimates are multiplied by the standard deviation of each ratio. In regressions with the full set of controls, a standard deviation increase in the GVAR is equivalent to an increase in the z score of 0.09. Since the mean z score is -0.34 (Table 2), the average increase in the improvement in sons' short-term health is around 27%. The effect of the GVAR on boys' weight for age z scores is lower, with an overall magnitude of 17%, and the coefficient estimate is not statistically significant when additional controls are included. In regressions on the sample of girls alone, the coefficient on the GVAR is -0.004 and not statistically significant with any additional regressors. A standard deviation increase in either sex ratio implies a 14% average decline in girls' weight for age z scores.

Such differences between the sample of boys and girls may be explained by women having a stronger preference for sons. This may be because they expect to outlive their husbands and therefore rely on their children for help in old age.²⁷ Once a daughter marries, she moves to her husband's village and cares for his parents rather than her own, so that women must rely on their sons more in old age.

Other covariates do not significantly impact weight for age z scores, including competitive pressures from an individual's own or neighboring cohorts, the number of younger or older women available for marriage, average female education levels, or the degree of exposure to the famine. However, perhaps not surprisingly, fathers' health, as measured by his height, is a significant predictor of both boys' and girls' weight for age z scores.

Father's height is also a significant predictor of children's height for age z scores, a commonly used measure of children's long term health. However, availability ratios are not significant predictors of this measure of long term health (see Table 4). Coefficient estimates are not statistically significant once any additional regressors are included, for both boys and girls. Magnitudes are also relatively low, as an increase of one standard deviation in the GVAR raises boys' height for age z scores by 1.5% when fixed effects are included, and by 7.4% when the full set of controls are included. Magnitudes are higher using the unweighted ratio, though estimates are not statistically significant with wild clustering of standard errors. Estimated effects are lower for girls than for boys, as with weight for age z scores. Thus, the availability ratio influences short term health, but not long term health of children. Other factors, such as father's health play a more important role in predicting height for age than do sex ratios.

That higher sex ratios result in healthier sons (in terms of weight for age z scores) may be explained by two potential factors. They may be a result of negotiations at the time of the marriage contract, in which a potential wife can extract more of the marriage surplus because of an advantageous position in the marriage market. Alternatively, men who find themselves in a relatively more difficult marriage market may be more likely to settle for a "less desirable" match. To the extent that we can measure such characteristics with observables such as education, height, and age at marriage, we can examine whether the characteristics of the chosen spouse influence child health.

²⁷According to World Bank World Development Indicators, in China, male life expectancy at birth in 2004-2008 was 71 years of age, while female life expectancy at birth was 75 years of age in 2008.

Table 5 provides the results of regressions on boys' weight for age z scores, which incorporate the characteristics of both a child's father and mother. The first column in the table is the same as in Table 3. Previous regressions do not include the mother's characteristics because they are endogenous choices made by the father.

In controlling for parent characteristics (parental schooling, age at marriage, age and schooling differences between spouses, and mother's height), the GVAR faced by the father continues to have a positive and statistically significant effect on boys' weight for age z scores, and the magnitude is unchanged. This indicates that the effect of the GVAR on boys' health cannot be entirely explained by observed characteristics of the match made in the marriage market, and suggests that less favorable marriage market conditions for men give women a better position to negotiate for a greater share of the marriage surplus at the time of the initial marriage contract.

A priori, we might expect that a woman who is increasingly younger or less educated than her husband might play less of a role in household decisions (Beegle et al. 2001). The coefficient on the age difference between spouses is negative but not statistically significant and the magnitude is relatively small. However, the difference in education between spouses is correlated with weight for age z scores, and the coefficient estimate is statistically significant. As men are increasingly more educated than their wives, their sons are less healthy. This points towards evidence of perhaps more educated women having greater bargaining power vis a vis their husbands. Indeed, the coefficient on mother's education is statistically significant in predicting sons' health. Though, this may also be explained by the fact that more educated women have greater access to information via television, radio, and newspapers, as such access on the part of the mother has been found to improve children's health (Thomas et al. 1991). Among other parent characteristics, the heights of both parents are positive and statistically significantly related to boys' short-term health.

4.2 How marriage market conditions affect male consumption of tobacco and alcohol

Another way in which women might exert bargaining power once married is by encouraging husbands to smoke less or drink less alcohol. Until perhaps most recently, tobacco use has been generally confined to men, with mean daily tobacco consumption for women in the CHNS at zero. Both tobacco and alcohol are almost exclusively consumed by men. Regression results indicate men facing a higher sex ratio indeed smoke less and drink less alcohol (see Table 6). This may either be because women in a more favorable marriage market are able to exert their improved bargaining position to encourage their husbands to smoke less or drink less alcohol, or a result of men trying to make themselves more attractive mates in a competitive marriage market.

Regression results are fairly similar across the different specifications. Focusing on regressions with the complete set of controls, a unit increase in the GVAR or GUAR reduces alcohol consumption respectively by 0.008 or 0.005. With a mean alcohol consumption of 1.38 (Table 2), this implies men reduce average daily alcohol consumption by 0.6% (GVAR) or 0.4% (GUAR). To compare the overall magnitude of the effect of the ratios for the population, we multiply the coefficient estimate by the ratio's standard deviation. A standard deviation increase in the ratio implies an average population decrease

in alcohol consumption of about 7% using the GVAR and 4% using the GUAR. Not only is the magnitude of the effect considerably lower with the un-weighted ratio, estimates are only statistically significant using the weighted ratio, and only in regressions with the full set of control variables.

For tobacco consumption, both sex ratios similarly reduce male smoking, but only the GVAR results in statistically significant coefficients. A unit increase in the GVAR or GUAR reduces tobacco consumption by 0.003 units. Since the mean daily tobacco consumption is 0.07 (Table 2), this implies a 4% average reduction in tobacco consumption per unit increase in the sex ratio. A standard deviation increase in the GVAR or GUAR reduces average tobacco consumption for the population by 50%.

Finally, coefficient estimates on the degree of famine exposure are statistically significant when the dependent variable is tobacco consumption. Perhaps some men may be countering health challenges due to famine exposure.

4.3. How marriage market conditions affect age at marriage and mate selection

In addition to reducing tobacco or alcohol consumption, men in China might adjust to restrictive marriage market conditions by delaying age at marriage.²⁸ Tables 7 and 8 summarize results from OLS regressions which show that the age at which men marry is significantly influenced by the sex ratio. Interestingly for this outcome, coefficients do not differ between regressions with fixed effects and those with the full set of regressors. In the latter case, a standard deviation increase in the GVAR is equivalent to an average increase in the husband's age at marriage of roughly 5 months, and an increase in a woman's age at marriage of 3 months. Only the coefficient estimate on man's age at marriage is statistically significant at the 5% level under both clustering methods. Similarly, a standard deviation increase in the GUAR raises male age at marriage by about 5 months, and female age at marriage by about 1.5 months.

The positive effect of the sex ratio on male age at marriage may be explained by the fact that men who have the potential to become financially successful wait to marry until they can attract a wife otherwise more difficult to attain (Bergstrom and Bagnoli 1993). This is not to say that men who face higher sex ratios would become more financially successful as a result. Rather, since men earn more with age, those who face more competition in the marriage market may wait until they are older to marry, so that potential wives can then observe their success. This allows them to compete with younger men earning less.

Though the ratio is also slightly positively related to the age at which women marry and the age difference between husband and wife, magnitudes are significantly lower compared to male age at marriage. Differences across specifications result in the greatest differences in coefficient estimates of ratios when the outcome is female age at marriage. Ratios do not predict female age at marriage at all, once birth year and province fixed effects, hukou, and related interaction terms are included in the regression. Thus, female age at marriage may be

²⁸Marriage has been virtually universal for these cohorts (Zeng 1995). According to the 1982 census, by the time men reach the age of 46-50, 96% of them have been married, and 99% of women in this age group have married. While nearly all eventually marry, the interest here is at what age they marry and who they marry.

primarily determined by variation across regions and cohorts not specific to variation in availability ratios. In determining women's age at marriage, other factors may be more important than ratios, including cultural factors, and education and labor market opportunities. Cohort-specific factors may be particularly important here, as for example in the case of women born during the Great Famine. While these women are relatively scarce in the market, and may be in greater demand because of preferred ages at marriage and greater labor market opportunities, they may also be less in demand because women born during this period experienced adverse health and education outcomes (Meng and Qian 2006; Almond et al. 2007). Thus, in controlling for such factors, variation in women's age at marriage can be explained by variation across provinces and cohorts, as well as possibly by variation in cohort-level education averages, and cohort-level measures of labor market competition.²⁹

Perhaps not surprisingly, a man's height, often an attractive feature to women and an indicator of long-term health, has the opposite impact on age at marriage to high sex ratios. Taller men marry earlier and their wives are both closer to them in age and less likely to be more educated than them. Relatedly, those who had been more exposed to the famine while in-utero marry women who are considerably younger than themselves. As men exposed to the famine may be less attractive in the marriage market (Brandt et al. 2008, Almond et al. 2010), they may need to marry women increasingly younger than themselves.

Men might also adjust to competitive marriage market conditions by marrying more or less educated women. Men may prefer to marry women less educated than them, because more educated women have more employment opportunities and therefore can demand more of the marital surplus (Grossbard-Shechtman 1993; Grossbard and Amuedo-Dorantes 2007). For a more educated woman, the outside option of remaining unmarried may be preferable to marrying someone who may reduce her independence. For marriage to become an attractive option for a more educated woman, she might demand more bargaining power. Thus, men with more favorable marriage market conditions may choose to marry less educated women if given the option, whilst higher sex ratios faced by men would drive them towards marrying more educated women. Educated women in China have consistently had lower marriage rates than less educated women (Lavelly et al. 1990). In fact, recent media accounts have reported that women who attained a level of higher education are finding it difficult to find a spouse in China (Lee and Subramanian 2011, Ou 2011, *The Economist* 2011). Indeed, men who face more competitive marriage market conditions are slightly more likely to marry women who have attained a higher level of education than they have, and this is robust across different specifications (see Tables 7 and 8). From the results summarized in Table 8, a standard deviation increase in the GVAR raises the likelihood of a husband marrying a more educated wife by around 20%, whereas that of the GUAR raises the same probability by nearly 13%. Though, this coefficient estimate is not statistically significant under wild clustering.

²⁹A marriage matching model which incorporates these different matching characteristics and constraints may help to distinguish between these different dynamics (e.g., see Brandt et al. 2008).

4.4. Sensitivity analysis

In the sensitivity analysis below, results are qualitatively similar to those outlined above, but in comparing several different ratios, the resulting magnitudes are often sensitive to the assumptions used in estimating marriage market conditions (see Table 9). Each coefficient in the table is from an individual regression. For each dependent variable, the first rows are the coefficient estimates and standard errors on the GVAR outlined in the previous tables, and the second set of rows are for the GUAR also shown previously. GVAR1 is based on the assumption that men marry women who are anywhere from three years older to nine years younger than them. These results are then compared to that of GVAR2 (similar to GVAR1 but based on the assumption that men marry women who are between two years older and six years younger). GVAR3 is based on the assumption that men marry women who are anywhere from zero to five years younger. In order to be comparable with GVAR1, GUAR1 is based on 15 respective cohorts of men and women, with the median male cohort being a given individual's cohort, and the median female cohort being two years younger. GUAR2 is based on nine contiguous cohorts, enabling a comparison with GVAR2, while GUAR3 is based on five contiguous cohorts, enabling comparison with GVAR3.

Since ratios are available for slightly different sets of birth cohorts, sample sizes differ. A potential concern would be that differences in the estimates result from different samples rather than differences in the ratios. To test for this, regressions are also estimated using the smallest sample: those birth cohorts that are included in every regression. Results are unchanged from those shown in Table 9.

In regressions on boys' weight for age z scores, all sex ratios increase this outcome measure, and coefficient estimates are statistically significant for all GVARS and for GUAR3. Overall magnitudes (estimated by a standard deviation increase in the ratio) are slightly more varied across the different GUARs than across the GVARS. For the weighted availability ratio, the effect ranges from 17% to 27%. The un-weighted availability ratio results in an overall increase of 17% for GUAR2 to 30% for GUAR3.

For tobacco consumption, coefficient estimates are more similar, varying between -0.002 and -0.003. An increase of one standard deviation in GVAR2 would yield a slightly greater estimate since there is more variation in this ratio. While the average percent decrease in tobacco consumption for the population is 50% for GVAR1, it is 58% for GVAR2. Similar differences arise for the other ratio estimates.

When the dependent variable is the amount of alcohol consumed by the husband, coefficient estimates across the three different GVARS differ. Only coefficient estimates on GVAR1 are statistically significant, at the 10% level under the wild clustering method. Other coefficient estimates are considerably lower, where all but GUAR1 are zero.

When the dependent variable is the age at which men marry, effects are strongest for the ratios based on the largest number of cohorts (GVAR1 and GUAR1). When a tighter age interval is used to estimate the GUAR as in GUAR2 and GUAR3, magnitudes are somewhat lower. GVAR3 also yields lower estimates.

To confirm the importance of incorporating competition from other cohorts, additional regressions are estimated in which the sex ratio is the number of men born in one's cohort divided by the denominator in equation (1). Estimates are close to zero and standard errors are high. Finally, regressions are estimated to test whether the rural GVAR affects the urban sample, and vice versa; coefficients are considerably low, while standard errors are high.³⁰

5. Conclusion

We have seen that when fertility fluctuates as much as it has in China over the past several decades, consideration of a population's general preferences regarding spouse age is particularly important in estimating the degree of scarcity of women relative to men. The GVAR developed in this article approximates the sex ratio that applies to each individual at marriage, and reflects the availability of potential spouses among preferred cohorts. Such an estimation of the sex ratio improves upon those used in prior studies because the number of available potential spouses reflects the preferences an individual has for potential spouses from various age groups. It is an enhancement on the availability ratio, because it measures potential marriage possibilities based solely on one's circumstances of birth and the overall distribution in spousal age differences observed in the population as a whole. Such methods may be useful in future research on marriage markets, as they point to the importance of considering alternative measures of the marriage market to understand how sensitive one's findings are to ratio construction.

This article provides an important contribution to this literature, indicating further evidence of commitment to a sharing rule determined at the time of marriage. Historical fertility fluctuations in China provide a unique context for identifying such effects. Nonetheless, conditions at the time of the initial marriage contract may become less important as possibilities for renegotiation become more readily available, with divorce becoming an increasingly credible outside option. Future work is required to determine the degree to which conditions at the time of entering a marriage contract matter relative to such possibilities for renegotiation.

In recent years, as population growth in China has declined and sex ratios at birth have risen, we can presume that marriage market conditions for Chinese women have been improving significantly. Nonetheless, this imbalance may be mitigated by a rise in the immigration of women from neighboring countries, who may serve as substitutes for Chinese women to the extent that men are willing to marry them. *Ceteris paribus*, results from this article indicate that today's imbalanced sex ratio of 120, relative to the usual average ratio of 107, increases boys' short-term health by roughly 31%, raises the age at which men marry by nearly six months, and reduces men's daily tobacco consumption by 56%. It is important to note, however, that while immigration of women may dampen these effects, declining population growth further augments these effects to the extent that women continue to marry older men. In this paper, we have seen that one consequence of such a demographic shift is greater resources for women.

³⁰Results are available upon request from the author.

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Figure 1. Fertility and Generalized Weighted Availability Ratios in Rural China

Figure 2. Fertility and Generalized Unweighted Availability Ratios in Rural China

Figure 3. Fertility and Generalized Weighted Availability Ratios in Urban China

Figure 4. Fertility and Generalized Unweighted Availability Ratios in Urban China

Table 1

Distribution in age differences between husband and wife

Age gap	No. Obs.	Men born 1932-64			No. Obs.	Men born 1932-37		
		% where $i \in [-3, 9]$	% where $i \in [-2, 6]$	% where $i \in [0, 5]$		% where $i \in [-3, 9]$	% where $i \in [-2, 6]$	% where $i \in [0, 5]$
		Rural						
-3 Years	3,946	2.92		613	3.68			
-2 Years	6,950	5.15	5.90	925	5.55	6.62		
-1 Years	11,082	8.20	9.39	1,278	7.67	9.15		
0 Years	16,240	12.03	13.78	17.66	1,794	10.76	12.84	16.93
1 Year	17,995	13.33	15.26	19.56	1,865	11.19	13.35	17.60
2 Years	17,804	13.19	15.10	19.35	1,962	11.77	14.04	18.52
3 Years	16,093	11.92	13.65	17.49	1,823	10.94	13.05	17.21
4 Years	13,295	9.85	11.28	14.45	1,669	10.01	11.94	15.75
5 Years	10,570	7.83	8.97	11.49	1,481	8.88	10.60	13.98
6 Years	7,861	5.82	6.67		1,177	7.06	8.42	
7 Years	5,764	4.27			905	5.43		
8 Years	4,297	3.18			670	4.02		
9 Years	3,119	2.31			509	3.05		
Total	135,016	100.00	100.00	100.00	16,671	100.00	100.00	100.00
Urban								
-3 Years	540	2.07			94	2.33		
-2 Years	983	3.77	4.33		122	3.02	3.60	
-1 Years	1,851	7.10	8.16		250	6.20	7.38	
0 Years	3,018	11.58	13.30	16.59	346	8.57	10.21	12.85
1 Year	3,530	13.54	15.56	19.40	451	11.18	13.31	16.75
2 Years	3,426	13.14	15.10	18.83	491	12.17	14.49	18.24
3 Years	3,148	12.08	13.88	17.30	516	12.79	15.23	19.17
4 Years	2,794	10.72	12.32	15.36	476	11.80	14.05	17.68
5 Years	2,276	8.73	10.03	12.51	412	10.21	12.16	15.30
6 Years	1,658	6.36	7.31		325	8.05	9.59	

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Age gap	No. Obs.	Men born 1932-64			Men born 1932-37		
		% where $i \in [-3,9]$	% where $i \in [-2,6]$	% where $i \in [0,5]$	No. Obs.	% where $i \in [-3,9]$	% where $i \in [0,5]$
7 Years	1,240	4.76			266	6.59	
8 Years	895	3.43			170	4.21	
9 Years	708	2.72			116	2.87	
Total	26,067	100.00	100.00	100.00	4,035	100.00	100.00

Notes: Estimates are based on 1982 Chinese census and author's calculations. Only couples in their first marriages were included.

Summary statistics

Table 2

Variable	No.Obs.	Mean	Std.Dev.	Min	Max
Independent Variables:					
Generalized Weighted Availability Ratio 1 e [-3,9]	2910	101.49	11.67	79.34	164.88
Generalized Weighted Availability Ratio 1 e [-2,6]	3367	104.53	13.63	79.34	176.11
Generalized Weighted Availability Ratio 1 e [0,5]	3518	105.04	15.88	75.02	183.54
Generalized Un-weighted Availability Ratio, 15 cohorts	3269	104.59	11.61	84.40	164.46
Generalized Un-weighted Availability Ratio, 9 cohorts	3537	105.42	14.13	76.08	176.36
Generalized Un-weighted Availability Ratio, 5 cohorts	3669	106.55	17.13	70.51	187.25
1% of number men born in husband's cohort	3668	215.78	167.53	13.00	837.00
Numerator of GWARD (competing men)	2910	221.65	156.31	25.97	724.63
Numerator of GUARD (competing men)	3269	3267.6	2349.12	351.00	10920
Average female education (from 2000census)	2855	3.83	0.50	3.08	4.99
Number of women from younger cohorts	2910	255.74	198.38	17.00	891.00
Number of women from older cohorts	2910	187.08	144.86	12.00	774.00
Husband's famine exposure while in utero	3654	1.26	4.11	0.00	44.73
Husband's height (cm)	3182	166.85	6.35	148.00	189.00
Husband's year of birth	3669	1954.0	10.29	1934.00	1977.00
Husband's hukou = 1 if urban, =2 if rural	3669	1.56	0.50	1.00	2.00
CHNS survey year	3669	2003.5	4.42	1991.00	2006.00
Age of child (ages 10 and under)	2370	7.40	1.99	1	10
Survey year of child (ages 10 and under)	2370	1996.0	4.72	1991	2006
Mother's height (boys ages 10 and under)	1223	155.90	5.64	141	174
Father's education (boys ages 10 and under)	1253	1.95	1.07	0	5
Mother's education (boys ages 10 and under)	1260	1.56	1.15	0	6
Parent education gap (boys ages 10 and under)	1233	0.40	1.07	-3	4
Dependent Variables:					
Husband's age at marriage	2223	24.63	4.07	15	55
Wife's age at marriage	2213	22.86	3.28	15	50
Age difference between spouses	3662	1.82	2.84	-5	15

Variable	No.Obs.	Mean	Std.Dev.	Min	Max
Husband's tobacco smoked per day	3669	0.07	0.87	0	20
Husband's alcohol consumption per day	3669	1.38	1.66	0	9
Husband's schooling was past primary = 1	3415	0.66	0.47	0	1
Wife completed secondary school = 1	3539	0.22	0.42	0	1
Weight for age z-scores (boys age 10 & under)	1285	-0.34	0.48	-2.72	3.76
Weight for age z-scores (girls age 10 & under)	1085	-0.34	0.41	-1.71	2.47
Height for age z-scores (boys age 10 & under)	1285	-0.79	0.72	-4.41	1.85
Height for age z-scores (girls age 10 & under)	1083	-0.76	0.69	-4.21	1.51

Notes: Married couples in China Health and Nutrition Survey. Ratios are based on the author's calculations, using the 1982 Chinese census unless otherwise noted.

Table 3
Availability ratios as determinants of children's short term health

	Boys' weight for age z-scores						Girls' weight for age z-scores					
	GWAR ^a	GWAR	GWAR	GUAR ^a	GUAR	GUAR	GWAR ^a	GWAR	GWAR	GUAR ^a	GUAR	GUAR
Availability Ratio	0.007 (0.003)	0.006 (0.002)	0.008 (0.003)	0.008 (0.003)	0.006 (0.003)	0.005 (0.004)	0.006 (0.002)	-0.004 (0.003)	0.008 (0.003)	-0.001 (0.003)	-0.004 (0.004)	
Robust P	0.003***	0.001***	0.015**	0.012**	0.063*	0.207	0.017**	0.898	0.018**	0.794	0.333	
Wild P	0.030**	0.044**	0.032**	0.032**	0.110	0.240	0.062*	0.920	0.074*	0.858	0.378	
Male Birth			-0.001**			0.000		0.000			0.000	
Cohort Size			(0.000)			(0.000)		(0.000)			(0.000)	
Competing			0.001*			0.000		-0.001			0.000	
Cohorts Size			(0.001)			(0.000)		(0.001)			(0.000)	
Avg. Female Education			-0.243			0.243		0.517**			0.262	
			(0.288)			(0.302)		(0.192)			(0.344)	
Num			0.000			0.000		0.000			0.000	
Younger Women			(0.000)			(0.000)		(0.000)			(0.000)	
Num Older Women			0.000			0.000		0.000			0.000	
			(0.000)			(0.000)		(0.000)			(0.000)	
Famine Exposure			-0.007			-0.006		0.002			0.002	
			(0.007)			(0.007)		(0.008)			(0.008)	
Husband's Height			0.012***			0.009***		0.011***			0.013***	
			(0.002)			(0.003)		(0.002)			(0.003)	
Child Age	-0.022**	-0.016	-0.021	-0.018*	-0.012	-0.013	-0.029***	-0.025**	-0.028***	-0.019**	-0.020**	
	(0.011)	(0.011)	(0.014)	(0.009)	(0.010)	(0.011)	(0.008)	(0.009)	(0.007)	(0.007)	(0.008)	
Num Obs	1,138	1,138	956	1,236	1,236	1,036	968	806	1,047	1,047	872	
R-squared	0.0383	0.2463	0.2962	0.0322	0.2431	0.2795	0.0413	0.2986	0.0477	0.2568	0.3024	

 Notes: p<0.01,

**
 p<0.05,

*
 p<0.1.

OLS regressions. Robust standard errors in parentheses, clustered by father's hukou and birth province. Sample is restricted to children ages 10 and under. Ratios are based on the author's calculations using the 1982 Chinese census. Additional controls include: father's hukou, father's birth year and province fixed effects, interaction terms between hukou status and birth province fixed effects, and CHNS survey year dummy variables.

^aNo additional controls in these regressions.

Table 4
Availability ratios as determinants of children's long term health

	Boys' height for age z-scores						Girls' height for age z-scores					
	GWAR ^a	GWAR	GWAR	GUAR ^a	GUAR	GUAR	GWAR ^a	GWAR	GWAR	GUAR ^a	GUAR	GUAR
Availability Ratio	0.010 (0.004)	0.001 (0.003)	0.005 (0.004)	0.011 (0.005)	0.005 (0.005)	0.010 (0.005)	0.010 (0.002)	0.003 (0.006)	0.001 (0.006)	0.013 (0.004)	0.006 (0.009)	0.002 (0.009)
Robust P	0.018**	0.816	0.204	0.036**	0.343	0.036**	0.000***	0.574	0.933	0.002***	0.502	0.806
Wild P	0.050**	0.858	0.482	0.056*	0.410	0.132	0.000***	0.674	0.944	0.012**	0.570	0.790
Male Birth Cohort Size			-0.001 (0.001)			0.000 (0.001)			0.000 (0.001)			0.000 (0.000)
Competing Cohorts Size			0.001 (0.001)			0.000 (0.000)			0.001 (0.006)			0.000 (0.000)
Avg. Female Education			-0.375 (0.381)			-0.564 (0.359)			0.035 (0.453)			-0.195 (0.462)
Num Younger Women			0.000 (0.000)			0.000 (0.000)			0.000 (0.001)			0.000 (0.001)
Num Older Women			0.000 (0.001)			0.000 (0.001)			0.000 (0.000)			0.000 (0.001)
Famine Exposure			-0.005 (0.011)			-0.005 (0.011)			-0.017* (0.008)			-0.017* (0.008)
Husband's Height			0.031*** (0.003)			0.031*** (0.004)			0.033*** (0.004)			0.032*** (0.004)
Age of Child	-0.033** (0.015)	-0.018 (0.014)	-0.020 (0.013)	-0.034** (0.014)	-0.024 (0.014)	-0.017 (0.013)	-0.035*** (0.010)	-0.030** (0.011)	-0.030*** (0.009)	-0.028** (0.010)	-0.025** (0.012)	-0.024** (0.011)
Num Observations	1,138	1,138	956	1,236	1,236	1,036	966	966	804	1,045	1,047	870
R-squared	0.0267	0.2189	0.2983	0.0289	0.2202	0.2968	0.0300	0.2462	0.3266	0.0322	0.2568	0.3266

Notes: p<0.01,

**
p<0.05,

*
p<0.1.

OLS regressions. Robust standard errors in parentheses, clustered by father's hukou and birth province. Sample is restricted to children ages 10 and under. Ratios are based on the author's calculations using the 1982 Chinese census. Additional controls include: father's hukou, father's birth year and province fixed effects, interaction terms between hukou status and birth province fixed effects, and CHNS survey year dummy variables.

^aNo additional controls in these regressions.

Table 5
Availability ratio effects on boys' health remain while controlling for parent characteristics

	Boys' weight for age z-scores					
GWAR	0.008 (0.003)	0.008 (0.003)	0.009 (0.003)	0.009 (0.003)	0.008 (0.003)	0.008 (0.003)
Robust P	0.015** (0.003)	0.027** (0.003)	0.008*** (0.003)	0.013** (0.003)	0.030** (0.003)	0.030** (0.003)
Wild P	0.032** (0.003)	0.07* (0.003)	0.034** (0.003)	0.038** (0.003)	0.064* (0.003)	0.064* (0.003)
Male Birth Cohort Size	-0.001** (0.000)	-0.001** (0.000)	-0.001** (0.000)	-0.001** (0.000)	-0.001** (0.000)	-0.001** (0.000)
Size of Competing Cohorts	0.001* (0.001)	0.002** (0.001)	0.001** (0.001)	0.001* (0.001)	0.001** (0.001)	0.001** (0.001)
Avg. Female Education	-0.243 (0.288)	-0.295 (0.313)	-0.382 (0.268)	-0.351 (0.261)	-0.221 (0.349)	-0.221 (0.349)
Number Younger Women	0.000 (0.000)	0.000 (0.000)	0.000 (0.000)	0.000 (0.000)	0.000 (0.000)	0.000 (0.000)
Number Older Women	0.000 (0.000)	0.000 (0.000)	0.000 (0.000)	0.000 (0.000)	0.000 (0.000)	0.000 (0.000)
Degree of Famine Exposure	-0.007 (0.007)	-0.012 (0.008)	-0.007 (0.007)	-0.007 (0.007)	-0.01 (0.008)	-0.01 (0.008)
Husband's Height	0.012*** (0.002)	0.009*** (0.002)	0.011*** (0.002)	0.011*** (0.002)	0.011*** (0.002)	0.011*** (0.002)
Mother's Height	0.009*** (0.003)	0.009*** (0.003)	0.009*** (0.003)	0.009*** (0.003)	0.009*** (0.003)	0.009*** (0.003)
Father's Education	-0.003 (0.012)	-0.003 (0.012)	-0.002 (0.011)	-0.002 (0.011)	-0.002 (0.011)	-0.002 (0.011)
Mother's Education	0.019 (0.013)	0.019 (0.013)	0.031** (0.011)	0.031** (0.011)	0.031** (0.011)	0.031** (0.011)
Father's Age at Marriage	-0.005 (0.007)	-0.005 (0.007)	-0.005 (0.007)	-0.005 (0.007)	-0.005 (0.007)	-0.005 (0.007)
Mother's Age at Marriage	0.010 (0.006)	0.010 (0.006)	0.010 (0.006)	0.010 (0.006)	0.010 (0.006)	0.010 (0.006)

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Boys' weight for age z-scores					
Parent Age Gap					
		-0.008			
		(0.005)			
Parent Education Gap		-0.016 [*]			
		(0.008)			
Observations	956	853	935	939	896
R-squared	0.2962	0.3061	0.2968	0.2977	0.3001

*** Notes: p<0.01,
** p<0.05,
* p<0.1.

OLS regressions. Robust standard errors in parentheses, clustered by father's hukou and birth province. Sample is restricted to boys ages 10 and under. Ratios are based on the author's calculations using the 1982 Chinese census. Additional controls include: father's hukou, father's birth year and province fixed effects, interaction terms between hukou status and birth province fixed effects, and CHNS survey year dummy variables.

Availability ratios as determinants of male consumption of alcohol and tobacco

Table 6

	Amount of Alcohol Consumed by Husband				Amount of Tobacco Consumed by Husband			
	GWAR ^a	GWAR	GUAR ^a	GUAR	GWAR ^a	GWAR	GUAR ^a	GUAR
Availability Ratio	-0.002 (0.003)	-0.006 (0.004)	-0.001 (0.003)	-0.005 (0.004)	-0.005 (0.004)	-0.002 (0.001)	-0.009 (0.001)	-0.003 (0.002)
Robust P		0.137 0.032 **		0.191 0.243		0.024 ** 0.043 **		0.055 * 0.077 *
Wild P		0.212 0.088 *		0.208 0.324		0.132 0.090 *		0.128 0.140
Male Birth Cohort Size		0.001 (0.001)		0.001 (0.001)		0.000 (0.000)		0.000 (0.000)
		(0.002)		(0.000)		(0.000)		(0.000)
Avg. Female Education		-1.10 *** (0.315)		-0.412 (0.505)		0.306 (0.298)		0.347 (0.210)
Number Younger Women		-0.001 * (0.000)		-0.001 * (0.0000)		0.000 (0.000)		0.000 (0.000)
Number Older Women		0.001 (0.001)		0.000 (0.001)		0.000 * (0.000)		0.000 (0.000)
Degree of Famine Exposure		0.030 (0.024)		0.033 (0.022)		-0.004 ** (0.002)		-0.004 * (0.002)
Husband's Height		0.001 (0.007)		0.003 (0.006)		0.000 (0.002)		0.001 (0.001)
Number of Observations	2,910	2,910	3,269	3,269	2,744	2,910	3,269	2,744
R-squared	0.0002	0.0289	0.0292	0.0307	0.0294	0.0460	0.0004	0.0530

Notes: p<0.01,

**
p<0.05,

*
p<0.1.

OLS regressions. Robust standard errors in parentheses, clustered by husband's hukou and birth province. Ratios are based on the author's calculations using the 1982 Chinese census. Sample is restricted to married couples in the CHNS. Additional controls include: husband's hukou, husband's birth year and province fixed effects, interaction terms between hukou status and birth province fixed effects, and CHNS survey year dummy variables.

^aNo additional controls in these regressions.

Table 7
Availability ratios as primary determinants of marriage-related behavior

	Age at Marriage						Husband Marries a More Educated Wife ^a					
	Male			Female			Difference					
	GWAR	GUAR	GWAR	GUAR	GWAR	GUAR	GWAR	GUAR	GWAR	GUAR	GWAR	GUAR
Availability Ratio	0.082 *** (0.026)	0.090 *** (0.022)	0.044 *** (0.014)	0.043 *** (0.011)	0.037 *** (0.009)	0.039 *** (0.007)	0.015 *** (0.005)	0.011 * (0.006)	0.015 *** (0.005)	0.011 * (0.006)	0.015 *** (0.005)	0.011 * (0.006)
Number of Observations	1,967	2,129	1,959	2,119	2,904	3,262	2,611	2,938	2,611	2,938	2,611	2,938
R-squared	0.0443	0.0481	0.0200	0.0168	0.0242	0.0269						
Availability Ratio ^b	0.035	0.036	0.005	0.003	0.016	0.011	0.012	0.004	0.012	0.004	0.012	0.004
	(0.013)	(0.013)	(0.013)	(0.014)	(0.004)	(0.006)	(0.009)	(0.013)	(0.009)	(0.013)	(0.009)	(0.013)
Robust P	0.007 ***	0.006 ***	0.704	0.836	0.000 ***	0.053 *	0.161	0.779	0.161	0.779	0.161	0.779
Wild P	0.002 ***	0.034 **	0.690	0.814	0.008 ***	0.114	0.234	0.812	0.234	0.812	0.234	0.812
Number of Observations	1,967	2,129	1,959	2,119	2,904	3,262	2,594	2,918	2,594	2,918	2,594	2,918
R-squared	0.2164	0.2260	0.1004	0.1083	0.1082	0.1119						

*** Notes: p<0.01,

** p<0.05,

* p<0.1.

Robust standard errors in parentheses, clustered by husband's hukou and birth province. Sample is restricted to married couples in the CHNS. Ratios are based on the author's calculations using the 1982 Chinese census.

^a Logit regressions, otherwise OLS regressions.

^b Additional controls include: husband's hukou, husband's birth year and province fixed effects, interaction terms between hukou status and birth province fixed effects, and CHNS survey year dummy variables.

Table 8
Availability ratios and other determinants of marriage-related behavior

	Age at Marriage						Husband Marries a More Educated Wife ^d		
	Male			Female			Difference		
	GUAR	GUAR	GUAR	GUAR	GUAR	GUAR	GUAR	GUAR	GUAR
Availability Ratio	0.037 (0.017)	0.036 (0.016)	0.023 (0.013)	0.012 (0.014)	0.015 (0.007)	0.013 (0.006)	0.014 (0.008)	0.01 (0.015)	
Robust P	0.042**	0.039**	0.100	0.387	0.041**	0.064*	0.083*	0.514	
Wild P	0.036**	0.054*	0.152	0.386	0.168	0.230	0.202	0.628	
Male Birth Cohort Size	0.001 (0.004)	0.001 (0.003)	0.002 (0.003)	0.001 (0.003)	0.002 (0.001)	0.001 (0.001)	-0.001 (0.002)	0.000 (0.002)	
Size of Competing Cohorts	-0.002 (0.005)	0.000 (0.000)	-0.001 (0.005)	0.000 (0.000)	-0.001 (0.003)	0.000 (0.000)	-0.002 (0.003)	0.000 (0.000)	
Avg. Female Education	1.671 (1.623)	0.796 (1.790)	-1.018 (1.285)	-0.462 (1.526)	1.065 (0.986)	0.599 (1.166)	-2.110***	-0.999 (1.315)	
Number Younger Women	0.001 (0.002)	0.001 (0.002)	0.003 (0.002)	0.002 (0.002)	0.000 (0.001)	0.000 (0.001)	0.000 (0.001)	0.000 (0.001)	
Number Older Women	0.000 (0.002)	0.002 (0.002)	-0.001 (0.002)	0.002 (0.002)	-0.001 (0.001)	0.000 (0.001)	0.001 (0.001)	0.002 (0.002)	
Degree of Famine Exposure	0.063 (0.053)	0.049 (0.051)	0.029 (0.060)	0.021 (0.059)	0.049** (0.017)	0.038** (0.015)	0.007 (0.044)	0.012 (0.040)	
Husband's Height	-0.034*	-0.031*	0.008	0.008	-0.049***	-0.050***	-0.021*	-0.020**	
Number of Observations	1,608	1,741	1,602	1,733	2,429	2,737	2,338	2,638	
R-squared	0.2246	0.2354	0.1023	0.1100	0.1203	0.1235			

 Notes: p<0.01,

**
 p<0.05,

*
 p<0.1.

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Robust standard errors in parentheses, clustered by husband's hukou and birth province. Sample is restricted to married couples in the CHNS. Additional controls include: husband's hukou, husband's birth year and province fixed effects, interaction terms between hukou status and birth province fixed effects, and CHNS survey year dummy variables. Ratios are based on the author's calculations using the 1982 Chinese census.

^a Logit regressions, otherwise OLS regressions.

Table 9
Comparing empirical results across different availability ratios

	Boys' weight for age z-score ^a	Tobacco Husband Smokes	Alcohol Husband Consumes	Husband's age at marriage
Generalized Weighted Availability Ratio 1 $\in [-3, 9]$ (GWAR1)				
Coeff.	0.008	-0.003	-0.008	0.037
Std. Err.	(0.003)	(0.001)	(0.004)	(0.017)
Robust P	0.015 **	0.043 **	0.032 **	0.042 **
Wild P	0.032 **	0.090 *	0.088	0.036 **
Generalized Un-weighted Availability Ratio 1, 15 cohorts (GUAR1)				
Coeff.	0.005	-0.003	-0.005	0.036
Std. Err.	(0.004)	(0.002)	(1.168)	(0.016)
Robust P	0.207	0.077 *	0.243	0.039 **
Wild P	0.240	0.140	0.324	0.054 *
Generalized Weighted Availability Ratio 2 $\in [-2, 6]$ (GWAR2)				
Coeff.	0.006	-0.003	0.000	0.028
Std. Err.	(0.003)	(0.001)	(0.046)	(0.013)
Robust P	0.035 **	0.078 *	0.963	0.043 **
Wild P	0.064 *	0.098 *	0.984	0.104
Generalized Un-weighted Availability Ratio 2, 9 cohorts (GUAR2)				
Coeff.	0.004	-0.003	-0.001	0.012
Std. Err.	(0.003)	(0.002)	(0.265)	(0.016)
Robust P	0.183	0.174	0.791	0.474
Wild P	0.230	0.182	0.854	0.576
Generalized Weighted Availability Ratio 3 $\in [0, 5]$ (GWAR3)				
Coeff.	0.007	-0.002	-0.001	0.020
Std. Err.	(0.003)	(0.001)	(0.286)	(0.009)
Robust P	0.024 **	0.059 *	0.775	0.035 **
Wild P	0.024 **	0.058 *	0.808	0.108
Generalized Un-weighted Availability Ratio 3, 5 cohorts (GUAR3)				
Coeff.	0.006	-0.002	0.000	0.019
Std. Err.	(0.003)	(0.002)	(0.102)	(0.009)
Robust P	0.030 **	0.151	0.918	0.052 *
Wild P	0.008 ***	0.164	0.948	0.120

 Notes: $p < 0.01$,

**
 $p < 0.05$,

*
 $p < 0.1$.

^a Sample is restricted to boys ages 10 and under. Robust standard errors in parentheses, clustered by husband's hukou and birth province. Ratios are based on the author's calculations using the 1982 Chinese census. Sample is restricted to married couples in the CHNS. Additional controls include:

husband's hukou, husband's birth year and province fixed effects, interaction terms between hukou status and birth province fixed effects, and CHNS survey year dummy variables.

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