The Competitive Saving Motive: Evidence from Rising Sex Ratios and Savings Rates in China

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The high and rising household savings rate in China is not easily reconciled with the traditional explanations that emphasize life cycle factors, the precautionary saving motive, financial development, or habit formation. This paper proposes a new competitive saving motive: as the sex ratio rises, Chinese parents with a son raise their savings in a competitive manner in order to improve their son's relative attractiveness for marriage. The pressure on savings spills over to other households. Both cross-regional and household-level evidence supports this hypothesis. This factor can potentially account for about half the actual increase in the household savings rate during 1990–2007.

I. Introduction

High savings rates in certain countries are said to be a major contributor to the recent housing price bubbles and the global financial crisis by depressing global long-term interest rates in the last decade (Greenspan

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2009). Because the Chinese household savings as a share of disposal income nearly doubled from 16 percent in 1990 to 30 percent in 2007, it has caught special attention.

The purpose of this paper is to test a new hypothesis regarding household savings behavior using Chinese household and regional data. Before doing that, we note that there is no shortage of theories of savings behavior in the literature. First, the life-cycle theory (Modigliani 1970; Modigliani and Cao 2004) predicts that the savings rate rises with the share of working age population in the total population. This explanation does not appear to be consistent with the profile of savings at the household level (Chamon and Prasad 2010). The second explanation has to do with a precautionary savings motive in combination with a rise in income uncertainty, which is favored by Blanchard and Giavazzi (2005) and Chamon and Prasad (2010). The problem is that while both pension systems and the public provision of health care in China have been improving since 2003, household savings as a share of disposable income continued to rise sharply during the same period. This time series pattern contradicts the precautionary motive theory. The third explanation is a low level of financial development. This has the same difficulty as the last explanation since the financial system was most likely more efficient today than a few years ago, yet the savings rate still rose. The fourth explanation is cultural norms. But cultural norms tend to be persistent and therefore are unlikely to be capable of explaining the visible rise in the savings rate over the last 2 decades.[1]

In this paper, we suggest that an alternative savings motive may be at work: people save in order to improve their relative standing in the marriage market. When the sex ratio (the number of men per woman in the premarital cohort) rises, families with sons compete with each other to raise their savings rate in response to ever-rising pressure in

[1] We do not study corporate savings behavior in this paper. The Chinese corporate savings rate has risen sharply in recent years, accounting now for about half the national savings rate (Kuijs 2006). Corporate savings behavior is a separate puzzle to be explained. The existing explanation suggests that a combination of state ownership and windfalls in resource sectors is the primary driver. We note, however, that a recent paper by Bayoumi, Tong, and Wei (2010) discusses China's corporate savings rate in a global context and casts doubt on the usual interpretation. Because corporate savings rates have been rising globally (International Monetary Fund 2005; Bates, Kahle, and Stulz 2009), it turns out that the Chinese corporate savings rate is only modestly higher than those of other countries (by 2 percentage points above the global average savings rate when comparing listed companies across countries). So differences in corporate savings are unlikely to be a big part of cross-country differences in national savings rates. Moreover, comparing corporate savings by Chinese firms of different ownership, in resource and nonresource sectors, Bayoumi et al. (2010) do not find evidence that state-owned firms save more than private firms or that the corporate savings rate is unusually high in resource sectors. This casts doubt on the notion that a high Chinese corporate savings rate is mainly a result of poor corporate governance and inefficiencies tied either to state ownership or to windfalls in resource sectors.

the marriage market. Families with daughters may not decrease their savings due to the existence of two offsetting motives. On the one hand, they are tempted to reduce their savings in order to take advantage of the higher savings rates of their future sons-in-law. On the other hand, they wish to avoid erosion of bargaining power by their daughters in their marriages if the relative wealth level of the husband and wife affect their relative bargaining power within a family. These two effects go in opposite directions. In addition, households without a son may also be induced to save more if the competition among families with a son bids up the price of housing.

The increased pressure in the marriage market comes from China's rising sex ratio imbalance, which has made it progressively more difficult for men to get married. As far as we know, we are the first in the literature to propose this hypothesis as an explanation for a rising savings rate in China. Toward the end of the paper, we argue that the explanation is likely applicable to many other countries as well.

We provide a series of evidence. First, we examine household-level data that cover 122 rural counties and 70 cities. While households with a son typically save more than households with a daughter, we do not regard this per se as supportive evidence of our hypothesis, since other channels could account for this difference. Instead, the evidence that we find compelling is that savings by otherwise identical households with a son tend to be greater in regions with a higher local sex ratio. This is something clearly predicted by our hypothesis but not directly by any other existing explanations. In addition, we find that savings by households with a daughter do not decline in regions with a high sex ratio, which is consistent with our interpretation that there are offsetting incentives on savings faced by households with a daughter. We discuss reasons that these patterns are unlikely to be the outcome of a selection bias in the data.

Second, across provinces, we show that the local savings rate tends to be higher in regions and years in which the local sex ratio (for the premarital age cohort) is higher. This continues to be true after we control for local income, social safety net, the age profile of the local population, and province and year fixed effects.

Third, in recognition of possible endogeneity of and measurement error in local sex ratios, we employ instrumental variables where the local sex ratio for the premarital age cohort is shown to be linked to local financial penalties for violating family planning policies set more than a decade earlier (as Ebenstein [2008] and Edlund et al. [2008] also document). With two-stage least squares estimation, the effect of local sex ratios on local savings rates remains positive and statistically significant. In fact, the point estimate becomes larger. This suggests that an increase in the sex ratio causes a rise in the savings rate. By our estimation, the sex ratio effect can explain more than half the actual rise in the household savings rate from 1990 to 2007.

Because the sex ratio imbalance at birth has been increasing steadily since the mid-1980s, the imbalance for the premarital age cohort will almost surely be higher over the next decade than in the last decade, even if the sex ratio at birth starts to be reversed soon. This implies that the incremental savings rate that is stimulated by the competitive savings motive will almost surely rise in importance in the near future.

A theoretical model (Du and Wei 2010) that has been developed as a result of the current paper suggests that the competitive savings triggered by a rise in the sex ratio can produce a significant amount of current account imbalances. Once one recognizes that the sex ratio imbalance is a structural factor behind a rising savings rate, it should be clear that a discussion of global imbalances that focuses narrowly on exchange rates or even social safety nets is incomplete. Furthermore, policy actions that improve the economic status of women could potentially reduce the sex ratio imbalance by reducing parental preference for sons (Qian 2008). A relaxation of family planning policy could also reduce the imbalance. Because of their important implications for aggregate savings and current account imbalances, these changes deserve more attention than they receive now.

The paper is organized in the following way. In Section II, we review the relevant literature. In Section III, we provide statistical evidence for our hypothesis. Finally, in Section IV, we conclude and discuss possible future research. A Data Appendix explains the sources and definitions of the main variables.

II. Review of the Existing Literature

A relevant literature is the work on status goods and social norms (e.g., Cole, Mailath, and Postlewaite 1992; Hopkins and Kornienko 2004, 2009; Hopkins 2009). When allowing certain goods to offer utility beyond their direct consumption value (i.e., through "status," which in turn could affect the prospect of finding a marriage partner), it is easy to show that consumption and savings behavior can be altered. However, none of these papers formally features a sex ratio imbalance. The imbalance leads to nontrivial general equilibrium questions. In particular, while men may react to a rise in the sex ratio by raising their savings rate, could the women do the reverse to take advantage of the higher savings of their future husbands? In addition, could men (or their parents) compete by increasing their conspicuous consumption as a signal of their attractiveness? This could result in a decline in the savings rate. Our response is that, while conspicuous consumption may increase the frequency of dating, the probability of securing a marriage partner may depend more on showing substantial wealth than on showing off a few flashy goods. In any case, it is an empirical question as to whether savings rates are positively or negatively related to sex ratio imbalances.

Another relevant literature is the economics of family. Several papers have explored the effects of a sex ratio imbalance on marriage prospects by gender and (female) labor market participation (e.g., Edlund 2001; Angrist 2002; Chiappori, Fortin, and Lacroix 2002). One interesting finding that is particularly relevant for this paper is that higher sex ratios (more men than women) tend to increase female bargaining power in the marriage market and within households. Siow (1998) studies the consequences of the relative shortage of fecund women in the marriage market for gender roles. However, these papers do not study directly the implications for aggregate savings.

The model in Du and Wei (2010), developed concurrently with this paper, and Bhaskar and Hopkins (2011) are the only ones that we are aware of that study the effect of a rise in the sex ratio on the aggregate savings rate in general equilibrium. They consider an overlapping generations (OLG) model with two sexes and a desire to marry. Everyone lives two periods. She or he works and saves in the first period. Marriage can take place only at the beginning of the second period and only between men and women in the same generation. All men (and women) are identical ex ante.

There are two benefits associated with marriage. First, the couple can pool their savings, and their consumption has a partial public good feature (e.g., the same car and furniture can be used by both). In other words, the sum of the husband's and wife's consumption can be more than their combined wealth. Second, in a marriage, one obtains emotional utility (or "love") from his or her partner. "Love," or the amount of emotional utility any person can offer to his or her spouse, is a random variable in the first period; only its distribution is known. It becomes public information once one enters the marriage market in the second period. Everyone ranks members of the opposite sex by a combination of two things: the level of wealth (which is determined by the first-period savings rate) and the value of emotional utility. This implies that, in partial equilibrium, raising the savings rate is a channel for a man (or a woman) to improve his (her) standing relative to his (her) competitors in the marriage market.

The matching of men and women in the marriage market is assumed to follow the Gale and Shapley (1962) "deferred acceptance" algorithm: first, each man proposes to his first choice. When a woman obtains multiple proposals, she rejects all unacceptable proposals and "holds" the most preferred one in this round. Second, every man who is rejected in one round proposes in the next round to his next best woman among those who haven't rejected him. Each woman "holds" the most preferred suitor so far and rejects the rest. Third, the process repeats itself until no new proposal is made. With this assumption, there exists a unique and stable equilibrium (pairs of men and women), and the equilibrium features a positive assortative matching—the best man and the best woman are matched, the second best man and the second best woman are matched, and so on. The lowest ranked men are not married.

The key proposition in Du and Wei (2010) can be summarized as follows (with intuition provided where appropriate). As the sex ratio increases, a representative man raises his savings rate (with the hope of improving his chance of success in the marriage market). In the benchmark model where there is no intrahousehold bargaining within a marriage, a representative woman decreases her savings rate in response to a rise in the sex ratio, because she expects to free ride on her future husband's higher savings rate. However, in an extension when they consider intrahousehold bargaining (the relative bargaining power is partly a function of the relative premarriage wealth level between husband and wife), the effect of a rise in the sex ratio on a representative woman's savings rate is ambiguous. The desire to avoid erosion in bargaining power within marriage may induce the woman to raise her savings rate in response to a rise in the sex ratio; this offsets the desire to free ride on her future husband's higher savings rate. In both the benchmark model and the extension, the aggregate savings rate unambiguously increases in response to a rise in the sex ratio.

The last result is easy to understand when there is intrahousehold bargaining and women do not decrease their savings (or at least not by too much) in order to avoid major erosion in bargaining power. But when there is no intrahousehold bargaining, and women do decrease their savings, what ensures that the aggregate savings rate goes up in general equilibrium? The answer is that men anticipate that women may decrease their savings, and raise their savings rate even more to compensate.[2]

In calibrations of the model, a rise in the sex ratio from a balanced level to 1.15 (which is approximately the level of the sex ratio for the premarital age cohort in China in 2007) would lead to an increase in the aggregate private savings rate by 6–10 percentage points, which is

[2] The proposition requires that the sex ratio is below some threshold. The intuition for the threshold is this: when the sex ratio becomes extremely unbalanced, any additional increase in the sex ratio may lead men to give up (by switching to a mixed strategy). When the probability of marriage is already sufficiently small, a further marginal improvement in marriage probability is not worth the additional sacrifice that men have to make from increasing the savings rate even more. If men derive enough utility from having a marriage partner, then the threshold value of the sex ratio imbalance is relatively high. In calibrations, Du and Wei find the threshold to be greater than 2. No real economy in the data has a sex ratio that is higher than 1.3. In other words, no real economy is likely to have reached this threshold.

about 30–60 percent of the actual increase in the household savings rate observed in the data.

In this model, it is clear that the high savings rate is inefficient for men: men raise their savings rate in the hope of improving their relative competitiveness in the marriage market. Yet, in the aggregate, the number of men who cannot get married is independent of the individual savings rate. If a central planner could coordinate the men's savings behavior, they could all reduce their savings proportionately without a negative consequence on their marriage outcome. However, the model is not able to prove that a higher savings rate is socially inefficient in general, because women benefit from the higher savings rate of their husbands. In calibrations with transferable utilities, however, it is generally the case that a higher savings rate due to a higher sex ratio is socially inefficient.

Outside the Du-Wei model, it is useful to point out some other forces that could raise the savings rate even by parents without a son. First, there is a housing price channel. Parents with a son (or unmarried men) may attempt to increase their competitiveness by buying a larger house and may bid up housing prices in a region with an unbalanced sex ratio. As a consequence, even parents without a son have to save more in order to afford housing. Second, there is a "tournament effect." When men save more, the reward for savings by women or parents with a daughter increases, if wealthier men also prefer relatively wealthy women. As a result, parents with a daughter are also more willing to save. For a theoretical model that may deliver this result, see Peters and Siow (2002). Many papers have looked into the determinants of the Chinese savings rate (e.g., Qian 1988; Horioka and Wan 2006; He and Cao 2007). None has explored the role of a sex ratio imbalance.

This discussion has clear implications for empirical work. First, it would be useful to check how savings by households with a son or a daughter respond to local sex ratios. Second, because different households may respond differently to the same change in the sex ratio, it is useful to go beyond household data and estimate the net effect of higher local sex ratios on aggregate savings rates. Third, it would be informative to check if local housing prices are indeed linked to local sex ratios.

III. Statistical Evidence

Since 1980, both the sex ratio for marriage-age youths and the savings rate in China have been rising. In figure 1, we present a time series plot of standardized versions of both variables.[3] The sex ratio at birth is lagged by 20 years, since the median age of first marriage for Chinese

[3] Standardized variable = (raw variable − mean)/standard deviation.

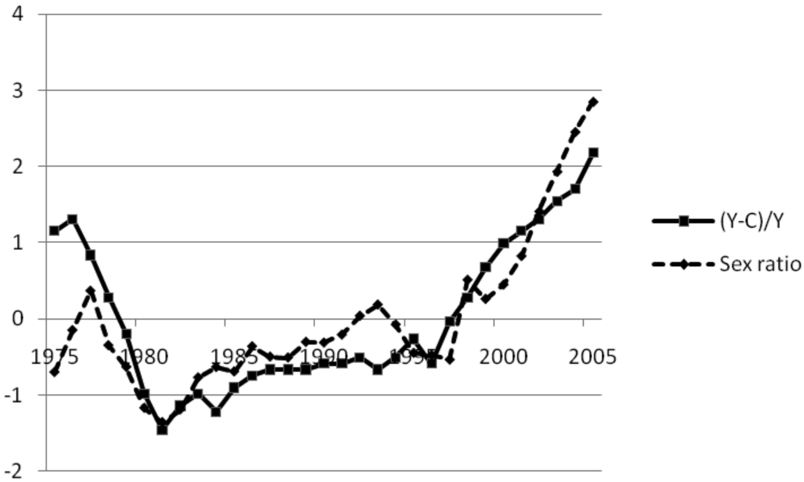

FIG. 1.—Sex ratio and saving rate. The sex ratio is defined as the ratio at birth 20 years earlier. See the note to table 12 for data sources. The saving rate is defined as the percentage of GDP-private and government consumption in total GDP, which is available from *China Statistical Yearbook 2007*. Both variables have been rescaled by subtracting the mean and dividing by the standard deviation.

women is 20. It is clear that the two variables are highly correlated (the correlation coefficient is 0.822). While this is consistent with the competitive savings hypothesis, it is not a rigorous proof by itself.

To set the stage for formal statistical testing, we first provide a series of background information in Section III.A. We describe the patterns of the sex ratio imbalance in the country and households' self-reported savings as a function of the gender composition of the children. We present some evidence on how household savings are heavily connected to children's marriage events and on how family wealth enhances a young person's marriage prospects. We also show evidence that the marriage market is fairly local for most people.

In Section III.B, we employ formal statistical tests of the competitive savings hypothesis based on household-level data. We are especially interested in examining how a combination of having a son and living in a region with a skewed sex ratio alters a family's savings behavior. In Section III.C, we turn to cross-regional evidence. Because savings by different households could respond differently to a given rise in the sex ratio, the regional analysis offers us a chance to estimate the general equilibrium effect of a higher sex ratio on the (province-level) aggregate savings rate.

A. *Background Information on the Marriage Market and Savings Behavior*

From Unbalanced Sex Ratios to "Bare Branches"

Left to nature, the sex ratio at birth (in a society without massive starvation) is generally around 106 boys per 100 girls (with human biology compensating for a slightly higher mortality rate for infant boys than girls). The sex ratio was balanced or slightly below normal in the 1960s and 1970s (most likely due to malnutrition). The sex ratio at birth in China was close to normal in 1980 (with 106 boys per 100 girls) but has climbed steadily since the mid-1980s to over 120 boys for each 100 girls in 2005 (Das Gupta 2005; Li 2007; Zhu, Lu, and Hesketh 2009) and is estimated to be 124 boys/100 girls in 2007. By 2005, men outnumbered women at age 25 or below by about 30 million. The excess men mathematically cannot be married. The number of unmarried men—sometimes referred to as "bare branches" in colloquial Chinese—continues to rise as the sex ratio imbalance deteriorates.[4] Some men may partner off as gays, and others may emigrate or marry women from other countries. Because the scale of the "bare branches" is so large—30 million men are more than the entire female population of Canada—and because most "bare branches" come from low-income households, actions by a small portion of men do not provide a practical solution to the problem. In any case, a rising sex ratio imbalance must imply a diminishing probability that a man will find a bride.[5]

Self-Reported Reasons for Savings

A survey of rural households (Chinese Household Income Project) in 2002 asked households why they save. There were seven possible reasons: (1) children's wedding, (2) children's education, (3) bequest to chil-

[4] If some newborn girls are not reported by their parents due to hopes of being pregnant again with a son, then the sex ratio imbalance at birth could be overstated in official statistics. To assess the quantitative importance of this possibility, we compare the sex ratio at birth in the 1990 population census with the sex ratio for 10-year-olds in the 2000 census. It is reasonable to assume that parents would not hide their 10-year-old girls from census takers in 2000. First, those parents who would like to try for another child in 1990 most likely would have done so already within 10 years and have paid a fine for violating family planning policy by 2000. In addition, there were also positive incentives to report girls who have reached school age since registration was required for (free) immunization shots and school attendance. Since the sex ratios for both newborns in 1990 and 10-year-old children in 2000 were 1 : 12, we conclude that the number of underreported infant girls, as a proportion of the total number of newborn girls, is not large enough to make a noticeable distortion to the reported sex ratio imbalance. Zhu et al. (2009) reached the same conclusion with a somewhat different methodology.

[5] A fraud has emerged in which some women pretend to be willing to marry bachelors in a rural area in return for a bride price (*cai li*) on the order of RMB 40,000 (about US$5,900, or "five years' worth of farm income"). The women then run away after the wedding with the bride price (Mei Fong 2009).

TABLE 1
WHY DO PEOPLE SAVE? SELF-REPORTS OF THE MOST OR THE SECOND MOST IMPORTANT REASON FOR SAVINGS (Percentage of Respondents)

	Three-Person Household		Four-Person Household			
	Girl (1)	Boy (2)	Only Girls (3)	Boy and Girl (4)	Only Boys (5)	All Households (6)
Total sample:						
Directly related to children	86.4	92.2	86.4	94.0	96.1	78.2
Children's wedding	18.3	29.8	22.0	34.0	37.4	33.0
Children's education	75.9	79.2	75.7	82.1	80.4	52.0
Bequest to children	12.5	11.9	10.2	8.9	6.8	13.8
Not directly related to children	69.6	59.2	72.3	56.0	55.9	69.5
To build a house	19.7	20.2	20.3	24.3	26.7	18.3
Retirement	45.5	37.3	45.8	27.9	22.8	47
Medical expenses	14.2	6.1	14.7	7.5	8.5	18.9
Others	8.7	7.1	2.3	8.2	11.7	9.5

SOURCE.—Authors' tabulation based on the Chinese Household Income Project (2002), available from http://www.icpsr.umich.edu/cocoon/ICPSR/STUDY/21741.xml.

dren, (4) building a house, (5) retirement, (6) medical expenses, and (7) others.

We group the first three reasons as "directly related to children." In table 1, we tabulate the percentage of households that designate a given category as either the most important or the second most important reason for savings. (Note that the sum of the numbers in a column can be more than 100 percent because a household can list one category as the most important and the other category as the second most important.) In columns 1 and 2, we focus on households with husband and wife plus a child. We tabulate the answers from households with a son in column 1 and those from households with a daughter in column 2. The table shows that 92.2 percent of the son-households give at least one factor directly related to their son as their primary or secondary reason for savings. This number is 5.8 percentage points higher than the percentage of daughter-households who give similar answers. It is also telling to look at savings for children's weddings: 29.8 percent of son-families list savings for their son's wedding as the primary or secondary most important reason for savings versus only 18.3 percent of daughter-families who give the same answer.

Interestingly, when it comes to children's education, a majority of both son-families and daughter-families save for this purpose (79.2 percent vs. 75.9 percent, respectively), and the difference is small. When it comes to bequests, there is virtually no difference between the two types of households. It is important to point out that these numbers do not directly reveal the intensity of savings for a given type of household. (For that, we will perform formal regression analyses.)

In the last three columns of table 1, we look at four-person households with a father, a mother, and two children. There are three types of such households: those with two girls, those with a girl and a boy, and those with two boys. The relative differences are similar to three-person households. More precisely, households with at least one son are more likely to report that their savings are primarily for their children, particularly for their children's wedding.

Note that among the factors labeled as "not directly related to children," "building a house" for families with a son could very well be motivated by a desire to help their sons improve their marriage prospects. Our personal interviews with families in rural areas suggest to us that most families with a son believe that having a house for a son is essential for their son's prospects in successfully finding a wife. Because the CHIP survey does not allow us to separate "building a house for their children's wedding" from building a house for other reasons, we do not include "building a house" as a part of the "savings reasons directly related to children."

Savings Rates and Wedding Events

Our hypothesis connects sex ratio to savings rate through pressure in the marriage market. It is therefore desirable to have some evidence on whether household savings rates actually vary with the timing of a wedding event in the family and whether the pattern differs between households with a son and those with a daughter. Unfortunately, most household surveys do not ask about the timing or costs of wedding events in a way that would allow one to trace out a time series profile of a household's savings rate with respect to wedding events.

Fortunately, for 26 natural villages (in three administrative villages) in Guizhou Province, two rounds of household census were conducted in 2005 and 2007 by the International Food Policy Research Institute (IFPRI; with one of the authors as a project leader). In each round, all households were asked their income and expenditure in the previous year. The second round survey also includes recall data of major events, such as weddings in a family, in the preceding 10 years. From this data set, we construct a time series profile of a "typical" household's savings rate with respect to the timing of a wedding, that is, the savings rate 2 years before the wedding, 1 year before the wedding, the year of the wedding, and so on, all the way to 4 years after the wedding. By a "typical" household, we mean that we take the average across households with the same number of years distant from a wedding. For example, for the

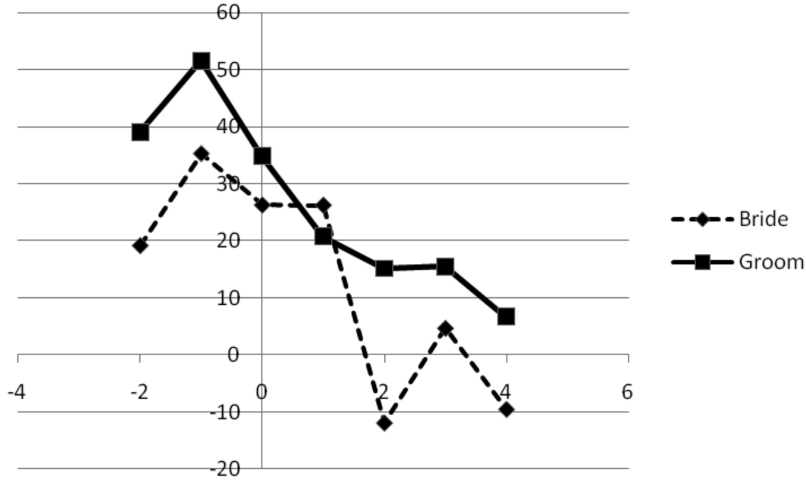

Fig. 2.—Time profile of household savings rate in relation to the timing of a wedding: evidence from 26 natural (three administrative) villages in Guizhou Province. Authors' calculation based on surveys designed by one of the authors and conducted by IFPRI.

savings rate in the year of a wedding, we average the savings rate across all households that have a wedding in that year. Because the number of households that have a wedding event is relatively small, we do not have enough statistical power to include control variables or to perform formal tests on the differences in the average savings rates across years or across household types. So the result should be interpreted with these limitations in mind.

Figure 2 plots the time profile of a representative household savings rate for groom and bride families, separately.[6] The horizontal axis stands for the number of years away from the time of a wedding. The vertical axis depicts the average household savings rate measured by the formula $100 \times (\text{income} - \text{expenditure})/\text{income}$. Three patterns in the graph are particularly suggestive. First, the two savings rate curves exhibit an inverse-V shape, peaking in the year before the wedding. The savings rate is as high as 50 percent for a groom's family. (The median wedding cost for the groom's family was 18,150 RMBs in 2006, exceeding 800 percent of the per capita income in the sample. Because a wedding event itself represents a major expenditure, it is not surprising that the savings rate for the year of a wedding is lower.) Second, household savings rates tend to be much lower after a wedding rather than recovering to the prewedding level, suggesting that a big part of household

[6] To reduce noise, the top and bottom of 5 percent outliers in terms of savings rates are dropped from the sample.

savings is motivated by wedding-related expenditures. Third, the savings rate curve for a groom's family lies almost everywhere above the curve for a bride's family (except in the year after a wedding). This suggests that savings for marriage is more important for a groom's family than for a bride's family. (The savings rate for a bride's family turns negative in some years after the wedding, possibly indicating that they consume more than their income due to a transfer from the groom's family.)

While this piece of evidence comes from a rural location, the patterns revealed are consistent with the cultural norms in both urban and rural areas. In particular, a groom's family is more likely than a bride's family to be expected to provide a house or an apartment for the newlyweds or at least to contribute the biggest chunk of the cost for a domicile. A groom's family is often responsible for paying his bride's family a one-time transfer that compensates the latter for rearing their daughter (Zhang and Chan 1999). In addition, the groom's family bears most of the financial cost of holding a wedding ceremony, although the bride's family may share some of the cost as well. Because weddings in China are occasions that call for significant cash outlays, families may have to save more before weddings. Brown, Bulte, and Zhang (2011) show that families of a groom expend more on both the wedding ceremony and the bride price over time, while the families of a bride do not.

While the inverse-V shape of the savings curve in figure 2 means that savings rates tend to decline after a wedding, it does not imply that the net consequence of a higher sex ratio on savings rates is zero. First, in a society with a positive population growth rate, the sum of the extra savings by families preparing for a wedding may grow faster than the sum of the dis-savings by postwedding families. Second, more importantly, in response to a rising sex ratio, the entire savings curve is likely to shift upward, especially for households with a son. In addition, parents wish to leave a positive bequest for their children.

Material Wealth and Marriage Likelihood

A key assumption of our story is that a higher level of family wealth improves a man's chances in the marriage market. We look for evidence in this regard. Note first that the social norm in China is such that most unmarried young men or women live with their parents. As a result, household surveys rarely capture households consisting of a single unmarried person. (Married children, on the other hand, have their own households and do not appear in their parents' households in the survey.) We make an indirect educated guess on the marital status of eligible bachelors from the CHIP data by looking at households whose head is 50–60 years old. Their children are likely to be in the 25–35-year age range. We check if family wealth reduces the likelihood that such a household has an unmarried adult child living with them.

In rural areas, virtually every household has a house, but the quality varies a great deal. A house built of concrete, brick, or stone is likely of higher quality (and more expensive) than a house built of mud and straw. The regression in column 1 of table 2 shows that households with a son are much less likely to have an unmarried adult son staying at home if they have a relatively higher-quality house. In comparison, in column 2 of table 2, the marriage status of a daughter (for daughter-families) is unrelated to the relative quality of the parents' house. A multinomial logit (reported in cols. 3–4 of table 2) confirms the same pattern.

In urban areas, all apartments or houses are built of concrete and bricks. However, some people are renters, not owners. Generally speaking, those who own their residence are wealthier than those who rent their residence. Column 5 shows that son-families are much less likely to have an unmarried adult son at home if they are a house owner (as opposed to a renter). This is consistent with the idea that a higher level of wealth makes a man more marriageable.

We also look at the likelihood that an adult child is married in rural Guizhou during 1996–2006. This is based on the same census of households in 26 natural villages used in figure 2. While this is a much smaller sample than the CHIPS data set, it contains an estimate of the value of the family house (likely the most important piece of household wealth). In addition, all the marriages that occurred during the period of 1996–2006 are recorded. Since there is no mortgage product in rural Guizhou, families tend to construct their houses and own them outright. Table 3 reports a set of logit regressions on the likelihood that an adult son (or daughter) is married. Among families with at least one adult son (older than 19), a higher level of housing wealth is a statistically significant predictor of the likelihood that the adult son gets married. Family income also has a positive coefficient but is not statistically significant. In comparison, among families with at least one adult daughter (older than 17), the level of housing wealth is not significant, but a higher level of family income is a positive and significant predictor for the likelihood that the daughter is married.

Determinants of Sex Ratio Imbalance

The sex ratio imbalance comes almost entirely from sex-selective abortions. This, in turn, results from a combination of three factors: (*a*) parental preference for sons; (*b*) some limit to the number of children a couple is allowed or wants to have, which for the Chinese is a strict family planning policy; and (*c*) availability of inexpensive technology to screen the sex of a fetus (Ultrasound B in particular) and to perform abortions.

Our empirical work will start with regressions that assume exogenous sex ratios. This assumption can be justified by recognizing that parental preference for sons is part of a culture and, as such, it changes only very slowly. Korea experienced a sustained increase in its sex ratio imbalance for about 25–30 years that has only recently started to decline; this evidence is consistent with our assumption (Guilmoto 2007).

We nonetheless also report instrumental variable regressions that allow for potential endogeneity of (and measurement error in) sex ratios. The instrumental variables for sex ratio in the premarital age cohort explore regional variations in the financial penalty for violating birth quotas, set by regional governments years before newborns grow to marriageable age.

Low Mobility for Marriage and the Invisible Hand

As we will explore variations across provinces in sex ratios and savings rates, it is useful to know how local the marriage market is. First, according to the China population census of 2000, 92 percent of rural residents live in their county of birth, and 62 percent of urban residents live in the city of their birth. Second, in rural China, 89 percent of marriages take place between husbands and wives from the same county. Since a county is a smaller geographic unit (a typical province may have more than 100 counties), the percentage is surely higher for marriages between men and women from the same province. Third, the CHIP migrant workers survey in 2002 shows that 82 percent of the migrant working families in cities report that the husbands and wives come from the same province. This suggests that migrant workers often get married before leaving their hometown to look for a job. To sum up, mobility for marriage reasons appears modest.

For the sex ratio to affect household savings rates, parents do not have to know local sex ratio statistics. There is an invisible hand at work. Consider two otherwise identical households with a son, one in a region with a high sex ratio and the other in a region with a low sex ratio. Parents in the first region would observe or be told by relatives or colleagues with a son that it is difficult for their sons to find a girlfriend and expensive to marry. The expectation for how much the parents (or the sons) need to contribute to their son's new household, given costs of housing, cars, furniture, or honeymoons, would differ in the two regions. The types of furniture, cars, and honeymoons, and local housing prices may reflect the degree of competition in a local marriage market and thus affect the savings required of parents with a son. In other words, even without the knowledge of local sex ratio statistics,

TABLE 2

MATERIAL WEALTH AND MARITAL STATUS: WHICH FAMILIES ARE MORE LIKELY TO HAVE AN UNMARRIED ADULT CHILD?

	RURAL SAMPLE				URBAN SAMPLE			
	Logit		Multinominal Logit		Logit		Multinominal Logit	
	Son (1)	Daughter (2)	Son (3)	Daughter (4)	Son (5)	Daughter (6)	Son (7)	Daughter (8)
Housing wealth	−.45**	.05	−.45**	.05	−.39*	.33	−.38*	.27
	(.18)	(.38)	(.18)	(.38)	(.21)	(.37)	(.21)	(.35)
Per capita income (log)	.26**	.17	.26**	.18	.33**	.17	.30**	.16
	(.12)	(.23)	(.12)	(.23)	(.15)	(.20)	(.15)	(.20)
Household size	.11**	.21**	.11**	.21**	.43**	.58**	.45**	.66**
	(.04)	(.07)	(.04)	(.07)	(.08)	(.08)	(.08)	(.09)
Household head age	.11**	.05	.11**	.05	.18**	.15**	.18**	.16**
	(.02)	(.04)	(.02)	(.04)	(.02)	(.03)	(.02)	(.03)
Household head gender (female = 1)	−.68	−.97	−.68	−.97	.61**	.10	.61**	.18
	(.47)	(1.04)	(.47)	(1.04)	(.17)	(.25)	(.17)	(.24)

Household head year of schooling	.02	.11**	.02	.11**	−.05*	.08**	−.05*	.08**
	(.03)	(.05)	(.03)	(.05)	(.03)	(.04)	(.03)	(.04)
Household head as a minority	−.15	−.29	−.14	−.29	.14	−.02	.12	−.04
	(.18)	(.33)	(.18)	(.33)	(.33)	(.49)	(.33)	(.49)
Poor health	.60**	.35	.60**	.32	.32	.11	.31	.13
	(.26)	(.53)	(.26)	(.53)	(.20)	(.29)	(.20)	(.29)
AIC	1,463.2	602.5	2,078.4		1,080.0	650.1	1,756.1	
Observations	2,575	2,423	2,640		1,621	1,525	1,712	

NOTE.—We restrict the sample to households whose heads are between 50 and 60 years old. For the logit regressions, the dependent variable is defined as one if a family has an unmarried son (or daughter) between the ages of 25 and 35, zero otherwise (all similar households without an adult child at home). For the multinomial logit regressions, we define the dependent variable as one for having an unmarried adult son, two for having an unmarried daughter, and zero otherwise (all similar households without an adult child at home). Columns 3–4 (or cols. 7–8) represent a single regression. The variable "housing wealth" in the rural sample is a dummy if the family house is made of concrete, bricks, or stones (as opposed to mud or other inferior material); it is a dummy in the urban sample if the family owns an apartment/house (as opposed to renting one). The data are from the Chinese Household Income Project (2002). AIC stands for Akaike Information Criterion.

* Significant at the 10 percent level.

** Significant at the 5 percent level.

TABLE 3
Wealth and Wedding in Rural Guizhou Province, China

	Household Head Age (45–65)		Whole Sample	
	Wedding for Son (1)	Wedding for Daughter (2)	Wedding for Son (3)	Wedding for Daughter (4)
House value (log)	.034**	.019	.022**	.009
	(.018)	(.275)	(.010)	(.350)
Household income (log)	.013	.059**	.021	.030*
	(.57)	(.05)	(.13)	(.07)
Household size	.001	−.01	.003	−.01
	(.93)	(.70)	(.67)	(.17)
Household head age	.009*	.003	.000	.000
	(.06)	(.64)	(.70)	(.95)
Household head gender (female = 1)	−.034	−.138	−.048	−.063
	(.63)	(.10)	(.27)	(.30)
Household head year of schooling	.013	.043**	.00	.046**
	(.44)	(.03)	(.74)	.00
Household head as a minority	−.354**	−.360**	−.373**	−.386**
	(.00)	(.00)	(.00)	(.00)
Adjusted R^2	.273	.25	.267	.238
AIC	126.1	228.7	389.4	651.8
Observations	225	225	664	664

Note.—Authors' calculation based on a household survey in Guizhou Province designed by one of the authors and conducted by IFPRI in 2007. The survey asked about weddings in the period 1996–2006. Columns 1 and 3 are for families with at least a son; columns 2 and 4 are all families with at least a daughter. Robust standard errors are in parentheses.

* Significant at the 10 percent level.

** Significant at the 5 percent level.

parents with a son may make savings decisions that reflect the local sex ratio.

B. Household-Level Evidence

The competitive saving motive predicts that household savings should respond systematically to local sex ratio imbalances, especially for families with a son. Our sample is drawn from the Chinese Household Income Project (CHIP) of 2002, which covers 122 rural counties and 70 cities.

One may be tempted to compare savings rates for households with sons and those with daughters. But this comparison is not particularly informative for our hypothesis. Under the hypothesis of a competitive saving motive, one expects that families with a son will save more, if

TABLE 4
Summary Statistics on Household Savings in 2002

Household Type	Mean	Median	Max	Min	SD	Observations
Rural:						
One son	.393	.394	2.462	−2.986	.625	580
One daughter	.318	.353	1.812	−3.559	.626	326
All families	.316	.316	2.846	−5.026	.582	9,199
Urban:						
One son	.312	.306	1.849	−1.816	.333	769
One daughter	.302	.308	2.153	−1.299	.356	766
All families	.304	.286	2.308	−2.432	.378	6,835

Note.—The savings rate is defined as log(income/consumption). The data come from the Chinese Household Income Project (2002), available from http://www.icpsr.umich.edu/cocoon/ICPSR/STUDY/21741.xml. To maximize comparability, we restrict the sample to nuclear households with both parents still alive and mother younger than 40. "All families" include families with at-home children and families with more than two children.

other things are held constant. However, other things cannot be held constant because parents may not expect to get the same degree of help from their daughters than from their sons as they age, especially in a rural area where a woman moves to live with a man's family after the marriage. In other words, families with only daughters may need to save more to prepare for old age. For this reason, a direct comparison of the savings rates between these two types of households is not informative for our purpose.

To ensure that we compare apples with apples, we focus on those three-person nuclear families with both parents still alive, mother younger than 40, and no grandparents, uncles, or aunts living at home. If a family has an adult child who has moved out of the family, the household survey does not capture this accurately (and doesn't have information on the gender of the adult child). By placing an age limit on the mother, we are likely to have true three-person families and, hence, to maximize the comparability across the households in the sample. These restrictions reduce the size of the sample relative to the universe of households in the survey. However, as we will see, our regressions do not suffer from a lack of statistical power. (As noted earlier, since most unmarried young people live with their parents, the survey does not contain many observations of an unmarried young man or woman as the household head. Therefore, we are not able to analyze such households directly.)

In any case, table 4 reports average savings rates for households with children of different genders. Following Chamon and Prasad (2010), we define the local savings rate by log (income net of taxes/living expenditure). This definition is less susceptible to extreme values and makes the error term more likely to satisfy the normality assumption. In both rural and urban areas, households with a son have a higher average savings rate than those with a daughter (39 percent vs. 32 percent in rural areas and 31 percent vs. 30 percent in urban areas). However, none of the differences in savings rates is statistically significant; the standard deviation of the savings rate within any type of household easily overwhelms the difference in the savings rates between any two types of households. Of course, large standard deviations also suggest the presence of considerable noise in the savings data. In any case, we cannot confirm or reject our hypothesis by comparing the savings rates across household types in this way.

Our hypothesis, however, implies a particular regional variation in savings rates: households with a son should save more in a region with a more unbalanced sex ratio, holding constant family income and other characteristics. Moreover, this pattern is not predicted by either the life cycle theory or by existing precautionary savings motive hypotheses. Therefore, examining the relationship between household savings rates and local sex ratios may be particularly informative for testing our hypothesis.

We can also examine the relationship between the savings rate by households with a daughter and the local sex ratio. Free riding on future husbands' wealth is not the only driver for savings by households with a daughter. The model of Du and Wei (2010) suggests that when women (or their parents) are concerned about the erosion of bargaining power within a household, they may respond to an anticipated increase in their future husbands' wealth by raising their own savings rates. In addition, there may be a spillover channel in savings pressure from households with a son to other households. If the spillover channel operates through housing prices, it is likely to be stronger in urban areas as houses are more likely to be purchased from the local market (as opposed to being built on the occupants' own land in rural areas).

In the household-level regressions, we infer local sex ratios for the cohort of ages 12–21 in 2002 from the 1990 population census (the cohort was 0–9 years old in 1990). There are substantial regional variations. In the rural sample, the average sex ratio is 1.09, with a standard deviation of 0.04. The smallest and largest values are 1.01 and 1.23, respectively. In the urban sample, the sex ratio ranges from 1.02 to 1.24 with a mean of 1.08 and a standard deviation of 0.04.

Household Savings in Rural Areas

The regression results for the rural sample are presented in table 5. The first two regressions are performed on a full sample. The regressions control for family income, children's ages, characteristics of the head of household (sex, education, and ethnic background), and local income inequality. They also control for health shocks to the family by a dummy denoting "poor health" if the family has a disabled or severely ill member.

Column 1 of table 5 relates savings by households with a son to local sex ratios and other determinants of the savings rate. We find that the local sex ratio has a strongly positive effect on the household savings rate (with a point estimate of 1.34 on the sex ratio, which is statistically significant at the 5 percent level). An increase in the local sex ratio from 1.05 to 1.14 (the mean increase in rural China from 1990 to 2007; see table 12) is associated with a higher savings rate by son-households by 12.1 percentage points, which is economically large (and is more than the actual increase in the average rural household savings rate during the period). In comparison, column 2 of table 5 reports the regression concerning savings by households with a daughter. The coefficient on the local sex ratio is negative but not statistically significant. This is consistent with the interpretation in Du and Wei (2010) that conflicting motives faced by daughter-families have offset each other, resulting in an ambiguous net effect.

Since the large standard deviations in the savings rates reported in table 4 likely reflect noise, we conduct a sequence of additional regressions where possible outliers are removed through three different filters. In columns 3 and 4 of table 5, we take out a small number of households either whose reported annual family income or reported annual expenditure is less than 2,000 Chinese yuan. (Such income or expenditure seems too low to be realistic, possibly due to underestimation of imputed nonmarket income or expenditure.) The same qualitative patterns are preserved. In particular, the savings rate by son-families rises with the local sex ratio, but the savings rate by daughter-families is insensitive to the local sex ratio. In fact, the point estimate on the sex ratio is now (moderately) bigger for the son-families and smaller for the daughter-families.

In columns 5 and 6 of table 5, we remove the top and bottom 5 percent of households (within each household type) in terms of their savings rates. In columns 7 and 8, we remove those households which do not explicitly give marital status for their child (in addition to removing the top and bottom 5 percent of households in terms of their savings rates). In both cases, we observe the same patterns: the coefficient on the sex ratio is positive and significant for son-families but not different from zero for daughter-families.

In table 6, we report additional robustness checks. First, we expand the sample to include all households whose mothers are younger than 45 (as opposed to 40 in table 5). (One important caveat is that the newly added households could contain many multiple-children households that are erroneously classified as one-child families, because older children who have moved out after marriage are not recorded as mem-

TABLE 5
RURAL HOUSEHOLD-LEVEL SAVINGS FOR THREE-PERSON HOUSEHOLDS WITH A CHILD IN 2002

	FULL SAMPLE		SUBSAMPLE THAT REMOVES THE FOLLOWING POTENTIAL OUTLIERS: Income or Expenditure < 2,000 Yuan		SUBSAMPLE THAT REMOVES THE FOLLOWING POTENTIAL OUTLIERS: Bottom and Top 5% Savers		SUBSAMPLE THAT REMOVES THE FOLLOWING POTENTIAL OUTLIERS: Bottom and Top 5% Savers and No Explicit Marriage Status for Child	
	Son (1)	Daughter (2)	Son (3)	Daughter (4)	Son (5)	Daughter (6)	Son (7)	Daughter (8)
Local sex ratio (county level)	1.34**	−.17	1.38**	−.18	1.10**	−.23	1.20**	−.32
	(.52)	(.55)	(.51)	(.54)	(.44)	(.43)	(.43)	(.44)
Per capita income (log)	2.88**	2.49**	2.60**	2.76**	1.82**	2.44**	1.52**	3.10**
	(.63)	(.45)	(.91)	(.80)	(.56)	(.72)	(.56)	(.59)
Per capita income squared (log)	−.15**	−.12**	−.13**	−.14**	−.09**	−.13**	−.07**	−.17**
	(.04)	(.03)	(.06)	(.05)	(.04)	(.05)	(.04)	(.04)
Child aged 5–9	.01	−.03	−.02	.09	−.05	−.05	.00	−.06
	(.10)	(.08)	(.10)	(.07)	(.07)	(.07)	(.06)	(.07)
Child aged 10–14	−.07	−.14	−.08	−.02	−.05	−.11	−.02	−.12
	(.10)	(.09)	(.09)	(.08)	(.07)	(.07)	(.06)	(.08)
Child aged 15–19	−.23*	−.20*	−.23**	−.08	−.20**	−.08	−.19**	−.08
	(.12)	(.11)	(.11)	(.11)	(.08)	(.09)	(.07)	(.10)
Household head age	.00	.01	.00	.01	.00	.00	.00	.00
	(.01)	(.01)	(.01)	(.01)	(.00)	(.00)	.00	(.00)

Household head gender (female = 1)	−.06	−.36*	−.05	−.36*	−.09	−.10	−.12	−.09
	(.14)	(.20)	(.13)	(.20)	(.12)	(.14)	(.14)	(.14)
Household head year of schooling	.01	−.02**	.01	−.02**	−.01	−.02*	−.01	−.02
	(.01)	(.01)	(.01)	(.01)	(.01)	(.01)	(.01)	(.01)
Household head as a minority	−.15**	−.22**	−.14**	−.17**	−.04	−.11*	−.05	−.10
	(.06)	(.08)	(.06)	(.07)	(.05)	(.06)	(.05)	(.07)
Poor health	.03	−.72*	.06	−.72**	−.04	−.17**	−.03	−.02
	(.12)	(.37)	(.12)	(.37)	(.11)	(.05)	(.09)	(.16)
Gini at the county level	−1.01**	−.40	−1.16**	−.33	−.88**	−.53	−.71**	−.38
	(.45)	(.44)	(.40)	(.40)	(.34)	(.33)	(.35)	(.35)
Adjusted R^2	.30	.54	.27	.36	.20	.30	.20	.29
AIC	906	376.4	836.3	329.8	466.5	171.2	393.5	157.2
Observations	580	326	564	315	522	292	489	269

NOTE.—The savings rate is defined as log(income/consumption). To maximize comparability, we restrict the sample to those three-person nuclear families with both parents still alive, mother age younger than 40, and one child. For the first six regressions, a family member is defined as a child if younger than 20 years old, while for cols. 7 and 8, those unmarried and registered as a child of household head are counted as children. "Poor health" is a dummy that takes the value of one if a household has at least one member with disability or extreme bad health. The sex ratio at the county level is calculated by the authors for the age cohort of 0–9 from China population census 1990 (who became 12–21 in 2002). Robust standard errors are in parentheses.

* Significant at the 10 percent level.

** Significant at the 5 percent level.

TABLE 6

Robust Checks: Rural Household-Level Savings

	Full Sample		Subsample That Removes the Following Potential Outliers							
			Income or Ependiture < 2,000 Yuan		Bottom and Top 1% Savers		Bottom and Top 5% Savers		Bottom and Top 5% Savers and No Explicit Marriage Status for Child	
	Son (1)	Daughter (2)	Son (3)	Daughter (4)	Son (5)	Daughter (6)	Son (7)	Daughter (8)	Son (9)	Daughter (10)
	Mother Age Younger than 45 (Nuclear Family)									
OLS	1.29***	−.13	1.32***	−.31	1.03**	−.19	1.32***	−.32	1.07***	−.58
	(.46)	(.60)	(.46)	(.58)	(.42)	(.56)	(.39)	(.48)	(.41)	(.48)
Median regression	1.18**	.07	.97*	.34	.95**	.07	1.09***	−.36	1.20**	−.26
	(.52)	(.46)	(.56)	(.52)	(.41)	(.41)	(.36)	(.31)	(.49)	(.63)
Observations	705	387	688	372	689	379	633	347	634	321

	Mother Head Age Younger than 45 (Extended Family)									
OLS	.99**	.19	.96**	.15	.85**	.21	1.19***	−.01	1.30***	.23
	(.39)	(.47)	(.39)	(.47)	(.36)	(.44)	(.32)	(.39)	(.34)	(.38)
Median regression	1.10***	.19	1.00**	.16	1.00***	.27	1.14***	.26	1.15***	.21
	(.41)	(.43)	(.41)	(.41)	(.39)	(.42)	(.34)	(.50)	(.35)	(.58)
Observations	1,125	639	1,099	617	1,101	625	1,011	575	933	562

NOTE.—The nuclear family refers to households with two parents and one single child. The extended family sample includes all those households with one single child no matter. Therefore, the household size for extended family may be larger than three. In the regression for the extended family sample, we include an additional variable to control for household size. To maximize comparability, we restrict the sample to those three-person nuclear families with both parents still alive, mother age younger than 40, and one child. For the first six regressions, a family member is defined as a child if younger than 20 years old, while for cols. 9 and 10, those unmarried and registered as a child of household head are counted as children.

* Significant at the 10 percent level.

** Significant at the 5 percent level.

*** Significant at the 1 percent level.

bers of the household. This could be a problem since the data do not tell us if the older children are sons or daughters.) Second, we extend the sample to include extended families (those with grandparents, uncles, or aunts). Third, we report median regressions as well as ordinary least squares (OLS) regressions. Fourth, we apply four additional filters similar to those in table 5. These produce a total of 10 pairs of new regressions. For each pair, we always obtain the same patterns: the coefficient on the sex ratio is positive and statistically significant for the son-families but not significant (and sometimes negative) for daughter-families. We therefore conclude that a robust feature of the data is that savings by son-families tend to be higher in regions with a more skewed sex ratio. In contrast, the net effect of the sex ratio imbalance on savings by daughter families appears to be zero.

In table 7, we pool the two household types into the same regression(s). We include an interaction term between the local sex ratio and a dummy if the family has a son. An advantage of this specification is that a simple *t*-test on the interaction term can tell us whether the son-families and daughter-families react differently to a given rise in the sex ratio. A disadvantage is that (additional) estimation bias can be introduced if we inappropriately impose the condition that the parameters on all other variables be the same for different types of households. Columns 1 and 2 report OLS and median regressions for one-child households including extended families (e.g., families with grandparents). Columns 3 and 4 report OLS and median regressions on nuclear one-child families only. In all regressions, the coefficient on the interaction between the son dummy and the sex ratio is positive and significant, but the coefficient on the sex ratio itself is not significant. In other words, son-families do not intrinsically have a higher savings rate than daughter-families. Instead, it takes a combination of having a son and living in a region with a high sex ratio for the family to have a high savings rate. This is very much consistent with our story.

It might be interesting to see if the savings response to the local sex ratio varies with family income. In table 8, we create dummies for households in different income quartiles (within a region) and interact them with the local sex ratio. If one just looks at the point estimates, there is some evidence that the savings response by son-families is moderately weaker for lower-income families (possibly because some of these families give up their hope for marriage for their sons and therefore stop competing through savings). However, the difference in savings responses across income groups is not statistically significant. Similarly, there is no significant difference in the savings responses by daughter-families across income quartiles.

TABLE 7
POOLED SAMPLE: RURAL HOUSEHOLD-LEVEL SAVINGS IN 2002

	All Families with One Child		All Nuclear Families with One Child	
	OLS (1)	Median (2)	OLS (3)	Median (4)
Local sex ratio (county level)	−.45	−.43	−.03	−.12
	(.36)	(.35)	(.56)	(.51)
Sex ratio × dummy for son	1.26**	1.15**	1.39*	1.21*
	(.48)	(.46)	(.77)	(.65)
Son	−1.39**	−1.25**	−1.53*	−1.33*
	(.53)	(.50)	(.83)	(.70)
Per capita income (log)	2.37**	2.36**	2.59**	2.19**
	(.20)	(.15)	(.38)	(.19)
Per capita income squared (log)	−.12**	−.12**	−.13**	−.11**
	(.01)	(.01)	(.02)	(.01)
Child aged 5–9	.04	−.01	−.02	−.07
	(.04)	(.04)	(.07)	(.05)
Child aged 10–14	−.06	−.04	−.1	−.10*
	(.04)	(.04)	(.07)	(.05)
Child aged 15–19	−.12**	−.14**	−.22**	−.21**
	(.05)	(.04)	(.09)	(.06)
Household size	.06**	.06**		
	(.01)	(.01)		
Household head age	.00*	.00	−.01	−.01**
	(.00)	(.00)	(.00)	(.00)
Household head gender (female = 1)	−.09*	−.10**	−.20	−.08
	(.05)	(.05)	(.12)	(.08)
Household head year of schooling	−.01**	−.02**	−.01	−.01*
	(.00)	(.00)	(.01)	(.01)
Household head as a minority	−.13**	−.09**	−.18**	−.23**
	(.03)	(.03)	(.05)	(.04)
Poor health	−.12**	−.13**	−.15	−.12
	(.06)	(.05)	(.14)	(.11)
Gini at the county level	−.69**	−.74**	−.81**	−.92**
	(.19)	(.18)	(.32)	(.26)
Observations	2,616	2,616	906	906

NOTE.—For the one-child nuclear family sample, the mother is younger than 40.
* Significant at the 10 percent level.
** Significant at the 5 percent level.

TABLE 8

Rural Household-Level Savings in 2002 (with the Interactive Terms of Income Quartiles and Local Sex Ratio)

	Before Removing Outliers		After Excluding Top/Bottom 5% Observations	
	Son (1)	Daughter (2)	Son (3)	Daughter (4)
Quartile 1 × sex ratio	.90	−.32	.82*	−.18
	(.56)	(.56)	(.49)	(.45)
Quartile 2 × sex ratio	1.02*	−.27	.94*	−.22
	(.55)	(.56)	(.49)	(.45)
Quartile 3 × sex ratio	1.13**	−.21	1.02**	−.13
	(.55)	(.55)	(.49)	(.46)
Quartile 4 × sex ratio	1.19**	−.15	1.10**	−.09
	(.54)	(.55)	(.47)	(.45)
Per capita income (log)	2.35***	2.56***	.84	2.92***
	(.54)	(.46)	(.80)	(.55)
Per capita income (log) squared	−.12***	−.13***	−.04	−.17***
	(.03)	(.03)	(.05)	(.04)
Household head age	.01	−.02	.00	−.07
	(.09)	(.08)	(.07)	(.06)
Child aged 5–9	−.06	−.13	.00	−.12
	(.09)	(.09)	(.07)	(.07)
Child aged 10–14	−.23**	−.15	−.14*	−.05
	(.11)	(.12)	(.08)	(.11)
Child aged 15–19	.00	.00	−.01	.00
	(.01)	(.01)	(.00)	(.01)
Household head gender (female = 1)	−.06	−.36*	−.10	−.11
	(.16)	(.20)	(.11)	(.14)
Household head year of schooling	.01	−.02**	.00	−.01
	(.01)	(.01)	(.01)	(.01)
Household head as a minority	−.17**	−.21**	−.06	−.08
	(.07)	(.08)	(.06)	(.07)
Poor health	.07	−.09	.08	−.03
	(.11)	(.12)	(.07)	(.09)
Gini at the county level	−1.00**	−.32	−.79**	−.53
	(.46)	(.47)	(.36)	(.35)
Adjusted R^2	.31	.54	.21	.33
AIC	898.6	383.6	425	142.1
Observations	580	326	466	259

* Significant at the 10 percent level.
** Significant at the 5 percent level.
*** Significant at the 1 percent level.

Household Savings in Urban Areas

We now turn to urban household savings. Ex ante, the relative strength of the savings response to a rise in the sex ratio between urban and rural samples is ambiguous. First, the more educated segment of the urban population could be more mobile (e.g., college students from other regions may stay on to work in the city of their college after graduation). This would make the local sex ratio statistics inferred from the population census less accurate in describing the true sex ratio in the local marriage market. The noise in the sex ratio measure could induce a downward bias in the estimated coefficient. Second, since urban residents have a higher income than rural residents on average, the marriage challenge generated by a sex ratio imbalance might be solved by importing brides from adjacent rural areas. However, mobility for most urban residents is still limited. (As noted earlier, according to the 2000 population census, most urban residents live and work in the same cities where they were born.) Furthermore, the practice of importing brides from rural areas is not widespread. As noted earlier, during the 5-year period 1995–2000, only 4 percent of marriage-age people changed their place of residency due to marriage.

There are reasons for urban residents to react more strongly to a given rise in the sex ratio. In particular, because the housing market is organized differently between urban and rural areas, a given rise in the sex ratio may bid up housing prices more in urban areas (a possibility that we will check later). Since parents of a son are often expected to help out with the cost of purchasing an apartment for the newlyweds, this would translate into greater pressure to raise their savings rates. Because parents of a daughter (and indeed all other households) also need to buy an apartment, they may have to raise their own savings rate (cutting down nonhousing consumption) in response to a rise in the local sex ratio, especially if the benefit of greater savings by men mainly accrues to women instead of their parents.

In the first two columns of table 9, we contrast the savings behavior between households with a son and those with a daughter. The coefficients on the sex ratio are positive and significant for both types of households. In fact, the point estimate in the second column is bigger than in the first column (1.85 vs. 1.54), although the difference is not statistically significant. Because of concerns for noise in the data, we attempt to remove possible outliers through a number of filters, similar to what we do with the rural sample. In columns 3–4, we exclude those households whose reported annual income or expenditure is less than 3,000 yuan (any annual income or expenditure below the threshold

TABLE 9

Urban Household-Level Savings for Three-Person Households in 2002

	Full Sample		Subsample That Removes the Following Possible Outliers					
			Income or Expenditure < 3,000 Yuan		Bottom and Top 5%		Bottom and Top 5% and No Explicit Marriage Status	
	Son (1)	Daughter (2)	Son (3)	Daughter (4)	Son (5)	Daughter (6)	Son (7)	Daughter (8)
Local sex ratio (county level)	1.54**	1.85**	1.16**	1.07**	.74**	.65**	.98**	.47
	(.29)	(.33)	(.30)	(.37)	(.24)	(.26)	(.31)	(.32)
Per capita income (log)	.10	.69	1.60	.80	.45	.46	1.38**	.37
	(.73)	(.53)	(1.02)	(.85)	(.62)	(.43)	(.45)	(.50)
Per capita income squared (log)	.00	−.03	−.07	−.03	−.02	−.02	−.07**	−.01
	(.04)	(.03)	(.06)	(.05)	(.04)	(.02)	(.03)	(.03)
Child aged 5–9	−.01	−.05	.01	−.04	−.02	−.05*	−.01	.00
	(.04)	(.03)	(.04)	(.03)	(.03)	(.03)	(.04)	(.04)
Child aged 10–14	−.01	−.03	−.01	−.01	−.02	.01	−.02	.04
	(.04)	(.04)	(.05)	(.04)	(.03)	(.03)	(.05)	(.05)
Child aged 15–19	−.19**	−.17**	−.17**	−.14*	−.11**	−.05	−.11	−.01
	(.06)	(.07)	(.07)	(.08)	(.05)	(.05)	(.08)	(.07)
Household head age	.00	.00	.01	.00	.00	.00	.00	.00
	(.01)	(.00)	(.01)	(.00)	(.01)	(.00)	(.00)	(.00)

Household head gender (female = 1)	−.05**	−.08**	−.01	−.05*	−.03*	−.02	−.02	.00
	(.03)	(.03)	(.03)	(.03)	(.02)	(.02)	(.03)	(.03)
Household head year of schooling	−.01**	−.01*	−.01	−.01*	−.01**	.00	−.01**	.00
	(.00)	(.00)	(.00)	(.00)	(.00)	(.00)	(.00)	(.00)
Household head as a minority	.00	.06	.01	.05	−.07**	.04	−.03	−.07
	(.04)	(.06)	(.03)	(.06)	(.02)	(.04)	(.03)	(.06)
Poor health	−.06	−.03	−.01	−.04	−.03	.01	−.02	.00
	(.05)	(.05)	(.04)	(.05)	(.03)	(.03)	(.05)	(.04)
Own a house	.09**	.03	.09**	.03	.08**	.05*	.09**	.08**
	(.03)	(.03)	(.03)	(.03)	(.02)	(.03)	(.02)	(.04)
Gini at the county level	.35	−.05	.69**	−.21	.33	−.18	.41	−.33
	(.30)	(.34)	(.29)	(.37)	(.24)	(.25)	(.28)	(.30)
Adjusted R^2	.11	.14	.16	.14	.07	.06	.08	.07
AIC	418	489.7	235.9	310.9	−81.1	−2.7	−78.7	−21.8
Observations	769	766	604	605	691	688	384	399

NOTE.—The savings rate is defined as log(income/consumption). To maximize comparability, we restrict the sample to those three-person nuclear families with both parents still alive, mother age younger than 40, and one child. For the first six regressions, a family member is defined as a child if younger than 20 years old, while for the last two columns, those unmarried and registered as a child of household head are counted as children. "Poor health" is a dummy that takes the value of one if a household has at least one member with disability or extreme bad health. The sex ratio at the county level is calculated by the authors for the age cohort of 0–9 from the China population census 1990 (who became 12–21 in 2002). Robust standard errors are in parentheses.

* Significant at the 10 percent level.

** Significant at the 5 percent level.

seems implausibly low).[7] This results in a reversal of the relative size of the two coefficients. That is, the sex ratio has a larger effect on savings by son-families than by daughter-families, although the difference is still not statistically significant.

Similar to the rural household sample, an uncommon one-time expenditure such as a big hospital bill could make a household's savings rate appear unusually low. Conversely, an uncommon one-time income such as a lottery win could make the savings rate appear unusually high. Neither is representative of how much a household intends to save under normal circumstances. In columns 5–6, we exclude the top and bottom 5 percent of households in terms of their savings rate. In columns 7–8, we exclude families with no explicit information on the marital status of their children (in addition to removing the top and bottom 5 percent of households in terms of their savings rate). In both cases, the coefficient on the local sex ratio is greater for son-families than for daughter-families. In fact, in column 8, the coefficient for daughter-families is no longer statistically different from zero.

In table 10, we report a sequence of additional robustness checks (similar to the set of robustness checks reported in table 6 for the rural sample). First, we extend the sample to include all three-person households with the mother younger than 45 years old (with the caveat that the newly added households may not be true one-child families). Second, we consider both extended families (e.g., those with grandparents) as well as nuclear families. Third, we conduct both median regressions and ordinary least squares. Fourth, we apply four (independent) filters to the expanded sample. In all, there are 20 pairs of new regressions. In all cases, the coefficient on the sex ratio for son-families is always positive and statistically significant. The coefficient for daughter-families is mostly but not always statistically significant. Once a filtering rule to remove outliers is applied, the coefficient is mostly greater for son-families than for daughter-families. We conclude that there is robust evidence that son families save more in cities with higher sex ratios. Daughter-families are also likely to save more in cities with more skewed sex ratios, but the evidence is weaker.

We have also pooled all families with a child into a common sample and include an interaction term between the local sex ratio and a dummy for son families (similar to table 7 for the rural sample). We do not include a table of the regression results to save space. Generally speaking, the coefficient on the sex ratio is always positive and significant, but the interaction term is not significant. In other words, we

[7] The nominal threshold is higher in the urban sample than in the rural sample because the cost of living is also higher.

TABLE 10
ROBUST CHECKS ON URBAN HOUSEHOLD-LEVEL REGRESSIONS FOR THREE-PERSON HOUSEHOLDS IN 2002

			Subsample That Drops the Following Observations							
	Full Sample		Income or Expenditure < 3,000 Yuan		Bottom and Top 1% Savers		Bottom and Top 5% Savers		Bottom and Top 5% Savers and No Explicit Marriage Status for Child	
	Son (1)	Daughter (2)	Son (3)	Daughter (4)	Son (5)	Daughter (6)	Son (7)	Daughter (8)	Son (9)	Daughter (10)
	Mother Age Younger than 45 (Nuclear Family)									
OLS	1.32***	1.53***	1.07***	.98***	1.31***	1.18***	.65***	.59***	.84***	.35
	(.26)	(.25)	(.26)	(.27)	(.24)	(.23)	(.20)	(.20)	(.24)	(.22)
Median regression	.89***	1.31***	.75**	.56**	1.00***	1.26***	.70**	.78***	.90**	.37
	(.30)	(.27)	(.34)	(.25)	(.28)	(.24)	(.32)	(.24)	(.37)	(.36)
Observations	1,188	1,145	971	931	1,164	1,121	1,068	1,029	675	657
	Mother Age Younger than 45 (Extended Family)									
OLS	1.28***	1.51***	.96***	.89***	1.16***	1.21***	.61***	.57***	.76***	.50**
	(.25)	(.25)	(.25)	(.25)	(.23)	(.22)	(.19)	(.19)	(.23)	(.22)
Median regression	.88***	1.20***	.67***	.57**	.89***	1.10***	.56**	.67***	.72*	.44
	(.29)	(.19)	(.25)	(.25)	(.24)	(.22)	(.25)	(.22)	(.42)	(.37)
Observations	1,347	1,294	1,090	1,041	1,319	1,268	1,211	1,164	761	744

NOTE.—To maximize comparability, we restrict the sample to those three-person nuclear families with both parents still alive, mother age younger than 40, and one child. For the first eight regressions, a family member is defined as a child if younger than 20 years old, while for cols. 9 and 10, those unmarried and registered as a child of household head are counted as children. Only the coefficients for the sex ratio variable are presented in the table.

* Significant at the 10 percent level.
** Significant at the 5 percent level.
*** Significant at the 1 percent level.

cannot formally reject the null that son-families and daughter-families in urban areas react equally to a rise in the local sex ratio.

For the urban sample, we have more information about the employer characteristics of family members. We therefore construct a set of additional proxies to detect a precautionary savings motive. We create an indicator variable for households with no access to public health insurance, one for those with at least one family member who has been laid off, one for those with at least one family member employed in a state-owned company, one for those with at least one family member working in a company that has recently experienced a reorganization (and hence at risk of being laid off), and another one for households with a member working for an employer that has been losing money. In addition, we create a dummy for households that currently rent, rather than own, an apartment. All these variables provide a richer set of descriptions of the vulnerability of a family to income uncertainty. We do not report the table to save space, but generally speaking, the positive relationship between household savings rates and local sex ratios remains unchanged.

Discussion of Alternative Interpretations

Could the local sex ratio be correlated with some omitted or unobserved variable that also affects household savings decisions? We have already controlled for a long list of variables that may reflect life cycle considerations (e.g., age of head of household and age of children) and precautionary savings motives (e.g., dummies for poor health of family members, job loss by family members, employment in the public sector, enrollment in public insurance, or employment in firms that experience losses or reorganization). Nonetheless, there may be certain dimensions of quality of the local social safety net, growth potential, or income uncertainty that may affect household savings decisions but have not been included in our long list of controls. One may imagine that a region with more intrinsic income uncertainty, or a greater local aversion to a given uncertainty, may simultaneously exhibit a higher local sex ratio imbalance and a higher local savings rate. There may be a positive association between household savings and local sex ratios, but it does not reflect a causal relationship from the sex ratio to the savings rate. Can we rule this out?

If the local sex ratio is suspected of reflecting an unobserved location-specific shock, we can rule this out relatively easily. A pure location-specific shock should affect savings by all households in the same region in the same way. But that is not what we find in the rural sample. Instead, only the savings by those rural households with a son react strongly and positively to a rise in the local sex ratio, while savings by households with a daughter do not.

The next possibility is far more challenging: could a sex ratio imbalance reflect something that is both location and household specific? For example, a region may have an unusually high level of income uncertainty that is common to all households, but some households care about this uncertainty more than others. Those households with a stronger aversion to uncertainty may engage in sex-selective abortions more aggressively and save more at the same time. By construction, selection at both household and location level is much harder to rule out since our unit of observation is at the same level.

However, there are good reasons to think that if we focus on households with a single child, such selection is unlikely to be quantitatively significant. Ebenstein (2008) shows that sex ratio imbalance is overwhelmingly a result of sex-selective abortions at higher orders of birth. That is, the sex ratio for firstborn children is close to normal. This is particularly true in rural areas: since a second child is officially permitted if the first child is a girl and since many families exhibit a preference for a balanced sex ratio (one boy and one girl) over having two boys (Ebenstein 2008), there is very little reason to perform sex-selective abortions on the first pregnancies. In contrast, the sex ratio at birth goes up substantially over time for the second-born children and becomes even more skewed for higher-order births. Similarly, Zhu et al. in an article on sex ratio imbalance in China published in the *British Medical Journal* (2009, 1) conclude: “The sex ratio at birth was close to normal for first order births but rose steeply for second order births, especially in rural areas, where it reached 146 (in 2005).” This suggests that the first son (or daughter) is unlikely to result from a sex-selective abortion. Going back to tables 5–6, where we restrict attention to households with only one child, we clearly see that those with a son exhibit a strongly positive elasticity of savings with respect to the local sex ratio, but those with a daughter do not.

Sex Ratios and Housing Values

In explaining the pattern that savings by households with a daughter may also rise with local sex ratios, we have suggested several stories. First, parents of a daughter wish to preserve their daughter’s bargaining power after marriage. This offsets the desire to take advantage of the higher savings rate of their son-in-law. Second, the higher savings from the groom’s family may accrue mostly to the bride and groom (e.g., in the form of a larger house), which would not directly help parents of the bride.[8] In addition, we have suggested a possible spillover channel through the price of housing: if a higher sex ratio leads to a higher cost of housing due to intensified competition by households with a son, then all other households also have to raise their savings in order to afford local houses. This effect is exacerbated in a country with an underdeveloped financial market. We now look for some direct evidence on a connection between local housing values and sex ratios. Because we do not have individual housing information with detailed housing characteristics, we cannot answer this question with hedonic price adjustment. Instead, we report some suggestive evidence by making use of data in the China population census 2000 on average housing value and house size on November 1, 2000 (as specified in the census), across 2,088 rural counties and 671 cities on November 1, 2000. We control for local income, average household size, and age profile of the local population.

The results are reported in table 11. The first four regressions are on the rural sample. Generally, a higher sex ratio is associated with both a larger average house size and a higher housing value. Based on the point estimate in column 4, a 10 basis point increase in the local sex ratio is associated with a higher cost of housing by about 4 percent. As important, the elasticity of the housing value with respect to the local sex ratio is more than twice as big as that for the average house size. This implies that the cost of a housing unit, holding its size constant, is also higher in regions with a more unbalanced sex ratio.

Columns 5–8 of table 11 examine the urban sample. We obtain similar patterns but even bigger point estimates. In particular, based on the estimate in column 8 (0.74), a 10 basis point increase in the local sex ratio is associated with a higher housing cost by 7.4 percent. Furthermore, the elasticity of the housing value with respect to the sex ratio is twice as big as that for average housing size. Therefore, the increase in total housing cost is likely to be evenly split by a 3.7 percent increase in the average housing size and another 3.7 percent increase in the unit cost.

While future research with individual housing data would have to adjust for other determinants of housing size and value, table 11 is consistent with the interpretation that a higher sex ratio leads to more competition for bigger and more expensive houses. Since a house tends to be the single most expensive purchase for most families, households with no son still have to save more in a region with a high sex ratio. (In addition, local norms may induce them to want to buy a bigger house to fit in with their friends and relatives in the same social strata—to keep up with the Zhangs [or Joneses].)

[8] We thank a referee for suggesting this possibility.

TABLE 11
Sex Ratios and Housing Values
Left-Hand-Side Variable = Per Capita Living Space or Average Housing Value (in Log) at the County or City Level in 2000

	County				City			
	Space (1)	Space (2)	Value (3)	Value (4)	Space (5)	Space (6)	Value (7)	Value (8)
Sex ratio for age cohort 10–19 in 2000	.22**	−.02	.54**	.37**	.70**	.37**	1.46**	.74**
	(.09)	(.08)	(.16)	(.13)	(.16)	(.14)	(.32)	(.23)
Per capita GDP in 1999 (log)	−.06	−.04	−1.32**	−.65**	−.35	−.30	−3.24**	−1.78**
	(.15)	(.13)	(.22)	(.21)	(.24)	(.22)	(.46)	(.37)
Per capita GDP in 1999 (log) squared	.01	.01	.11**	.06**	.03*	.02	.21**	.11**
	(.01)	(.01)	(.01)	(.01)	−.014	−.013	−.026	−.021
Household size (log)	−.43**	−.44**	.70**	1.00**	1.50**	1.55**	.77**	.28*
	(.07)	(.07)	(.16)	(.21)	(.12)	(.11)	(.19)	(.16)
Share of population aged 0–19	−4.35**	−2.57**	−4.01**	−4.11**	−3.66**	−2.42**	−1.85**	−.30
	(.32)	(.29)	(.68)	(.70)	(.33)	−.314	(.65)	−.567
Share of population aged 20–59	−2.93**	−1.32**	−3.30**	−2.52**	−2.89**	−1.05**	−2.28**	−.28
	(.30)	(.28)	(.54)	(.57)	(.37)	(.33)	(.68)	(.56)
Province fixed effect		Yes		Yes		Yes		Yes
Adjusted R^2	.41	.64	.33	.57	.39	.69	.43	.70
AIC	31.73	−958.00	2,584.46	1,704.71	−325.82	−718.74	404.35	48.26
Observations	2,088	2,088	2,088	2,088	671	671	671	671

Note.—Per capita living space and housing value are from the China population census 2000. The resident bank deposit and per capita GDP are from various issues of the China County Social and Economics Statistical Yearbooks. The sex ratio for age cohort 0–9 is inferred from the 1990 census (who became age cohort 10–19 in 2000). The share of population aged 0–19 and 25–59 is calculated from the 2000 census.

* Significant at the 10 percent level.

** Significant at the 5 percent level.

C. *Evidence across Provinces*

We have seen that the savings response to a rise in the sex ratio could vary by household types. This raises the question of what the general equilibrium effect is from a rise in the sex ratio. We address this by examining provincial-level data from 1980–2007 for any association between local total savings rates and sex ratios. This exercise is valid because, as documented earlier, mobility across provinces for the purpose of marriage is low. (If mobility were high, then even if our story is correct, we might not find any association between local sex ratios and savings rates.)

Summary statistics for local savings rates, sex ratios, and a few other key variables are provided in table 12. The average sex ratio for age cohort 7–21 across provinces rose from 1.045 in 1990 to 1.136 in 2007. Provinces with extreme values of sex ratios at birth are tabulated in table 13. For example, in 2005, Jiangxi, Shannxi, and Anhui had sex ratios of 1.37, 1.32, and 1.32, respectively. Ignoring migration, reporting errors, and differential mortality rates between girls and boys for the moment, these numbers suggest that more than one out of every six boys born in 2005 in these provinces will end up being single in their adulthood.

Panel Regressions across Provinces

We perform a panel regression that links a location j's savings rate in year t with the sex ratio for the appropriate age cohort in that same location and year, controlling for location fixed effects, year fixed effects, and other factors. To be precise, our specification is the following:

$$\begin{aligned}\text{Savings\_rate}_{j,t} = {} & \beta \text{Sex\_ratio}_{j,t} + X_{j,t}\Gamma_{j,t} \\ & + \text{province fixed effects} + \text{year dummies} + e_{j,t}.\end{aligned}$$

The savings rate is defined as local income (Y) minus consumption (C), divided by income (Y). We cluster the standard errors by province.

Ideally, we would like to know sex ratios for a fixed age cohort in every region and in every year. However, such data are not available as the Chinese population census is carried out only once every few years (in 1982, 1990, and 2000). Moreover, only the 2000 census offers public data for individual age groups at the provincial level. Given these constraints, we make the following shortcut: we focus on the sex ratios for the age cohort 7–21 in all years, inferred from the 2000 population census. For example, for the age cohort 7–21 in 2007, we infer their sex ratio from the age cohort 0–19 in the 2000 census, since the two groups should theoretically be the same. Similarly, for the age cohort

TABLE 12
Summary Statistics for Key Variables in Provincial Panel Regressions

Variables	China	Province Mean	Province Median	Province SD
	1980			
Savings rate	.159	.137	.141	.049
Sex ratio for age cohort 7–21	1.059	1.059	1.059	.038
Sex ratio at birth in 1982	1.083	1.048	1.070	.124
Per capita income (log)	5.331	5.444	5.362	.223
Share of SOEs in total employment	.189	.190	.142	.103
	1990			
Savings rate	.162	.147	.150	.048
Sex ratio for age cohort 7–21	1.045	1.045	1.047	.057
Sex ratio at birth	1.147	1.114	1.117	.029
Per capita income (log)	6.600	6.715	6.684	.252
Share of SOEs in total employment	.162	.185	.150	.100
	2000			
Savings rate	.262	.274	.258	.076
Sex ratio for age cohort 7–21	1.079	1.080	1.082	.048
Sex ratio at birth	1.199	1.180	1.160	.080
Per capita income (log)	7.868	8.087	8.046	.343
Share of SOEs in total employment	.114	.131	.116	.063
Share of labor force enrolled in social security	.191	.174	.144	.107
	2007			
Savings rate	.302	.310	.304	.056
Sex ratio for age cohort 7–21	1.136	1.136	1.130	.041
Sex ratio at birth in 2005	1.200	1.200	1.200	.077
Per capita income (log)	8.743	9.028	8.898	.337
Share of SOEs in total employment	.082	.086	.070	.035
Share of labor force enrolled in social security	.256	.295	.257	.174

Note.—Savings rate is defined as log(income/consumption). The sex ratios for age cohort 7–21 are inferred from the 2000 population census. For example, the cohort 7–21 in 2007 was the cohort 0–14 in the 2000 census, since the two groups should theoretically be the same. The sex ratios at birth in 1982, 1990, and 2000 at the national level are published figures from China population censuses. Since the disaggregate sex ratios at birth in 1980 and 1990 are not publicly available, we use sex ratios for cohort 20 and 10 years old, respectively, from the 2000 census to approximate them. The sex ratios at birth in 2005 are from Zhu et al. (2009), which are calculated based on *China 1% Population Survey 2005*. Zhu et al. report sex ratios at birth for urban, town, and rural areas. To keep consistency with early years in the table, we combine town and rural areas as rural based on the weights provided in table 1 of the paper. SOE = state-owned enterprise.

TABLE 13
Top and Bottom Five Provinces in Terms of Sex Ratios at Birth

	Highest Ratios		Lowest Ratios	
	Province	Ratio	Province	Ratio
1982:				
1	Anhui	1.11	Tibet	.99
2	Guangxi	1.11	Qinghai	1.02
3	Guangdong	1.1	Xinjiang	1.04
4	Henan	1.1	Yunnan	1.04
5	Shandong	1.09	Ningxia	1.05
1990:				
1	Guangxi	1.22	Guizhou	1
2	Zhejiang	1.18	Tibet	1
3	Henan	1.17	Xinjiang	1.04
4	Shandong	1.16	Qinghai	1.04
5	Jiangsu	1.15	Shanghai	1.05
2000:				
1	Hainan	1.37	Tibet	1.03
2	Guangdong	1.31	Xinjiang	1.06
3	Anhui	1.29	Guizhou	1.08
4	Hubei	1.29	Ningxia	1.09
5	Guangxi	1.27	Inner Mongolia	1.09
2005:				
1	Jiangxi	1.37	Tibet	1.05
2	Shaanxi	1.32	Liaoning	1.09
3	Anhui	1.32	Jilin	1.09
4	Hunan	1.28	Xinjiang	1.09
5	Guizhou	1.28	Heilongjiang	1.10

Source.—Sex ratios in 1990 and 2000 are from the China population censuses 1990 and 2000. The 1982 values are inferred from the 1990 census. The figures in 2005 are from Zhu et al. (2009), which are calculated based on *China 1% Population Survey 2005*.

7–21 in 1990, we infer their statistics from the cohort 17–31 in the 2000 census; and so on.

A caveat with this method is that the actual sex ratio is likely to be different from the inferred one for all years other than 2000. In particular, because the mortality rates for boys and young men are generally slightly higher than those for girls and young women, we may underestimate the true sex ratios for years before 2000 and overestimate them for years after 2000. However, under the assumption that measurement errors are common across all regions in any given year (but may vary from year to year), we can eliminate the effect of measurement errors by including year fixed effects in regressions.

In column 1 of table 14, we report regression results with only sex ratio, log income, and proportions of the local population in the age brackets of 0–19 and 20–59, respectively, as the regressors (plus province and year fixed effects). The effect of local income on local savings rates is positive: a 1 percent increase in local income is associated with a higher local savings rate by 0.20 percentage points. The coefficient on local sex ratio is 0.28 and statistically different from zero at the 5 percent level. In other words, the local savings rates tend to be higher in regions with a more unbalanced sex ratio.

The age profile of local populations produces some puzzling patterns relative to the prediction of the life cycle hypothesis. The share of working age population (the age cohort of 20–59) has a negative coefficient. This means that old-age households and households with children save more than do households in between. We note that this is consistent with the notion that parents save more when they have children and that old people save more either because they have a strong bequest motive or because they realize their financial need in old age is greater. In any case, we note that similar patterns are documented in Chamon and Prasad (2010).

Because men and women may have different savings rates and different income levels, one might worry that income inequality could affect local savings rates directly and that a sex ratio imbalance is simply a proxy for earnings inequality. For a subset of provinces and years, we can measure local income inequality directly by the Gini coefficient.[9] In column 2 of table 14, we add the local Gini coefficient as an additional control. The coefficient is not statistically different from zero (and the point estimate is negative). With this much-reduced sample, the coefficient on the local sex ratio is still positive and significant. The point estimate jumps to 0.58.

Since the skewed sex ratio comes in part from the strict version of the family planning policy and since the family planning policy also produces a lower fertility rate, one wonders if the sex ratio is simply a proxy for the fertility rate. In column 3 of table 14, we include the local fertility rate in the same premarital age cohort as an independent variable. It turns out that the new regressor is not statistically significant. In contrast, the local sex ratio is still positive and highly significant. In column 4 of table 14, we include local life expectancy as an additional regressor to account for the possibility that households save more when they expect to live longer.[10] The coefficient on the new regressor is positive but not statistically significant.

The existing literature has hypothesized that rising income and job

[9] The Gini coefficients are obtained from Ravallion and Chen (2007) and are available for 29 provinces in 1988, 1990, and 1993 and 28 provinces in 1996 and 1999 for the urban sample; and 27 provinces in 1988 and 28 provinces in all other years for the rural sample. The Gini coefficients at the province level are approximated by weighted average of rural and urban coefficients.

[10] Life expectancy at the provincial level is only available in three census years (1982, 1990, and 2000).

TABLE 14
SEX RATIOS AND SAVINGS RATES ACROSS PROVINCES: PANEL REGRESSION, 1980–2007
LEFT-HAND-SIDE VARIABLE = $(Y - C)/Y$

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)
Sex ratio for age cohort 7–21	.28**	.58**	.73**	.28**	.32**	.24**			
	(.05)	(.18)	(.15)	(.05)	(.06)	(.07)			
Sex ratio for age cohort 7–14							.19**		.17**
							(.04)		(.04)
Sex ratio for age cohort 15–21								.17**	.14**
								(.04)	(.04)
Per capita income (log)	.20**	.29**	.31**	.20**	.20**	.19**	.19**	.20**	.20**
	(.02)	(.04)	(.04)	(.02)	(.02)	(.02)	(.02)	(.02)	(.02)
Share of population aged 0–19	.01	.09	.08	0	.04	−.03	.01	−.01	.04
	(.06)	(.21)	(.21)	(.06)	(.08)	(.09)	(.07)	(.07)	(.08)
Share of population aged 20–59	−.24**	.22	.27	−.25**	−.23*	−.26*	−.24*	−.30**	−.2
	(.11)	(.43)	(.29)	(.11)	(.13)	(.15)	(.13)	(.12)	(.13)
Gini coefficient		−.08							
		(.25)							

Average birth rate for age cohort 7–21			.04						
			(.23)						
Life expectancy (log)				1.29					
				(.97)					
Share of SOE employment in total labor force					−.10**	−.03	−.06*	−.03	−.10**
					(.04)	(.04)	(.04)	(.04)	(.04)
Share of labor force enrolled in social security					−.03*	−.03*	−.04*	−.02	−.03
					(.02)	(.02)	(.02)	(.02)	(.02)
Growth rate from $t-5$ to t						−.03*			
						(.02)			
Province fixed effects?	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Year fixed effects?	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Adjusted R^2	.78	.82	.72	.79	.78	.8	.78	.78	.78
AIC	−3,324	−596	−457	−3,270	−3,326	−2,840	−3,311	−3,296	−3,322
Observations	844	141	112	829	844	710	844	844	844

NOTE.—Robust standard errors are in parentheses. Sex ratios are inferred from the 2000 population census. SOE = state-owned enterprise.
* Significant at the 10 percent level.
** Significant at the 5 percent level.

uncertainty is what motivates the Chinese to save more. In column 5, we use the proportion of the local labor force that works for state-owned firms or government agencies as a proxy for degree of job security. We also include the share of local labor force enrolled in social security as a proxy for the extent of the local social safety net. Under the precautionary savings hypothesis, the savings rate should decline with better job security or better social security coverage. Both coefficients are negative and statistically significant. These patterns support the precautionary saving motive. In any case, the coefficient on sex ratio is still positive and significant. In fact, the point estimate (0.32) becomes larger. This suggests that the competitive saving motive has independent explanatory power from the precautionary saving motive.

Carroll, Overland, and Weil (2000) point out that, assuming habit formation in consumption, the savings rate should be higher in a fast-growing economy since consumption growth may not catch up with income growth immediately. In column 6, we add the local growth rate of the preceding 5 years. It turns out to have a negative coefficient, which is consistent with a more standard model of consumption smoothing without habit formation. In any case, the coefficient on local sex ratio is still positive (0.24) and statistically significant. To summarize the discussion so far, across the first six columns in table 14, the positive association between a high local sex ratio and a high local savings rate is robust and statistically significant.

To see whether parents start to save for their children's marriage when their children are relatively young or when they are already very close to the marriage age, we break down the sex ratio into two components: one for the age cohort 7–14 and the other for the cohort 15–21. When we enter the two sex ratios one by one (cols. 7 and 8), each has a positive and significant coefficient. When we enter them simultaneously (col. 9), both are significant. The difference between the two point estimates ($0.17 - 0.14 = 0.03$) is not significant (less than the standard deviation of either estimate). The sum of the two coefficients is equal to the coefficient on the sex ratio for the wider age cohort 7–21 in column 5.[11] We interpret these patterns as evidence that parents start to save for their children when they are relatively young, and the sex ratio imbalance in the different segments of the premarital age cohort matters approximately equally.

Instrumental Variable Regressions

Local sex ratios for the premarital age cohort are predetermined by parental decisions—whether to undertake a sex-selective abortion be-

[11] We obtain similar results if the sex ratio is broken down into those for four age cohorts, 7–10, 11–13, 14–17, and 18–21.

fore a child is born—taken many years prior to the corresponding savings variables in the regressions. This helps to justify the panel regressions specification. Nonetheless, as we have noted earlier, local sex ratios may be measured with errors since they are inferred from the 2000 population census. In this case, the estimated elasticity of local savings with respect to the sex ratio may be biased downward.

A solution to both measurement errors and possible endogeneity is to employ an instrumental variable approach, which we turn to now. A key determinant of the sex ratio imbalance is the strict family planning policy introduced in the early 1980s.[12] We explore two determinants of local sex ratios that are unlikely to be affected by local savings rates and for which we can get data. First, while the goals of family planning are national, the enforcement is local. Eberstein (2008) proposes to use regional variations in monetary penalties for violating the family planning policy as an instrument for the local sex ratio. Using data collected by Scharping (2003) and extending them to more recent years, Eberstein focuses on two dimensions of penalties: (*a*) a monetary penalty for violation of the policy, expressed as a percent of annual income in the province, and (*b*) a dummy for the existence of an extra penalty for having higher order unsanctioned births (e.g., for having a third child in a one-child zone or for having a fourth child in a two-child zone). These two variables are part of our set of instruments, too.

Second, while the Han ethnic group faces a strict birth quota, the rest of the population (i.e., the 50-some ethnic minority groups) do not face or face much less stringent quotas. (The government allowed the exemption, possibly to avoid criticism for using the family planning policy to marginalize minority groups.) Since non-Han Chinese are not uniformly distributed across space, this variation offers one more possible instrument (Bulte, Heerink, and Zhang 2011). In tabulations not reported here, we show evidence that interracial marriages involving Han and most of the 55 other ethnic groups are common in China. (The primary exceptions are marriages between Han and Uigers and between Han and Tibetans.)[13]

[12] China's family planning policy, commonly known as the "one-child policy," has many nuances. Since 1979, the central government has stipulated that Han families in urban areas should normally have only one child (with some exceptions). The Han is the majority ethnic group, accounting for about 95 percent of the Chinese population at the time the policy was introduced. Families in rural areas can generally have a second child if the first one is a daughter (this is referred to as the "1.5 children policy" by Eberstein [2008]). Ethnic minority (i.e., non-Han) groups are generally exempted from birth quotas. Non-Han groups account for a relatively significant share of local populations in Xinjiang, Yunnan, Ganshu, Guizhou, Inner Mongolia, and Tibet.

[13] In principle, variations in the cost of sex-screening technology, especially the use of an Ultrasound B machine, and the economic status of women (such as that documented in Qian [2008]) could also be candidates for instrumental variables. Unfortunately, we do not have the relevant data. Note, however, that for the validity of the instrumental variable

TABLE 15
First-Stage Regressions for Savings Rate at the Provincial Level
Left-Hand-Side Variable = Sex Ratio

	R1 (1)	R2 (2)	R3 (3)	R4 (4)	R5 (5)
Penalty for violating family planning policy (% of local yearly income)	.010**			.010**	.010**
	(.002)			(.00)	(.00)
Dummy for extra penalty for higher order births		.008**		.009**	.003
		(.002)		(.002)	(.002)
Share of local population exempted from birth quotas			−.283**		−.278**
			(.04)		(.04)
Per capita income (log)	−.034**	−.034**	−.050**	−.035**	−.050**
	(.01)	(.01)	(.01)	(.01)	(.01)
Share of population aged 0–19	−.286**	−.274**	−.158**	−.273**	−.156**
	(.07)	(.07)	(.08)	(.07)	(.08)
Share of population aged 20–59	−.430**	−.396**	−.240*	−.421**	−.266**
	(.12)	(.13)	(.13)	(.12)	(.12)
Share of SOE employment in total labor force	.422**	.384**	.343**	.416**	.375**
	(.05)	(.05)	(.05)	(.05)	(.05)
Share of labor force enrolled in social security	.007	.001	.009	.009	.017
	(.02)	(.02)	(.02)	(.02)	(.02)
Province fixed effects?	Yes	Yes	Yes	Yes	Yes
Year fixed effects?	Yes	Yes	Yes	Yes	Yes
Adjusted R^2	.72	.71	.73	.72	.74
AIC	−3,795	−3,777	−3,839	−3,799	−3,862
Observations	844	844	844	844	844

Note.—Robust standard errors are in parentheses. Sex ratios are inferred from the 2000 population census.
* Significant at the 10 percent level.
** Significant at the 5 percent level.

In table 15, we report the first-stage regressions for the instrumental variables (IV) regressions that link local sex ratios to their determinants, with both financial penalties and minority shares lagged by 14 years (to match the median birth year for the age cohort 7–21). In columns 1–3, we enter the three potential instrumental variables one by one. Each variable is statistically significant and has the expected sign. First, more severe financial penalties—set more than a decade earlier—are indeed associated with a more unbalanced sex ratio. Second, the proportion

regressions, we do not need a complete list of the determinants of the local sex ratio in the first stage.

of people not subject to birth quotas is negatively related to the local sex ratio imbalance.

In column 4 of table 15 we enter the two financial penalty variables jointly. The coefficients on both variables are positive and statistically significant, consistent with the interpretation that more stringent enforcement of the family planning policy has led to more aggressive sex-selective abortions and, as a result, a more skewed sex ratio. In column 5, we add the proportion of the local population that is legally exempted from the family planning policy. The new variable produces the (expected) negative sign and is statistically significant. With all three potential instruments in the same regression, the dummy for the existence of extra financial penalties for higher order births is no longer significant, but the "regular" financial penalty variable is still positive and significant. The adjusted R-squared for the first-stage regressions is in excess of 70 percent.

Results from the two-stage least squares (2SLS) estimation for local savings rates are reported in table 16. In column 1, the local sex ratio is instrumented by all three variables described above. The Durbin-Wu-Hausman test rejects the null that the bias induced by either measurement errors or endogeneity is not serious (with a p-value of .03). This indicates that the estimates from a (correctly specified) 2SLS procedure are more reliable than the straightforward panel estimates. The Hansen's J-statistic for overidentification fails to reject the null that the instrumental variables are valid (with a p-value of .11). With the 2SLS procedure, the local sex ratio continues to have a positive coefficient that is statistically different from zero. The point estimate, 0.61, is larger than five of the first six points estimates in table 14. This suggests that an attenuation bias induced by measurement errors might have outweighed any endogeneity bias (if the latter exists) in the original panel regressions.

The coefficients on the control variables are sensible. For example, regions with more people working in the state sector save less. An increase in the share of state sector employment by one percentage point is associated with a reduction in the local savings rate by 0.21 percentage points.[14] This is consistent with a precautionary savings motive. As before, we do not find supportive evidence for the life cycle hypothesis.

In column 2 of table 16, we replicate the 2SLS estimation but exclude the share of minorities in the local population from the set of instruments. Even though the Hansen's J-test in the first column suggests that all three instrumental variables, including the minority share, are valid in a purely statistical sense, we act conservatively here and allow for the

[14] The coefficient on the sex ratio variable is largely the same if we drop the state-owned enterprise (SOE) employment share variable.

TABLE 16
2SLS Regressions for Savings at the Provincial Level
Left-Hand-Side Variable = $(Y - C)/Y$

	IV Set		
	All Three IVs (1)	Two Financial Penalty Variables (2)	One Financial Penalty (3)
Local sex ratio	.61**	1.08**	1.17**
	(.18)	(.36)	(.47)
Per capita income (log)	.21**	.22**	.23**
	(.02)	(.03)	(.03)
Share of population aged 0–19	.12	.25*	.28
	(.09)	(.15)	(.18)
Share of population aged 20–59	−.11	.08	.12
	(.18)	(.28)	(.32)
Share of SOE employment in total labor force	−.21**	−.40**	−.43**
	(.08)	(.15)	(.20)
Share of labor force enrolled in social security	−.03	−.03	−.03
	(.02)	(.03)	(.03)
Province fixed effects?	Yes	Yes	Yes
Year fixed effects?	Yes	Yes	Yes
Adjusted R^2	.77	.71	.69
AIC	−3,224	−3,020	−2,965
Durbin-Wu-Hausman test	.03	.00	.00
Hansen's J-statistic for overidentification	.11	.59	
Observations	844	844	844

Note.—Robust standard errors are in parentheses. Sex ratios are inferred from the 2000 population census. The instruments used in the three regressions correspond to those presented in table 15.
* Significant at the 10 percent level.
** Significant at the 5 percent level.

possibility that non-Han Chinese have a different desired rate of savings and thus should not be in the set of instruments. In this case, the Hansen's J-statistic still fails to reject the hypothesis that the two instrumental variables are uncorrelated with the error term (with a p-value of .59). Hence, in a statistical sense, the instrumental variables are valid. In any case, the (instrumented) local sex ratio continues to exhibit a positive and statistically significant coefficient, with the point estimate becoming even larger (1.08). An increase in the sex ratio for the age cohort 7–21 from 1.05 in 1990 to 1.14 in 2007 would now lead to a rise in the local savings rate by 9.72 percentage points ($= 0.09 \times 1.08 \times 100$), accounting for 60 percent of the actual increase in the Chinese savings rate during this period ($= 9.72/[(0.31 - 0.15) \times 100]$).

In column 3 of table 16, we employ a single instrumental variable—

the financial penalty for violating the birth quota that was set 14 years earlier. We confirm the results in the earlier regressions: a higher local sex ratio raises the local savings rate. The point estimate (1.17) is even bigger than before. An increase in the sex ratio from the mean value across the provinces in 1990 to the mean value in 2007 would now lead to a rise in the savings rate by 10.53 percentage points ($= 0.09 \times 1.17 \times 100$), or about 66 percent of the actual observed increase in the mean savings rate across provinces during this period.

To summarize, the cross-province analysis provides a way to estimate the general equilibrium effect of a higher sex ratio on the savings rate. Because the Durbin-Wu-Hausman tests suggest that there are serious biases from measurement errors in the sex ratio statistics (or possible endogeneity), we put more trust in the 2SLS estimations. Among the three IV estimates, our preferred estimate is in the second column of table 16, because we do not need to be concerned with possibly different savings habit of ethnic minorities and because we can conduct an over-identification test with more than one instrument. By that estimate, an increase in the sex ratio from 1990 to 2007 can explain about 60 percent of the actual increase in the savings rate.

Additional Evidence: Sex Ratios and Bank Deposits

Since saving is unspent income, it may reflect both passive and active decisions. It is useful to take a look at bank deposits, which reflect an active household decision. This is especially true in rural areas where cash income often takes the form of physical currency notes, which need to be taken to a bank branch in person for deposits. We are able to compute actual bank deposits per person—or more precisely, local bank deposits in 2002 divided by local population in 2000—for 1,972 rural counties. In columns 1–3 of table 17, we regress per capita bank deposits by rural county on the local sex ratio and other controls. The first regression considers a linear log income term and the age structure of the local population; column 2 adds a quadratic log income term. The third regression adds province fixed effects. The coefficients on the key regressor, sex ratio imbalance, are positive and statistically significant across all three specifications. The elasticity estimates for bank deposits with respect to the sex ratio are large. Using column 3 as an example, the point estimate is 1.20.

It is noteworthy that the coefficients on the two variables describing the age structure of the local population are now supportive of the life cycle hypothesis. In particular, regions with a bigger share of working age population exhibit a higher bank deposit per capita.

In columns 4–5 of table 17, we regress the change over time in log bank deposits on the change in the sex ratio and other controls from

TABLE 17
Sex Ratios and Bank Deposits
Left-Hand-Side Variable = Log (Bank Deposit)

	Deposit in 2002 (1)	Deposit in 2002 (2)	Deposit in 2002 (3)	Growth in Deposit 1992–2002 (4)	Growth in Deposit 1992–2002 (5)
Sex ratio for age cohort 12–21 in 2002 (or change in the sex ratio in cols. 4 and 5)	2.54**	2.67**	1.20**	.43*	.40*
	(.89)	(.92)	(.49)	(.27)	(.22)
GDP/capita in 1999 (log) (or change from 1994 to 1999 in cols. 4 and 5)	.62**	−1.72**	−1.52**	.08**	.08**
	(.04)	(.54)	(.58)	(.02)	(.02)
GDP/capita in 1999 (log) squared		.14**	.12**		
		(.03)	(.04)		
Share of population aged 0–19	7.12**	7.21**	10.62**	1.12	.97
	(1.17)	(1.17)	(1.19)	(.75)	(.85)
Share of population aged 20–59	17.56**	17.64**	22.77**	.63	.55
	(1.25)	(1.25)	(1.21)	(.66)	(.73)
Provincial fixed effects			Yes		Yes
Adjusted R^2	.38	.39	.485	.02	.08
AIC	5,576	5,561	5,255	2,542	2,441
Observations	1,972	1,972	1,972	1,875	1,875

Note.—The per capita residential bank deposit and per capita GDP are from the China County Social and Economics Statistical Yearbooks (CNBS). For the first four regressions, the sex ratio is inferred from the age cohort 0–9 in the 1990 population census who aged to be 12–21 in 2002. The shares in population for age cohorts 0–19 and 20–59 are derived from the 2000 census. For the last two regressions, the sex ratio variable is defined as the change in sex ratio from age cohort 0–19 to age cohort 10–19 from the 1990 census, while the share of population variables refer to their changes from 1990 to 2000 inferred from the 1990 and 2000 censuses.

* Significant at the 10 percent level.

** Significant at the 5 percent level.

1992 to 2002. Due to missing values, the sample is smaller. The coefficient on the change in sex ratio is smaller (0.40) but continues to be positive and statistically significant.

IV. Conclusions

This paper proposes a competitive saving motive to explain China's high household savings rate. The trigger for competitive savings is the country's rising sex ratio imbalance. Due to intensified competition in the marriage market, households with a son ratchet up their savings rates in hopes of improving their son's odds of finding a wife. Families with a daughter may not reduce their savings rate, because a desire to avoid an erosion of bargaining power by their daughter after marriage may offset a desire to free ride on a future son-in-law's savings. The high savings rate by households with a son may also spill over to other households through higher housing prices. (Of course, in general equilibrium, elevated savings rates are futile because the aggregate number of unmarried men is not changed by individual savings decisions.)

Household data provide a way to test the hypothesis. Households with a son are found to save more in regions with a more skewed sex ratio, holding constant other household features. Importantly, having a son per se is not necessarily associated with greater household savings per se (as parents of a daughter may have to save more if a son is more likely than a daughter to provide for his or her aging parents). It takes a combination of having a son and facing a scarcity of women in the local marriage market for these families to raise their savings rates. This is exactly as predicted by the competitive saving motive. Interestingly, households with a daughter in rural areas do not reduce their savings in response to a rise in the local sex ratio. There is some evidence that daughter-families in the urban sample also have a higher savings rate in cities with a more skewed sex ratio.

We provide direct evidence that housing sizes and prices tend to be higher in regions with a higher sex ratio. This housing value channel is much stronger in urban areas than in rural areas. This lends further credence to the idea of a spillover effect.

Since savings responses to a given change in the sex ratio can vary by household types, it is useful to estimate the general equilibrium effects from a panel data set across Chinese provinces. There is clear evidence that local savings rates tend to be higher in regions with more unbalanced sex ratios. To go from correlation to causality, we implement an instrumental variable approach by exploring regional variations in the financial penalties for violating official birth quotas and in the proportion of the local population that is legally exempted from the family planning policy. This approach enhances confidence in the interpretation that a higher sex ratio has caused households to raise their savings rates. Based on our preferred IV estimate, the increase in the sex ratio from 1990 to 2007 can explain about 60 percent of the actual increase in the household savings rate during this period. Other explanations such as a precautionary savings motive may collectively explain the remaining 40 percent of the increase in savings.

Accumulating more wealth is not the only way for men or households with a son to compete in the marriage market. Parents may also invest more in the education of their sons and push them to work harder. There may also be spillover from a boy's education to a girl's education.

Furthermore, in response to a rise in the sex ratio, men or parents with a son may be more prepared to take on higher-risk and higher-return activities (Wei and Zhang 2011). A careful investigation of these possibilities is left for future research.

Finally, while the paper focuses on evidence from China, the basic mechanism can be applied to other countries. Indeed, other economies known to have a strong sex ratio imbalance include Korea, Taiwan, Hong Kong, Singapore, and India. These economies also happen to have high savings rates. We leave a systematic examination of international data to future research.

Appendix

Data Sources

Sex ratios.—Regional data on sex ratios are derived from the China population census in either 1990 or 2000. The method of inference is explained in the text. Summary statistics for selected years (1980, 1990, 2000, and 2007) are reported in table 12 (and see National Bureau of Statistics, China 1988). The data on national level sex ratios at birth (for cohorts born later than 1988), used to generate figure 1, are from Coale and Banister (1994, table 3). Earlier data (for the period of 1988–1993) are from Gu and Roy (1995).

Savings rates.—Following Chamon and Prasad (2010), the savings rate is defined as log (per capita disposable income/per capita living expenditure). The per capita disposable income and living expenditure in cities from 1985 to 1998 and the per capita rural net income and living expenditure for the period of 1978–1998 are from *Comprehensive Statistical Data and Materials on 50 Years of New China* (CNBS). The data for later years are from China Statistical Yearbooks, various issues.

The residential bank deposit and per capita GDP at the county level in 2002 are from China County Social and Economic Statistical Yearbooks (CNBS).

The living space per household and average housing values (purchase value or construction cost) at the county/city level are from the China population census 2000.

Other data used in household level regressions.—The rural and urban household survey data sets are obtained from the Chinese Household Income Project (2002), available from the Web site of the Inter-University Consortium for Political and Social Research (http://www.icpsr.umich.edu/cocoon/ICPSR/STUDY/21741.xml).

References

Angrist, Josh. 2002. "How Do Sex Ratios Affect Marriage and Labor Markets? Evidence from America's Second Generation." *Q.J.E.* 117 (August): 997–1038.

Bates, Thomas W., Kathleen M. Kahle, and René M. Stulz. 2009. "Why Do U.S. Firms Hold So Much More Cash than They Used To?" *J. Finance* 64 (5): 1985–2021.

Bayoumi, Tam, Hui Tong, and Shang-Jin Wei. 2010. "The Chinese Corporate Savings Puzzle: A Firm-Level Cross-Country Perspective." Working Paper no. 16432, NBER, Cambridge, MA.

Bhaskar, V., and Ed Hopkins. 2011. "Marriage as a Rat Race: Noisy Pre-marital Investments with Assortative Matching." Working paper, University College London and University of Edinburgh.

Blanchard, Olivier J., and Francesco Giavazzi. 2005, "Rebalancing Growth in China: A Three-Handed Approach." Working Paper no. 05-32, Massachusetts Inst. Tech., Dept. Econ.

Brown, Philip H., Erwin Bulte, and Xiaobo Zhang. 2011. "Positional Spending and Status Seeking in Rural China." *J. Development Econ.* 96 (1): 139–49.

Bulte, Erwin, Nico Heerink, and Xiaobo Zhang. 2011. "China's One-Child Policy and 'the Mystery of Missing Women': Ethnic Minorities and Male-Biased Sex Ratios." *Oxford Bull. Econ. and Statis.* 73 (1): 21–39.

Carroll, Christopher D., Jody R. Overland, and David N. Weil. 2000. "Saving and Growth with Habit Formation." *A.E.R.* 90 (3): 341–55.

Chamon, Marcos, and Eswar Prasad. 2010. "Why Are Saving Rates of Urban Households in China Rising?" *American Econ. J.—Macroeconomics* 2 (1): 93–130.

Chiappori, Pierre-Andre, Bernard Fortin, and Guy Lacroix. 2002. "Marriage Market, Divorce Legislation, and Household Labor Supply." *J.P.E.* 110 (1): 37–72.

Coale, Ansley J., and Judith Banister. 1994. "Five Decades of Missing Females in China." *Demography* 31 (3): 459–79.

Cole, Harold L., George J. Mailath, and Andrew Postlewaite. 1992. "Social Norms, Savings Behavior, and Growth." *J.P.E.* 100 (6): 1092–1125.

Das Gupta, Monica. 2005. "Explaining Asia's 'Missing Women': A New Look at the Data." *Population and Development Rev.* 31 (3): 529–35.

Du, Qingyuan, and Shang-Jin Wei. 2010. "A Sexually Unbalanced Model of Current Account Imbalances." Working Paper no. 16000, NBER, Cambridge, MA.

Ebenstein, Avraham. 2008. "Estimating a Dynamic Model of Sex Selection in China." Working paper, Harvard Univ.

Edlund, Lena. 2001. "Dear Son—Expensive Daughter: Do Scarce Women Pay to Marry?" Manuscript, Columbia Univ.

Edlund, Lena, Hongbin Li, Junjian Yi, and Junsen Zhang. 2008. "More Men, More Crime: Evidence from China's One-Child Policy." IZA Discussion Paper no. 3214, Inst. Study Labor, Bonn.

Gale, D., and L. S. Shapley. 1962. "College Admissions and the Stability of Marriage." *American Math. Monthly* 69:9–14.

Greenspan, Alan. 2009. "The Fed Didn't Cause the Housing Bubble: Any New Regulations Should Help Direct Savings toward Productive Investments." *Wall Street J.*, March 11.

Gu, Baochang, and Krishna Roy. 1995. "Sex Ratio at Birth in China, with Reference to Other Areas in East Asia: What We Know." *Asia-Pacific Population J.* 10 (3): 17–42.

Guilmoto, Christophe Z. 2007. "Sex-Ratio Imbalance in Asia: Trends, Consequences and Policy Responses." Working paper, LPED/IRD, Paris. http://www.unfpa.org/gender/docs/studies/summaries/regional_analysis.pdf.

He, Xinhua, and Yongfu Cao. 2007. "Understanding High Saving Rate in China." *China and World Econ.* 15 (1): 1–13.

Hopkins, Ed. 2009. "Job Market Signaling of Relative Position, or Becker Married to Spence." Working paper, Univ. Edinburgh.

Hopkins, Ed, and Tatiana Kornienko. 2004. "Running to Keep in the Same Place: Consumer Choice as a Game of Status." *A.E.R.* 94 (4): 1085–1107.

———. 2009. "Which Inequality? The Inequality of Endowments versus the Inequality of Rewards." Working paper, Univ. Edinburgh.

Horioka, Charles Y., and Junmin Wan. 2006. "The Determinants of Household Saving in China: A Dynamic Panel Analysis of Provincial Data." Working Paper no. 12723 (November), NBER, Cambridge, MA.

International Monetary Fund. 2005. "Awash with Cash: Why Are Corporate Savings So High?" In *World Economic Outlook*, chap. 4, 135–59. Washington, DC: IMF.

Kuijs, Louis. 2006. "How Would China's Saving and Investment Evolve?" Policy Research Working Paper no. 3958, World Bank, Washington, DC.

Li, Shuzhuo. 2007. "Imbalanced Sex Ratio at Birth and Comprehensive Interventions in China." Paper presented at the fourth Asia Pacific Conference on Reproductive and Sex Health and Rights, Hyderabad, India, October 29–31.

Mei Fong. 2009. "It's Cold Cash, Not Cold Feet, Motivating Runaway Brides in China." *Wall Street Journal*, June 5.

Modigliani, Franco. 1970. "The Life Cycle Hypothesis of Saving and Intercountry Differences in the Saving Ratio." In *Growth and Trade: Essays in Honor of Sir Roy Harrod*, edited by W. A. Elits, M. F. Scott, and J. N. Wolfe. Oxford: Oxford Univ. Press.

Modigliani, Franco, and Shi Larry Cao. 2004. "The Chinese Saving Puzzle and the Life Cycle Hypothesis." *J. Econ. Literature* 42:145–70.

National Bureau of Statistics, China. 1988. *China Population Statistical Material Compilation, 1949–1985*. Beijing: China Finance and Economic Press.

Peters, Michael, and Aloysius Siow. 2002. "Competing Premarital Investments." *J.P.E.* 110 (3): 592–608.

Qian, Nancy. 2008. "Missing Women and the Price of Tea in China: The Effect of Sex-Specific Earnings on Sex Imbalance." *Q.J.E.* 123 (3): 1251–85.

Qian, Yingyi. 1998. "Urban and Rural Household Saving in China." *IMF Staff Papers* 35 (4): 592–627.

Ravallion, Martin, and Shaohua Chen. 2007. "China's (Uneven) Progress against Poverty." *J. Development Econ.* 82 (1): 1–42.

Scharping, Thomas. 2003. *Birth Control in China, 1949–2000.* New York: Routledge Curzon.

Siow, Aloysius. 1998. "Differential Fecundity, Markets, and Gender Roles." *J.P.E.* 106 (2): 334–54.

Wei, Shang-Jin, and Xiaobo Zhang. 2011. "Sex Ratios, Entrepreneurship, and Economic Growth in the People's Republic of China." Working Paper no. 16800, NBER, Cambridge, MA.

Zhang, J., and W. Chan. 1999. "Dowry and Wife's Welfare: A Theoretical and Empirical Analysis." *J.P.E.* 107 (4): 786–808.

Zhu, Wei Xing, Li Lu, and Therese Hesketh. 2009. "China's Excess Males, Sex Selective Abortion, and One Child Policy: Analysis of Data from 2005 National Intercensus Survey." *British Medical J.* 2009 (338): b1211.