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Mating Competition and the Pursuit of Bigger and More Costly Homes

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Abstract

If housing is a "positional good" that enhances one's status in the marriage market, then competition for marriage partners might motivate people to pursue a bigger and more expensive house/apartment beyond its direct consumption (and financial investment) value. To test the empirical validity of the hypothesis, we explore regional variations in both the relative shortage of brides and housing market characteristics across China. First, we verify that house (apartment) ownership does generally enhance a young person's marriage prospect. Second, we show that houses (apartments) tend to be systematically bigger and more expensive in regions with a greater sex ratio imbalance, beyond what can be explained by local income and other local population characteristics. Third, when using local implementation of the family planning policy as instruments for the local sex ratio, the effects are still significant, suggesting a causal relationship from higher sex ratio imbalances to bigger and more expensive houses.

Key words: housing price, price-rental ratio, positional good, excess men, missing women

JEL codes: F4, O1, O53

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1. Introduction

Housing is sometimes referred to as a "positional good," meaning that the owner derives utility from comparing its value with the values of the houses owned by members of his comparison group (Robert Frank, 2007; Marsh, 2011). This positional good may be especially important in the marriage market. That is, the value of a home belonging to the family of a young man relative to those of other young men may be a key determinant of how attractive he is considered to be relative to his competitors in the eyes of women (and vice versa). For example, in a survey of mothers with young daughters by Shanghai Daily in March 2010, 80% of the mothers indicate that they would object to their daughters marrying a young man who does not own an apartment (house). In other words, those young men with home ownership are considered more attractive than those without. Presumably even among those owning an apartment, those with a more expensive unit are also considered relatively more attractive.

This implies that the intensity of competition in the marriage market could have consequences for housing value and housing size. The stronger the biological desire to have a marriage partner, the stronger the impact of the mating competition for housing market equilibrium. The literature on positional goods has a long tradition (Veblen, 1898; Hirsch, 1977; and Frank, 1985). The notion that concerns for one's (or one's children's) relative status in the marriage market could affect the desire for home ownership and the types of homes being built in equilibrium is therefore not revolutionary. Yet, such a notion is not universally accepted as part of the standard economics of the housing market. One important reason is that marriage market competition is not easily quantifiable, which makes it difficult to formally confirm or reject such a hypothesis. Another important reason is that average housing prices in a region tend to be correlated with other regional attributes, such as better schools and lower crime. In the United States, a rise in housing prices translates into more property tax revenue, and therefore more funding for local public schools and local police force. This makes it difficult to disentangle the pursuit of expensive housing purely because it enhances status from the pursuit of expensive housing in order to obtain other functionally useful attributes that happen to be correlated with housing prices.

In this paper, we empirically investigate this hypothesis by exploring regional variations in the ratio of marrying age men to women in China. This is made possible by the high and varying sex ratios across different parts of the country in recent years. Left to nature, the sex ratio at birth should be in the neighborhood of 106 boys per 100 girls. Starting from mid-1980s, however, increasing availability of ultrasound B machines and a strict family planning policy have led to progressively more aggressive sex selective abortions in China, which in turn have generated progressively more skewed sex ratios. The national average sex ratio at birth rose to about 115 boys/girls in 2000 and about 120 in 2005 (Li, 2007; Zhu, Lu, and Hesketh, 2009). This implies that, when the cohort born in 2005 grows up, roughly one out of every six men may not be able to find a bride[1]. The excess males at the age of 25 and below are estimated to be on the order of 30 million, which are greater than the entire female population in Canada. There are regional variations in the sex ratio, partly due to uneven enforcement of the family planning policy. The most skewed sex ratios at the province level are on the order of 130 boys per 100 girls.

It is important to note that China does not have property taxes (at least not until 2011), and higher housing prices are not mechanically associated with better school quality. Similarly, if demand for expensive homes is found to be correlated with higher local sex ratios, it is unlikely because both are correlated with lower crime rates. If anything, higher sex ratios may be correlated with higher crime rates (Edlund, Li, Yi, Zhang, 2007).

The rising cost of buying an apartment (or house) is a dominant topic in Chinese internet chat rooms and is considered by the Chinese government as a potential source of political instability. The ratio of median housing price to median monthly household income is over 80 (see Table 1). In comparison, the ratio in the United States even before the 2008-2009 crisis is about 50 (?). The rise in housing price cannot be explained by rising demand for living space alone, since demand factors should push up both the rent

[1] The ratio of men to women in the marriage market may differ from the sex ratio at birth. On one hand, because boys and young men have a slightly higher mortality rate than girls and young women, the sex ratio at age 20 would be somewhat better than the ratio at birth. On the other hand, because the cohort size tends to be progressively smaller over time as a result of the strict family planning policy and because husbands tend to be a few years older than their wives, the ratio of young men to young women in the marriage market is more skewed than the sex ratio at the age of 20. These two factors offset each other to some extent. In any case, mathematically speaking, the rising sex ratio at birth over time must imply increasing difficulty for young men to find a wife.

and the housing price but the housing price has risen much faster than rent. By now, the price to rental ratio in Chinese coastal cities is higher than most other major economies in the world. According to Wei and Allen (2010), while the ratio of the price of a house to the monthly rental of a comparable house (apartment) in major US cities is on the order of 250 in 2010, it is 348 in China in 2005 (see Table 1) and 500 in coastal Chinese cities in 2010.

The combined forces of housing ownership as a positional good for marriage purposes and an intensification of competition in the marriage market can logically produce the outcome of a rising housing cost. As the sex ratio for the marriage-age cohort varies across regions, this gives us an opportunity to check if housing market characteristics (housing value and housing size) also vary across regions in a way that is consistent with the hypothesis.

To preview the results, we find evidence that home ownership by a young man or his parents improves his chance of marriage (or more precisely, reduces the probability that he remains unmarried when he reaches the age bracket of 25-40). The value of the home that a household owns tends to be higher if the household has a son and lives in a region with a higher sex ratio. When looking at regional average characteristics of houses (apartments), we find clear evidence that the average home tends to be more expensive in regions with a more skewed sex ratio, beyond what can be explained by the local average household income and other characteristics of the local population. The increased home value in regions with a strong sex ratio imbalance comes from a combination of people buying larger houses on average and people paying a higher price per square meter.

The rest of the paper is organized in the following way. In Section 2, we develop the argument more fully and connect to related literatures. In Section 3, we provide statistical evidence. Finally, in Section 4, we conclude and discuss possible future research.

2. The connection between the hypothesis and the existing literature

The hypothesis that a higher sex ratio can be an important driver for housing market characteristics is related to four sets of literature: (a) status goods, (b) economics of family, (c) economics of housing market, and (d) causes and consequences of sex ratio imbalance. Each of them is too vast to be referenced comprehensively here. Instead, we selectively discuss some of them, with a view to highlight some insight most relevant for our empirical investigation.

Several theoretical papers have pointed out a connection between concerns for status (one's relative position in a society), the savings rate, and the economic growth rate (Cole, Mailath and Postlewaite, 1992; Cornero and Jeanne, 1999; and Hopkins and Kornienko, 2009). When wealth defines one's status in the marriage market, a greater concern for status may lead to an increase in the growth rate. In principle, concerns for status could also produce the opposite effect on savings and growth. In particular, if status is enhanced by conspicuous consumption, then a greater concern for status can translate into a reduction in savings (Frank, 1985 and 2005). It is interesting to note that, while many papers on the topic of status use competition in the marriage market to illustrate the idea, the sex ratio is always assumed to be balanced. In other words, no explicit comparative statistics are derived in terms of a rise in sex ratio imbalance.[2]

Du and Wei (2010) develop a model that explores the effect of a higher sex ratio on household savings rate. They do not directly assume positional goods – goods whose utility depends on the difference between own consumption and the group average. Instead, by incorporating matching in the marriage market, they endogenously generate the result that savings is a sorting variable in the marriage market. Moreover, they show under general conditions that a rise in the sex ratio not only leads to more savings by men but also a rise in the aggregate savings.

Wei and Zhang (2009) provide the first systematic empirical evidence that higher sex ratios lead to higher savings rates in China. They estimate that about half of the increase in household savings from 1990 to 2007 can be attributed to a rise in the sex

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[2] Edlund (1996) showed that a higher sex ratio imbalance may have a nonlinear impact on women's status and dowry price.

ratio. In other words, the effect of the sex ratio imbalance is economically significant.

The desire to enhance one's relative standing in the marriage market can also induce people to be more entrepreneurial and to work longer and harder. Wei and Zhang (2011) provide evidence supporting this hypothesis. They estimate that the rise in the Chinese sex ratio has contributed 20% to the overall observed growth rates in recent years.

Robert Frank (1985 and 2004) points out that people tend to over-spend on positional goods (such as housing) and consequently under-spend on non-positional goods. The arms race in the consumption of positional goods in general equilibrium would bid up the price of positional goods, potentially causing large and preventable welfare losses.

There is an extensive literature in demography that documents the phenomenon of unbalanced sex ratios in Asia (for example, Gu and Roy, 1995; Guilmoto, 2007; and Li, 2007). Several papers have examined the determinants of sex ratio imbalance (including Das Gupta, 2005; Ebenstein, 2009; Edlund, 2009; Li and Zheng, 2009; and Bulte, Heerink and Zhang, 2011). In an influential paper, Oster (2005) proposes that the prevalence of Hepatitis B is a significant cause of the sex ratio imbalance in Asia. But this conclusion is later shown to be incorrect, including by Lin and Luoh (2008) and Oster, Chen, Yu and Lin (2008). In a paper with a clever instrumental variable approach, Qian (2008) shows that an improvement in the economic status of women tends to reduce the sex ratio imbalance. Her instrument for the economic status of women is the world price of tea, whose production is apparently particularly suitable for women laborers.

3. Statistical Evidence

We organize our evidence in three steps. First, we show that ownership of a house (apartment) improves a young man's chance on the marriage market. Second, we show that the demand for housing varies, based simultaneously on whether a family has a son and whether it lives in a region with a skewed sex ratio. Third, we turn to local general equilibrium and investigate whether the average value of a house (apartment) is higher in regions with a higher sex ratio, holding constant local income level and other characteristics.

Data

We combine data from two principal sources: data on housing are from the 1% survey of the population in China in 2005 (of which we were given access to a 20% random sample); local sex ratios are constructed from the 2000 population census. The survey asks respondents to report the purchase or construction cost and construction area of a house, as well as the year of construction if a household owns a house/apartment. Unfortunately, it does not ask when the household bought the home. If a household does not own a house, it asked for monthly rent. To ensure the year of construction and the year of purchase are the same (or very close), we compute the average values of residential homes that were constructed in 2004 and transacted on residential markets.

The summary statistics for the key variables are reported in Table 1. The sex ratio for the age cohort of 5-19 was 107 males per 100 females in 1985, which was moderately unbalanced, but reached 112 males per 100 females in 2005. Ignoring differential mortality rates between the two sexes and the age difference between husbands and wives in most marriages, the sex ratio in 2005 implies that one out of every nine young men cannot get married, mathematically speaking. In other societies when the cohort size grows over time, that fact that husbands are a few years older than their wives would imply that the prospect for men to find a wife is slightly better than the raw sex ratio. However, since China's age cohort is shrinking over time due to its strict family planning policy, the reverse is true. (Even though men may wish to marry an even younger wife, it is an entirely different matter as to whether younger women would want to marry the surplus old men.) In any case, with a rise in the sex ratio, it must be the case that the collective prospect for young men to find a wife declines.

Sex ratios vary across regions. The standard deviation for the sex ratio in 2005 is 0.070 in rural areas (compared to a mean of 1.122), and it is 0.073 in urban areas (compared to a mean of 1.117). Due to both policy restrictions on internal migration (through the household registration system) and culture, marriages by and large are local affairs. (The 2000 population census indicates that 90% of marriages take place between husbands and wives from the same county.) This means that the competition among men for marriage partners is more intense in regions with a more skewed sex ratio.

In Figure 1, we plot the time series of both sex ratios (for 18 years old) and the national housing price index from 1990 and 2010. Both show a strong upward trend and the two are visibly positively correlated. Of course, this is only suggestive and far from being a proof of anything, since presumably many other variables could also have a similarly upward trend.

In Figure 2, we plot the average ratio of the local home value to the local household income against the local sex ratio across rural areas (the left graph) and across urban areas (the right graph) in 2004. More precisely, for every location, we compute the sex ratio for the age cohort of 5-19 years old. We also compute the average house value for those homes whose construction was completed in 2004 and whose purchases were made in 2004 or 2005. For every basis point of the sex ratio (which may correspond to multiple rural prefectures or cities), we compute the average of the ratio of home value to household income. It is clear that across both rural prefectures and urban areas, there is a strong positive association between the local sex ratio and the ratio of home value to household income. In other words, in regions with a higher sex ratio, the ratio of home value to household income tends to be higher too.

Similarly, in Figure 3, we plot the average ratio of home value to rental rate against the sex ratio across rural prefectures (the left graph) and urban areas (the right graph). There is a strong positive association between these two variables: in regions with a higher sex ratio, the ratio of home value to rent also tends to be higher.

<u>House and honey</u>

A key assumption of our hypothesis is that home ownership as a visible form of wealth raises one's status in the marriage market. To check the validity of this assumption, we go to the population survey data (the 1% population survey in 2005) and look at marital status of adults between 25 and 40 years old. We do a sequence of Probit regressions, separately for men and women, and for the rural and urban populations.

The regression results are reported in Table 2. The left-hand-side variable is the probability a given adult is unmarried. On the right hand side, we link it to a host of individual characteristics. The first key one is a dummy for whether the person or his/her family owns a house (or an apartment). The second key variable is an interaction between that dummy and the log value of the house (apartment). If our assumption is right, we should expect a negative sign on the first regressor, meaning that house ownership reduces the chance of not being married (or improves the chance of a marriage). We should also expect a negative sign on the interaction term, meaning that among those owning a house, those owning a more expensive one are also more likely to be married.

As controls, we include the individual's own income, his/her family income, number of people in the household (say, parents and siblings), age, educational level, ethnic background (whether he/she is an ethnic minority), and health status of the individual. We also control for the presence of a family member who is 65 or older. This is designed to capture the idea that the presence of old people implies a need for physical care, which may discourage a potential suitor. In addition, we control for the local sex ratio. Intuitively, a man living in a region with a greater relative surplus of men would have a harder time finding a wife. Conversely, a woman living in such a region should have an easier time finding a husband.

The first column of Table 2 looks at the marital status of men in the age cohort of 25-40 in rural areas. The coeffients on the first two regressors are both negative and statistically significant. This is consistent with the notion that, other things equal, house ownership improves one's chance to be married. Among house owners, a more valuable home further raises the chance to be married.

Other control variables mostly have sensible signs. Higher own income reduces the chance of being single. Given home ownership, value of the house, and one's own income, family income does not appear to have an additional effect. The negative sign on own age implies that there are more unmarried young men than unmarried middle-aged men. The negative sign on "years of schooling" implies that more educated men are less likely to be single. Men of poorer health are more likely to be unmarried. The positive coefficient on the local sex ratio is consistent with the idea that a greater ratio of men to women reduces the chance for any given man to be married, but the coefficient is not statistically significant.

The second column examines women's marital status in rural areas. The negative sign on "home ownership" means that a woman's marriage prospect is also brightened with home ownership; (This is consistent with a micro survey that suggests men also prefer women who own a home. The smaller coefficient (-1.64) in absolute value than the corresponding coefficient for men in the first column implies that house ownership is more important for men than for women in terms of one's status relative to one's same sex competitors in the marriage market. (According to the 2010 Marriage Market Survey, 71% of women prefer that their husbands-to-be own a house, compared with 48% of men having the same preference on house ownership by their wives-to-be.)

The sign on the interaction term (between house ownership and the log value of the house) is positive and significant. This is puzzling as it literally means that a woman whose family owns a more expensive house is less likely to be married. As we will see later, this particular pattern is not a robust feature of the data. In terms of control variables, it is noteworthy that greater household income (especially a woman's parental income), rather than her own income, improves her chance of marriage. A higher ratio of men to women also improves her chance of marriage.

The next two columns report similar Probit regressions for the urban sample. Home ownership improves both a man and a woman's chance for marriage. Conditional on owning a home, a more expensive home also increases the chance of marriage.

It is important to point out that the Probit regressions simply describe a set of associations. In particular, the negative coefficient on the dummy for home ownership indicates both married men and married women are more likely to own a home than single men and single women. One may be tempted to say that the association could simply arise from the need for space by a married couple. However, the need for space in principle can be separated from the rent versus buy decision. In other words, if a married couple needs more space, they could simply rent a bigger place. In addition, the negative coefficient on the interaction term between home ownership and home value is also telling. This means conditional on owning a home, and holding constant household size (the need for physical space) and other individual and family characteristics, a married man is more likely to own a more expensive home than a single man. In comparison, this is less true for rental units when one compares a married woman to a single woman.

Who wants a big house?

We now turn to the value of a home owned by parents as a joint function of the gender (and the number) of their children and the local sex ratio. We implement a series of Tobit estimation where the dependent variable is the value of the home. The variable is left-censored as the value of owned home is zero for those families that rent a house/apartment.

The key regressor is an interaction term between a dummy for having a son in the family and the local sex ratio. Control variables include family income and size, the household head's age, educational level, gender and ethnicity, presence of severe health problems by some members of the family, the number of children by gender, and age brackets of the children. We include location (city or prefecture) fixed effects. It is well-known that adding fixed effects to non-linear panel models could make the estimates biased and inconsistent. However, Greene (2003) shows through Monte Carlo simulations that slope coefficients in a Tobit model, unlike those in probit and logit models, are unaffected by the 'incidental parameters problem.'(In Appendix A, we also report a set of almost identical Tobit regressions without the fixed effects, and obtain the same qualitative results.)

Under the hypothesis that a family with a son has a greater need to own an expensive home, we expect to see a positive coefficient on the interaction term. Note that if the gender of a child makes a difference for the type of home a family needs but the local sex ratio plays no additional role, this would be captured by the coefficients on the number of sons and the number of daughters in a family, and the coefficient on the interaction between the sex ratio and the dummy for having a son would be zero.

The regression results are reported in the first columns of Table 3. The coefficient on the interaction term is positive and statistically significant for both the rural and the urban sample. This is consistent with our hypothesis. Interestingly, the coefficient on the number of sons is negative for the rural sample and indifferent from zero for the urban sample. This means that having sons per se does not lead a family to buy a more expensive home. Rather, it takes a combination of having sons and living in a region with a skewed boy/girl ratio for a family to go after a more expensive home.

Other regressors have sensible coefficients. For example, a family with a higher income or with more members tends to own a more expensive home.

In columns 3 and 4 of Table 3, we do similar Tobit regressions on the physical size (in square meters) of a home. The qualitative results are similar to the previous table. In particular, it is not surprising that richer and larger families tend to own a large home. It is also interesting that, holding constant the family size, having more sons per se is not associated with a larger home. On the other hand, a combination of having sons and living in a region with a more skewed sex ratio does lead to a larger home.

In the last two columns of Table 3, we look at the level of rent as a placebo test. If a combination of sons and local sex ratio reflects a need for more space unrelated to mating competition, one would expect to see the same sign pattern on the coefficients. In fact, the estimated coefficients on the interaction term are statistically indifferent from zero.

In sum, these results suggest that a combination of the presence of male children and a high local sex ratio motivates parents to buy bigger and more expensive homes. Consistent with our hypothesis, there is no comparable pattern for rental units.

The general equilibrium effect

We now turn to the general equilibrium effects of higher sex ratios on housing market characteristics. In principle, the net effect could be ambiguous. If parents with daughters or women do the opposite things from parents with sons or men (e.g., reduce home ownership or buy a smaller home when the sex ratio goes up), the net effect could be zero. However, the net effect could also be stronger than the partial equilibrium effect from the behavior of men or their parents alone, if mating competition among men induces women to do the same thing on the housing market. This could happen through a "tournament channel." For example, as a higher sex ratio shifts the distribution of home size and value owned by men to the right, the reward for a woman to be matched with the most attractive man becomes bigger too. If a woman owning a larger home is more likely to be matched with a man owning a larger home, women may respond to a higher sex ratio by increasing their home ownership and by pursuing larger and more expensive homes in order to improve their status in the marriage market. In sum, the general equilibrium effect of higher sex ratios on housing market characteristics can only be settled empirically.

Using the 1% population survey in 2005, we examine the relationship between the characteristics of houses bought in 2004 or 2005 by prefecture or city and the local sex ratio. The regression specification is of the following:

$$\text{House}(k) = \beta \,\text{sexratio}(k) + X(k)\,\Gamma + e(k)$$

where house(k) refers to the average value (the average physical size in square meters, or the average price per square meter) of houses newly constructed in 2004[3] in location k, sexratio(k) is the sex ratio for the age cohort of 5-19 in location k (inferred from the 2000 population census), and X(k) is a vector of control variables including the average local household income, and the size and the age structure of local population. β and Γ parameters to be estimated.

We implement the regressions separately for the rural sample (331 prefectures) and the urban sample (259 cities). Table 4 reports the regression results. Not surprisingly, regions with higher household income tend to have more expensive homes. The first column indicates that, on average, higher sex ratios are associated with more expensive homes. This is true in both rural and urban areas. The elasticity of home value to the sex ratio is greater in rural areas, although the difference between the two samples is not statistically significant.

The second column looks at the physical space of homes. The coefficient on the sex ratio is positive and significant for both rural and urban samples. This implies that the average house (apartment) size tends to be bigger in locations with a more skewed sex ratio. The elasticity is bigger in cities than in rural areas. The third column examines the home price per square meter. The results indicate that a higher sex ratio is associated with a higher price in rural areas but not in the urban area.

One could consider the three columns together, and regard column 1 as the sum of columns 2 and 3. (Since house price = house value/size by construction, in terms of the

[3] Since the survey was conducted in October 2005, the number of houses built in 2005 is far less than that in 2004. We focus on value of houses built in 2004. We do not include houses built before 2004 because we do not have the information on year of purchase.

dependent variable, log(house value) = log(house space) + log(price per square meter).) Collectively, the results suggest that in rural areas, a higher sex ratio is simultaneously associated with a bigger house and a higher price per square meter. This gives rise to a more expensive home in a region with a higher sex ratio. Bigger space and higher price contribute roughly equally to higher overall house value.

In urban areas, a higher sex ratio is associated with a bigger physical space but not necessarily a higher price. So the association between more expensive homes and higher sex ratios in 2004 comes almost entirely from the association between home sizes and sex ratios.

<u>Instrumental variable regressions</u>

For our research question, we do not think endogeneity of sex ratios is a serious problem, since we are comparing the values of homes in 2004 with sex ratios of the age cohort born many years earlier. Nonetheless, sex ratios may be measured with errors. For example, in spite of the household registration system, a small amount of migration for marriage purpose adds noise to the local sex ratios as a gauge of the tightness of the local marriage market.

In any case, a strategy to address both the measurement error and the endogeneity problems is to employ an instrumental variable approach. A key determinant of the sex ratio imbalance is a strict family planning policy introduced at the beginning of the 1980s[4]. We explore three determinants of local sex ratios that are unlikely to be affected by the growth of local private firms, and for which we can get data. First, while the goals of family planning are national, the enforcement is local. Ebenstein (2009) proposes to use regional variations in the monetary penalties for violating the birth quotas, originally collected by Scharping (2003), as instruments for the local sex ratio. The idea is that, in regions with stiff penalties, parents may engage in more sex-selective abortions, rather than paying a penalty and having more children. The monetary penalty is often on the

[4] 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 child (with some exceptions). Ethnic Han families in rural areas can have a second child if the first one is a daughter (this is referred to as the "1.5 children policy" by Ebenstein, 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.

order of between one to five times the local average annual household income. In addition, Ebenstein (2008) coded a dummy for the existence of extra fines for violations at higher-order births. For example, an additional penalty may kick in on a family for having the 3rd or 4th child in a one-child zone, or the 4th or 5th child in a two-child zone. Such a non-linear financial penalty scheme was introduced by different local governments in different years (if at all), generating variations across regions and over time. These two monetary penalty variables constitute the first two candidates for our instrumental variables.[5]

The third instrumental variable explores the legal exemptions in the family planning policy. While the policy imposes a strict birth quota on the Han ethnic group (the main ethnic group in the country), the rest of the population (i.e., some 50 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 the minority groups.) As a result, the share of non-Han Chinese in the total population has risen from 6.7% in 1982 to 8.5% in 2000 (Bulte, Heerink, and Zhang, 2011). Non-Han Chinese are not uniformly distributed across space. In regions with relatively more ethnic minorities, marriages between Han and non-Han peoples are not uncommon, reducing the competitive pressure for men in the marriage market (Wei and Zhang, 2009). Therefore, the share of non-Han Chinese in the local population offers another possible instrument.[6]

The first stage regressions for the urban and rural samples are reported in Columns 1 and 2 of Table 5, respectively. The coefficients on the share of the local population not subject to birth quotas are negative and statistically significant in both regressions. This is consistent with the notion that sex selective abortions are less prevalent when birth quotas apply to less people.

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[5] Edlund et al. (2007) conduct some diagnostic checks and conclude that the level of financial penalties is uncorrelated with a region's current economic status. We will perform and report a formal test on whether the proposed instruments and the error term in the second stage regressions are correlated.

[6] In principle, variations in the cost of sex screening technology especially the use of an Ultrasound B machine (as documented by Li and Zheng, 2009), 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, for the validity of the instrumental variable regressions, we do not need a complete list of the determinants of the local sex ratio in the first stage.

The financial penalties for violating birth quotas generate a positive and significant coefficient in both regressions. The dummy for the existence of extra penalties for violations at higher-order births also produces a positive coefficient in both regressions (and significant for the rural sample). These results imply that a more severe penalty for violating legal birth quotas tends to induce parents to more aggressively abort girls, resulting in a higher sex ratio imbalance. In other words, when the penalties are light, many couples with daughters may opt to keep the daughter, pay the penalties, and have another child, rather than abort the female fetus.

The adjusted R^2's are in the range between 0.15-0.27. The F statistics (for the null that all slope parameters are jointly zeros) ranges from 14.2 to 15.0. The Stock-Yogo critical values for the Kleibergen-Paap statistics (for weak instruments) are 11.6 at the 15% level and 19.9 at the 10% level. This means that the three instruments are somewhat weak.

The second stage regressions are reported in Table 6. The Durbin-Wu-Hausman test rejects the null that the 2SLS and OLS estimates are the same in eight out of twelve regressions. Interestingly, in most cases in which the 2SLS and OLS estimates are not the same, the Hansen's J statistics do not reject the null that the instruments and the error term are uncorrelated. The point estimates in Table 6 are generally much larger than their OLS counterparts in Table 4. This suggests that the downward bias in Table 4 generated either by missing regressors or by measurement errors is substantial.

In any case, the IV results suggest that the data patterns from the OLS estimate carry over. In particular, higher sex ratios tend to systematically generate larger and more expensive homes. Both the ratio of average home value to average income and the ratio of average home value to average rent tend to rise with the local sex ratio.

Urban housing prices during 2003-2009

China's urban house prices have increased dramatically in recently years, prompting street protests and repeated government announcements to do something about them. Our data from the 1% population survey do not allow us to examine house prices after 2005. Fortunately, the Chinese Statistical Yearbooks report average house prices in 35 major cities (including all provincial capitals) since 2003. We now examine if the local sex ratio imbalance has any predicting power over the evolution of local housing prices beyond the growth of local income and local population.

Table 7 presents the summary statistics on variables used in the panel regression. Table 8 reports the first-stage regression on the sex ratio variables by three instrument variables — share of minority population, penalty for violating family planning policy, and a dummy variable for extra penalty for higher order births. The three instrument variables all have the expected signs and are mostly significant. The large F-statistic shown in columns 2-4 indicates these instrument variables are not weak instruments when fixed effects are included.

The panel regressions with fixed effects are reported in the top panel of Table 9. Without any fixed effects (Column 1), there is a strong positive association between local housing prices and local sex ratios. This is beyond the effects that richer and more populous cities tend to have more expensive homes. When we add just city fixed effects (Column 2), we see the same strong positive association between sex ratios and home prices. When we add both city and year fixed effects (Column 3), the coefficient on the sex ratio continues to be positive and statistically significant, although the point estimate becomes smaller. As a robustness check, in the last column, we drop three big cities, Beijing, Shanghai, and Shenzhen, where housing price has become extremely expensive in the past several years. The coefficient on the sex ratio variable remains significant. Using the coefficient in the third column (0.89), an increase in the sex ratio by 6 basis points (about the actual increase from 2003 to 2009) is associated with a cumulative increase in the (real) home price by five percentage points over this period, accounting for 30% of the real home price increase in the period.

In the lower panel of Table 9, we instrument the sex ratio by the three variables as shown in Table 8. The strong positive relationship between the sex ratio and the home price survive. In fact, the IV estimates are bigger than the corresponding OLS estimates, consistent with the notion that there may be measurement errors associated with the sex ratio in the OLS estimates. Using the point estimate (1.39) in the third regression with both year and city fixed effects, an increase in the sex ratio by 6 basis points contributes to 48% increase in the home price over 2003-2009 (6*1.39/(0.371-0.198)).

4. Concluding Remarks

House prices in China and some other economies appear to rise too fast relative to the growth of income. Rising need for housing due to urbanization or other factors does not seem to be a complete explanation by itself since the same factors should also push up rental rates, yet the ratio of house price to rent also tends to rise substantially. One possibility is that the increasing competition in the marriage market since the turn of this century, triggered by a rise in the ratio of men to women in the pre-marital age cohort, is another fundamental source of the increases in housing value. Since ownership of a house is a more visible form of wealth than alternative components of wealth, it may be a positional (or status) good in the marriage market.

That mating competition induces people to pursue ever bigger and more expensive homes can be true in all societies even without a sex ratio imbalance. But such a hypothesis is hard to test as it is difficult to measure variations in the intensity of mating competition. In this paper, we explore regional variations in the sex ratio in China and link them to regional variations in housing characteristics (average size, price per square meter, and value). We find robust evidence that housing values vary systematically with local sex ratios. As a placebo test, we find no such pattern for rental rates. Based on the more conservative OLS regression, a rise in the sex ratio from 1.05 to 1.12 (corresponding to the actual rise in the national sex ratio for the age cohort of 5-19) from 2000 to 2005 would contribute to 36 % of the observed rise in the average home value in Chinese cities in the same period. A higher sex ratio explains over 15% of the increase in housing value in rural China from 2000 to 2005. A greater sex ratio imbalance accounts for between 30-48% of the real increase in housing prices in 35 major cities in the period of 2003-2009.

The hypothesis that a significant fraction of the observed rise in home prices is due to mating competition has important policy implications. In particular, the hypothesis suggests that some of the increases in home size and home cost are socially inefficient. People pursue larger and more costly homes and suppress their consumption of non-positional goods with the hope of improving their status in the marriage market. But in the aggregate, the number of men who cannot be married is not altered. If there is social coordination so that every household can cut down demand for housing proportionally, all households could consume more non-positional goods and the marriage market outcome is not affected. Property tax and stamp tax on house transactions in this context are different from a situation in which home ownership is not a positional good or mating competition is not as intense. We leave a thorough investigation of these issues to future research.

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Table 1: Summary Statistics on Sex Ratios, Housing and Other Variables

Variables	China	Rural				Urban			
	Mean	Mean	Median	Std	N	Mean	Median	Std	N
Sex ratio for age cohort 5-19 in 2005	1.120	1.122	1.110	0.070	331	1.117	1.092	0.073	259
Sex ratio for age cohort 5-19 in 2000	1.079	1.102	1.093	0.065	331	1.052	1.051	0.077	259
Sex ratio for age cohort 5-19 in 1985	1.068	1.077	1.077	0.091	331	1.041	1.049	0.099	259
Housing value per unit (1,000RMB) in 2004	77	53	46	46	331	107	86	81	259
Housing size (square meters) in 2004	125	116	113	33	331	137	121	63	259
Housing price per square meter (RMB) in 2004	628	450	398	270	331	856	667	713	259
Household monthly income (RMB) in 2005	871	721	661	267	331	1061	999	372	259
Monthly rent (RMB) per house in 2005	248	206	183	109	331	302	259	171	259
Housing value to monthly income ratio	84	73	70	35	331	98	89	57	259
Housing value to monthly rent ratio	348	308	262	219	331	399	355	247	259
Share of primary age population (20-59) in 2000	0.593	0.570	0.570	0.048	331	0.622	0.628	0.053	259

Note: The sex ratio variable at the prefecture/city level is inferred from China Population Census 2000. All other variables are calculated by authors based on a 20 percent random sample of the China Population 1% Sampling Survey 2005. CPI is from China Statistical Yearbooks.

Table 2: Which Adults (between 25 and 40) Are Likely to Be Unmarried?

	Rural sample		Urban sample	
	Single man	Single woman	Single man	Single woman
Dummy for owning a house/apartment	-0.17**	-0.23**	-0.11**	-0.32**
	(0.04)	(0.05)	(0.05)	(0.06)
Housing value (log)*dummy for owning a house	-0.10**	0.05**	-0.13**	-0.04**
	(0.01)	(0.02)	(0.01)	(0.02)
Individual income (log)	-0.13**	0.09**	-0.04**	0.13**
	(0.01)	(0.01)	(0.01)	(0.01)
Family income (including parents) (log)	0.00	-0.13**	0.02	-0.18**
	(0.02)	(0.01)	(0.03)	(0.02)
Household size	-0.09**	0.05**	-0.02	0.03**
	(0.02)	(0.01)	(0.01)	(0.01)
Age	-0.12**	-0.14**	-0.17**	-0.16**
	(0.00)	(0.00)	(0.00)	(0.00)
Year of schooling	-0.05**	0.04**	0.06**	0.10**
	(0.00)	(0.01)	(0.01)	(0.01)
Minority	0.01	0.25**	0.14**	0.14**
	(0.05)	(0.08)	(0.05)	(0.07)
Poor health	0.84**	1.90**	1.39**	2.38**
	(0.07)	(0.11)	(0.13)	(0.15)
Having a family member at 65 or older	0.57**	0.22**	0.61**	0.42**
	(0.02)	(0.02)	(0.03)	(0.03)
Local sex ratio for the cohort 5-19 in 2005	0.10	-0.46**	-0.03	-0.37**
	(0.21)	(0.22)	(0.15)	(0.13)
Pseudo R square	0.17	0.20	0.24	0.27
N	218457	219109	105340	100189

Note: The sex ratio for the age cohort 5-19 in 2005 is inferred from the China Population Census 2000. The dependent variable is multiplied by 100. The average housing sale values (or construction costs) and ownership as well as other right hand variables are from a 20 percent random sample of the China Population 1% Sampling Survey 2005. Standard errors are clustered at the city/prefecture level. * and ** denote statistically significant at the 10% and 5% levels, respectively.

Table 3: Who Wants an Expensive and Bigger Home and Who Wants to Rent? Tobit Estimation of Housing Values, Size of House, Monthly Rent and Family Characteristics

	Housing value		Housing space		Rent	
	Rural	Urban	Rural	Urban	Rural	Urban
Sex ratio * having a son aged 5-19	0.06**	0.12**	0.06**	0.14**	-0.26	-0.53**
	(0.01)	(0.05)	(0.01)	(0.06)	(0.16)	(0.18)
Household income (log)	0.13**	0.12	-0.18**	-0.22*	3.38**	-0.38*
	(0.01)	(0.11)	(0.01)	(0.12)	(0.14)	(0.22)
Household size	0.09**	0.50**	0.18**	0.64**	-1.66**	-0.23**
	(0.01)	(0.05)	(0.01)	(0.06)	(0.10)	(0.11)
Household head age	-0.00**	-0.01**	0.01**	0.00	-0.06**	0.08**
	(0.00)	(0.01)	(0.00)	(0.01)	(0.01)	(0.01)
Household head year of schooling	0.00	-0.02	-0.05**	-0.08**	0.27**	-0.34**
	(0.00)	(0.02)	(0.00)	(0.02)	(0.03)	(0.03)
Female household head	-0.18**	-0.18**	-0.53**	-0.42**	3.99**	1.02**
	(0.02)	(0.07)	(0.03)	(0.07)	(0.24)	(0.14)
Minority household head	-0.09**	(0.08)	0.02	-0.08	-1.76**	0.20
	(0.03)	(0.12)	(0.03)	(0.13)	(0.35)	(0.31)
Poor health among at least one family member	-0.03*	0.09	0.04**	0.22**	-2.75**	-2.31**
	(0.01)	(0.08)	(0.01)	(0.09)	(0.41)	(0.49)
Having a child 10-14	0.06**	0.24**	0.04**	0.26**	-0.86**	-1.31**
	(0.01)	(0.06)	(0.01)	(0.05)	(0.13)	(0.14)
Having a child 15-19	0.24**	0.34**	0.24**	0.37**	-3.02**	-2.63**
	(0.01)	(0.05)	(0.01)	(0.05)	(0.15)	(0.37)
Number of sons	-0.02**	0.00	-0.03**	0.00	0.66**	0.88**
	(0.01)	(0.05)	(0.01)	(0.06)	(0.15)	(0.21)
City/prefecture fixed effects	Yes	Yes	Yes	Yes	Yes	Yes
Pseudo R square	0.03	0.03	0.03	0.04	0.11	0.05
N	229567	74012	229567	74012	229567	74012

Notes: The dependent variable is the purchased housing price or construction cost (log), the construction area of a house (log) which is owned by the household who live there, and monthly rent (log) for those who rent a house. The sample is restricted to those with a child aged 5-19. The sex ratio for the age cohort 5-19 is inferred from the age cohort 0-14 in the 2000 population census at either the city or the prefecture level. Other data are from a 20 percent random sample of the China 1% Population Survey in 2005. Standard errors are clustered at the city (or prefecture) level. * and ** denote statistically significant at the 10% and 5% levels, respectively.

Table 4: Sex Ratios and Housing Market in 2004

	House value (1,000RMB per unit)	House space (m^2 per unit)	House price (RMB per m^2)	House value / monthly rent	House value / income	Rent (RMB)
Rural areas						
Local sex ratio for age cohort 5-19 in 2005	1.59**	0.82**	0.76**	2.51**	1.59**	-0.92*
(prefecture level)	(0.38)	(0.26)	(0.28)	(0.72)	(0.38)	(0.55)
Household monthly income (log)	0.91**	0.22**	0.69**	0.52**	-0.09	0.39**
	(0.09)	(0.05)	(0.07)	(0.11)	(0.09)	(0.08)
Total population (log)	0.12**	0.07**	0.05**	0.09*	0.12**	0.03
	(0.03)	(0.02)	(0.02)	(0.05)	(0.03)	(0.04)
Share of primary age population (20-59 years old)	0.96**	-0.04	1.00**	2.52**	0.96**	-1.55**
in 2000 (log)	(0.33)	(0.18)	(0.26)	(0.43)	(0.33)	(0.30)
Adj. R-squared	0.42	0.16	0.38	0.23	0.18	0.08
N	331	331	331	331	331	331
Cities						
Local sex ratio for age cohort 5-19 in 2005	1.28**	1.21**	0.07	1.37**	1.28**	-0.09
(city level)	(0.51)	(0.39)	(0.50)	(0.54)	(0.51)	(0.39)
Household monthly income (log)	1.03**	0.33**	0.70**	0.38**	0.03	0.65**
	(0.12)	(0.09)	(0.13)	(0.15)	(0.12)	(0.10)
Total population (log)	0.06	-0.02	0.08*	0.08*	0.06	-0.02
	(0.04)	(0.03)	(0.04)	(0.05)	(0.04)	(0.04)
Share of primary age population (20-59 years old)	-0.22**	-0.03	-0.20**	-0.14**	-0.22**	-0.09**
in 2000 (log)	(0.04)	(0.03)	(0.06)	(0.03)	(0.04)	(0.03)
Adj. R-squared	0.28	0.09	0.15	0.06	0.04	0.14
N	259	259	259	259	259	259

Note: The sex ratio for the age cohort 5-19 in 2005 is inferred from the China Population Census 2000. The total population and share of primary age population are from China Population Census 2000. The average housing sale values (or construction costs) and construction areas are computed based on those houses/apartments built in 2004 from a 20 percent random sample of the China Population 1% Sampling Survey 2005. All the dependent variables are in logarithmic form. Robust standard errors are in parentheses. * and ** denote statistically significant at the 10% and 5% levels, respectively.

Table 5: First Stage Regressions – Instrumenting for Local Sex Ratios

	R1	R2
Share of minority population (log)	-0.81**	-1.56**
	(0.17)	(0.29)
Penalty for violating family planning policy	2.12**	2.41**
	(0.57)	(0.68)
Dummy for extra penalty for higher order births	3.74**	1.88
	(1.46)	(1.91)
Household monthly income (log)	-1.25	2.25
	(1.29)	(1.63)
Total population (log)	1.38**	-1.28**
	(0.52)	(0.49)
Share of primary age population (20-59 years old)	-22.27**	-0.84
in 2000 (log)	(4.30)	(1.44)
Adj. R-squared	0.27	0.15
Kleibergen-Paap rk Wald F statistic	14.23	14.98
N	331	259

Notes: (a) The dependent variable is local sex ratio for age cohort 5-19 in 2005 expressed in percentage, inferred from the 2000 Population Census. (b) The two family planning variables are averaged over the years of the age cohort 5-19. (c) Stock-Yogo weak ID test critical values: 13.91 for 5% maximal IV relative bias and 9.08 for 10% maximal IV relative bias. (d) Robust standard errors are in parentheses. (e) * and ** denote statistically significant at the 10% and 5% levels, respectively.

Table 6: Instrumental Variable Regressions for Housing Characteristics and Sex Ratios (2SLS)

	House value (1,000RMB per unit)	House space (m^2 per unit)	House price (RMB per m^2)	House value / monthly rent	House value / income	Rent (RMB)
Rural areas						
Local sex ratio for age cohort 5-19 in 2005	8.57**	3.91**	4.66**	9.56**	8.57**	-0.99
(prefecture level)	(1.55)	(0.89)	(1.03)	(1.92)	(1.55)	(1.15)
Household monthly income (log)	0.85**	0.19**	0.65**	0.46**	-0.15	0.39**
	(0.11)	(0.07)	(0.07)	(0.14)	(0.11)	(0.09)
Total population (log)	-0.04	0.00	-0.04	-0.08	-0.04	0.03
	(0.06)	(0.03)	(0.04)	(0.08)	(0.06)	(0.04)
Share of primary age population (20-59 years old)	2.70**	0.73**	1.97**	4.27**	2.70**	-1.57**
in 2000 (log)	(0.59)	(0.33)	(0.37)	(0.76)	(0.59)	(0.41)
AIC	592	196.43	337.27	740.43	592	431.35
Durbin-Wu-Hausman test for endogeneity (p-value)	0.00	0.00	0.00	0.00	0.00	0.95
Hansen's J statistic for over-identification (p-value)	0.71	0.07	0.26	0.01	0.71	0.00
N	331	331	331	331	331	331
Cities						
Local sex ratio for age cohort 5-19 in 2005	2.44*	2.81**	-0.36	5.23**	2.44*	-2.79**
(city level)	(1.29)	(0.96)	(1.28)	(1.71)	(1.29)	(1.18)
Household monthly income (log)	0.99**	0.27**	0.72**	0.24	-0.01	0.75**
	(0.13)	(0.11)	(0.14)	(0.17)	(0.13)	(0.13)
Total population (log)	0.07	-0.01	0.08*	0.10**	0.07	-0.03
	(0.04)	(0.03)	(0.04)	(0.05)	(0.04)	(0.04)
Share of primary age population (20-59 years old)	-0.22**	-0.02	-0.20**	-0.12	-0.22**	-0.10**
in 2000 (log)	(0.06)	(0.03)	(0.05)	(0.09)	(0.06)	(0.04)
AIC	463.55	317.84	473.36	598.82	463.55	387.81
Durbin-Wu-Hausman test for endogeneity (p-value)	0.33	0.07	0.72	0.01	0.33	0.00
Hansen's J statistic for over-identification (p-value)	0.03	0.84	0.03	0.54	0.03	0.17
N	259	259	259	259	259	259

Note: Robust standard errors are in parentheses. * and ** denote statistically significant at the 10% and 5% levels, respectively.

Table 7: Summary Statistics on Sex Ratios, Housing Price and Other Variables in 35 Big Cities

	Mean	Median	Std
Housing price per square meter (RMB) for 35 big cities in 2003	2426	2131	1019
Housing price per square meter (RMB) for 35 big cities in 2009	5706	4463	3055
Change in housing price from 2003 to 2009 (log)	0.371		
Change in urban consumer price index from 2003 to 2009	0.198		
Urban sex ratio aged 9-24 at the provincial level in 2003	1.071	1.068	0.032
Urban sex ratio aged 9-24 at the provincial level in 2009	1.132	1.116	0.061
Per capita GDP (RMB) in 2003	22100	13231	31735
Per capita GDP (RMB) in 2009	44644	43161	18103
Population in 2003 (million)	650	602	526
Population in 2009 (million)	773	717	565

Note: The housing price, GDP and population data are from various issues of China Statistical Yearbooks. The sex ratio refers to average urban sex ratio at the provincial level from the 2000 Population Census. The age cohort is chosen so that the data are available from the 2000 population census (anyone younger than 9 in 2009 was not born in 2000 yet). The cohort aged 9-23 in 2009 was aged 0-14 in 2000.

Table 8: First Stage Regressions – Instrumenting for Provincial-Level Sex Ratios

	Full Sample	Full Sample	Full Sample	Excluding three cities
Share of minority population (log)	-0.45**	-15.06**	-14.46**	-11.51**
	(0.23)	(1.17)	(1.24)	(1.91)
Penalty for violating family planning policy	1.98**	3.01**	1.18	2.42**
	(0.51)	(1.06)	(1.08)	(1.03)
Dummy for extra penalty for higher order births	3.93**	10.22**	6.25**	4.74*
	(0.99)	(2.73)	(2.79)	(2.58)
Per capita GDP (log)	-0.62	-0.37	-1.11*	-0.93*
	(0.45)	(0.45)	(0.59)	(0.54)
Population (log)	-0.55	1.37	-0.16	13.55**
	(0.56)	(1.37)	(1.27)	(3.79)
City effects	No	Yes	Yes	Yes
Year effects	No	No	Yes	Yes
Adj. R-squared	0.10	0.76	0.31	0.79
Kleibergen-Paap rk Wald F statistic (weak identification test)	9.13	72.55	50.77	26.88
N	245	245	245	224

Notes: The sex ratio, which is in percentage, refers to average urban sex ratio at the provincial level from the 2000 Population Census. The age cohort is chosen so that the data are available from the 2000 population census (anyone younger than 9 in 2009 was not born in 2000 yet). The cohort aged 9-23 in 2009 was aged 0-14 in 2000.). In the last column, Beijing, Shanghai and Shenzhen are excluded from the sample. The share of minority population and two family planning variables are averaged over the years of the age cohort 9-23. Stock-Yogo weak ID test critical values: 19.93 for 10% maximal IV size and 7.25 for 25% maximal IV size. Robust standard errors are in parentheses. * and ** denote statistically significant at the 10% and 5% levels, respectively.

Table 9: Sex Ratios and Housing Prices in 35 Major Cities (2003-2009)

	Full Sample	Full Sample	Full Sample	Excluding three cities
Panel A: OLS				
Sex ratio for age cohort 9-23	3.02**	2.98**	0.89**	1.34**
	(0.28)	(0.46)	(0.25)	(0.32)
Per capita GDP (log)	0.51**	0.38**	0.00	0.01
	(0.04)	(0.06)	(0.03)	(0.04)
Population (log)	0.19**	0.51**	0.03	-0.17
	(0.02)	(0.08)	(0.05)	(0.18)
City fixed effects	No	Yes	Yes	Yes
Year fixed effects	No	No	Yes	Yes
Adj. R-squared	0.77	0.90	0.96	0.96
Panel B: 2SLS				
Sex ratio for age cohort 9-23	8.99**	4.39**	1.39**	1.77**
	(1.54)	(0.88)	(0.36)	(0.61)
Per capita GDP (log)	0.47**	0.34**	0.01	0.01
	(0.04)	(0.06)	(0.03)	(0.03)
Population (log)	0.21**	0.40**	0.03	-0.24
	(0.04)	(0.10)	(0.05)	(0.18)
City fixed effects	No	Yes	Yes	Yes
Year fixed effects	No	No	Yes	Yes
Adj. R-squared	0.44	0.89	0.96	0.95
Durbin-Wu-Hausman test for endogeneity (p-value)	0.00	0.00	0.06	0.30
Hansen's J statistic for over-identification (p-value)	0.33	0.00	0.81	0.21
N	245	245	245	224

Note: The dependent variable is log(average housing sale price per square meter). In the last column, Beijing, Shanghai and Shenzhen are excluded from the sample. The housing price, GDP and population data are from various issues of China Statistical Yearbooks. The sex ratio refers to average urban sex ratio at the provincial level from the 2000 Population Census. The age cohort is chosen so that the data are available from the 2000 population census (anyone younger than 9 in 2009 was not born in 2000 yet). The cohort aged 9-23 in 2009 was aged 0-14 in 2000. In the 2SLS regressions in Panel B, three instrument variables, share of minority population (log), penalty for violating family planning policy (% of local yearly income) and a dummy for extra penalty for higher order births, are included as instrument variables. Robust standard errors are in parentheses. * and ** denote statistically significant at the 10% and 5% levels, respectively.

Appendix Table A: Housing Characteristics and Family Characteristics (Tobit Without Location Fixed Effects)

	Housing value		Housing space		Rent	
	Rural	Urban	Rural	Urban	Rural	Urban
Local sex ratio for the cohort 5-19 in 2005	1.10**	1.94*	1.36**	1.83*	-3.15	2.15
	(0.34)	(1.10)	(0.30)	(1.18)	(3.89)	(4.21)
Sex ratio * having a son aged 5-19	0.05**	0.12**	0.06**	0.16**	-0.37**	-0.53**
	(0.01)	(0.06)	(0.01)	(0.06)	(0.16)	(0.21)
Household income (log)	0.18**	0.24**	-0.18**	-0.24**	3.98**	0.77**
	(0.01)	(0.11)	(0.01)	(0.10)	(0.15)	(0.24)
Household size	0.09**	0.57**	0.17**	0.75**	-1.77**	-0.79**
	(0.01)	(0.04)	(0.01)	(0.04)	(0.10)	(0.14)
Household head age	-0.01**	0.00	0.01**	0.01*	-0.05**	0.08**
	0.00	(0.01)	0.00	(0.01)	(0.01)	(0.01)
Household head year of schooling	0.00	-0.07**	-0.05**	-0.13**	0.23**	-0.36**
	0.00	(0.02)	0.00	(0.03)	(0.04)	(0.05)
Female household head	-0.16**	-0.19**	-0.54**	-0.45**	4.30**	1.21**
	(0.03)	(0.08)	(0.03)	(0.09)	(0.23)	(0.16)
Minority household head	-0.26**	-0.45**	-0.09**	-0.51**	-1.19**	0.24
	(0.04)	(0.15)	(0.04)	(0.15)	(0.32)	(0.35)
Poor health among at least one family member	(0.02)	0.16*	0.07**	0.28**	-2.99**	-1.85**
	(0.02)	(0.09)	(0.02)	(0.10)	(0.43)	(0.47)
Share of primary age population (20-59) in 2000 (log) at the city level	0.51*	-0.03	0.91**	0.05	0.25	-0.82**
	(0.31)	(0.04)	(0.25)	(0.04)	(3.63)	(0.13)
Having a child 10-14	0.06**	0.24**	0.03**	0.27**	-0.82**	-1.40**
	(0.01)	(0.06)	(0.01)	(0.06)	(0.14)	(0.16)
Having a child 15-19	0.27**	0.30**	0.26**	0.34**	-3.37**	-2.69**
	(0.01)	(0.06)	(0.01)	(0.06)	(0.16)	(0.37)
Number of sons	-0.02*	0.02	-0.04**	0.01	0.83**	0.81**
	(0.01)	(0.06)	(0.01)	(0.06)	(0.16)	(0.27)
Pseudo R square	0.01	0.01	0.01	0.02	0.07	0.02
N	229567	74012	229567	74012	229567	74012

Notes: The dependent variable is the purchased housing price or construction cost (log), the construction area of a house (log) which is owned by the household who live there, and monthly rent (log) for those who rent a house. The sample is restricted to those with a child aged 5-19. The sex ratio for the age cohort 5-19 is inferred from the age cohort 0-14 in the 2000 population census at either the city or the prefecture level. Other data are from a 20 percent random sample of the China 1% Population Survey in 2005. Standard errors are clustered at the city (or prefecture) level. * and ** denote statistically significant at the 10% and 5% levels, respectively.

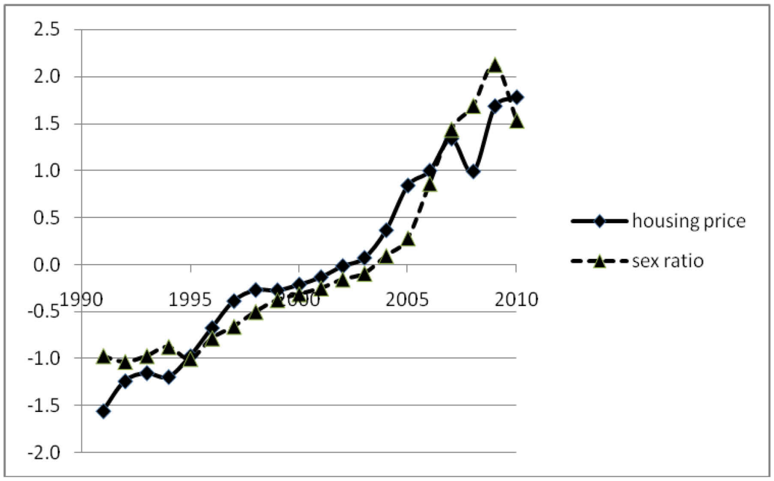

Figure 1: (Standardized) Sex Ratios and Housing Prices

Notes: Sex ratios for 18 years old, inferred from the sex ratio at birth lagged by 18 years. Average house sale prices (per square meter) are deflated by national fixed asset price index, available from various issues of China Statistical Yearbooks. Both variables are rescaled by subtracting the mean and dividing by the standard deviation.

Figure 2: Sex Ratios and House Price-to-Income Ratios

Note: On the horizontal axis is the sex ratio for the age cohort 5-19 inferred from the *China Population Census 2000*. On the vertical axis is the ratio of housing value to household income in 2004, averaged over all cities that had the same value of sex ratio (up to a basis point). The housing value refers to either sale price or construction cost computed from a 20 percent random sample of China 1% Population Survey in 2005.

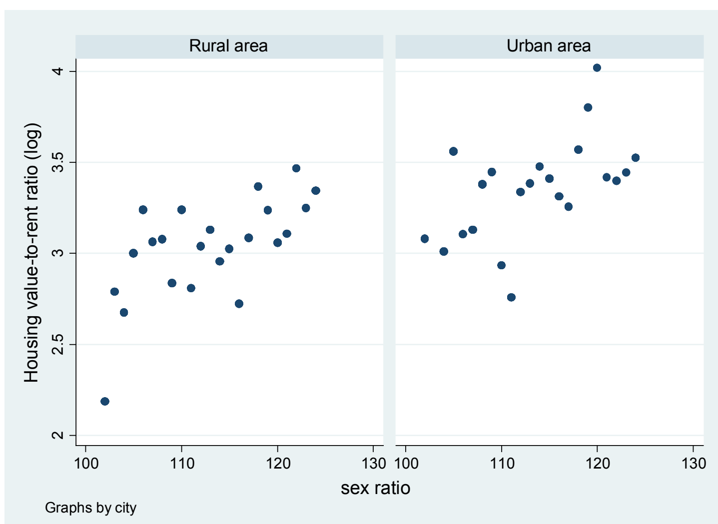

Figure 3: Sex Ratios and House Value-to-Rent Ratios

Note: On the horizontal axis is the sex ratio for the age cohort 5-19 inferred from the *China Population Census 2000*. On the vertical axis is the ratio of housing value in 2004 to annual rent in 2005, averaged over all cities that had the same value of sex ratio (up to a basis point). Both housing value and rent are computed from a 20 percent random sample of China 1% Population Survey in 2005.