NBER WORKING PAPER SERIES

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Working Paper 18000
http://www.nber.org/papers/w18000

NATIONAL BUREAU OF ECONOMIC RESEARCH
1050 Massachusetts Avenue
Cambridge, MA 02138
April 2012

This research is supported by U.S. National Science Foundation grant SES-1024574, which we gratefully acknowledge. The authors would like to thank Kyle Bagwell, V. Bhaskar, Qingyuan Du, Neng Wang and seminar and conference participants at Columbia University, Hong Kong University of Science and Technology, Shanghai Advanced Institute of Finance, University of Hong Kong, University of Southern California, and University of Virginia for helpful discussions, Joy Glazener and Jin Yang for assistance. All errors are the authors' responsibilities. The views expressed herein are those of the authors and do not necessarily reflect the views of the National Bureau of Economic Research.

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NBER Working Paper No. 18000
April 2012
JEL No. G12,R2,R3

ABSTRACT

While in standard housing economics housing is regarded as an asset and a consumption good, we study in this paper the consequences for housing prices if housing is also a status good. More concretely, if a family's housing wealth relative to others is an important marker for relative 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 have to overcome the usual difficulty of not being able to observe the intensity of status competition. Our innovation is to explore regional variations in the sex ratio for the pre-marital age cohort across China, which likely has triggered variations in the intensity of competition in the marriage market. The empirical evidence appears to support this hypothesis. We estimate that due to the status good feature of housing, a rise in the sex ratio accounts for 30-48% of the rise in real urban housing prices in China during 2003-2009.

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1. Introduction

In standard housing economics, housing is regarded as an asset and as a consumption good. However, housing is also sometimes referred to as a "status good," meaning that the owner derives utility from comparing its value with the values of the houses owned by members of his comparison group. This paper studies the equilibrium consequences for housing prices if housing has the status good feature.

This notion of status goods (also known as positional goods) can trace its intellectual heritage to a literature on the economics of social status (Veblen, 1898; Hirsch, 1977; and Frank, 1985). This possibility that housing represents a positional good has been invoked as an explanation for why Americans have chosen to live in ever larger and ever more expensive homes over the last five decades, even though the average household size also has declined over the same period (Robert Frank, 2007; Marsh, 2011)[1]. While this hypothesis is intriguing, we are not aware of any rigorous empirical studies confirming or rejecting it. A conceptual challenge is that if housing is a normal good (without being a "status" good), the demand naturally rises as a household's income rises. This presumably can be controlled for. A more important empirical challenge 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.

The status good feature of housing can be especially important in the marriage market. That is, the value of a home belonging to the family of a young man 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 Chinese mothers with young daughters by *Shanghai Daily* in March 2010, 80% of the mothers indicate that they

[1] For media accounts about the phenomenon of ever larger houses in the United States, see "Living ever larger; how wretched excess became a way of life in Southern California," *Los Angeles Times*, June 9, 2002; "In land of giants, smallest houses larger than ever," *Washington Post*, July 9, 2006; "Parade of homes; houses in Madison are larger than ever," *The Daily (Wisconsin)*, June 14, 2007.

would object to their daughters marrying a man who does not own a home. 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. The underlying reason could be that wealth improves one's perceived attractiveness, other things equal, and housing is a very visible form of wealth and can be observed with more precision than a bank deposit, a stock trading account, and many other forms of wealth.

This implies that the intensity of competition in the marriage market can 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. We note, however, that this notion is not universally accepted as part of the standard economics of the housing market. One important reason is that status competition is not easily quantifiable, which makes it difficult to formally confirm or reject such a hypothesis.

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 the mid-1980s, however, increasing availability of ultrasound B machines and a strict family planning policy have led to 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[2]. The excess males at the age of 25 and below are estimated to be on the order of 30 million, which is almost double the entire female population in Canada. There are regional variations in the sex

[2] 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.

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 useful 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 combined forces of housing ownership as a status good for marriage purposes and an intensification of competition in the marriage market can logically produce the outcome of rising housing prices. 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 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 connect the hypothesis to related literatures. In Section 3, we present a simple model that links the housing price to the sex ratio. In Section 4, we provide statistical evidence. Finally, in Section 5, 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 has not been established in the existing literature. It is indirectly 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 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 the sex ratio imbalance.[3]

Du and Wei (2010) develop a model that explores the effect of a higher sex ratio on the household savings rate and a country's current account. 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 aggregate savings.

Wei and Zhang (2011a) provide the first systematic empirical evidence from China that higher sex ratios lead to higher savings rates. They estimate that about half of the increase in household savings from 1990 to 2007 can be attributed to a rise in the sex 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 (2011b) provide evidence supporting this hypothesis. They estimate that the rise in the Chinese sex ratio has contributed 20% to overall growth rates in recent years.

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,

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

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. A Model

While this is primarily an empirical paper, we construct a model to illustrate the underlying logic for how status competition in the marriage market affects housing prices. We aim to demonstrate two points. First, even with a balanced sex ratio, equilibrium housing prices are higher when status competition is present than when it is not. Second, when the sex ratio is unbalanced, a higher sex ratio can lead to a rise in both the housing price and the ratio of housing price to rent. While the first point is likely to be relevant for all countries in the world, it is hard to test empirically since we do not observe the counterfactual – a world in which people do not care about status. In comparison, the second point can be put to the test if we can find a country with varying sex ratios across regions. Since both points come from a common model, evidence in favor of the second point also lends some credibility to the first point.

We consider an overlapping generation small open-economy model with two genders (men and women) and two goods (a tradable good and non-tradable housing). Both men and women live two periods. Each person is endowed with an initial wealth y . In the first period, agents could purchase a certain quantity of house (which is a non-negative continuous variable) and a certain amount of a composite perishable good, and save the rest of their endowment for the second period. The small open economy assumption means that both the price of the tradable good and the interest rate are set exogenously. We set the price of the tradable good at one. Because housing is non-tradable, its price is determined endogenously.

Each person can enter the marriage market at the beginning of his/her second period. A marriage can take place only between a man and woman in the same cohort. There are two benefits associated with a marriage. First, some of the within-marriage consumption has a public good feature, meaning that a household member's consumption does not exclude the consumption by other household members. We let housing service assume this public good feature. Second, a married person derives additional utility from his/her spouse, which we follow Du and Wei (2010) and label as "emotional utility," but can also be called "love." Everyone is endowed with an exogenous ability to generate some emotional utility to her/his spouse. The exact value of emotional utility is unknown in the first period (although its distribution is common and known across all agents), and is realized and becomes public information when agents enter the marriage market.

For simplicity, we assume zero population growth, and each generation is characterized by an exogenous ratio of men to women $\phi(\geq 1)$. All men are identical *ex ante*, and all women are also identical *ex ante*. Men and women are symmetric in all respects except that the sex ratio may be unbalanced.

Due to zoning requirements, restrictions on building height, and a fixed land supply in a given location, housing supply tends to be much less elastic than say, the supply of electronic appliances or automobiles. Trading houses across space is also very costly, making housing an arch type of a non-tradable good. To capture these features, we make an extreme assumption that the housing supply is completely inelastic.

Agents purchase a continuous amount of house in the first period, and sell it in the second period as there is no reason (within the model) for them to leave a non-zero bequest. The amount of house one buys can be interpreted as physical size. In equilibrium, each young generation obtains housing as owner-occupiers, and rents out a fraction of the house to the old cohort. In standard housing economics, housing plays a dual role, as a consumption good and a means for financial investment. Both functions are present in our model. However, we wish to capture a third feature that is not usually present in standard housing economics. In particular, housing is a status variable in the marriage market. To be more concrete, with competition for mating partners, relative wealth becomes a sorting variable in the marriage market. While both housing wealth and non-housing wealth (say, regular savings) are components of wealth, housing wealth has a distinct advantage, namely, that it can be observed with less noise. To capture this idea, we assume that the value of a person's house can be observed perfectly, whereas the value of his/her non-housing wealth can only be observed with a noise.

Note that we do not directly assume that utility is a function of relative wealth or relative housing value. Instead, through competition in the marriage market, relative wealth will emerge endogenously as an important status variable in equilibrium, and a typical person's indirect utility will turn out to be a function of his/her relative wealth in the same gender-cohort.

We describe our model in five steps. First, we start with the representative agent's optimization problem. Second, we describe how the marriage market works. Third, we derive the equilibrium conditions. Fourth, we solve the model analytically under some special assumptions. Finally, we calibrate the model for a more general setting.

The Agents' Optimization Problems

The representative man and woman make the saving and housing decision in the first period, taking as given the house price and rent, which will be determined endogenously through market clearing conditions. A representative woman chooses the amount of saving and house purchase to maximize her life time value function

$$\max_{s_w, H_w} \ u(c_{1w}) + \beta\delta_w E\left[u(c_{2w}) + \eta_m\right] + \beta(1-\delta_w)u(c_{2ws})$$

where E is the expectation operator, η_m is the emotional utility from her husband, and δ_w is the probability that she will be married. Her second period utility is a weighted average of her utility if she is married and her utility if she is single. The key thing to note is that her probability of marriage, δ_w , is an endogenous variable that depends not only on her house purchase and savings decisions, but also on those of her competitors in the same gender-cohort and those of members of the opposite sex.

She is subject to a set of budget constraints. Her first period budget constraint is

$$z_{1w} + r_1 h_{1w} = y - s_w - p_1 H_w + r_1 H_w \tag{3.1}$$

Where z_{1w} is her consumption of a composite non-housing consumption good, y is her endowment at the beginning of the first period, s_w is her savings in the first period, h_{1w} and H_w are the physical quantity of housing service she consumes and the physical quantity of house she purchases, respectively, and p_1, r_1 denote the house price and rent in the first period, respectively. The left hand side (LHS) of (3.1) is her expenditure (where we normalize the price of numeraire goods to 1), and the right hand side (RHS) of (3.1) is her disposable wealth. By this formulation, we allow her to consume a fraction of the house that she owns and rents out the rest. Home ownership would cost her an expenditure outlay of $p_1 H_w$ but give her a total rental revenue of $r_1 H_w$.

Her second period constraint depends on whether she is single or married.

$$\text{single:} \quad z_{2ws} + r_2 h_{2ws} = Rs_w + p_2(1-d)H_w \tag{3.2}$$

$$\text{married:} \quad z_{2w} + z_{2m} + r_2 h_2 = R(s_w + s_m) + p_2(1-d)(H_w + H_m) \tag{3.3}$$

where d is the house depreciation rate (which is set to zero in the benchmark model), R is the (exogenous) gross interest rate, and p_2, r_2 are the house price, house rent in the second period respectively. In the second period, the representative household sells its house and obtains housing consumption as a tenant. The LHS of (3.3) is the total household consumption expenditure, while the RHS is the household's combined disposable wealth.

Even with these simplifying assumptions, the optimization problem is nontrivial as the representative woman needs to take into account how her savings and house purchase decisions affect her probability of being married in the second period; she has to think through what her life would be like if she remains single versus if she is married; and she has to anticipate what men would do as a function of the economic environment.

To make the analysis tractable, we make two additional simplifying assumptions. First, her period utility function takes the log form $u(c) = \log(c)$. Second, her consumption in each period c_{1w}, c_{2w} is a composite of a numerical good z and housing service h in a Cobb-Douglas form

$$c_{t,w} = \frac{z_{t,w}^{\gamma} h_{t,w}^{1-\gamma}}{\gamma^{\gamma}(1-\gamma)^{1-\gamma}}, \quad t = 1,2$$

A representative man's optimization problem is identical to the representative woman's problem in terms of the structure, except that we would swap the subscripts "*m*" (denoting man) and "*w*" (denoting woman). In other words, we intentionally make men and women completely symmetric. Men and women do not have intrinsic reasons to have different demands for housing.

The Marriage Market[4]

When men and women enter the marriage market at the beginning of the second period, the values of their respective emotional utility are realized. In the marriage market, every woman ranks all men by a combination of men's wealth (with housing wealth being its most visible and significant component) and emotional utility. The weights are determined by her utility function. To be precise, woman i prefers a higher ranked man to a lower ranked one, where the rank on man j is given by $u\left[c_{2,w,i}(j)\right]+\eta_{m,j}$ ($c_{2,w,i}(j)$ denotes woman $i's$ consumption in the second period if married to man j). Symmetrically, man j assigns a rank to woman i based on the sum of utility $u\left[c_{2,m,j}(i)\right]+\eta_{w,i}$.

As stated earlier, we assume that housing wealth H can be observed perfectly but non-housing wealth can only be observed with a noise. That is, saving s is observed as a noisy signal $\tilde{s}$. This turns out to make a big difference for how the two variables enter the first-order conditions. Note that the choice of housing investment and savings decisions made in the first period are meant to demonstrate one's relative attractiveness to potential marriage partners at the beginning of the second period (in addition to consumption smoothing). One is playing a Stackelberg game vis-à-vis potential marriage partners (while playing a Nash game vis-à-vis everyone else in the first period). An important (and somewhat surprising) insight from Bagwell (1995) is that, in a Stackelberg game with imperfect observability of the action of the players, a player's best response to a noisy signal, if we restrict attention to only pure strategies, is to ignore the actual signal value and simply calculate what the opponent would have done if the variable (i.e.,

[4] We use the term "marriage market" informally. Since matching of men and women is not done through prices, marriage is, strictly speaking, a non-market outcome in the model.

savings here) is not observed at all. A similar reasoning works in our model. Namely, an agent will pay attention to the value of housing investment by potential marriage partners (since it can be observed perfectly), but ignores the noisy signal about savings. So in deciding the optimal amount of savings, agents will assume it has no impact on the probability of marriage and simply choose a value that is optimal for his/her consumption smoothing purpose. Any deviations from that would not produce any advantage for him or her in the marriage market. [Since this is primarily an empirical paper, we choose simplicity and focus only on pure strategies.] In comparison, in deciding on the optimal amount of housing investment in the first period, the agent would take into account its effect on his/her probability of marriage in the second period (as well as the standard consumption smoothing considerations.)

We assume that the pairing of men and women follows the Gale-Shapley algorithm, which generates a unique and stable equilibrium of matching (Gale and Shapley, 1962; and Roth and Sotomayor, 1990)[5]. Since all men (women) in the marriage market have identical problems, they make the same savings and house owning decisions in their first period. Following Du and Wei (2011), this implies that there exists a mapping M from η_w to η_m in equilibrium such that

$$1 - F(\eta_w) = \phi\left(1 - F(M(\eta_w))\right) \Leftrightarrow M(\eta_w) = F^{-1}\left(\frac{F(\eta_w)}{\phi} + \frac{\phi - 1}{\phi}\right)$$

We assume that the lowest possible values of emotional utilities of both man and woman are sufficiently small, so they cannot get married unless their realized emotional utilities are above certain threshold values, denoted by $\bar{\eta}_w$ and $\bar{\eta}_m$, which make the opposite sex indifferent between being married and being single.

$$\bar{\eta}_w = \max\left\{u(c_{2ms}) - u(c_{2m}), M^{-1}(\bar{\eta}_m)\right\} \text{ and } \bar{\eta}_m = \max\left\{u(c_{2ws}) - u(c_{2w}), M(\bar{\eta}_w)\right\} \qquad (3.4)$$

The Equilibrium Conditions

In equilibrium, woman i 's total utility in the second period, given all her rivals' and men's strategies, is

[5] One could also model the matching process using search with frictions. This would appear to add considerable complexities with no obvious benefits.

$$\delta_{w,i}\, u(c_{2w,i}) + \int_{\bar{\eta}_w} M(\tilde{\eta}_{w,i}) d\tilde{F}(\tilde{\eta}_{w,i}) + (1-\delta_{w,i}) u(c_{2ws,i})$$

where $\tilde{\eta}_{w,i} = u\left[c_{2m}(i)\right] - u(c_{2m}) + \eta_w$ and the probability that she will be married is:

$$\delta_{w,i} = Prob(u\left[c_{2m}(i)\right] - u(c_{2m}) + \eta_{w,i} > \bar{\eta}_w \left| s_w, s_m, \lambda_w, \lambda_w \right.) = 1 - F\left(\bar{\eta}_w - u\left[c_{2m}(i)\right] + u(c_{2m})\right) \tag{3.5}$$

In the steady state we have $p_1 = p_2,\ r_1 = r_2$. With a bit of algebra (available upon request), we can work out the first order conditions of her optimization problem as:

$$-\frac{1}{y_{1w}} + \beta R \delta_w \bullet \frac{1}{y_2} + \beta R(1-\delta_w) \bullet \frac{1}{y_{2ws}} = 0 \tag{3.6}$$

$$\frac{-p+r}{y_{1w}} + \beta\left(\int_{M(\bar{\eta}_w)} f\left(M^{-1}(x)\right) dx + f(\bar{\eta}_w)[M(\bar{\eta}_w) + u(c_{2w}) - u(c_{2ws})] + \delta_w\right) \bullet \frac{p(1-d)}{y_2} + \beta(1-\delta_w)\frac{p(1-d)}{y_{2ws}} = 0 \tag{3.7}$$

where we use y_{1w} to denote the representative woman's disposable wealth in the first period, y_{2ws}, her disposable wealth if she is single in the second period, and y_2, her family's disposable wealth if she is married in the second period. That is:

$$y_{1w} = y - s_w - pH_w + rH_w,\ y_{2ws} = Rs_w + p(1-d)H_w$$
$$y_2 = R(s_w + s_m) + p(1-d)(H_w + H_m)$$

When it comes to the representative man j, his probability of getting married is:

$$\delta_{m,j} = 1 - F\left(M(\bar{\eta}_w) - u\left[c_{2w}(j)\right] + u(c_{2w})\right)$$

The same process of obtaining (3.6), (3.7) gives the first order conditions for the representative man:

$$-\frac{1}{y_{1m}} + \beta R \delta_m \bullet \frac{1}{y_2} + \beta R(1-\delta_m) \bullet \frac{1}{y_{2ms}} = 0 \tag{3.8}$$

$$\frac{-p+r}{y_{1m}} + \beta\left(\int_{\bar{\eta}_w} f\left(M(y)\right) dy + f(M(\bar{\eta}_w))[\bar{\eta}_w + u(c_{2m}) - u(c_{2ms})] + \delta_m\right) \bullet \frac{p(1-d)}{y_2} + \beta(1-\delta_m)\frac{p(1-d)}{y_{2ms}} = 0 \tag{3.9}$$

where we define y_{1m}, y_{2ms} and y_2 similarly to the corresponding variables for the representative woman.

As stated earlier, in each period, the young cohort buys houses from the old cohort. The young consumes a portion of their owned house and rents out the rest to the old. In each period, the total demand for house ownership must sum up to the house stock:

$$\frac{1}{1+\phi}H_w+\frac{\phi}{1+\phi}H_m=H \tag{3.10}$$

And the total housing consumption demand equals to the house stock in each period:

$$\left(\frac{1}{1+\phi}h_{1w}+\frac{\phi}{1+\phi}h_{1m}\right)+\left(\frac{\delta_w}{1+\phi}h_2+\frac{1-\delta_w}{1+\phi}h_{2ws}+\frac{\phi-\delta_m}{1+\phi}h_{2ms}\right)=H \tag{3.11}$$

The equilibrium is determined by a non-linear system of six equations, (3.6) - (3.11). We are not able to solve this system analytically.

<u>The Analytical Solution of a Special Case</u>

It is useful to inspect the analytical solution to a special case in which the savings rates are exogenously set to be zero and house purchase is the only means for a cohort to transfer a portion of the first period income to the second period. With the assumption of no savings, the first order conditions (Equations (3.6)-(3.9)) are reduced to

$$\frac{-p+r}{y_{1w}}+\beta p\left\{\begin{array}{l}\frac{1}{y_2}\left(\delta_w+\frac{1}{\phi}(1-F(\overline{\eta}_w))+f(\overline{\eta}_w)\overline{\eta}_m\right)+(1-\delta_w)\frac{1}{y_{2ws}}\\+\frac{1}{y_2}f(\overline{\eta}_w)[u(c_{2w})-u(c_{2ws})]\end{array}\right\}=0 \tag{3.12}$$

$$\frac{-p+r}{y_{1m}}+\beta p\left\{\begin{array}{l}\frac{1}{y_2}\left(\delta_m+\frac{1}{\phi}(1-F(\overline{\eta}_m))+f(\overline{\eta}_m)\overline{\eta}_w\right)+(1-\delta_m)\frac{1}{y_{2ms}}\\+\frac{1}{y_2}f(\overline{\eta}_m)[u(c_{2m})-u(c_{2ms})]\end{array}\right\}=0 \tag{3.13}$$

If we further assume that the emotional utility is uniformly distributed, $\eta_w,\eta_m\sim[\eta^{\min},\eta^{\max}]$ and the average value of emotional utility, $E(\eta)$, is large enough (which causes agents to have a strong *ex ante* desire for marriage), then we obtain the following proposition:

Proposition *With the assumptions stated above, as the sex ratio rises, up to a threshold, both the house price and the price/rent ratio increase.*

Proof: See the appendix.

While the zero savings assumption is imposed in this case, it could be rationalized. Because housing wealth is more useful as a competitive weapon than non-housing wealth in the mating market (due to the better observability and verifiability of the housing wealth), agents generally find it optimal to borrow money (i.e., having a negative savings rate) in order to finance the purchase of a larger house. (If everyone borrows up to the maximum amount allowed by the law, then one can still rank different agents' relative wealth based on the relative values of the houses.) In this context, if we introduce a no-borrowing constraint, which is a way to capture financial under-development in developing economies, then zero savings would be an optimal outcome anyway. [In a survey of 10,043 urban households across 19 provinces and province-level municipalities in China in 2009, Wei (2011) found that the average amount of equity in owned housing dominates any other form of savings and financial investment. For example, the average household savings is about 13% of the equity value of the house. Another way to see it is through relative prevalence of various investment vehicles. 20% of the households invest in the stock market (buying individual stocks), and no more than 10% invest in mutual funds and/or bonds. In comparison, 85% of the households in the sample own a home. In other words, housing appears to be the overwhelming choice of financial investment vehicle for most urban households.]

A few other remarks are in order that explain the intuition behind the proposition and some extensions. First, as the sex ratio goes up, a representative man's demand for housing (as a given house price) goes up. If a representative man has a strong *ex ante* desire to be married, as the exogenous prospect of not being able to be married rises (due to a higher sex ratio), he is willing to compete harder by accumulating more housing asset that is valued by potential marriage partners.

Second, as the sex ratio goes up, the change in a representative woman's housing demand in the first period is ambiguous. On the one hand, since her future husband is likely to own a larger house, she could free ride by buying a smaller house (and consuming more in the first period). On the other hand, precisely because men have accumulated more housing wealth, their cutoff point in the type of women they are willing to marry (based on a combination of a woman's housing wealth and emotional utility) also goes up. This means that the probability that a given low-type woman would not be married also goes up. Since a representative woman does not know her precise type in the first period, this gives her an incentive to increase her demand for housing. The net effect of a higher sex ratio on a woman's housing demand depends on the relative importance of these two forces.

Third, the aggregate demand for housing (at a given house price) goes up unambiguously in response to a higher sex ratio. If the optimal response of the representative woman is to raise her demand for housing, it is not surprising that the aggregate demand for housing would rise unambiguously. Interestingly, even if the optimal response of the representative woman is to reduce her demand for housing, the incremental housing demand by men always dominates. This happens for two reasons. First, from a representative man's point of view, in addition to trying to improve his relative standing in the marriage market, he raises his desired level of housing wealth in the first period to compensate for any reduction in housing wealth of his future wife. The more his wife is expected to cut down the housing wealth, the more he would have to make it up for. This ensures that, if the conditions are such that a representative woman chooses to decrease her demand for housing, a representative man always raises his demand for housing sufficiently to offset the reduction in housing demand by his future wife. Second, as the sex ratio goes up, the share of men in the population also rises. Since men have a higher demand for housing whenever the sex ratio exceeds one, this change in the composition of the population also contributes to a higher aggregate demand for housing. Therefore, in spite of an ambiguous effect of a higher sex ratio on a representative woman's housing demand, the sum of all men's and all women's housing demand goes up unambiguously.

Fourth, we have considered an extension in which the bargaining power within a marriage depends on the relative wealth that a husband and a wife bring into the marriage. The smaller the share of a wife, the smaller claim she could have over their joint marital wealth. In this setting, a representative woman's housing wealth accumulation in the first period may rise in response to a higher sex ratio. This of course makes it even easier for the aggregate demand for housing to go up. (The extension is available upon request.)

Fifth, since the aggregate housing stock is assumed to be fixed, an increase in the aggregate demand for houses translates into a rise in the house price. In general, if the housing stock is not fixed, but with low supply elasticity, we conjecture that a rise in the sex ratio would generate a simultaneous increase in the house price and in the average size of a house. In other words, pressure from an increase in the aggregate demand for houses would be spread between house size and house price.

Sixth, because to be able to rent a large home is not the same thing as to be able to own a large home in the mating market, this model naturally generates a rise in the ratio of housing price to rent as the sex ratio rises.

Finally, the comparative statistics in the proposition are valid for a rise in the sex ratio up to a threshold. The intuition behind the existence of a threshold is this. When the sex ratio is sufficiently high, any further increase in the sex ratio may not motivate a representative man to increase his already very high housing wealth both because the chance of getting married is very low anyway and because the additional sacrifice in the form of a lower first-period consumption is too painful. In the extreme case in which the sex ratio is close to infinity, the probability for a man to be married in the second period is essentially zero. In this case, the desire to cut down first period consumption in order to improve a very slim chance of being married is weak. In other words, for high enough sex ratios, any additional increase in the sex ratio may cause men to give up. While the threshold exists in theory, we find that, in calibrations, such a threshold never emerges for reasonable values of other parameters in the model. This suggests that the theoretical threshold must be higher than any realistic sex ratios one observes in the data.

Numerical Examples

We start with the benchmark model which imposes a borrowing constraint on the agents: $s_m, s_w \geq 0$. This is meant to capture the idea that people face a limit on the amount of mortgage they can borrow to finance their house purchase, due to the under-development of the financial market. Such constraints are binding for Chinese households, the subject of the empirical part of this paper. In the open economy, we fix $R = \beta^{-1}$, where β is set to $0.97^{20} \approx 0.54$, since we take 20 years as one period in our calibration and pick an annual discount factor of 0.97. For most parameters, we follow the existing literature as much as possible (see Du and Wei, 2011, for more discussion). In the calibration of the benchmark model, we set the house depreciation to zero.

Perhaps the most unusual variable in the model is emotional utility η, which is assumed to follow a normal distribution, with the mean $\mu_\eta = 2$ and the standard deviation of $\sigma_\eta = 0.05$. These values are inferred from Blanchflower and Oswald (2004), who have estimated that, on average, a lasting marriage is equivalent to augmenting one's income by $100,000 (in 1990 dollars) per year in the United States during 1972-1998. Since the average income per working person was about $48,000 during that period, a sustained marriage is roughly worth twice the average income (on a permanent basis). They have also found that, controlling for other variables such as income, age, and education, the t statistics of the coefficient on the dummy variable for "never married" is around -20. Then the implied standard deviation of emotional utility σ_η can be pinned down at 0.05.

Table 1: Parameter Values in the Simulations

Parameters	Benchmark Values
Discount factor	$\beta = 0.54$
Initial endowment	$y = 1$
House stock	$H = 1$
Emotional utility, mean	$\mu_\eta = 2$
Emotional utility, standard deviation	$\sigma_\eta = 0.05$
Depreciation rate	$d = 0$
Sensitivity of bargaining power to relative wealth	$\varepsilon = 1$
Share of consumption on numeraire goods	$\gamma = 0.5$

We simulate the benchmark model with the borrowing constraint. Note that if the world interest rate is at a moderate level, agents, in the absence of the borrowing constraint, would optimally borrow money to finance an even greater demand for housing. Thus, the borrowing constraint effectively induces a zero savings rate to emerge endogenously. If one thinks that credit constraints are an important feature for some economies (e.g., China in the empirical section of this paper), one might regard the borrowing constraint as a realistic condition.

We let the sex ratio vary from 1 (balanced) to 1.5 (very unbalanced). Table 2 traces out the size of house purchase by the representative man and representative women (Columns 2 and 3), and house price and rent (Columns 4 and 5), respectively. As the sex ratio rises, housing demand by the representative man rises, while that by the representative woman declines. Because the aggregate demand for (owned) housing rises, the equilibrium housing price increases.

In Figure 1a, the solid dark line and broken dark line trace out the house price and rent as a function of the sex ratio, respectively. In Figure 1b, the dark line traces out the house price/rent ratio. It is clear that as the sex ratio rises, both the house price and the price/rent ratio go up. The relationship is monotonic at least for sex ratios up to 1.5 (which is higher than any actual country-level sex ratio in the data).

To see how concerns for mating competition affects the house price, it is useful to consider a counter-factual in which agents do not regard housing wealth as something that would affect their probability of marriage in the second period. (In the aggregate, this belief is correct since every man will end up having the same housing purchase decision, and the total number of unmarried men is not affected by the size of houses men buy.) We implement this counter-factual by assuming that the probability of marriage for a man is equal to the ratio of the number of women to the number of men (while the probability of marriage for a woman is 1). Table 3 traces out the housing price and the rent under this counter-factual. Compared to Table 2, it is clear that the house price would have been substantially lower had there not been a desire to use housing wealth to gain competitive advantage in the mating market.

It is interesting to note that, even when the sex ratio is balanced (φ=1), the housing price is higher when mating market competition is considered than when it is not. In that case, if agents believe that a higher level of housing wealth helps to be matched with a more attractive mate (someone of the opposite sex with a higher value of the emotional utility), they would raise their demand for housing. In equilibrium, this would result in a higher housing price than would otherwise be the case.

Back in Figure 1a, the light solid line and light broken line trace out the house price and rent under the counter-factual case of no concerns for mating market competition. (Note that the line for rent is the same whether or not people regard owned housing as a competitive weapon in the mating market.) The figure makes it clear that concerns for mating competition not only have raised the house price, the gap between the two price lines also widens as the sex ratio rises. This pattern is easy to understand. When there were no concerns for mating competition, a higher sex ratio would raise the house price mildly. This reflects a hedging motive – a decline in the objective probability of marriage means a decline in the probability of pooling housing resources with a spouse in the second period, which would motivate people to accumulate a somewhat larger housing wealth than would otherwise be the case. Once concerns for mating competition are present, men aggressively increase the demand for a house because a larger house is also regarded as a weapon in the war for mating partners. This leads to a more rapid increase in the house price as the sex ratio goes up.

Some additional remarks are in order. First, we have varied the parameter values in the neighborhood of the benchmark model (not reported to save space). The qualitative results always stay the same.

Second, we have also considered an extension in which the bargaining power within a marriage depends on the relative wealth that a husband and a wife bring into the marriage (not reported). The smaller the share of a wife in the initial wealth, the smaller claim she would have over their joint marital wealth. In this setting, a representative woman's housing wealth in the first period may rise (or decline less) in response to a higher sex ratio. This of course enhances the positive response of the equilibrium house price to a rise in the sex ratio.

To summarize, our theory suggests that the equilibrium house price is higher in a world in which agents treat housing as an instrument for mating competition (in addition to being a consumption good and financial investment) than in a world in which agents do not link housing to mating competition. Our theory also suggests that the housing price tends to be higher when the mating competition becomes fiercer (or when the sex ratio rises). The first result in principle applies to all countries in the world, but is difficult to test empirically since we do not observe a world in which concerns for mating competition is absent. The second result can be tested in countries where the sex ratio in the pre-marital age cohort is seriously unbalanced. Because the two results come from a common model, evidence in support of the second prediction also lends some credence to the first prediction.

We now turn to testing how variations in the sex ratio are linked to variations in housing market characteristics. We dissect the Chinese experience for this purpose.

4. Background on the Chinese housing market and the marriage market

Before presenting formal statistical tests, we briefly introduce some institutional background for both the housing market and the marriage market in China.

The Chinese housing market

In the three decades before 1988, there was no housing market in the People's Republic of China. In urban areas, every working person worked for either the government, a state-owned entity, or a collectively owned entity. A typical household lived in an apartment assigned to them by the work unit of one of the household members. They paid a nominal rent and could keep the place virtually forever. There was no market for them to sell the existing home or to buy a new one. In rural areas, a typical household lived in a house built by them or their extended family.

An amendment to China's constitution in 1988 formally laid the legal foundation for private housing development. The restructuring of state-owned firms (including closures or contractions) was one of the impetuses for housing reform. Some companies started to hire employees without providing an apartment. New homes (apartments) started to be built and sold on the market. While the state always retains the ownership of the land, home buyers automatically acquire the right of land use, bundled together with the apartment, for 70 years. It is not clear what will happen when the 70-year term expires. In rural areas, a majority of households build their own homes. Because private housing development was in a nascent stage, the secondary market for housing in rural areas was not very liquid in the first ten years of the reform.

The Chinese State Council's 23rd Degree of 1998 formally mandated work units (employers broadly defined to include companies, non-profit organizations, and government agencies) to monetize all housing benefits (i.e., making them a part of wages and salaries) rather than to provide employees with an apartment for life-time usage. The existing stock of housing units that were assigned to employees was asked to be sold to their current occupants within some time frame. These measures were designed to promote the commercial development of the private housing market (called "the commodity housing market"), which in turn was expected to provide a key source of economic growth. As a result, 1998 marked the beginning of the modern Chinese housing market. See Wu, Gyourko and Deng (2011) for additional information on China's urban housing market.

A housing construction boom ensued. In 1998, 1,178 million square meters of residential housing space was constructed, a majority of which took place in rural areas, mostly because that was where most people lived (773 million square meters in rural areas versus 406 million square meters in urban areas). Construction of new houses continued afterwards. In 2008, 1,546 million square meters of new residential housing space was built. With the faster rise in urban income, the pace of urban housing construction has almost caught up with that of rural construction in spite of the former's smaller population, with 760 million square meters of new housing space in urban areas versus 786 million square meters of new housing in the rural areas (Chinese Statistical Yearbook, 2010). With the rapid annual additions to the housing stock, a striking feature of the Chinese urban housing market is that most housing transactions (about 70%) involve newly built homes, even though the secondary market for housing has also begun to be active and liquid.

<u>The marriage market</u>

The marriage market is "local" for most Chinese. First, according to the China population census of 2000, 92% of rural residents live in their county of birth, and 62% of urban residents live in their city of birth. Second, in rural China, 89 % 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 China household and income project (CHIP) migrant workers survey in 2000 shows that

82% of the migrant worker 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 hometowns to look for a job. To sum up, there appears to be limited mobility for marriage reasons.

A recently developed social norm is that the family of a groom is supposed to supply an apartment for the young couple before the marriage. A survey of mothers with a daughter in eight major cities by *China Economic Daily* in March 2010 reported that 80% of the respondents would object to their daughter marrying a young man who can only rent rather than own a home. This prompts the headline that "it is fundamentally true that mothers-in-law (mothers of wives) are pushing up the housing prices."[6] Many mothers with a daughter express a willingness to help their future sons in law in the purchase of a home (although the majority of the funding has to come from the groom's family).

According to the 2010 Marriage Market Survey in China, 71% of unmarried women prefer that their future husbands own a home. This is consistent with the pattern reported above. Interestingly, 48% of unmarried men express a preference for their future wives to own a home as well. In other words, a home-owning woman is considered more desirable in the eyes of about half of the men.

In rural areas, a young couple after the marriage often lives in the same house built by the parents of the groom. In contrast, in urban areas, while it is possible for the young couple to live with the groom's parents, a more common arrangement would be for the parents of the groom to buy a home for the couple, or at least to pay for the down payment portion of the cost of the new home.

5. Statistical Evidence

We examine two types of evidence. First, at the household level, we show that the demand for housing varies in a way that is consistent with our hypothesis. In particular, a combination of having an unmarried son at home and living in a region with a skewed

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[6] *China Economic Daily* (中国经济时报), "It is fundamentally true that mothers in law are pushing up the housing prices; less than 20% accept sons in law who can only rent," March 8, 2010, http://msn.china.ynet.com/view.jsp?oid=63869864, accessed on March 15, 2010.

sex ratio raises a household's demand for owned housing. Second, at a regional level, we turn to local general equilibrium and investigate whether the average value of a house 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, the survey asked for monthly rent. To ensure that 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. (We ignore houses older than 2004 for this purpose.)

The summary statistics for the key variables are reported in Table 4. The sex ratio for the age cohort of 5-19 was 107 males per 100 females in 1985, which was moderately unbalanced, but reached a more skewed ratio of 112 males per 100 females in 2005. Ignoring differential mortality rates between the two sexes and the age difference between husbands and wives, 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, the 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 same-age sex ratio. However, since China's young 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 language and culture barriers, marriages by and large are local affairs. This means that the competition among men for marriage partners is more intense in regions with a more skewed sex ratio.

Who wants a bigger home?

We investigate whether and how the value of a home owned by parents depends jointly on 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. To maximize comparability across households, we restrict the sample to three-person nuclear families whose head of household is younger than 45 (inclusive) and whose child is between 5 and 19 (inclusive).

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 Table 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 5. 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 5, we examine the role of the sex ratio in determining the physical size of a home (in square meters). The qualitative results are similar to the previous columns. In particular, it is not surprising that richer and larger families tend to own a larger home. It is also interesting that, holding constant family size, having more sons per se is not associated with a larger home. Most noteworthy for us, a combination of having sons and living in a region with a more skewed sex ratio does lead to owning a larger home.

In the last two columns of Table 5, 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 coefficient on the interaction term is essentially zero for the rural sample and even negative for the urban sample. In sum, a higher sex ratio induces families with an unmarried son to want to buy a bigger and more expensive home. Consistent with our hypothesis, there is no comparable pattern for rental units.

The regressions described above pool together families with a son and families with a daughter. The specification requires that the coefficients on all variables be the same for these two types of families. We can relax this requirement by running separate regressions for son-families and daughter-families without imposing the equality restriction on other regressors between the two subsamples. To further strengthen the comparability of the two subsamples, we choose to focus on three-person families with two living parents and an unmarried son or daughter. This rules out families with multiple siblings.

The regression results for urban nuclear families are reported in Table 6. For both son-families and daughter-families, a higher household income is associated with a more expensive home and a larger home. The positive coefficient on the local sex ratio in the first column means that son-families tend to own a more expensive home if they happen to live in a region with a higher sex ratio. For daughter-families, their home value is uncorrelated with the local sex ratio. In Columns 3 and 4 where the dependent variable is the size of a home, the same patterns are observed. In particular, son-families in a region with a higher sex ratio imbalance appear to want to own a larger home. For daughter families, their home sizes are uncorrelated with the local sex ratio.. This suggests that, for daughter families, the incentive to free ride and the incentive to catch the richest men roughly balance each other out.

In Columns 5 and 6, the dependent variable is the rent (for those families that do not own a home). This could be viewed as a placebo test. If a higher sex ratio or having a son represents extra need for housing services, one would have observed a positive coefficient on the sex ratio for the renter sample. But we don't. This is consistent with the notion that being able to rent an apartment is not the same thing as being able to own an apartment as far as mating competition is concerned.

We perform similar regressions for separate son-families and daughter-families in rural areas. The results are reported in Table 7 and are broadly similar to the rural sample. One interesting difference is that the housing value for daughter families is also positively related to the local sex ratio. One possible explanation is that the "tournament effect" dominates the free-ride incentive. Since a higher local sex ratio makes all families with a son to accumulate more housing wealth, the reward for a daughter-family to be matched with the richest son-family also rises with the local sex ratio. If a rich son-family also prefers a rich daughter-family, then a typical daughter family may choose to accumulate more housing wealth in response to a higher sex ratio. Another possible explanation is that the desire to protect a daughter's bargaining power within her marriage outweighs the desire to free ride on the higher wealth of one's future son-in-law.

<u>Extension: Interactions between the sex ratio and household income</u>

So far we have focused on how a typical family with a son versus that with a daughter would react to higher sex ratios. We suggest that the response by a daughter-family depends on the net effect of several opposing forces: a combination of the tournament effect and a desire to protect the daughter's bargaining power within a marriage versus a desire to free ride on the higher wealth of their future son-in-law.

To allow for more richness in the family's response as a function of both the family income and the local sex ratio, we trace out the relationship between the (log) house value and (log) family income both by the local sex ratio and by the gender of the children. More precisely, we first define a high sex ratio region as the agglomeration of ten cities (or rural prefectures) with the highest sex ratios in the sample. Their sex ratios all turn out to be greater than 115. Similarly, we define a low sex ratio as the agglomeration of ten cities (or rural prefectures) with the lowest sex ratios in the sample. Their sex ratios all turn out to be less than 107. There are four cases of regions: (a) a high sex ratio rural region, (b) a low sex ratio rural region, (c) a high sex ratio urban region, and (d) a low sex ratio urban region. Within each region, we choose two types of families: (a) families with a single son between 5 and 19, and (b) families with a single daughter between 5 and 19.

For each type of households in each region, we use a locally weighted polynomial regression (LOWESS), originally developed by Cleveland (1979) and Cleveland and Devlin (1988), to trace out a non-parametric relationship between the local value of a home and the log household income.

The four non-parametric curves for the rural regions are plotted in Figure 2A. As one would expect, the house value generally rises with the family income (richer households own a larger home). A number of other interesting features are noteworthy. First, a family with a son, at a given level of family income, tends to own a more expensive home if it is located in a high sex ratio region. This is true over the entire range of income, but the gap becomes smaller at higher incomes. One possible interpretation is that once a family is rich enough, there is very low probability that their son will not be able to find a bride. As a result, the need to use housing to signal attractiveness in the marriage market also declines.

Second, a similar pattern is observed for families with a daughter. That is, while the home value rises with the family income, a daughter family also tends to own a more expensive home at a given income level, if it lives in a high sex ratio region. The latter pattern is consistent with the interpretation that a desire to protect the daughter's bargaining power within her marriage, or a desire to use own housing wealth to increase the chance that the daughter can be matched with a young man from a rich household, dominates a desire to free ride on the future son-in-law's wealth.

Third, when we compare the two types of families, we see that the curve for families with a son tends to be higher than families with a daughter, and the difference tends to be wider in the high sex ratio region than in the low sex ratio region. (A formal test is captured by the regression coefficients in the first row of Table 5.)

Similarly, four curves on home value and family income for the urban regions are plotted in Figure 2B. The patterns are somewhat more nuanced. First, while the overall relationship between the two is positive, we see some non-monotonicity on the far left of the curves: the poorest home owners appear to own slighter more expensive homes than the moderately less poor home owners. Second, on the other hand, the richest families in the low sex ratio region tend to own homes as expensive as (or more expensive than) their counterparts in the high sex ratio region. We do not have a rigorous explanation for the differences in the patterns between the rural and urban regions. One conjecture is that at the high levels of income, households no longer consider the need for the marriage market competition in their choice of home purchase since the probability of involuntary singlehood by their children is low. Notice that the highest household incomes in the urban region are much higher than their counterparts in the rural region. In other words, we do not observe similarly rich households in the rural region.

In Figures 3A and 3B, we trace out the relationship between house size and household income, separating son-families versus daughter-families, high sex-ratio regions versus low sex-ratio regions, and urban versus rural areas, in a fashion similar to Figures 2A and 2B. The broad patterns are also similar. Generally speaking, richer families own a larger home. At a given income level, son-families in a high sex-ratio region own a larger home than their counterparts in a low sex-ratio region. An exception to this rule is for families at the high end of the income distribution – the size of their homes is basically unrelated to the local sex ratio. This seems sensible. For a rich family with a son, it is very unlikely that their son cannot get married even if they live in a high sex-ratio region, and therefore their home size is not affected by the local sex ratio. Interestingly, the same is true for daughter families. There is something funny going on at the lowest end of the income distribution especially for families in the cities. We do not have a good explanation for this particular feature.

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. Separately, families with no unmarried children have their own demand for housing. Because of these possibilities, 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. For a given location (a city or a rural prefecture), we compute the average value of a home (among the newly constructed homes in a rural area or newly transacted homes in a city) in 2005 and the average household income for that location.

Basic Patterns in Pictures

In Figure 4, 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 occurred 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 5, we plot the average ratio of home value to rental rate against the sex ratio across rural prefectures (the left graph) and across cities (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.

Regression results

While Figures 4 and 5 suggest the existence of a positive relationship that is not driven by one or two outliers, they do not adjust for house size and do not control for other local characteristics. We now turn to regressions with the following specification:

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

where house(k) refers to the average value (or the average physical size in square meters, or the average price per square meter) of houses newly constructed in 2004[7] 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 are parameters to be estimated.

We implement the regressions separately for the rural sample (331 prefectures) and the urban sample (259 cities). Table 8 reports the regression results. Not surprisingly, regions with a higher household income tend to have more expensive homes. More interestingly, 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

[7] 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 information on the year of purchase.

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 per unit of space = house value/size, in terms of the 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 per square meter. 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, endogeneity of sex ratios might not be 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 is a strict family planning policy introduced at the beginning of the 1980s[8]. We explore three determinants of local sex ratios 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. 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, generating variations across regions and over time. These two monetary penalty variables constitute the first two candidates for our instrumental variables.[9]

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, 2011a). Therefore, the share of non-Han Chinese in the local population offers another possible instrument.[10]

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[8] 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). 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.

[9] 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.

[10] In principle, variations in the cost of sex screening technology especially the use of ultrasound B machines (as documented by Li and Zheng, 2009), and the economic status of women (such as that documented in Qian, 2008) could

The first stage regressions for the urban and rural samples are reported in Columns 1 and 2 of Table 9, 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.

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 0.27 and 0.15, respectively. The Kleibergen-Paap Wald 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 13.9 at the 5% level and 9.08 at the 10% level, respectively. This means that the three variables are not weak instruments.

The second stage regressions are reported in Table 10. 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. In other words, in most cases in which IV estimation is needed, the instruments are valid in a statistical sense (as they pass the over-identification test). (IV estimation is not needed in a few cases in which the IVs do not pass the over-identification test.) The point estimates in Table 10 are generally much larger than their OLS counterparts in Table 8. This suggests that the downward bias in Table 8 generated either by missing regressors or by measurement errors is substantial.

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.

In any case, the IV results suggest that the data patterns from the OLS estimate carry over. In particular, higher sex ratios tend to generate systematically 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 housing 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 predictive power over the evolution of local housing prices beyond the growth of local income and local population.

We pursue the following specification:

$$\text{LogPrice}(k, t) = \beta\ \text{sexratio}(k, t) + \sum \delta(j)\text{City}(j) + \sum \theta(s)\text{Year}(s) + X(k, t)\ \Gamma + e(k)$$

where LogPrice(k, t) refers to the log average price per square meters in city k and year t, sexratio(k) is the sex ratio for the age cohort of 5-19 in city k and year t (inferred from the 2000 population census), City(j) is a fixed effect for city j, Year(s) is a fixed effect for year s, and X(k, t) is a vector of control variables including the average local household income. β, δ, θ and Γ are parameters to be estimated.

The city fixed effects are meant to capture local regulations on migrant workers, local industrial structures, taxes and other time-invariant (or slow-moving) features that may affect a city's house price. The year fixed effects are meant to control for nationwide common shocks to housing prices such as movements in the interest rate and inflation rate. While it is unclear if the average housing price data adjust for house size and other amenities, we expect that location and year fixed effects would control for (most of the) systematic differences in housing qualities across locations and over time.

Table 11 presents the summary statistics on variables used in the panel regressions. While (unweighted) average house price was 2426 yuans per square meters in 2003, it grew to 5706 yuans in nominal terms by 2009. The rate of housing price appreciation is much higher than the rate of CPI inflation. The average sex ratio during the period rose from 1.07 to 1.13 across all provinces.

Table 12 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 13. 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 non-local (including foreign nationals) are said to be active speculators in the local housing market. The coefficient on the sex ratio variable remains significant, and the point estimate becomes larger (1.34 instead of 0.89). This means that the connection between the local sex ratio and local housing prices is stronger in China outside the three largest cities. To be conservative, using the coefficient in the third column of Table 13 (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 13, we instrument the sex ratio by the three variables in Table 12. 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 a 48% increase in home price over 2003-2009 (= 6*1.39/(0.371-0.198)).

6. 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 from 2000 to 2005 (corresponding to the actual rise in the national sex ratio for the age cohort of 5-19) would contribute to 36 % of the observed rise in the average home value in Chinese cities in the same period. A higher sex ratio imbalance accounts for between 30-48% of the increase in real housing prices in 35 major cities during 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.

Optimal property tax and stamp tax on house transactions should in principle take into account local sex ratios. Since sex ratio imbalance is also a problem in several other economies (e.g., Singapore, Hong Kong, India, and Vietnam), it would be interesting to investigate its role in other housing markets. We leave a thorough investigation of these issues to future research.

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Table 1 (in the text) is the table on parameter values in the calibrations of the model

Table 2: House Price and Rent in the Special Case, with borrowing constraint (Calibration Results)

Sex Ratio ϕ	Housing demand by representative man	Housing demand by representative woman	**House Price**	Rent
1	1.000	1.000	**1.351**	1.000
1.01	1.332	0.665	**1.393**	1.000
1.02	1.459	0.532	**1.412**	1.000
1.03	1.532	0.452	**1.424**	1.000
1.04	1.581	0.396	**1.433**	1.000
1.05	1.616	0.353	**1.440**	1.000
1.1	1.699	0.231	**1.461**	1.000
1.15	1.724	0.168	**1.474**	1.000
1.2	1.728	0.127	**1.485**	1.000
1.25	1.754	0.057	**1.494**	1.000
1.3	1.712	0.075	**1.503**	1.000
1.35	1.736	0.007	**1.511**	1.000
1.4	1.684	0.043	**1.518**	1.000
1.45	1.658	0.047	**1.525**	1.000
1.5	1.653	0.020	**1.532**	1.000

Table 3: Assuming No Concerns for Marriage Market Competition
with borrowing constraint (Calibration Results)

Sex Ratio ϕ	Housing demand by representative man	Housing demand by representative woman	**House Price**	Rent
1	1.000	1.000	**1.213**	1.000
1.01	1.017	0.982	**1.213**	1.000
1.02	1.033	0.966	**1.214**	1.000
1.03	1.047	0.952	**1.215**	1.000
1.04	1.060	0.938	**1.216**	1.000
1.05	1.071	0.925	**1.216**	1.000
1.1	1.116	0.872	**1.220**	1.000
1.15	1.147	0.831	**1.223**	1.000
1.2	1.169	0.797	**1.225**	1.000
1.25	1.185	0.769	**1.228**	1.000
1.3	1.196	0.745	**1.230**	1.000
1.35	1.205	0.724	**1.233**	1.000
1.4	1.211	0.705	**1.235**	1.000
1.45	1.215	0.688	**1.237**	1.000
1.5	1.218	0.673	**1.240**	1.000

Table 4: 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 5: Who Wants a Bigger Home? (Tobit regressions with locational fixed effects)

	Housing value (Tobit)		Housing space (OLS)		Rent (Tobit)	
	Rural	Urban	Rural	Urban	Rural	Urban
Sex ratio * having a son aged 5-19	0.06**	0.12**	0.02**	0.02**	0.00	-0.09**
	(0.01)	(0.05)	(0.00)	(0.01)	(0.00)	(0.03)
Household income (log)	0.14**	0.14	0.07**	0.12**	0.09**	-0.03
	(0.01)	(0.11)	(0.00)	(0.01)	(0.00)	(0.03)
Household size	0.09**	0.50**	0.04**	0.07**	-0.03**	-0.01
	(0.01)	(0.05)	(0.00)	(0.01)	(0.00)	(0.01)
Household head age	-0.01**	-0.01**	-0.00**	-0.01**	-0.00**	0.01**
	(0.00)	(0.01)	(0.00)	(0.00)	(0.00)	(0.00)
Household head year of schooling	0.00	-0.02	0.00**	0.02**	0.01**	-0.05**
	(0.00)	(0.02)	(0.00)	(0.00)	(0.00)	(0.01)
Female household head	-0.18**	-0.17**	-0.05**	-0.05**	0.16**	0.15**
	(0.02)	(0.07)	(0.01)	(0.01)	(0.01)	(0.02)
Minority household head	-0.10**	-0.08	-0.03**	-0.03*	-0.04**	0.00
	(0.03)	(0.12)	(0.01)	(0.02)	(0.01)	(0.04)
Poor health among at least one family member	-0.03**	0.08	0.05**	0.09**	-0.02**	-0.27**
	(0.01)	(0.09)	(0.01)	(0.02)	(0.01)	(0.05)
Having a child 10-14	0.06**	0.24**	0.02**	0.09**	-0.03**	-0.19**
	(0.01)	(0.06)	0.00	(0.01)	0.00	(0.03)
Having a child 15-19	0.25**	0.34**	0.12**	0.20**	-0.08**	-0.37**
	(0.01)	(0.06)	(0.00)	(0.01)	(0.01)	(0.05)
Number of sons	-0.02**	0.00	-0.02**	-0.02*	0.02**	0.16**
	(0.01)	(0.05)	(0.00)	(0.01)	(0.00)	(0.04)
City/prefecture fixed effects	Yes	Yes	Yes	Yes	Yes	Yes
R square (Pseudo R square for Tobit)	0.03	0.03	0.14	0.16	0.03	0.02
N	229567	74012	229342	73964	229567	74012

Notes: The dependent variable is the purchased housing price or construction cost (log) of a house which is owned by the household who live there, the construction area of a house (log), 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 6: Who Wants an Expensive and Bigger Home and Who Wants to Rent? Urban Nuclear Families

	Housing value (Tobit)		Housing space (OLS)		Rent (Tobit)	
	Son	Daughter	Son	Daughter	Son	Daughter
Local sex ratio for the cohort 5-19 in 2005	2.51*	-0.36	0.50*	0.34	0.88	0.97
	(1.48)	(1.56)	(0.27)	(0.25)	(0.86)	(0.72)
Household income (log)	0.50**	0.69**	0.08**	0.11**	0.16**	0.08**
	(0.14)	(0.18)	(0.02)	(0.02)	(0.05)	(0.04)
Household head age	-0.04**	-0.04**	-0.01**	-0.01**	0.02**	0.02**
	(0.01)	(0.01)	(0.00)	(0.00)	(0.00)	(0.00)
Household head year of schooling	-0.07	-0.11**	0.03**	0.03**	-0.10**	-0.09**
	(0.05)	(0.04)	(0.01)	(0.00)	(0.01)	(0.01)
Female household head	-0.10	-0.03	-0.01	-0.02*	0.03	0.04
	(0.15)	(0.14)	(0.01)	(0.01)	(0.03)	(0.03)
Minority household head	-0.27	-0.53**	-0.03	-0.04	-0.06	0.08
	(0.23)	(0.26)	(0.03)	(0.03)	(0.06)	(0.07)
Poor health among at least one family member	0.13	0.34	0.09**	0.12**	-0.12	-0.16
	(0.31)	(0.36)	(0.05)	(0.05)	(0.13)	(0.11)
Having a child 10-14	0.45**	0.03	0.17**	0.13**	-0.36**	-0.29**
	(0.12)	(0.11)	(0.02)	(0.01)	(0.04)	(0.04)
Having a child 15-19	0.44**	-0.10	0.27**	0.21**	-0.50**	-0.41**
	(0.15)	(0.12)	(0.02)	(0.02)	(0.05)	(0.06)
Share of primary age population (20-59) in 2000 at the city level	-0.11*	-0.10*	0.04**	0.05**	-0.09**	-0.11**
	(0.06)	(0.05)	(0.01)	(0.01)	(0.02)	(0.01)
Pseudo R square	0.00	0.01	0.06	0.07	0.01	0.01
N	15298	12650	15292	12647	15298	12650

Notes: The sample is limited to nuclear three-person families with household head's age younger than 45 and child age between 5 and 19. The dependent variable is the purchased housing price or construction cost (log) of a house which is owned by the household who live there, the construction area of a house (log), 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 7: Who Wants an Expensive and Bigger Home and Who Wants to Rent? Nuclear Families in Rural Areas

	Housing value (Tobit)		Housing space (OLS)		Rent (Tobit)	
	Son	Daughter	Son	Daughter	Son	Daughter
Local sex ratio for the cohort 5-19 in 2005	1.59**	1.27**	0.42**	0.30	0.03	-0.16
	(0.46)	(0.50)	(0.20)	(0.20)	(0.22)	(0.18)
Household income (log)	0.14**	0.17**	0.07**	0.07**	0.15**	0.14**
	(0.02)	(0.03)	(0.01)	(0.01)	(0.01)	(0.01)
Household head age	0.00	0.00	-0.01**	-0.00**	-0.00**	0.00
	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)
Household head year of schooling	-0.03**	-0.05**	0.00**	0.00	0.01**	0.01**
	(0.01)	(0.01)	(0.00)	(0.00)	(0.00)	(0.00)
Female household head	-0.33**	-0.31**	-0.05**	-0.06**	0.14**	0.17**
	(0.07)	(0.09)	(0.02)	(0.02)	(0.03)	(0.04)
Minority household head	-0.31**	-0.24**	-0.03	-0.03	-0.04**	-0.04**
	(0.06)	(0.06)	(0.02)	(0.02)	(0.02)	(0.02)
Poor health among at least one family member	-0.01	0.03	0.09**	0.09**	0.01	-0.03*
	(0.05)	(0.07)	(0.02)	(0.02)	(0.02)	(0.01)
Having a child 10-14	0.26**	0.15**	0.12**	0.09**	-0.12**	-0.10**
	(0.03)	(0.04)	(0.01)	(0.01)	(0.02)	(0.02)
Having a child 15-19	0.47**	0.43**	0.21**	0.19**	-0.20**	-0.18**
	(0.04)	(0.04)	(0.01)	(0.01)	(0.02)	(0.02)
Share of primary age population (20-59) in 2000 at the city level	1.62**	0.70	0.15	0.02	-0.05	0.01
	(0.41)	(0.45)	(0.15)	(0.15)	(0.24)	(0.20)
Pseudo R square	0.01	0.01	0.03	0.03	0.01	0.01
N	29043	20128	29005	20102	29043	20128

Notes: The sample is limited to nuclear three-person families with household head's age younger than 45 and child's age between 5 and 19. The dependent variable is the purchased housing price or construction cost (log) of a house which is owned by the household who live there, the construction area of a house (log), 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 8: 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	1.60**	0.82**	0.78**	2.50**	1.60**	-0.89
(prefecture level)	(0.38)	(0.26)	(0.28)	(0.72)	(0.38)	(0.55)
Household monthly income (log)	0.91**	0.24**	0.67**	0.53**	-0.09	0.37**
	(0.09)	(0.06)	(0.07)	(0.12)	(0.09)	(0.09)
Total population (log)	0.13**	0.08**	0.05**	0.09*	0.13**	0.03
	(0.03)	(0.02)	(0.02)	(0.06)	(0.03)	(0.04)
Share of primary age population (20-59 years old)	0.92**	-0.05	0.97**	2.50**	0.92**	-1.58**
in 2000 (log)	(0.33)	(0.18)	(0.26)	(0.43)	(0.33)	(0.30)
Adj. R-squared	0.42	0.16	0.37	0.23	0.19	0.08
N	327	327	327	327	327	327
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 9: First Stage Regressions – Instrumenting for Local Sex Ratios

Dependent variable = local sex ratio	Rural Areas	Cities
Share of minority population in 2000 (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 in 2005 (log)	-1.25	2.25
	(1.29)	(1.63)
Total population in 2000 (log)	1.38**	-1.28**
	(0.52)	(0.49)
Share of primary age population (20-59 years old) in 2000	-22.27**	-0.84
	(4.30)	(1.44)
Adj. R-squared	0.27	0.15
Kleibergen-Paap rank Wald F statistic	14.23	14.98
N	331	259

Notes: (a) The dependent variable is the local sex ratio for the age cohort 5-19 in 2005 expressed in percentage term, inferred from the 2000 Population Census. (b) The two family planning variables are averaged over the corresponding years of birth for the age cohort 5-19 in 2005. (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 10: 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)
Durbin-Wu-Hausman test for endogeneity (p-value)	0.00	0.00	0.00	0.00	0.00	0.87
Hansen's J statistic for over-identification (p-value)	0.08	0.01	0.75	0.00	0.08	0.02
N	327	327	327	327	327	327
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)
Durbin-Wu-Hausman test for endogeneity (p-value)	0.64	0.16	0.59	0.08	0.64	0.04
Hansen's J statistic for over-identification (p-value)	0.38	0.90	0.30	0.79	0.38	0.03
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 11: 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 12: 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 13: 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 (Tobit)		Housing space (OLS)		Rent (Tobit)	
	Rural	Urban	Rural	Urban	Rural	Urban
Local sex ratio for the cohort 5-19 in 2005	0.77**	1.27*	0.45**	0.39	-0.11	0.54
	(0.33)	(0.73)	(0.19)	(0.29)	(0.12)	(0.75)
Sex ratio * having a son aged 5-19	0.03**	0.08**	0.02**	0.03**	-0.01	-0.09**
	(0.01)	(0.04)	(0.00)	(0.01)	(0.00)	(0.03)
Household income (log)	0.32**	0.48**	0.07**	0.08**	0.11**	0.15**
	(0.01)	(0.08)	(0.01)	(0.02)	(0.01)	(0.03)
Household size	0.02**	0.25**	0.04**	0.09**	-0.04**	-0.08**
	(0.01)	(0.03)	(0.00)	(0.01)	(0.00)	(0.01)
Household head age	-0.01**	-0.01**	-0.00**	-0.01**	-0.00**	0.01**
	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)
Household head year of schooling	0.03**	0.00	0.00**	0.01**	0.01**	-0.05**
	(0.00)	(0.01)	(0.00)	(0.00)	(0.00)	(0.01)
Female household head	0.09**	0.04	-0.03**	-0.04**	0.17**	0.18**
	(0.02)	(0.05)	(0.01)	(0.01)	(0.01)	(0.02)
Minority household head	-0.29**	-0.25**	-0.02	-0.03	-0.02**	0.00
	(0.04)	(0.11)	(0.02)	(0.03)	(0.01)	(0.04)
Poor health among at least one family member	-0.02	0.01	0.07**	0.09**	-0.03**	-0.19**
	(0.02)	(0.06)	(0.01)	(0.02)	(0.01)	(0.04)
Having a child 10-14	0.07**	0.15**	0.03**	0.09**	-0.02**	-0.20**
	(0.01)	(0.04)	(0.00)	(0.01)	(0.00)	(0.03)
Having a child 15-19	0.22**	0.21**	0.12**	0.19**	-0.09**	-0.38**
	(0.01)	(0.04)	(0.01)	(0.01)	(0.01)	(0.05)
Number of sons	0.00	0.00	-0.02**	-0.03**	0.02**	0.14**
	(0.01)	(0.04)	(0.00)	(0.01)	(0.00)	(0.04)
Share of primary age population (20-59) in 2000 at the city level	0.24	-0.05**	0.32**	0.05**	0.01	-0.10**
	(0.32)	(0.03)	(0.14)	(0.01)	(0.12)	(0.02)
R square (Pseudo R square for Tobit)	0.03	0.02	0.03	0.06	0.01	0.01
N	229567	74012	229342	73964	229567	74012

Notes: The dependent variable is the purchased housing price or construction cost (log) of a house which is owned by the household who live there, the construction area of a house (log), 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.

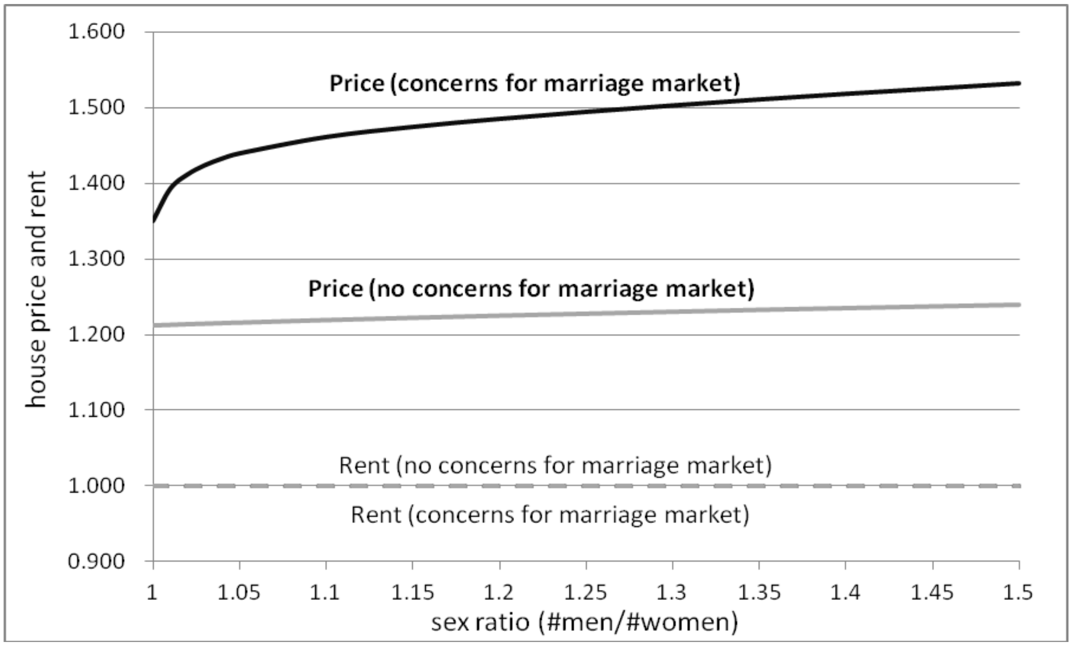

Figure 1a: House price and rent, with borrowing constraint, with and without concerns for marriage

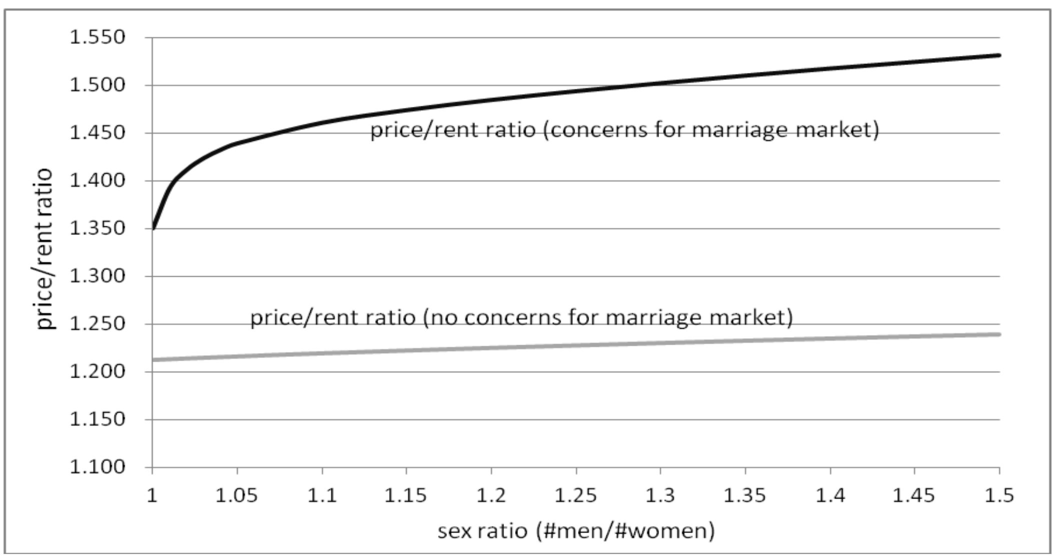

Figure 1b: Price/rent Ratio, with borrowing constraint, with and without concerns for marriage

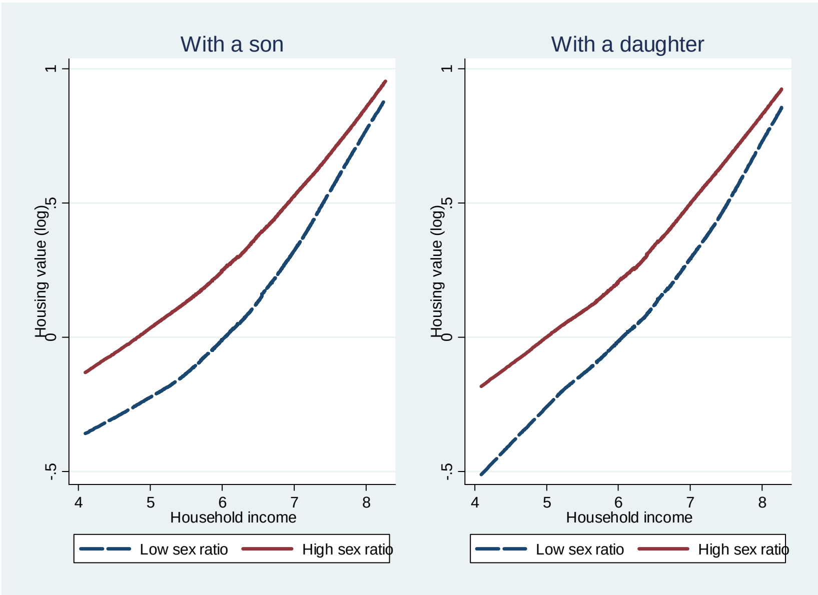

Figure 2A: Log housing value versus log household income in rural areas

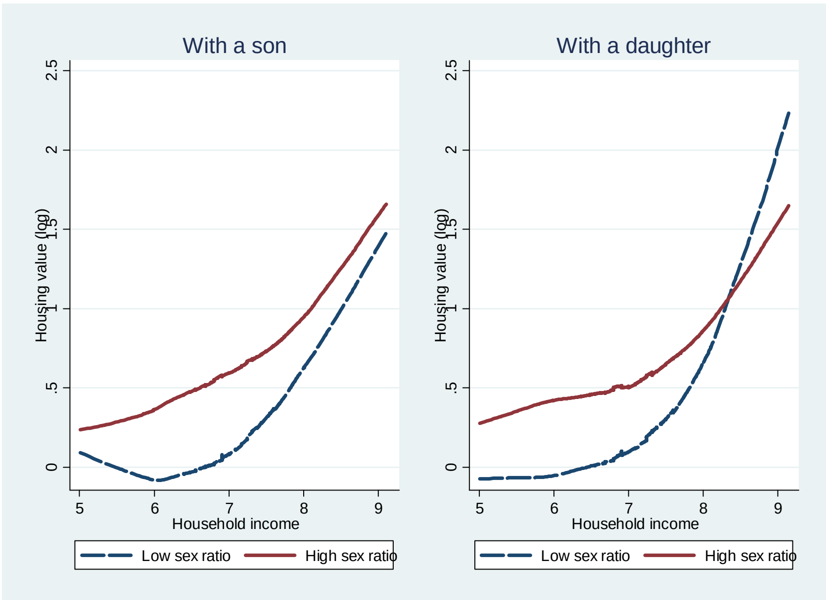

Figure 2B: Log housing value versus log household income in urban areas

Source: Plotted by the authors based on a 20 percent random sample of China 1% Population Survey in 2005. The low sex ratio region is a collection of locations with the lowest sex ratios for age cohort 5-19 in 2005 (below 107); while the high sex ratio region is a collection of regions with the highest sex ratio for age cohort 5-19 in 2005 (greater than 115). On the horizontal axis is log household income in 2005. On the vertical axis is log housing value in 2005.

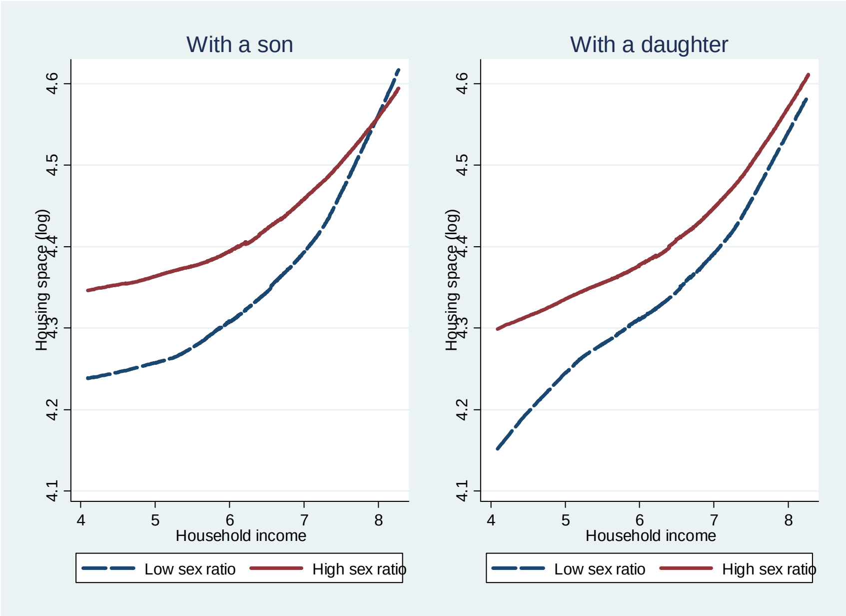

Figure 3A: Log housing space versus log household income in rural areas

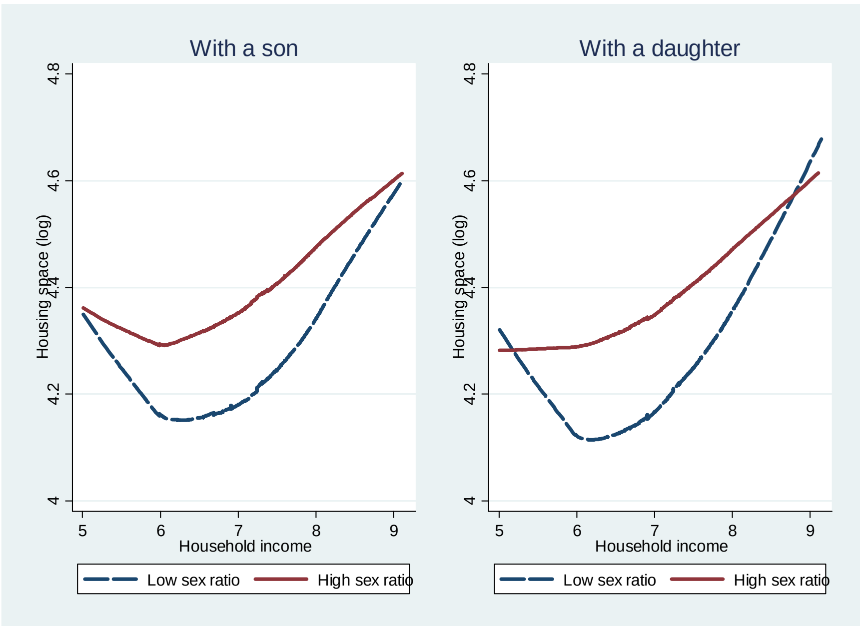

Figure 3B: Log housing space versus log household income in cities

Source: Plotted by the authors based on a 20 percent random sample of China 1% Population Survey in 2005. The low sex ratio region is a collection of locations with the lowest sex ratios for age cohort 5-19 in 2005 (below 107); while the high sex ratio region is a collection of regions with the highest sex ratio for age cohort 5-19 in 2005 (greater than 115). On the horizontal axis is log household income in 2005. On the vertical axis is log housing space in 2005.

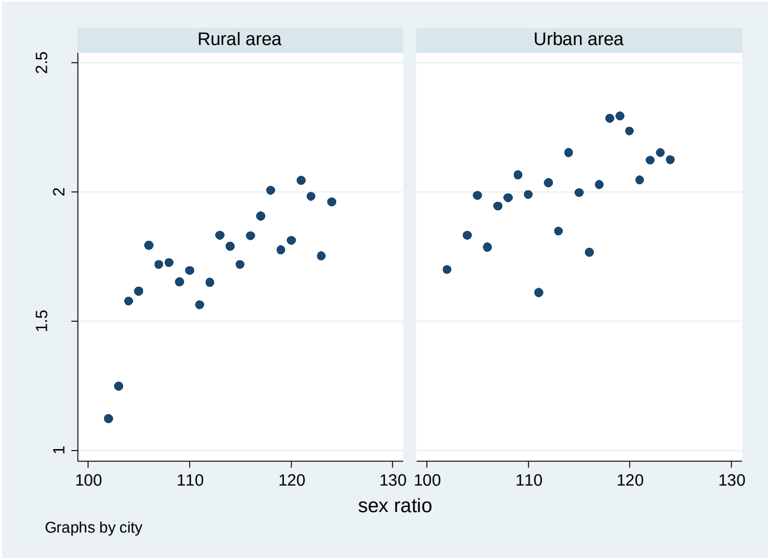

Figure 4: Sex Ratios and House Price-to-Income Ratios

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.

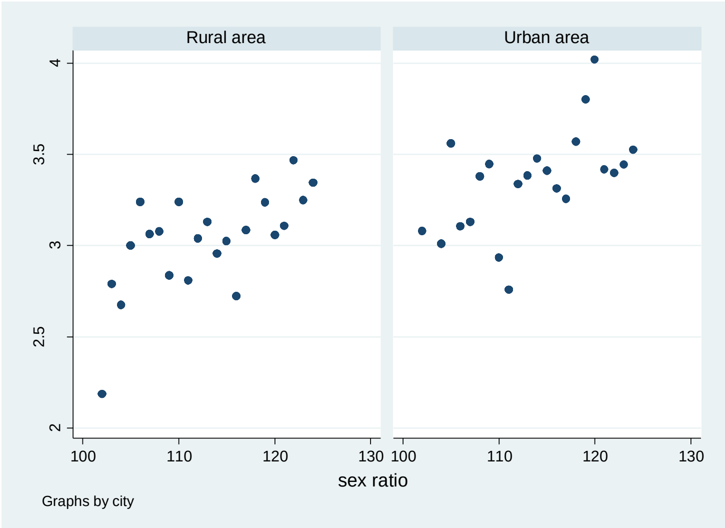

Figure 5: Sex Ratios and House Value-to-Rent Ratios

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.

Appendix A: Proof of Proposition (For online posting, not for publication in print)

The proof is somewhat involved, and can be broken down into two broad parts. Part 1 shows that the aggregate demand for housing (at a given price) goes up as the sex ratio rises (up to a threshold value). Part 2 shows that this translates into a rise in the house price and a rise in the price/rent ratio.

Part 1 itself is broken down into several pieces. The first order conditions for representative women and men in section 2.4 are:

$$\frac{-p+r}{y_{1w}}+\beta p\left\{\begin{array}{l}\frac{1}{y_2}\left(\delta_w+\frac{1}{\phi}(1-F(\bar{\eta}_w))+f(\bar{\eta}_w)\bar{\eta}_m\right)+(1-\delta_w)\frac{1}{y_{2ws}}\\+\frac{1}{y_2}f(\bar{\eta}_w)[Eu(c_{2w})-u(c_{2ws})]\end{array}\right\}=0 \qquad \text{(A.1)}$$

$$\frac{-p+r}{y_{1m}}+\beta p\left\{\begin{array}{l}\frac{1}{y_2}\left(\delta_m+\frac{1}{\phi}(1-F(\bar{\eta}_m))+f(\bar{\eta}_m)\bar{\eta}_w\right)+(1-\delta_m)\frac{1}{y_{2ms}}\\+\frac{1}{y_2}f(\bar{\eta}_m)[Eu(c_{2m})-u(c_{2ms})]\end{array}\right\}=0 \qquad \text{(A.2)}$$

where

$$y_{1w}=y-(p-r)H_w,\;\; y_{2ws}=pH_w$$
$$y_{1m}=y-(p-r)H_m,\;\; y_{2ms}=pH_m$$
$$y_2=p(H_w+H_m)$$

1) We first show by contradiction that $\bar{\eta}_w=u(c_{2ms})-Eu(c_{2m})$ and $\bar{\eta}_m=M(\bar{\eta}_w)$. Suppose not, then

$$\bar{\eta}_m>M(\bar{\eta}_w)\geq\bar{\eta}_w$$

where the first inequality holds because $\bar{\eta}_m:\;=\max\{u(c_{2ws})-Eu(c_{2w}),M(\bar{\eta}_w)\}$ and we have supposed that $\bar{\eta}_m\neq M(\bar{\eta}_w)$. The second inequality holds because $\phi>1$.

Then we have

$$\bar{\eta}_m=u(c_{2ws})-Eu(c_{2w})>\bar{\eta}_w\geq u(c_{2ms})-Eu(c_{2m}) \;\;\Rightarrow\;\; u(c_{2ws})>u(c_{2ms}) \qquad \text{(A.3)}$$

and hence, $H_w>H_m$. From (A.1) we know

$$\begin{aligned}\frac{p-r}{y_{1w}}&=\frac{\beta p}{y_{2ws}}+\delta_w\beta p\left(\frac{1}{y_2}-\frac{1}{y_{2ws}}\right)+\frac{\beta p}{y_2}\left(\frac{1}{\phi}(1-F(\bar{\eta}_w))+f(\bar{\eta}_w)\bar{\eta}_m\right)+\frac{\beta p}{y_2}f(\bar{\eta}_w)[Eu(c_{2w})-u(c_{2ws})]\\&<\frac{\beta p}{y_{2ws}}+\delta_m\beta p\left(\frac{1}{y_2}-\frac{1}{y_{2ws}}\right)+\frac{\beta p}{y_2}\left(\frac{1}{\phi}(1-F(\bar{\eta}_w))+f(\bar{\eta}_w)\bar{\eta}_m\right)+\frac{\beta p}{y_2}f(\bar{\eta}_w)[Eu(c_{2w})-u(c_{2ws})]\\&<\frac{\beta p}{y_{2ms}}+\delta_m\beta p\left(\frac{1}{y_2}-\frac{1}{y_{2ms}}\right)+\frac{\beta p}{y_2}\left(\phi(1-F(\bar{\eta}_m))+f(\bar{\eta}_w)\bar{\eta}_w\right)+\frac{\beta p}{y_2}f(\bar{\eta}_m)[Eu(c_{2w})-u(c_{2ms})]\\&=\frac{p-r}{y_{1m}}\end{aligned}$$

which means $y_{1w} > y_{1m}$, and hence, $H_w < H_m$, contradiction. The first inequality is due to $\delta_w < \delta_m$. The second inequality holds because:

i. $\frac{1}{\phi}(1-F(\bar{\eta}_w)) + f(\bar{\eta}_w)\bar{\eta}_m = \phi(1-F(\bar{\eta}_m)) + f(\bar{\eta}_w)\bar{\eta}_w$ by using the assumption that η is uniformly distributed.

ii. $\frac{\beta p}{y_{2ws}} + \delta_m \beta p\left(\frac{1}{y_2} - \frac{1}{y_{2ws}}\right) < \frac{\beta p}{y_{2ms}} + \delta_m \beta p\left(\frac{1}{y_2} - \frac{1}{y_{2ms}}\right)$ by using (A.3), which implies $H_w > H_m$ and hence $\frac{1}{y_{2ws}} < \frac{1}{y_{2ms}}$.

iii. $Eu(c_{2w}) - u(c_{2ws}) < Eu(c_{2w}) - u(c_{2ms})$ by using (A.3).

2) Totally differentiating (A.1) and (A.2) gives that:

$$\Omega \bullet dH = Z$$

where

$$\Omega = \begin{pmatrix} \Omega_{11} & \Omega_{12} \\ \Omega_{21} & \Omega_{22} \end{pmatrix},\ dH = \begin{pmatrix} dH_w \\ dH_m \end{pmatrix},\ Z = \begin{pmatrix} 0 \\ A \end{pmatrix}$$

$$\Omega_{11} = -\frac{(p-r)^2}{y_{1w}^2} + \beta p^2 \begin{bmatrix} -\frac{1}{y_2^2}(1+\frac{1}{\phi})(1-F(\bar{\eta}_w)) - (1-\delta_w)\frac{1}{y_{2ws}^2} \\ -\frac{f(\bar{\eta}_w)}{y_2^2}(\bar{\eta}_m + E(u_{2w}) - u_{2ws}) + 2f(\bar{\eta}_w)\frac{1}{y_2}(\frac{1}{y_2} - \frac{1}{y_{2ws}}) \end{bmatrix}$$

$$\Omega_{12} = \beta p^2 \left[-\frac{1}{y_2^2}(1+\frac{1}{\phi})(1-F(\bar{\eta}_w)) - \frac{f(\bar{\eta}_w)}{y_2^2}(\bar{\eta}_m + E(u_{2w}) - u_{2ws}) + f(\bar{\eta}_w)\frac{1}{y_2^2} + f(\bar{\eta}_w)(\frac{1}{y_{2ws}} - \frac{1}{y_2})(\frac{1}{y_{2ms}} - \frac{1}{y_2}) \right]$$

$$\Omega_{21} = \beta p^2 \left[-\frac{1}{y_2^2}(1+\frac{1}{\phi})(1-F(\bar{\eta}_m)) + f(\bar{\eta}_w)\frac{1}{y_2}\left((1+\frac{1}{\phi})\frac{1}{y_2} - \frac{1}{\phi y_{2ms}} \right) \right]$$

$$\Omega_{22} = -\frac{(p-r)^2}{y_{1m}^2} + \beta p^2 \left[-\frac{1}{y_2^2}(1+\phi)(1-F(\bar{\eta}_m)) - (1-\delta_m)\frac{1}{y_{2ms}^2} + \frac{f(\bar{\eta}_w)}{y_2^2}(\frac{1}{y_2} - \frac{1}{y_{2ms}})\left((1+\frac{1}{\phi})\frac{1}{y_2} - \frac{1}{\phi y_{2ms}} \right) \right]$$

$$A = \frac{1}{\phi^2}(1-F(\bar{\eta}_w))(\frac{1}{y_2} - \frac{1}{y_{2ms}})$$

Note that:

$$y_2 = p(1-d)H_w + p(1-d)H_m \geq y_{2ms},\ y_{2ws} \Rightarrow \frac{1}{y_2} - \frac{1}{y_{2ws}} \leq 0,\ \frac{1}{y_2} - \frac{1}{y_{2ms}} \leq 0$$

$$\bar{\eta}_m \geq u_{2ws} - E(u_{2w}) \Rightarrow \bar{\eta}_m + E(u_{2w}) - u_{2ws} \geq 0$$

$$(1+\phi)y_{2ms} > y_{2ms} + y_{2ws} \Rightarrow (1+\frac{1}{\phi})\frac{1}{y_2} - \frac{1}{\phi y_{2ms}} > 0$$

Thus it is easy to see that: $\Omega_{11}<0,\ \Omega_{22}<0,\ \Omega_{11}<\Omega_{21},\ A<0$

$$\det(\Omega)=\textit{positive terms}+\frac{(p-r)^4}{y_{1w}^2y_{1m}^2}+2\beta^2p^4\frac{\delta_m}{\eta^{\max}-\eta^{\min}}\frac{1}{y_{2ms}^2y_2}(\frac{1}{y_{2ws}}-\frac{1}{y_2})$$
$$+\beta^2p^4\frac{\delta_w}{\eta^{\max}-\eta^{\min}}\frac{1}{y_{2ws}^2}(\frac{1}{y_{2ms}}-\frac{1}{y_2})\left((1+\frac{1}{\phi})\frac{1}{y_2}-\frac{1}{\phi y_{2ms}}\right)$$
$$-\beta^2p^4\frac{1}{(\eta^{\max}-\eta^{\min})^2}\frac{1}{y_2}\left((1+\frac{1}{\phi})\frac{1}{y_2}-\frac{1}{\phi y_{2ms}}\right)\left(\frac{1}{y_{2ws}y_{2ms}}-(\frac{1}{y_{2ms}}+\frac{1}{y_{2ms}})\frac{1}{y_2}+\frac{2}{y_2^2}\right)$$

We can notice that all terms, except the last one, are positive, and we can give a sufficient condition to ensure that $\det(\Omega)>0$ is $E(\eta)>c$, where c is some constant. This can be interpreted as that the emotional utility should be large enough to make people willing to compete in the marriage market. We prove $\det(\Omega)>0$ by contradiction:

i. Suppose not, $\det(\Omega)<0$, then we must have:

$$\left\{\begin{aligned}&2\frac{\delta_m}{\eta^{\max}-\eta^{\min}}\frac{1}{y_{2ms}^2y_2}(\frac{1}{y_{2ws}}-\frac{1}{y_2})\ +\ \frac{\delta_w}{\eta^{\max}-\eta^{\min}}\frac{1}{y_{2ws}^2}(\frac{1}{y_{2ms}}-\frac{1}{y_2})\left((1+\frac{1}{\phi})\frac{1}{y_2}-\frac{1}{\phi y_{2ms}}\right)\\&-\frac{1}{(\eta^{\max}-\eta^{\min})^2}\frac{1}{y_2}\left((1+\frac{1}{\phi})\frac{1}{y_2}-\frac{1}{\phi y_{2ms}}\right)\left(\frac{1}{y_{2ws}y_{2ms}}-(\frac{1}{y_{2ms}}+\frac{1}{y_{2ms}})\frac{1}{y_2}+\frac{2}{y_2^2}\right)\end{aligned}\right\}<0$$

Let Δ denote the expression in the brace, and let $\Delta_1,\Delta_2,\Delta_3$ denote the three terms in the brace respectively:

$$\Delta_1=2\frac{\delta_m}{\eta^{\max}-\eta^{\min}}\frac{1}{y_{2ms}^2y_2}(\frac{1}{y_{2ws}}-\frac{1}{y_2})$$
$$\Delta_2=\frac{\delta_w}{\eta^{\max}-\eta^{\min}}\frac{1}{y_{2ws}^2}(\frac{1}{y_{2ms}}-\frac{1}{y_2})\left((1+\frac{1}{\phi})\frac{1}{y_2}-\frac{1}{\phi y_{2ms}}\right)$$
$$\Delta_3=-\frac{1}{(\eta^{\max}-\eta^{\min})^2}\frac{1}{y_2}\left((1+\frac{1}{\phi})\frac{1}{y_2}-\frac{1}{\phi y_{2ms}}\right)\left(\frac{1}{y_{2ws}y_{2ms}}-(\frac{1}{y_{2ms}}+\frac{1}{y_{2ms}})\frac{1}{y_2}+\frac{2}{y_2^2}\right)$$

Note that $\Delta_1,\Delta_2>0,\ \Delta_3<0$. If the sex ratio is not too high, the probability of being married is not too low, so it is reasonable to say that:

$$\eta^{\max}-\eta^{\min}>\frac{1}{\delta_w}\ \text{and}\ \eta^{\max}-\eta^{\min}>\frac{1}{2\delta_m}$$

Since $\Delta=\Delta_1+\Delta_2+\Delta_3<0$, and $\Delta_1,\Delta_2>0,\ \Delta_3<0$, Δ would be greater than

$$\left.\begin{aligned}&\Delta=\Delta_1+\Delta_2+\Delta_3<0\\&\Delta_1,\Delta_2>0,\ \Delta_3<0\\&\eta^{\max}-\eta^{\min}>1\\&\eta^{\max}-\eta^{\min}>\frac{1}{\delta_w},\frac{1}{2\delta_m}\end{aligned}\right\}\ \Rightarrow\ \Delta>\frac{1}{2\delta_m}\Delta_1+\frac{1}{\delta_w}\Delta_2+(\eta^{\max}-\eta^{\min})\Delta_3$$

ii. Next we show that $\frac{1}{2\delta_m}\Delta_1+\frac{1}{\delta_w}\Delta_2+(\eta^{\max}-\eta^{\min})\Delta_3$ is greater than 0, which contradicts $\Delta<0$.

First we sum up the last two terms:

$$\frac{1}{\delta_w}\Delta_2+(\eta^{\max}-\eta^{\min})\Delta_3=\frac{1}{\eta^{\max}-\eta^{\min}}\frac{1}{y_2}\left((1+\frac{1}{\phi})\frac{1}{y_2}-\frac{1}{\phi y_{2ms}}\right)\left(\begin{array}{l}\frac{y_2}{y_{2ws}^2}(\frac{1}{y_{2ms}}-\frac{1}{y_2})\left((1+\frac{1}{\phi})\frac{1}{y_2}-\frac{1}{\phi y_{2ms}}\right)\\ -\left(\frac{1}{y_{2ws}y_{2ms}}-(\frac{1}{y_{2ms}}+\frac{1}{y_{2ms}})\frac{1}{y_2}+\frac{2}{y_2^2}\right)\end{array}\right)$$

$$>\frac{1}{\eta^{\max}-\eta^{\min}}\frac{1}{y_2}\left((1+\frac{1}{\phi})\frac{1}{y_2}-\frac{1}{\phi y_{2ms}}\right)\frac{y_{2ws}-y_{2ms}}{y_{2ws}y_2^2}$$

Then we have:

$$\begin{aligned}&\frac{1}{2\delta_m}\Delta_1+\frac{1}{\delta_w}\Delta_2+(\eta^{\max}-\eta^{\min})\Delta_3\\ &>\frac{1}{\eta^{\max}-\eta^{\min}}\frac{1}{y_2}\left[\frac{1}{y_{2ms}y_2}(\frac{1}{y_{2ws}}-\frac{1}{y_2})+\left(\frac{(1+\phi)y_{2ms}-y_2}{\phi y_2 y_{2ms}}\right)\frac{y_{2ws}-y_{2ms}}{y_{2ws}y_2^2}\right]\\ &>\frac{1}{\eta^{\max}-\eta^{\min}}\frac{1}{y_2}\left[\frac{1}{y_{2ms}}\frac{y_2-y_{2ws}}{y_{2ws}y_2^2}+\frac{1}{y_{2ms}}\frac{y_{2ws}-y_{2ms}}{y_{2ws}y_2^2}\right]\\ &=\frac{1}{\eta^{\max}-\eta^{\min}}\frac{1}{y_2}\bullet\frac{y_{2ws}}{y_{2ms}y_{2ws}y_2^2}\\ &>0\end{aligned}$$

Thus we have shown $\Delta>0$, contradiction.

3) Now we can proceed to investigate how the demand for house ownership reacts to a rise in the sex ratio.

$$\frac{dH_m}{d\phi}=\frac{A\Omega_{11}}{\det(\Omega)}>0,\quad \frac{dH_w}{d\phi}=-\frac{A\Omega_{21}}{\det(\Omega)}$$

The sign of $\frac{dH_w}{d\phi}$ is ambiguous since the sign of Ω_{21} is ambiguous. One additional result we can obtain here is that $H_m>H_w$ because $\Omega_{11}<\Omega_{21}$. The aggregate demand for house ownership is:

$$H_{agg}=\frac{1}{1+\phi}H_w+\frac{\phi}{1+\phi}H_m$$

The response of the aggregate demand for house is:

$$\begin{aligned}\frac{dH_{agg}}{d\phi}&=\frac{H_m-H_w}{(1+\phi)^2}+\frac{\phi-1}{1+\phi}\frac{dH_m}{d\phi}+\frac{1}{1+\phi}(\frac{dH_w}{d\phi}+\frac{dH_m}{d\phi})\\ &=\frac{H_m-H_w}{(1+\phi)^2}+\frac{\phi-1}{1+\phi}\frac{dH_m}{d\phi}+\frac{1}{1+\phi}\frac{A(\Omega_{11}-\Omega_{21})}{\det(\Omega)}\\ &=\frac{H_m-H_w}{(1+\phi)^2}+\frac{\phi-1}{1+\phi}\frac{dH_m}{d\phi}+\frac{A}{(1+\phi)\det(\Omega)}\bullet\end{aligned}$$

$$\left(-(1-\delta_w)\frac{1}{y_{2ws}^2}-\frac{f(\bar\eta_w)}{y_2^2}(\bar\eta_m+E(u_{2w})-u_{2ws})+2f(\bar\eta_w)\frac{1}{y_2}(\frac{1}{y_2}-\frac{1}{y_{2ws}})-f(\bar\eta_w)\frac{1}{y_2}\left((1+\frac{1}{\phi})\frac{1}{y_2}-\frac{1}{\phi y_{2ms}}\right)\right)$$

It is easy to see that the first two terms on the RHS are positive. Since $A<0$ and all terms in the bracket are less than zero, the third term is also positive. Therefore, the aggregate demand for house increases as the sex ratio rises.

We now come to the task of showing that the house price and the ratio of price to rent go up as the sex ratio increases. The house market clears if

$$\frac{1}{1+\phi}H_w+\frac{\phi}{1+\phi}H_m=H \tag{A.4}$$

The rental market clears if

$$\left(\frac{1}{1+\phi}h_{1,w}+\frac{\phi}{1+\phi}h_{1,m}\right)+\left(\frac{\delta_w}{1+\phi}h_2+\frac{1-\delta_w}{1+\phi}h_{2,w,s}+\frac{\phi-\delta_m}{1+\phi}h_{2,m,s}\right)=H \tag{A.5}$$

Totally differentiating (A.4) with respect to sex ratio ϕ yields

$$\frac{1}{(1+\phi)^2}(H_m-H_w)+\frac{\partial H_{agg}}{\partial\phi}+\frac{\partial p}{\partial\phi}\frac{\partial H_{agg}}{\partial p}+\frac{\partial r}{\partial\phi}\frac{\partial H_{agg}}{\partial r}=0 \tag{A.6}$$

where the aggregate demand for house is: $H_{agg}=\frac{1}{1+\phi}H_w+\frac{\phi}{1+\phi}H_m$

A higher price means a lower return for housing investment, so $\frac{\partial H_{agg}}{\partial p}\le 0$. On the other hand, a higher rent means a higher dividend from housing investment, so $\frac{\partial H_{agg}}{\partial r}\ge 0$. Moreover, we have already proved that $H_m-H_w>0$ and $\frac{\partial H_{agg}}{\partial\phi}\ge 0$. Now we turn to equation (A.5). The demand for housing service is always a certain fraction $1-\gamma$ of the total disposable wealth divided by rental rate. Summing up all terms on the LHS gives:

$$\frac{1-\gamma}{r}\left[y-(p-r)H+p(1-d)H\right]=H$$

$$\frac{1-\gamma}{r}(y-dpH+rH)=H$$

$$(1-\gamma)(y-pdH)=\gamma rH \tag{A.7}$$

where $y-(p-r)H$ is the wealth of the young cohort (note that we have set the savings to zero), and $p(1-d)$ is the wealth of old cohort. The rental rate can be set so that the demand to obtain living space by the old cohort equals the supply of living space by young cohort. Or equivalently, the total demand for living space by the two generations equals the fixed house supply. Differentiating with respect to ϕ:

$$\gamma\frac{\partial r}{\partial\phi}=-d(1-\gamma)\frac{\partial p}{\partial\phi} \tag{A.8}$$

So if there is no depreciation, the house rent should be a constant. Since $d>0, 0<\gamma<1$, we have $\frac{\partial r}{\partial\phi}\bullet\frac{\partial p}{\partial\phi}<0$.

Substituting (A.8) into (A.6) gives:

$$\textit{positive terms}=\textit{positive terms}\bullet\frac{\partial p}{\partial\phi}\;\Rightarrow\;\frac{\partial p}{\partial\phi}>0$$

Therefore, the price to rent ratio goes up as the sex ratio goes up.