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ABSTRACT

Household Portfolio Choice, Reference Dependence, and the Marriage Market

This paper bridges the financial market and the marriage market using a reference-dependent mechanism. Male-biased sex ratios induce families with sons to hold more risky assets, since competitive marital payment in a tight market raises the reference level of marriage expenditure for such families. Using the 2013 China Household Finance Survey data, we find that a 0.1 increase in the sex ratio raises the probability of participating in the stock market by 25.7 percent, or the stock share of liquid wealth by 42.7 percent for families with a son; there appears no effect for families with a daughter.

JEL Classification: D03, G02, G11

Keywords: household portfolio choice, reference dependence, prospect theory, sex-ratio imbalance, difference-in-differences estimate

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1 Introduction

Household portfolio choice is of great financial and socioeconomic significance, and has profound implications for welfare (Campbell, 2006, 2016). Previous literature has shown that household portfolio choice depends on many factors, such as participation costs, wealth and labor income, and housing. We study the interaction between the financial market and the marriage market, and propose a new determinant of household portfolios: competition in the marriage market for children. Using a reference-dependent mechanism, we bridge the financial market and the marriage market.

We assume that altruistic parents derive utility from expenditure on children's marriage contrasting with a reference level, which is the expected amount that is just sufficient for the child to find a partner. Parents regard it as a loss when current feasible marriage expenditure is below the reference level, and correspondingly consider it as a gain when above. Thus, parents evaluate marriage expenditure for children by processing gains and losses relative to this reference level, rather than the absolute amount.

The reference level of marriage expenditure represents the shadow price of finding a partner, which depends on marriage market competition (Becker, 2009; Browning et al., 2014). When the sex ratio is biased towards boys, the competition in the marriage market intensifies among males, which in turn increases the reference level of marriage expenditure for families with a son. This shift moves a fraction of families with a son into the loss domain; thus these families are more risk-tolerant and hold more risky assets in liquid portfolios as implied by loss aversion and diminishing sensitivity in the prospect theory (Kahneman and Tversky, 1979). We predict that the effect is the strongest for families with a son who can afford modest marriage expenditure, whose position on the value scale relative to the reference level is most likely to be altered by a change in the sex ratio.

Families with a daughter, on the other hand, are unlikely to respond strongly to the sex-ratio imbalance in their portfolios. Their reference level most probably remains unchanged, maybe because the desire to avoid erosion in bargaining power within the future marriage offsets the desire to free ride on marital payment from the prospective groom (Zhang and Chan, 1999).

To test the theoretical predictions, we explore a quasi-experiment based on the differential treatment between families with a son and those with a daughter under male-biased sex ratios in China. The sex ratio, which refers to the number of males to that of females, has risen drastically in recent decades throughout the country. The sex ratio at birth reached 1.12 in 1990, rose to 1.20 in 2000, and has remained at that level (Ebenstein, 2010).[1] Figure 1 shows a steep rise in the fraction of male births among second- and higher-order births, although the sex ratio for first-order births is balanced. The excess supply of males leads to keen competition for brides: more than 20 percent of men in China aged 30–39 would have never been married by the year 2030 (Sharygin et al., 2013).

Our empirical tests use the 2013 China Household Finance Survey (CHFS). The CHFS is both nationally and provincially representative of China, because of its large sample size, advanced sampling design, and low refusal rate. It provides rich contents in household portfolios, supplemented with information on various family background characteristics. We construct four portfolio-choice measures using the survey data: (i) stock market participation indicator; (ii) risky asset market participation indicator; (iii) the share of liquid wealth held in stocks; and (iv) the share of liquid wealth held in all categories of risky assets.

We regress portfolio measures on first-son dummy, the local sex ratio, and their interaction term, besides other controls. The interaction term indicates how son and daughter families react differently to the sex-ratio imbalance, comparable to a difference-in-differences (DID) strategy. Focusing on the interaction term is not subject to serious identification issues, given the elimination of potential confounding factors. Consequently, we rely primarily on the ordinary least squares (OLS) estimates. In particular, first-child gender is plausibly random, because the sex ratio for first-order births, unlike that for higher-order births, appears unbiased as shown in Figure 1. Moreover, we control for cross-sectional differences by including prefectural fixed effects, thereby exploiting within-prefecture variation in sex ratios for identification. We also highlight the independent role of the sex-ratio imbalance, by examining whether portfolio choice differs between son and daughter families if the sex ratio were balanced.

The estimation results are consistent with our predictions; the sex-ratio imbalance has a significantly stronger impact on risky asset holdings for son families than for daughter families. The portfolio differences are driven by both the extensive and intensive margins: among son families, we find that a 0.1 increase in the sex ratio (from 1.08 in 2002 to 1.18 in 2010) would raise the stock market participation rate by approximately 1.2 percentage points or 25.7 percent

1. In a society without social intervention, the sex ratio at birth is generally in the biologically normal range of 1.03–1.08.

relative to the sample mean, and the stock share of portfolios by 0.7 percentage point or 42.7 percent.

Stockholdings for daughter families, in contrast, do not seem to vary with the local sex ratio. In addition, there is virtually no difference in portfolios between son and daughter households when the sex ratio is balanced. The variation in portfolios is indeed due to changes in parental attitude towards risk, as suggested by consistent risk attitude results elicited from households' hypothetical financial decisions. We also show the minimal role played by identification issues.

We further investigate the mechanism of our results. A reference-dependent model predicts that the sex-ratio effect is the strongest for families with a son who can afford modest marriage expenditure, since they are likely to move from the gain domain to the loss domain after a change in sex ratio. We show that son families in the middle class behave more responsively to a rising sex ratio than the poorer and wealthier. This result cannot be rationalized by confounding interpretations (such as a negative income shock for families when a son is born in high-sex-ratio regions), and thus lends strong support to our conceptual framework. Supplementary results on housing and earning incentive indicate that parental expected expenditure on their son's marriage is significantly raised by the unfavorable marriage market situations.

This paper proceeds as follows. The next section reviews the literature. Section 3 describes China's sex-ratio imbalance and marriage market. Section 4 presents a reference-dependent model. Section 5 describes the data and empirical strategy. Section 6 reports the empirical results. Section 7 discusses the underlying mechanism. The paper concludes with a brief summary and policy discussions.

2 Literature

Our analysis is an application of the prospect theory, which implies that outcomes are evaluated in a manner in which the location of a reference level emerges as a critical factor (Kahneman and Tversky, 1979; Tversky and Kahneman, 1992). Kőszegi and Rabin (2006, 2007, 2009) endogenize the reference level as rational expectations; Pagel (2017) incorporates the expectation-based loss aversion into the macroeconomic model to study consumption from a life-cycle perspective. Recent laboratory experiments provide supporting evidence for an expectation-based reference level (Abeler et al., 2011; Song, 2015; Sprenger, 2015). Field research also supports such reference levels in various domains, including taxi drivers' labor supply (Crawford and Meng, 2011), high-stakes gambles (Post et al., 2008), insurance decision (Barseghyan et al., 2013), and professional golf (Pope and Schweitzer, 2011).

Our setting is closely related to Kőszegi and Rabin (2007). Parents form their expectation or reference level according to the children's marriage market situations, independent of the relevant choice set of their actual feasible marriage expenditures. This could be rationalized by the fact that the reference level of marriage expenditure is determined by a general equilibrium in the marriage market instead of individual choices (Browning et al., 2014). In this sense, our analysis is the limiting case of the Unacclimating Personal Equilibrium (UPE) behavior.[2]

An important application of the prospect theory in stock trading settings is to explain the disposition effect: the propensity of investors to sell stocks that have risen in value rather than stocks that have fallen since purchase (Feng and Seasholes, 2005; Grinblatt and Keloharju, 2001; Meng and Weng, 2016; Odean, 1998; Shefrin and Statman, 1985; Weber and Camerer, 1998). Our study adds to this literature by examining a family's general choice to hold wealth in risky financial assets, taking into account children's marriage prospects.

Research documents how portfolio choice is affected by participation costs (Campbell et al., 2001; Cocco, 2005), borrowing costs (Davis et al., 2006), nonstandard preferences such as habit formation (Gomes and Michaelides, 2003; Polkovnichenko, 2007) and heterogeneous risk preferences (Gomes and Michaelides, 2005), labor income (Benzoni et al., 2007; Campbell et al., 2001; Cocco, 2005; Cocco et al., 2005), personal experience (Malmendier and Nagel, 2011), and housing (Chetty et al., 2016; Cocco, 2005). This paper adds to the literature by investigating how marriage market conditions for children affect household portfolio choice through a reference-dependent mechanism. Specifically, we estimate that a 0.1 increase in the sex ratio has a much larger impact for an average son household on the stock share of liquid wealth than a one standard deviation reduction in mortgage debt (Chetty et al., 2016), or a one standard deviation rise in log financial wealth (Calvet et al., 2007), in terms of percentage increases to the mean.

2. See Kőszegi and Rabin (2006, 2007, 2009) for rational expectations equilibrium concepts. The UPE is the personal equilibrium in which subjects' choices correspond to expectations. The Preferred Personal Equilibrium (PPE) is the UPE with the highest ex-ante expected utility. The Choice-acclimating Personal Equilibrium (CPE) is the UPE in which choices are committed well in advance of the resolution of uncertainty. If households were in the CPE, the reference level would adjust to their own choices; they would always be around the reference level facing balanced or unbalanced sex ratios. We in this case could not observe the sex-ratio effect on portfolios for all families.

Our study is also closely related to the recent burgeoning literature on the sex-ratio imbalance in China and other Asian societies (Bethmann and Kvasnicka, 2013; Bhaskar and Hopkins, 2016; Chiappori et al., 2002; Ebenstein, 2010, 2011; Edlund et al., 2013; Das Gupta, 2005; Das Gupta et al., 2013; Hu and Schlosser, 2015; Huang and Zhou, 2016; Sharygin et al., 2013). In particular, our study builds on Wei and Zhang (2011), who find that parents decrease consumption to enhance their sons' competitiveness facing a male-biased sex ratio. We further show that families with a son allocate more liquid wealth to risky assets relative to those with a daughter when the sex ratio is more biased towards males.

3 Background

3.1 Sex-ratio imbalance

The sex-ratio imbalance, first noted by Sen (1992), is a subject of extensive studies, given its large scale and wide-ranging implications. Sen (2003) predicts that about 100 million women worldwide would be missing in the 2000s. In China in particular, the sex ratio at birth reached 1.12 in 1990, rose to 1.20 in 2000, and has remained at that level (Ebenstein, 2010). For children below the age of 15, the number of males exceeds the number of females by 13 percent. The literature extensively investigates the socioeconomic, institutional, and cultural factors affecting the demand for sons.[3] For example, sons are expected to shoulder more responsibility of caring for elderly parents, and these expectations may contribute to the persistency of son preferences (Das Gupta et al., 2013). Ebenstein and Leung (2010) and Chew et al. (2016) show that a shortage of old-age pension programs increases parental demand for sons. Additionally, parental perceptions that sons have a higher earning capacity than daughters exacerbate preferences for sons. Historically, physical labor was valued as most jobs were in agriculture; men have comparative advantages in these brawn-intensive activities and hence anticipate higher wages (Pitt et al., 2012).

In China, the sex ratio is further distorted by its family planning policy, commonly known as the one-child policy, that stipulates the number of children a couple is allowed to have (Li et al., 2011). Parents may undertake gender selection to satisfy their dual interest in complying

3. See, for instance, Rosenzweig and Schultz (1982) and Ahn (1995). Leung (1988, 1991) proposes a method to detect son preferences using fertility data. Davies and Zhang (1995) distinguish modelling son preferences in budget constraints and utility functions.

with the stipulated birth quota and in having a son. This exercise is made viable by improved access to gender-screening technology (in particular, ultrasound B) and abortions (Ebenstein, 2010).[4]

3.2 Marriage market competition

In China, the sex-ratio imbalance has led to a prevalent oversupply of marriage-age males, which foreshadows a sizable bride shortage. Consequently, competition in the marriage market has intensified, bidding up marriage expenditure. Data from the 2010 China Family Panel Studies (CFPS) survey, which is regarded as nationally representative of Chinese communities, families, and individuals, show that household-level expenditure on wedding ceremonies increases by about 24 percent as the local sex ratio rises by 0.1;[5] see Figure 2 for a graphical elucidation. In particular, as the sex ratio rises, the burden of marriage expenditure is increasingly falling on grooms' families; see Figure 3 and Brown et al. (2011). Such expenditure is significant. In 2006, for example, the average wedding cost for a groom's family was approximately 5.5 times per capita income.

Marriage expenditure is mainly transfers to the bride and her parents from the groom's family, in the form of a bride price. The negotiation of a bride price is traditionally required, and continues to be the norm in most areas today.[6] As the sex-ratio imbalance deteriorates, marital transfers are used to compete for prospective brides. Competing marital transfers typically include a one-time payment to the bride's parents compensating them for raising their daughter (Zhang and Chan, 1999), and are usually accompanied by gifts such as major durable goods. Grooms' families also bear most of the financial burden for wedding ceremonies, which are occasions that call for a large amount of cash outlay. In China, the most crucial component of marriage expenditure is housing, which is also the largest lifetime financial commitment for most families. Social norms and marriage customs demand that a groom's family provides an apartment for the newlyweds, or at least shoulders more than half the cost. Households are much less likely to have an unmarried adult son in rural areas if they have a relatively higher-quality house; in urban areas, households are less likely so if they are a homeowner as opposed

4. Prior to the widespread use of ultrasound B machines in the 1980s, infanticide was one of the methods used to exercise gender selection (Coale and Banister, 1994; Scharping, 2013).

5. We control for regional characteristics such as local average household income and population composition, as well as province fixed effects.

6. Although dowries may coexist in China, they tend to be voluntary and typically financed with a return portion of the bride price (Anderson, 2007; Engel, 1984).

to a renter (Wei and Zhang, 2011).

4 Conceptual framework

4.1 A simple model of reference-dependent preferences

As illustrated in the preceding section, a large marriage expenditure enhances a man's marriage prospects in a competitive marriage market. Given that parents are altruistic and care about children's marriage, we assume for simplicity that they derive utility from marriage expenditure for their children, x. Specifically, parental utility over the child's marriage expenditure follows the reference-dependent value function proposed by Tversky and Kahneman (1992),

$$v(x|r) = \begin{cases} (x-r)^{\alpha} & \text{for } x \geq r, \\ -\lambda \cdot (r-x)^{\alpha} & \text{for } x < r, \end{cases}$$

where r denotes the reference level of marriage expenditure for the child, and α and $\lambda > 1$ are parameters. [7]

The reference level, r, is defined as the expected amount of marriage expenditure that just enables a young person (the child) to be married in the local marriage market. Parental assessment of gains or losses is determined by this level. Parents regard it as a loss when current feasible marriage expenditure is below the reference level, and correspondingly consider it as a gain when above. Intuitively, parents evaluate marriage expenditure for their children by processing gains and losses relative to this amount, rather than based on the absolute level. Attaining the reference level means parents spending just enough to ensure the child has a successful marriage. Meanwhile, exceeding the reference level indicates a sharp decline in parental marginal utility of any additional marriage expenditure for the child. This is the notion of loss aversion, which is shown by the kink at the reference level in the S-shaped value function in Figure 4. The kink emerges because the value function is steeper in losses than in gains, as indicated by $\lambda > 1$. In addition, the value function is convex in losses and concave in gains, in line with the notion of diminishing sensitivity.

Based on the marriage matching theory in Becker (2009) and Browning et al. (2014), the

7. We assume linear probability weighting for simplicity. Tversky and Kahneman (1992) estimate that $\alpha = 0.88$ and $\lambda = 2.25$ from experimental data.

reference level of marriage expenditure, r, represents the shadow price of finding a partner. It is an equilibrium outcome that depends on the supply and demand in the marriage market, which is proxied by the sex ratio. Moreover, the reference level is gender specific. An increase in the sex ratio increases the marriage pressure for young males and thus their reference level; there might be no effect or a negative effect on the reference level for young females. Therefore, a reduced-form relationship exists between the reference level and child gender, such that $r = r(sexratio, gender, v)$, where v is a random error when parents form their expectation of the reference level. In this way, we assume that the reference level is parental expectation of their children's marriage expenditure based on the marriage market situations (Kőszegi and Rabin, 2007).

As the sex ratio is biased towards young-generation males, their parents have to compete by offering large marriage payments representing a substantial fraction of family wealth. The reference level would thus increase for son families in regions with higher sex ratios, as demonstrated by a right shift of their reference level from the dashed to the solid plot of the value function in Figure 4.

4.2 Model predictions

To study the effect of an increase in the reference level of marriage expenditure on household portfolios for families with a son, we divide these families into three groups based on the comparison between the reference level and current feasible marriage expenditure, depending on whether the latter is far below, around, or much higher than the former; see Figure 4 for hypothetical bounds over the horizontal axis of the three regions.

Region I includes son families whose actual marriage expenditure for their sons would be far below the reference level. For these families, parents are unlikely to be satisfied with the actual marriage expenditure for their sons; they lie in the loss domain regardless of whether the sex ratio is biased (the solid line) or unbiased (the dashed line).[8] An increase in the reference level has little effect on their risk attitude, as in neither case can they reach the reference level by taking more risk in portfolios. Region III contains son families for whom marriage expense for their sons does not represent a heavy financial burden. A higher reference level of marriage

8. This is consistent with Siow's (1998) view that some men cannot be married even at a balanced sex ratio, because females biologically are fecund for a shorter period of their lives than males, and thus fecund women are relatively scarce.

expenditure has little effect on their risk attitude, since they are always in the gain domain regardless of the local level of marriage competition.[9]

Region II represents the group of son families in the middle, whose actual marriage expenditure for their sons is modest and in close proximity to the reference level. They are most likely in the gain domain when the sex ratio is balanced (the dashed line), and in the loss domain when the sex ratio increases and the reference level moves right (the solid line). Stepping from the gain domain to the loss indicates a shift in risk attitude. These families will try to reach the reference level by taking more risk in their portfolios, because the marginal utility of additional marriage expenditure is much higher in the current loss domain than in the previous gain domain, as suggested by loss aversion. A higher reference level leads more son households to region II, who move from the gain domain to the loss domain, and hence adopt a more aggressive risky asset holding strategy. The overall effect for son families of an increase in the reference level on household portfolios would be driven by families in region II, since those in regions I and III are unlikely to respond.

In contrast, a high sex ratio likely has an insignificant net impact on the reference level for families with a daughter, and correspondingly, their risk tolerance. The insignificant effect may be due to two conflicting motives. On one hand, parents possibly expect to spend less on their daughters' marriages, as females expect to receive substantial marital payments from their future husbands. This motive for daughter households potentially shifts the reference level left. On the other hand, parents may wish to spend more on their daughters' marriages to avoid erosion of their bargaining position, which is affected by each spouse's relative contribution to the marriage (Zhang and Chan, 1999). This motive potentially shifts the reference level right.

In summary, our model gives the following two predictions.

Prediction 1: Higher sex ratios make families with a son (but not families with a daughter) more risk-taking, which would be revealed by a more aggressive investment strategy, *i.e.*, holding more risky assets.[10]

Prediction 2: The sex-ratio effect is the strongest for families with a son who can afford modest marriage expenditure, since they are more likely to change from gains to losses relative

9. More precisely, in regions I and III where son households stay on the same side regardless of the sex ratio, an increase in the reference level of marriage expenditure has no effect on their risk attitude when the value function satisfies constant absolute risk aversion; they would become (slightly) less risk averse when it satisfies increasing absolute risk aversion, and more risk averse when it satisfies decreasing absolute risk aversion.

10. While intuitively plausible, there is no *a priori* guarantee of an insignificant sex-ratio effect on portfolios for daughter families; we test this empirically in the following section.

to the reference level.

5 Data and empirical strategy

5.1 China Household Finance Survey

The data used in our empirical analysis are drawn from the 2013 CHFS. It is both nationally and provincially representative, mainly due to its large sample size, advanced sampling design, and low refusal rate. Gan et al. (2013) show that the survey data on household income as well as many important demographic characteristics, such as the distribution of household size and age structure, are consistent with the China Population Census, which is widely recognized to be representative. We can thus generalize our empirical findings to the Chinese population. The CHFS was launched in 2011, and expanded its sample size drastically in 2013, making it the largest household survey in China. The 2013 wave consists of a sample of 28,228 households, covering 29 designated provinces, 262 counties, and 1,048 communities.[11]

The CHFS implements a scientific three-stage probability-proportional-to-size sampling design. The first-stage sampling unit is administrative county, the second-stage is neighborhood community, and the third-stage is household. In the first stage, sampling units are sorted by local GDP per capita, in which the local population is used as sampling weights. The second stage basically follows the same principle, but sampling units are sorted by the proportion of nonagricultural population. The third sampling stage is a systematic selection of households with equal probability.

The CHFS has the lowest refusal rate among internationally comparable surveys. The overall refusal rate in the 2013 wave is around 11 percent: 15.4 percent in urban areas and one percent in rural areas. In comparison, in the most recently conducted 2010 and 2013 Survey of Consumer Finances (SCF) in the U.S., about 30 percent of households selected for the area-probability sample refused to participate. The refusal rate in the list sample was about one-third; in the high-wealth sample, the refusal rate was double that level.[12] In the 2010 Eurosystem Household Finance and Consumption Survey, the refusal rate can be as high as 69.7 percent in Germany;

11. Note that the administrative divisions of China, due to its large population and area, consist of four practical levels: the province-level unit, prefecture, county, and community. There are 34 province-level units in total; Hong Kong, Macao, Taiwan, Xinjiang, and Tibet are excluded in the survey.

12. The SCF conducted by the Federal Reserve Board is acknowledged to be the U.S. survey with the most detailed data on financial wealth. Data on refusal rates are from the Federal Reserve Bulletin (2014).

countries with some of the lowest refusal rates are Portugal (10.3 percent) and Finland (11.1 percent).[13]

Additionally, the CHFS has comprehensive data on household portfolios. The questionnaire is designed to collect detailed information on households' financial activities, including stock and other financial asset market participation, housing value and mortgage, as well as income, consumption, and wealth. The CHFS also has information on demographic and labor-market characteristics of all family members, such as age, gender, schooling years, occupation, and geographic location. It offers a household identification number, which allows us to group individuals by living unit. Parent-child relationship can also be precisely identified in each family. Household portfolio choice is thus readily linked with potential covariates, enabling a systematic empirical analysis.

For each family involved in the CHFS, questions regarding household portfolios are answered by a main respondent, who has the best knowledge of the family's financial status and is most likely to be the household head. Information on demographics is obtained from the respondent as well as family members present during the interview. The face-to-face interviews are aided by the computer-assisted-personal-interviewing technology, which has been widely used in large-scale surveys such as the SCF in the U.S. This technology enables diversified questionnaire designs, efficient sample management in the field, and real-time supervision of the interviewer. It also automatically collects paradata during an interview, providing information of better quality than the traditional pen-and-paper method (Caeyers et al., 2012).

5.1.1 Sample restrictions and portfolio-choice measure definitions

Throughout this paper, the unit in the empirical analysis is household. For our purpose, we extract a main sample that includes all households in the 2013 CHFS in which the first-born child was aged 0–17 and both parents were aged 20–49.[14] We focus on families in which the eldest child was younger than 18 years old to rule out cases where children might participate actively in financial decisions. By placing age limits on parents and children, we also maximize comparability across families. Overall, the main sample contains 5,092 observations, and the

13. For more information on this dataset and the refusal rates, see Eurosystem Household Finance and Consumption Network (2013a, 2013b).

14. Because we are primarily interested in first-child gender, we exclude families in which first births are twins.

household head was either the father or mother in all these families.[15]

The key outcome variables in our analysis are four portfolio-choice measures. Specifically, we examine household allocation to risky financial assets, particularly stocks. Risky assets, besides stocks, include risky bonds, mutual funds, financial derivatives, etc. For stocks, the first measure is a binary variable for market participation, indicating whether the household owns a stock account by the survey date;[16] the second measure is the stock share, defined as the fraction of liquid wealth currently held in stocks. For the stock share, we impose no constraint on the household's stock market participation status; it is simply set to zero for nonparticipants as in Chetty et al. (2016). The resulting unconditional estimate gives the population average causal effect on the stock share without raising any selection bias (Angrist and Pischke, 2008).[17] Analogously, another two portfolio-choice measures are defined for all categories of risky assets: risky asset market participation dummy and the risky asset share of liquid wealth. Portfolio shares are specified to be zero for households whose reported liquid wealth by our definition is zero.

We define liquid wealth as the sum of holdings in risky assets plus holdings in riskless assets, which consist of cash and savings accounts. In our analysis below, we define gross financial wealth as the sum of liquid and illiquid assets. Illiquid assets, following common practice, primarily include residential real estate, major consumer durables such as cars, and agricultural properties. This gross financial wealth measure does not subtract mortgage or other household debt.

5.1.2 Summary statistics

Table 1 presents summary statistics for household portfolios and financial assets, in Panel A for the main sample and in Panel B for the subsample of stock market participants, *i.e.*, families that hold a stock account. As Panel A shows, the proportion of households participating in the stock market and in the risky asset market is 4.8 and 7.5 percent, respectively. These patterns

15. Trimming outliers, for example, by excluding the top and bottom one percent of households by wealth, has virtually no effect on our estimates.

16. The stock market participation measure equals one as long as the household has a stock account, regardless of the worth of stocks currently held; there is a negligible effect on the estimation results if this variable is instead defined to indicate a positive worth of stocks.

17. The overall effect can be decomposed into two parts: the effect on participation probability and the effect on the stock share conditional on participation. Many studies focus on the latter, yet the resulting estimates may not have a clean causal interpretation. Even when using censored regression models such as Tobit to deal with non-causality, the estimation still has to rely on a postulated latent variable and implausible distributional assumptions.

are consistent with the common finding that household investment in risk assets is generally limited. Since we apply the CHFS sampling weights, these numbers represent Chinese families, and the weighted estimates obtained in the empirical tests below are also representative of the population at large.[18] Households hold on average approximately 1.6 percent of their liquid wealth in the form of stocks and 2.8 percent in all kinds of risky financial assets. Mean liquid wealth is 37,878 Chinese *yuan*; yet this distribution is skewed and the median level is 5,000 *yuan*. Mean gross financial wealth is 289,200 *yuan* and median is 105,000 *yuan*.

A comparison of Panels A and B shows that stock market participants, with a median holding of 78,000 *yuan* in liquid assets, tend to be wealthier than nonparticipants. The pattern is similar for gross financial wealth. In addition, Panel B shows that among stockholders, the share of liquid assets held in stocks is 33.5 percent, while in all categories of risky assets is 41 percent. While a large fraction of households avoid participation at all, for those who invest, holdings as a share of their liquid wealth are significant.

Column (1) of Table 2 reports descriptive statistics for family demographic characteristics, *i.e.*, potential explanatory variables in the empirical estimations. In our main sample, nearly half of families have a first-born son while the other half have a first-born daughter. Of all households, 41.1 percent reside in rural areas, and 12.1 percent are ethnic minorities. An average first-born child is around 11 years old. Parental characteristics are also presented. For example, fathers, on average, are 38 years old and have 9.6 years of schooling, while mothers are 37 years old and have 8.9 years of schooling. Also, 29.7 percent of fathers and 28.1 percent of mothers are registered as urban residents according to China's *hukou* system, the household registration system that is used to differentiate permits of where people are allowed to live and work. Finally, 12.5 percent of fathers and 4.0 percent of mothers are members of the Chinese Communist Party.

18. The unweighted stock market participation rate in the CHFS is higher because high-income households that have a disproportionate influence on stock demands were oversampled, like in the SCF. Thus the number of observations in Panel B is larger than 4.8 percent of the number of observations in Panel A. Using an unweighted sample may result in unrepresentative average treatment effect; this concern outweighs the loss of efficiency associated with weighted estimates (Malmendier and Nagel, 2011). The robustness regressions based on the unweighted sample show, however, that weighting makes virtually no difference for our key results.

5.2 Regression specifications

We estimate the following regression equations:

$$y_{icp} = \beta_0 + \beta_1 first\_son_{icp} + \beta_2 sex\_ratio_c + \beta_3 first\_son_{icp} * sex\_ratio_c + X_{icp}\Gamma + \lambda_p + \epsilon_{icp}, \quad (1)$$

where the dependent variable, y_{icp}, measures portfolio holdings, and particularly stockholdings, of household i in county c of prefecture p. A key explanatory variable, $first\_son_{icp}$, is a binary indicator that equals one if the first-born child is a son. Another key explanatory variable, sex_ratio_{cp}, refers to the county-level sex ratio for the appropriate age cohort. A vector of additional control variables, X_{icp}, includes various parental (age, schooling years, occupation dummies, and party dummies) and household characteristics (a dummy indicating ethnic minorities, *i.e.*, either the father or mother is non-*Han* Chinese, a dummy indicating rural region of residence, and dummies for the age of the first child). In the robustness checks we add more regressors, and the pattern of results is preserved.[19] All specifications include prefecture fixed effects, λ_p. The error term, ϵ_{icp}, captures other sources of unobservable heterogeneity in portfolios, possibly including entrepreneurial risk (Heaton and Lucas, 2000), investment mistakes (Calvet et al., 2007; Odean, 1999), or measurement error in income (Cocco, 2005).

We treat each county as a local marriage market, and explore sex-ratio variation across counties. The marriage market is indeed local for most people, as China's *hukou* system presents a formidable obstacle to marriage migration (Davin, 2005). According to the 2000 census, 92 percent of rural residents and 62 percent of urban residents live in their county of birth; 89 percent of marriages take place between two people from the same county. Moreover, among migrant couples in cities, 82 percent are from the same province, suggesting that migrants often get married before leaving their hometown (Wei and Zhang, 2011).

The local sex ratio is obtained for the marriage-age cohort between the ages of 18 and 32 years in our main regressions. Parents do not necessarily know the exact local sex ratio statistics for the cohort of their own children, and are more likely to gather information based on the difficulty with which their relatives' or colleagues' adult sons found wives, or the prevailing large marriage expenditure that signals a fierce competition in the marriage market, and respond accordingly.

19. In the main regressions, we exclude regressors that may be endogenous, such as wealth and income controls and the number of children. Particularly, wealth and income levels are closely correlated with parental schooling years. Our results remain almost the same when we control for these variables.

Moreover, because of the persistence over time of the local level of the sex-ratio distortion, the estimation results are likely insensitive to the particular age bracket. Prior empirical studies show that using sex ratios in different segments of the cohort gives qualitatively similar findings; see Wei and Zhang (2011). Data on the local sex ratios are projected from the 2010 census: in each county, the sex ratio for the group of youths aged 18–32 in 2013 is calculated using the group aged 15–29 in 2010, since these two cohorts are supposed to be the same.

We estimate our regressions using the OLS method. Although almost every outcome in this paper is either binary (such as stock market participation dummy) or non-negative (such as the stock share), the estimation of average causal effect presents no special challenges using linear models, according to Angrist and Pischke's (2008) view of regressions as inheriting legitimacy from the conditional expectation function.[20] Only the interpretation of the coefficient estimates changes for different sorts of dependent variables. For example, when the outcome is a dummy for stock market participation, the average causal effect is on stock-holding probability or stock market participation rate; in this case, the regression becomes a linear probability model (LPM).

5.3 Empirical testing

To connect the empirical analysis to the predictions in Section 4, we examine the coefficient estimates in equation (1). The focus is on the sign, statistical significance, and economic magnitude. Testing Prediction 1 relies on coefficients on the sex ratio for different types of families: $\beta_2 + \beta_3$ determines the sex-ratio effect on household portfolios for families with a son and β_2 for those with a daughter, respectively. For families with a son, the sex-ratio imbalance is predicted to have a substantial impact on portfolio shares, because the reference level of marriage expenditure is higher and more difficult to arrive at for those in regions with higher sex ratios. On the contrary, daughter families may not react to a sex-ratio imbalance, since the local marriage market is always not unfavorable for them, and their reference level of marriage expenditure does not necessarily vary with the sex ratio. Therefore we expect $\beta_2 + \beta_3$ to be positive and statistically significant, and β_2 to be small and insignificant. In a DID sense, the cross-term coefficient β_3 measures the difference in the sex-ratio effect on portfolios between son and daughter families. We expect β_3 to be positive and statistically significant. To test

20. The differences between the standardized OLS estimates and marginal effects from nonlinear censored regression models, such as Probit or Tobit, seem unlikely to be of substantive importance in most empirical practices (Angrist and Pischke, 2008). This is also the case in our estimations.

Prediction 2 on the heterogeneity of the sex-ratio effect, we compare the estimated magnitudes of $\beta_2 + \beta_3$ for son families across different wealth levels.

We also examine the possibility that the differential in portfolio shares is purely due to child gender, by studying $\beta_1 + \beta_3$, which denotes the difference in portfolios between son and daughter families at a balanced sex ratio. Since it takes a combination of having a son and experiencing competition for brides to induce a change in household risk attitude, child gender in and of itself would not evidently influence family financial decisions in the absence of a sex-ratio distortion. We therefore expect $\beta_1 + \beta_3$ to be not significantly different from zero.

Finally, we perform F-tests that involve child gender, the sex ratio, and their interaction term to evaluate the joint explanatory power. The null hypothesis is that none of these variables have a considerable impact on risky asset holdings; *i.e.*, their coefficients are simultaneously equal to zero ($H_0 : \beta_1 = \beta_2 = \beta_3 = 0$). We expect to reject the null.

5.4 Identification assumptions

Obtaining unbiased OLS estimates of β_1, β_2, and β_3 in equation (1) requires that the error term is not substantially correlated with the key independent variables, first-son dummy and the local sex ratio. Note first that our specifications reduce the potential endogeneity problems as we control for prefecture fixed effects.[21] These prefecture dummies are expected to catch numerous sources of regional heterogeneities that could account for variation in household portfolios, including heterogeneity in average risk tolerance. They also greatly absorb other location-specific omitted or unobserved factors, such as cross-region difference in parental sophistication, that might simultaneously impact household fertility decision and portfolio choice. In this part, we discuss in detail the two crucial regressors to further stress the validity of our identification assumptions.

5.4.1 Randomness of first-child gender

The imbalance of the overall sex ratio in China is driven by gender selections on the second- or higher-order births, as can be seen from Figure 1. The sex ratio of first-born children, however, is rather stable and falls in the biologically normal range, as revealed by the four recent censuses

21. In China, each province comprises approximately ten prefectures, while each prefecture comprises approximately nine counties. Counties within a prefecture are more homogenous than those within a province. Therefore, geographical heterogeneities are better handled by the inclusion of prefecture dummies than province dummies.

(1982, 1990, 2000, and 2010). Because of the nuances of China's birth control policy, families are least likely to practice gender selection on the first birth even though some parents may prefer sons. Specifically, for households in most rural areas, a second child is officially permitted if the first one is a daughter. Son preferences appear more common among rural residents, and this "1.5 children" policy alleviates their motivation to abort the first daughter.[22] Consistently, statistics show that an average family in our sample has about 1.5 children, manifesting the representativeness of the sample. Based on these discussions, first-child gender can arguably be considered as random (Ebenstein, 2010, 2011; Li and Wu, 2011; Wei and Zhang, 2011).

More importantly, statistical evidence confirms the randomness of first-child gender in our main sample. For example, the mean of first-son dummy being 0.519 as reported in column (1) of Table 2 implies a sex ratio of first-born children of 1.079, which is within the biologically normal range. A more interesting support is that the standard deviation of first-son dummy is 0.5. For a Bernoulli random variable with a mean of 0.5, its standard deviation equals 0.5 ($\sqrt{0.5(1-0.5)}$). First-child gender is thus most likely a random Bernoulli trial in which there is equal probability of having a son or a daughter. In addition, we regress first-son dummy on the full list of controls used in our regression analysis, and find no significant effect of these variables.

Furthermore, predetermined family and parental characteristics for first-son and first-daughter households are statistically similar in our sample. Columns (2) and (3) of Table 2 present means as well as standard deviations of the main demographic variables for first-son and first-daughter families, respectively. Column (4) reports the differences between the means (son minus daughter). These differences are negligible in magnitude. For example, 40.8 percent of son households and 41.4 percent of daughter households are in rural regions, giving rise to a difference of only -0.6 percent. Indeed, a balance test shows that the differences across son and daughter families are not statistically significant at a ten percent level or less, as reported in column (5).

5.4.2 Sex ratio

It can be recognized that the sex ratio is not, strictly speaking, exogenous; see Section 3.1. Consequently, an important concern with our research design is that portfolio choice and the

22. Ebenstein (2011) shows that many families prefer one boy and one girl to two boys. Chu (2001) finds that too many boys are no longer desirable in modern China, given the fact that rearing a son is much more costly than rearing a daughter.

county-level sex ratio are simultaneously affected by certain omitted factors. Some studies attempt to deal with this possible endogeneity bias by instrumenting for local sex ratios using regional variation in financial penalty levied for unauthorized births, or regional variation in the cost of gender-screening technology. Nonetheless, these candidates may not be valid instruments, partly due to their potential correlation with household portfolios independent of the sex ratio. Moreover, birth quotas in China are stipulated by regional governments based on local conditions.[23] To avoid these problems with instruments, and for several reasons that we shall discuss below, the OLS estimates are of the first-order importance in our work, as standard in the empirical literature (Edlund et al., 2013; Wei and Zhang, 2011).[24]

Our regression specifications alleviate identification concerns with the OLS estimates involving the sex ratio. As indicated above, a variety of confounding factors related to the sex ratio, such as the local implementation of China's family planning policy, are fairly well handled by the inclusion of prefecture fixed effects. Specifically, we identify the parameters of interest purely from within-prefecture variation in sex ratios, and counties within a prefecture are more homogenous in many aspects. In addition, the estimates of the interaction-term coefficient β_3 are obtained from regressions that are comparable to a DID strategy: by implicitly contrasting the sex-ratio effect for son families with the effect for daughter families, the impacts of omitted sex-ratio confounders are substantially eliminated. In this sense, we refer to β_3 as a DID estimate in the rest of the paper. Moreover, our specifications enable us to compare portfolios across son and daughter households at a balanced sex ratio to bring out the independent explanatory power of the sex-ratio distortion.

We further discuss and try to rule out various factors that may confound local sex ratios and at the same time directly impact household portfolios; see Section 6.3.3. We address the concerns by including in the robustness regressions observed confounders such as local average household income, old-age support, gender differential in earnings, and fertility, as well as by excluding prefecture dummies. Hausman's general specification test for the equality of the interaction-term coefficient estimates reveals no significant differences between these estimations and the benchmark. The robustness analyses demonstrate that our findings are not mainly driven by

23. As remarked in Short and Zhai (1998), China's family planning policy is passed down the administrative chain of command until it is interpreted, adapted, and implemented to suit local needs.

24. An important caveat is that with possible measurement errors in local sex ratios, the estimated sex-ratio effect may be biased downward. So, our estimates should be interpreted as a lower bound of the sex-ratio effect for son families in terms of absolute values.

identification problems regarding the sex ratio.

6 Empirical results

This section presents the empirical results on the overall sex-ratio effect on portfolio choice for families with a son and those with a daughter, respectively, and the differential across families in such effect. Consistent findings on hypothetical household financial decisions are reported as supplementary evidence for changes in parental risk attitude. We first provide some preliminary evidence. We divide the main sample into four groups of families according to the potentially differential marriage market conditions they are faced with: (i) first-son families in a region with a balanced sex ratio; (ii) first-daughter families in a region with a balanced sex ratio; (iii) first-son families in a region with a high sex ratio; and (iv) first-daughter families in a region with a high sex ratio. Panel C of Table 1 shows summary statistics of portfolio-choice measures for these four subsamples, where we explicitly define balanced- or high-sex-ratio regions depending on whether the local sex ratio is below or above the median. Compared with son families in balanced-sex-ratio regions, those in high-sex-ratio regions appear more likely to hold risky assets. For daughter families, in sharp contrast, portfolio choice does not evidently vary with the sex ratio.

Figure 5 graphically illustrates the relationship between the local sex ratio and stockholdings for families with a son and those with a daughter. Local stock market participation rate among first-son families clearly rises with the sex ratio. In comparison, stockholdings among daughter families barely respond to the local sex ratio. The patterns in Panel C of Table 1 and Figure 5 are in line with Prediction 1 that son families hold more stocks in their portfolios as the sex ratio increases, whereas daughter families do not respond. However, using a regression framework instead of simple averages allows us to control for covariates, and also include prefecture fixed effects to absorb unobserved determinants of portfolio choice.

6.1 Household portfolio choice

We test our first prediction on the sex-ratio effect. The results from estimating equation (1) on our main sample are presented in Table 3. We only report, for brevity, the estimates of the key parameters, *i.e.* β_1, β_2, and β_3, at the top of the table.[25] In the middle we present the implied effect of the local sex ratio on portfolio choice of families with a son, reported in both percentage points, $\beta_2 + \beta_3$, and percentages, $(\beta_2 + \beta_3)$ divided by the mean of the dependent variable. We also report the effect at a balanced sex ratio of having a first son on household portfolios, $\beta_1 + \beta_3$, as well as the corresponding F-statistics for testing $\beta_1 = \beta_2 = \beta_3 = 0$.

6.1.1 Stocks

In the first column of Table 3, we regress stock market participation indicator on first-son dummy, the local sex ratio, their interaction term, as well as other controls mentioned above. Column (2) concerns the current stock share of liquid wealth. The estimate is 0.144 (standard error 0.055) for the coefficient on the interaction term in column (1) and 0.037 (standard error 0.019) in column (2). The positivity and statistical significance (at least at a ten percent level) of these DID estimates suggest that the sex-ratio imbalance has a substantially stronger impact on stockholdings for families with a first-born son than for those with a first-born daughter. These patterns are the same as predicted in Section 4.

As shown in the middle of columns (1) and (2), the local sex ratio has a positive and highly significant effect on stock market results for son families. This pattern is again consistent with our prediction. Column (1) shows that a 0.1 rise in the sex ratio would, on average, increase the probability that families with a son hold a stock account by roughly 1.2 percentage points. This estimate represents a 25.7 percent increase relative to the baseline stock ownership rate of 4.8 percent in our sample. Column (2) shows that the stock share of portfolios for son households would be 0.7 percentage point higher with a 0.1 increase in the local sex ratio. While this effect may appear small in absolute terms, it is substantial in percentage because many families do not hold any stock: the estimate constitutes a 42.7 percent increase to the sample mean stock share. In comparison, smaller and statistically insignificant coefficient estimates of the sex ratio suggest little sex-ratio effect on stockholdings for daughter families.

In addition, the estimates of $\beta_1 + \beta_3$ in the middle of columns (1) and (2) of Table 3 show a small and statistically insignificant effect of having a first son on stock market results when the local sex ratio is set to one. This finding, in accord with our theoretical framework,

25. Standard errors given in parentheses are clustered by the county category. The unreported estimates of coefficients on other control variables have the sign and magnitude that one would expect given the findings of prior literature.

suggests that in the absence of financial pressures arising from the marriage market, household portfolio decision is not influenced by child gender. Finally, large F-statistics indicate that first-child gender, the local sex ratio, and their interaction term are jointly significant, implying the rejection of the null hypothesis. Therefore, marriage market conditions, as measured by child gender together with the local level of sex-ratio imbalance, are substantively important in explaining household stockholdings.

6.1.2 All risky financial assets

The remaining two columns in Table 3 show estimates on households' holdings in risky assets. Similar to the estimates on stockholdings, the empirical evidence is strong, both in terms of statistical significance and economic magnitude. The estimates of the interaction-term coefficients at the top of columns (3) and (4) are positive and significant. This pattern means a considerably differential impact of the sex-ratio distortion on risky asset holdings between families with a son and those with a daughter.

As can be seen from the middle, a higher sex ratio is associated with larger holdings in risky assets for families with a son. Based on the estimates in column (3), a 0.1 increase in the sex ratio increases the probability of investing in risky assets by 2.1 percentage points for son families. This effect is substantial, translating to a 27.6 percent increase given the unconditional mean investment probability of 0.08. Column (4) shows that with a 0.1 increase in the sex ratio, the portfolio share invested in risky assets increases by about 0.9 percentage point or 31.1 percent to the mean, when the first child is a son. Note that these estimates for the sex-ratio effect for son families are larger compared with the estimates in columns (1) and (2), because risky assets considered here include stocks but also other financial products such as mutual funds. However, families with a daughter are not affected by the sex-ratio imbalance; the coefficient estimates of the sex ratio are statistically insignificant and smaller. Moreover, a balanced sex ratio mostly eliminates the differences between son and daughter families in risky asset holdings. Similar to the results on stockholdings, large F-statistics indicate that marriage pressures do matter in risky investments.

Across all columns in Table 3, the pattern of the sex-ratio effect is remarkably consistent. In line with the prediction, marriage pressures emerge as a major influence on portfolio choice for households, as reflected by the differential sex-ratio effect assessed for son and daughter families as well as large magnitudes of F-statistics. Note that holdings in stocks and other risky financial products such as derivatives might not be determined in a similar manner *per se*; financial market participation and portfolio shares may also be distinct investment decisions. The noteworthy consistency of the results supports the interpretation that our model captures a general structure of how marriage market conditions affect household portfolio choice.

6.2 Risk attitude

In our theoretical framework, variation in actual household portfolios is due to changes in attitude towards risk. A potential confounding channel may be that parents of sons in regions with a high sex ratio, expecting large marriage expenditure, would take more effort in learning various investment options and searching for opportunities with higher returns.[26] Admittedly, risky assets in most cases yield higher expected returns. To rule out this possibility, we directly examine household risk attitude elicited from hypothetical financial decisions, using an analogous empirical strategy. We construct two measures of households' self-reported decision-making under financial risk based on survey questions, in a manner similar to Malmendier and Nagel (2011). Such questions are directly targeted at measuring risk attitude. Specifically, the hypothetical return levels are explicitly given, and hence household choices mainly reflect their risk attitude. This analysis alleviates the concern that our basic results on risky asset holdings are driven by return considerations only.

In one question in the 2013 CHFS, respondents were asked which of the following statements most precisely described their degree of risk tolerance when making financial investments: (i) take substantial risks expecting to earn substantial returns; (ii) take above average risks expecting to earn above average returns; (iii) take average risks expecting to earn average returns; (iv) take below average risks expecting to earn below average returns; and (v) not willing to take any risk. As the answer to this question cannot be quantitatively interpreted in a cardinal sense, we code it as a binary variable, which equals one if the first or second option was chosen, *i.e.*, the household is willing to take a substantial amount of financial risk. We refer to this measure as elicited risk tolerance.

Another hypothetical-decision measure comes from a question in which respondents were asked to choose between two hypothetical lotteries: (i) a sure gain of 4,000 *yuan*; versus (ii)

26. Oswald and Powdthavee (2010) find that parents of daughters have different political attitude from those of sons.

10,000 *yuan* with a 50 percent chance. We code the answer as a binary variable, which equals one if the second option was chosen, *i.e.*, the household is willing to take more risk in investment. We refer to this measure as lottery playing inclination.

The results using similar specifications with hypothetical-decision measures as dependent variables are reported in Table 4. The patterns revealed are similar to the patterns using the portfolio-allocation measures. The positive estimates of the interaction-term coefficients imply that a skewed sex ratio has an evidently larger impact on risk tolerance for son families than for daughter families. The estimates of $\beta_2 + \beta_3$ show that for families with a son, the local sex ratio has a strongly positive effect on financial risk taking. In contrast, for families with a daughter, the local sex ratio always has an insignificant effect on elicited risk attitude. Furthermore, with a balanced sex ratio, the estimated coefficient on first-son indicator in both columns is negligible in magnitude and insignificant, indicating that son families do not intrinsically take more risk than daughter families.

The results in Table 4 present robust evidence that variation in portfolios indeed results from variation in risk attitude. Our findings are consistent with prior literature documenting that self-reported hypothetical financial decisions are strongly correlated with subjects' actual allocation to risky assets and risky behavior in field experiments (Dohmen et al., 2011).

6.3 Robustness

The following robustness checks verify that our basic findings on the overall sex-ratio effect are not likely to be driven by identification issues, including unobserved region-level factors and potential concerns with the key regressors. The robustness results are reported in Table 5.[27] The first column repeats the basic results in Section 6.1.1, *i.e.*, column (1) of Table 3, and serves as a benchmark. For each robustness regression in the remaining columns, we conduct a generalized Hausman's specification test to investigate whether the estimate of the interaction-term coefficient β_3 is essentially the same as in the benchmark. The corresponding p-values are presented at the bottom of the table.

27. We report the results using stock market participation dummy as the outcome variable. Choosing other portfolio-choice measures yields qualitatively similar patterns, which we do not present due to space constraints.

6.3.1 Geographic heterogeneities

We start by checking whether the existence of various location-specific confounding factors contaminates our results. In our main regressions, prefecture dummies are included to absorb these variables, hence mitigating potential endogeneity concerns associated with unobservable regional heterogeneity. To partially evaluate the importance of such geographic heterogeneities in driving our estimation results, we exclude prefecture dummies in column (2) of Table 5. The estimates replicate the pattern of the benchmark, indicating that unobserved geographic factors have a negligible influence on our findings. Specifically, the coefficient estimate of the cross term turns out to be statistically equivalent to that in the benchmark. A comparison of the first two columns shows that the sex-ratio effect for son families appears more significant, both statistically and economically, once cross-prefecture heterogeneities are controlled for. This finding suggests that our estimates for the sex-ratio effect for son families would possibly be lower bounds in the presence of geographical heterogeneities, which we cannot deal with directly at the county level.

6.3.2 Potential concerns with first-child gender

There are two potential concerns with the validity of our estimates using first-child gender as a regressor. First, although first-child gender is random, it may directly affect other family choices, which may be correlated with household portfolios. For example, since first-son and first-daughter families typically have differential subsequent fertility decisions, one might worry that the number of children in a family could play a direct role in portfolio allocations. Second, first-child gender may be an inappropriate proxy for child gender composition in a family, and therefore an inappropriate proxy for marriage market conditions. A first daughter followed by a second son, for instance, might also imply a larger marriage expenditure and therefore a right shift of the reference level.

To address the first concern, we include the number of children in a family as an additional control to account for different fertility decisions induced by first-child gender. Adding this regressor does not have a material effect on our estimates compared with the benchmark, as reported in column (3) of Table 5. The cross-term coefficient remains the same at a five percent level of significance.[28] Moreover, we observe qualitatively similar results when restricting the

28. Note that the coefficient on the number of children is negative, which is in line with the notion that parents

sample to one-child families; the pattern still remains when further restricting the sample to one-child families with a child older than two years old, in which the possibility is reduced that parents are preparing for a second child; see Table A1 in the Appendix. These results underscore the importance of child gender, rather than subsequent fertility decisions, in household portfolios.

To address the second concern, we construct two other child-gender measures: a dummy for having at least one son and the share of sons in a family. We conclude from the qualitatively similar results that child gender, no matter how it is measured, does affect how families allocate their portfolios in response to sex-ratio distortions; see Table A1. This analysis enhances our confidence in the presumption that child gender affects financial portfolios through the channel of parental considerations for their children's marriages. Unlike first-son dummy, these two alternative measures might be endogenous. So, we focus on the first-child gender in our main analyses.

6.3.3 Possible confounders related to sex ratio

Subsequently, we check whether abnormal sex ratios, in the presence of related confounders, indeed have a significant and independent role in accounting for the observed differential in portfolios across families with a son and those with a daughter. Male-biased sex ratios in China mainly result from a combination of preferences for sons, enforcement of the family planning policy, and access to gender-selection technology, as discussed in Section 3.1. Cross-region differences within the country along these dimensions may affect sex ratios, but also household portfolios directly. We examine these potential origins of the sex-ratio difference across regions, and test whether these factors confound the interpretation of our results.

In China, parents have historically preferred sons to daughters, and more so in remote or more traditional agricultural areas. Thus the degree of son preferences likely reflects the general level of economic development. As such, we additionally control for local average household income as an indicator of development levels in column (4) of Table 5. There is virtually no effect on the results compared to the benchmark, as Hausman's general specification test yields a p-value of 0.78, which does not suggest a significant difference in the interaction-term coefficients. This result casts doubt on the possibility that our findings are mainly driven by

allocate less wealth to financial investments if they have to meet greater daily expenses.

variations in the extent of son preferences.

One may also argue that China's traditional son preferences are accompanied by both skewed sex ratios and high earning or saving incentives when households have sons. As such, son families tend to be relatively wealthier than daughter families in regions with higher sex ratios, and mechanically retain more holdings in risky assets. This channel spuriously generates a positive association for son households between risky asset holdings and the local sex ratio. But since we control for parental education levels in our main regressions, this channel is not likely to be at play. We further add household wealth controls to account for this effect. Both liquid wealth and gross financial wealth are used, and the results are reported in columns (5) and (6) of Table 5, respectively. There is essentially no effect on our results, particularly on the interaction-term coefficients, as revealed by the relatively large p-values in Hausman's general specification test.[29]

Moreover, parental demand for sons may arise because sons are perceived to serve as a better source of insurance against old ages in a patriarchal and patrilocal society (Chew et al., 2016). In the sense that sons have more responsibility of caring for elderly parents in regions with higher sex ratios, parents of sons in these regions have more stockholdings presumably because they are less subject to pension issues. To address this concern, we take into account the effect of guaranteed social old-age support, with the presence of which parents rely less heavily on sons' support in old ages. This variable is measured by a dummy indicating whether either the father or mother has assured income after retirement, such as retirement payments and public insurances. We include this old-age pension dummy in column (7) of Table 5, and find negligible impact on the results.

To address the concern that son-preference variation may originate with different relative levels of adult female earnings in a traditional agrarian economy, we include in the regression county-specific gender differential in wages (male minus female). The new specification yields virtually identical results to the benchmark as shown in column (8) of Table 5.

Furthermore, the practice of China's family planning policy varies considerably from place to place. More stringent enforcement of the policy exacerbates the sex-ratio distortion. At the same time, the implementation of the policy may be based on unobservable local socioeconomic

29. We show that stock market participation probability increases with liquid assets, in line with the prediction of a standard model with fixed per-period participation costs (Malmendier and Nagel, 2011; Vissing-Jorgensen, 2004).

conditions, which may also correlate with local residents' financial decisions. As the policy directly targets fertility, we add the number of children in a household in our regression. The pattern of results stays almost invariant; see column (3) of Table 5.

Finally, variations in the sex ratio may also partly result from different access to gender-screening technology. Probably, a more advanced technology leads to more boys being born by enabling gender selection, and meanwhile eases families' access to financial markets, or alternatively, improves their investment skills. Risky investments then reflect the easiness to make such investments. Presumably, any factor that varies across counties would impact all local households in the region in similar ways, and child gender would not make a clear difference in household financial decisions. We would accordingly observe little difference in portfolios across son and daughter families, as both likely attach more importance to risky assets when facing higher sex ratios. This is, however, not what our results suggest. In addition, we show in column (4) of Table 5 that our results are robust to taking into account local average household income, which is closely correlated with the level of technological advancement. Therefore, our findings cannot be attributed to difference in technologies.

7 Underlying mechanism

This section discusses the mechanism for our results. We first show that the sex-ratio effect is heterogeneous across families with a son in different wealth classes, which particularly demonstrates the validity of the proposed reference-dependent mechanism. We also provide some suggestive evidence that the sex ratio affects the reference level of marriage expenditure to further support the reference-dependent mechanism. Moreover, we show that our empirical results cannot be interpreted by an alternative interpretation of the wealth effect.

7.1 Reference-dependent mechanism

Given that marriage market competition increases stock market participation through the reference-dependent mechanism, we should observe that (i) the sex-ratio effect is the strongest for families with a son who can afford modest marriage expenditure (Prediction 2), and (ii) the sex ratio influences the reference level of marriage expenditure.

In the sense that household financial wealth is an appropriate proxy for feasible expenditure on children's marriage, we divide the main sample into three groups using two cut-off points: the 50th (median) and 90th percentiles of gross financial wealth. According to our model, middle-class families with a son behave more responsively to a high sex ratio than poorer and wealthier counterparts. Table 6 reports the estimation results based on the three subsamples using the same specifications and the same outcome variables as in Table 3. Consistently, the sex ratio indeed has the largest impact on household portfolio choice for son families from the middle group. In particular, with a 0.1 increase in the sex ratio, the stock market participation rate among families with a son who are between the 50th and 90th percentiles of wealth rises by 2.5 percentage points, which is more than twice as large as the overall estimate obtained from the whole sample. Using other portfolio measures gives similar findings. Yet, the sex-ratio effect for son families below the median or above the 90th percentile of wealth is not as significant as for those in-between, and in some cases, the sign of the estimates is not preserved.

To provide closer insights into the proposed reference-dependent mechanism, we infer a finer prediction on the heterogeneity of the sex-ratio effect for middle-class families with a son (Region II in Figure 4). The kink generated by loss aversion in the value function implies a discontinuous decline in the marginal utility when crossing the reference level. This sharp change leads to behavior demonstrating excessive risk aversion (Rabin, 2000), suggesting that households closer to the reference level of marriage expenditure hold fewer stocks. When current potential marriage expenditure is relatively smaller, where the probability of crossing the reference level in the future is lower, the effect of the kink is weakened. Households become less risk averse and thus hold more stocks. Intuitively, families moving into losses are in the convex region of the value function, and are therefore willing to gamble almost as far as the edge of it. However, they are not willing to take a larger gamble than this, which may bring them to the right of the kink where the marginal utility of additional gains is significantly lower (Barberis and Xiong, 2009).

Accordingly, we divide the middle class between the 50th and 90th percentiles of gross financial wealth into two subgroups using the 70th percentile as the cut-off point, which graphically correspond to II(a) and II(b), respectively, in Figure 4. As reported in Table 6, the sex-ratio effect for son families below the 70th percentile is more significant and much larger than that for those above; the former effect always about doubles the latter effect in percentages using the four portfolio-choice measures. This pattern is in line with the argument that since son families in region II(a) are further from the reference level, it takes a larger share allocation for them to gamble to the edge of the convex region than it does for those in region II(b).

Sex-ratio effect heterogeneity across son households with different wealth levels shown in Table 6 cannot be easily rationalized by alternative interpretations other than reference-based utilities. Although the cut-off points in this empirical exercise may be somewhat artificial, the estimation results do not fully rely on the choice of these points; see Figure A1 for sex-ratio effect heterogeneity for son families across deciles of gross financial wealth.

Does the sex ratio influence the reference level of marriage expenditure? To answer this question, we test whether the sex ratio affects parental expectations about marriage expenditure for their children. However, we do not have data on these expectations. We provide several pieces of suggestive evidence about such reference level. First, Figure 2 shows that actual marriage expenditure increases more for grooms when the sex ratio increases. As parents learn this information from peers, it is likely that they will adjust their expectations about marriage expenditure. Second, we investigate additional household outcomes related to housing and parental earning incentive, using data from the 2010 CFPS survey. Housing is a fundamental part of marriage expenditure as well as the single most important piece of household wealth, as described in Section 3.2. We also examine earning incentive, which in China is mainly reflected by temporary or circular migration, *i.e.*, working in a different region from their hometown (Zhao, 1999). Parental migration substantially improves family income, so migrant parents can afford larger marriage expenditure for their children.

For housing, we construct two outcome variables: an ownership dummy and log mortgage debts. Ownership dummy indicates whether the family owns the property right of any house, and equals one if the property deed and other relevant contracts of one or more houses belong solely to this family; self-constructed houses in rural regions are also counted. Larger mortgages signal houses of higher quality, and imply larger expenditure for children's marriage. The first two columns of Table 7 present the results. Column (1) shows that son households appear more likely to possess a house in regions with a more male-biased sex ratio. Specifically, the probability that son families own any house, on average, rises by 2.1 percentage points when the sex ratio increases by 0.1. In column (2), using a reduced sample, the sex-ratio effect for son households is still positive and significant; a 0.1 increase in the sex ratio leads to a 27.4 percent increase in housing loans.

We next turn to migration results. As reported in columns (3) and (4) of Table 7, both the father and mother of sons are more likely to work outside their hometown when they face a higher sex ratio. Paternal and maternal migration probability for son families rises by 16 and 43 percent, respectively, with a 0.1 increase in the local sex ratio.

Migration remittances sent back by migrant family members can ease the household budget constraint and thereby increase marriage expenditure for children. A closer look at the intended purpose of such remittances in column (5) suggests that son families are more willing to spend remittances on children's marriage in the presence of a more biased sex ratio. This finding indicates that parental earning incentive is most probably motivated by the desire to help children improve their marital competitiveness.

In Table 7 we consistently show that the sex ratio does not have a significant effect for daughter families, and that first-son dummy does not receive a significant coefficient when the sex ratio is set to be balanced. Together, these results validate the reference-dependent mechanism, by suggesting that household outcomes reflect parental considerations of their children's marriages, and that expected expense on their sons' marriages increases with the sex ratio.

7.2 Wealth effect

A plausible simple explanation for a reduced-form association between high sex ratios and risky asset holdings for son families is that they are less risk averse as their relative wealth decreases, given the negative income shock of competing marriage expenditure. We have several pieces of evidence against this explanation. First, an agent's risk attitude is determined by the curvature of his utility function, which is not necessarily affected by income, except in an extreme case of households exhibiting increasing absolute risk aversion in their utilities. In general cases, a decrease in income does not necessarily increase risk tolerance. Second, a common finding in prior literature is that the tendency to take financial risk is positively, but not negatively, associated with wealth;[30] see the last column of Table 6 for consistent evidence that wealthier families hold more risky assets in their portfolios. Third, we control for parental education levels in our main regressions to take care of the wealth effect. We also explicitly add income and wealth controls in the robustness checks, and the pattern of results is preserved; see Section

30. This positive correlation may in part be explained by preferences with decreasing relative risk aversion (Calvet and Sodini, 2014; Calvet et al., 2009; Carroll, 2002; Wachter and Yogo, 2010). Alternatively, limited participation among the poorer may be rationalized by fixed costs (Gomes and Michaelides, 2005).

6.3.3. Last and most importantly, our theory predicts that the sex-ratio effect is the strongest for son families who can afford modest marriage expenditure (Prediction 2). However, the income-effect explanation implies that all son families would behave similarly in portfolio choice in the presence of a high sex ratio, regardless of their wealth.

8 Conclusion

This paper examines the role of marriage market conditions for children in household portfolio choice based on the prospect theory. We provide a conceptual framework that incorporates both reference-dependent preferences and marriage prospects for children. The model predicts that families with a son hold more risky assets when the sex ratio is higher, and that the effect is the strongest for son families that can afford modest marriage expenditure. The empirical results are consistent with our theoretical predictions. The reference-dependent channel captures and reconciles all our empirical findings, whereas confounders related to regressors or other possible channels have minimal effects, as suggested by a series of robustness checks and ancillary results, particularly the finding on the heterogeneity of the sex-ratio effect.

We acknowledge that part of the reference-dependent evaluation is possibly driven by status seeking, peer effects, or audience effects; see, for instance, Brown et al. (2011) and Allen et al. (2016). Specifically, an increase in the reference level of marriage expense for son families may come from incentives to fit in with their friends or relatives in the same social strata, instead of revealing an intrinsic adjustment of the reference level. The increase in the reference level may also be due to the perception that larger marital transfer would be evaluated more favorably by potential marriage partners. However, all these possibilities still reflect reference-dependent preferences, and merely suggest that such preferences partly originate with peers or potential matches.

Finally, the results in this paper are relevant for several policy issues. For example, our findings suggest that marriage market conditions likely change a household's reference point, which in turn impacts household portfolio choice. The resulting changes in participation in the risky asset market may have further implications for financial market outcomes, such as asset pricing. Moreover, the differential patterns in household investments in risky assets may contribute meaningfully to the evolution of wealth inequality, as highlighted in Campbell (2016).

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TABLE 1

Summary Statistics for Household Financial Characteristics

	(1)	(2)	(3)	(4)	(5)
	Mean	SD	Min	Median	Max
Panel A: All Families					
Stock market participation	0.048	0.214	0	0	1
Stock share	0.016	0.101	0	0	1
Risky asset market participation	0.075	0.264	0	0	1
Risky asset share	0.028	0.137	0	0	1
Liquid wealth	37,878	160,294	0	5,000	4,665,000
Gross financial wealth	289,200	662,608	0	105,000	1.08e+07
Observations	5,092				
Panel B: Stock-Market Participants					
Stock share	0.335	0.323	0	0.238	1
Risky asset share	0.410	0.339	0	0.353	1
Liquid wealth	224,304	554,225	0	78,000	4,665,000
Gross financial wealth	828,407	1,190,957	0	403,000	9,167,500
Observations	278				

Panel C: Families Facing Different Marriage Market Conditions	Mean	SD	Mean	SD	Observations
First-son families	Balanced-sex-ratio regions		High-sex-ratio regions		2,646
Stock market participation	0.032	0.177	0.058	0.233	
Stock share	0.011	0.081	0.022	0.120	
Risky asset market participation	0.062	0.242	0.098	0.297	
Risky asset share	0.024	0.127	0.038	0.160	
Observations	1,369		1,277		
First-daughter families	Balanced-sex-ratio regions		High-sex-ratio regions		2,446
Stock market participation	0.057	0.232	0.046	0.209	
Stock share	0.017	0.099	0.015	0.098	
Risky asset market participation	0.077	0.266	0.064	0.244	
Risky asset share	0.027	0.131	0.024	0.125	
Observations	1,235		1,211		

Notes: Data are from the 2013 CHFS. Descriptive statistics are weighted.

TABLE 2

Summary Statistics for Household Demographic Characteristics

	Mean (SD)				
	All families	First-son families	First-daughter families	Difference	SE
	(1)	(2)	(3)	(4)	(5)
First son	0.519	–	–	–	–
	(0.500)				
First-child age	10.72	10.67	10.77	-0.100	0.136
	(4.833)	(4.811)	(4.857)		
Region of residence (rural=1)	0.411	0.408	0.414	-0.006	0.014
	(0.492)	(0.492)	(0.493)		
Ethnicity (minority=1)	0.121	0.116	0.126	-0.010	0.009
	(0.326)	(0.320)	(0.332)		
Father's age	38.37	38.36	38.39	-0.030	0.167
	(5.957)	(5.922)	(5.997)		
Father's schooling years	9.643	9.677	9.606	0.071	0.097
	(3.460)	(3.483)	(3.436)		
Father's *hukou* (urban=1)	0.297	0.298	0.296	0.002	0.013
	(0.457)	(0.457)	(0.457)		
Father's political status (party=1)	0.125	0.124	0.127	-0.003	0.009
	(0.331)	(0.329)	(0.333)		
Mother's age	36.56	36.59	36.53	0.060	0.171
	(6.090)	(6.069)	(6.113)		
Mother's schooling years	8.882	8.920	8.841	0.079	0.107
	(3.817)	(3.794)	(3.843)		
Mother's *hukou* (urban=1)	0.281	0.287	0.274	0.013	0.013
	(0.449)	(0.452)	(0.446)		
Mother's political status (party=1)	0.040	0.044	0.036	0.007	0.006
	(0.197)	(0.205)	(0.187)		
Observations	5,092	2,646	2,446		

Notes: Data are from the 2013 CHFS. In the first three columns, standard deviations are given in parentheses. Standard errors for the difference between first-son and first-daughter families are reported in the last column, none of which are statistically significant at the ten percent level or less.

TABLE 3

Household Portfolio Choice

Dependent variable	Stocks: Participation (1)	Stocks: Share (2)	All kinds of risky assets: Participation (3)	All kinds of risky assets: Share (4)
First son (β_1)	-0.155**	-0.039**	-0.273**	-0.086*
	(0.062)	(0.015)	(0.095)	(0.037)
Sex ratio (β_2)	-0.020	0.032	-0.068	0.000
	(0.069)	(0.036)	(0.126)	(0.069)
First son * Sex ratio (β_3)	0.144**	0.037*	0.276**	0.088*
	(0.055)	(0.019)	(0.091)	(0.037)
Sex-ratio effect				
for first-son families ($\beta_2 + \beta_3$)	0.123***	0.069*	0.207**	0.088*
Percentage increase: sex ratio+0.1	25.7	42.7	27.6	31.1
Son effect at sex ratio=1 ($\beta_1 + \beta_3$)	-0.012	-0.001	0.002	0.003
F-statistic ($H_0 : \beta_1 = \beta_2 = \beta_3 = 0$)	16.067	12.167	6.116	11.224
Observations	5,092	5,092	5,092	5,092
R-squared	0.181	0.108	0.205	0.146
Dependent variable mean	0.048	0.016	0.075	0.028
Model	LPM	OLS	LPM	OLS
Other controls?	YES	YES	YES	YES
Prefecture fixed effects?	YES	YES	YES	YES

Notes: Data are from the 2013 CHFS. The sex-ratio effect for son families is reported in both percentage points and percentages. Standard errors given in parentheses are clustered by the county category.

***Significant at the 1 percent level.

**Significant at the 5 percent level.

*Significant at the 10 percent level.

TABLE 4

Hypothetical Financial Decisions

Dependent variable	Elicited risk tolerance (1)	Lottery playing inclination (2)
First son (β_1)	-0.204**	-0.306
	(0.073)	(0.183)
Sex ratio (β_2)	0.120	-0.126
	(0.121)	(0.174)
First son * Sex ratio (β_3)	0.193**	0.314
	(0.072)	(0.193)
Sex-ratio effect		
for first-son families ($\beta_2 + \beta_3$)	0.313**	0.187*
Percentage increase: sex ratio+0.1	22.3	6.3
Son effect at sex ratio=1 ($\beta_1 + \beta_3$)	-0.011	0.007
F-statistic ($H_0 : \beta_1 = \beta_2 = \beta_3 = 0$)	4.834	2.956
Observations	5,092	5,092
R-squared	0.114	0.075
Dependent variable mean	0.141	0.298
Model	LPM	LPM
Other controls?	YES	YES
Prefecture fixed effects?	YES	YES

Notes: Data are from the 2013 CHFS. The sex-ratio effect for son families is reported in both percentage points and percentages. Standard errors given in parentheses are clustered by the county category.

***Significant at the 1 percent level.

**Significant at the 5 percent level.

*Significant at the 10 percent level.

TABLE 5

Robustness Checks

Dependent variable	Stock-market participation, mean=0.048							
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
First son (β_1)	-0.155**	-0.129*	-0.156*	-0.154**	-0.117*	-0.149*	-0.155**	-0.159**
	(0.062)	(0.053)	(0.064)	(0.058)	(0.060)	(0.063)	(0.063)	(0.061)
Sex ratio (β_2)	-0.020	0.002	-0.017	-0.059	-0.008	-0.022	-0.024	-0.060
	(0.069)	(0.024)	(0.068)	(0.064)	(0.070)	(0.073)	(0.071)	(0.087)
First son * Sex ratio (β_3)	0.144**	0.119**	0.142**	0.142**	0.104*	0.137*	0.143**	0.147**
	(0.055)	(0.048)	(0.056)	(0.051)	(0.052)	(0.056)	(0.056)	(0.054)
Number of children			-0.008					
			(0.004)					
Local average household income, million				1.340**				
				(0.405)				
Liquid wealth, million					0.267***			
					(0.052)			
Gross financial wealth, million						0.036***		
						(0.006)		
Old-age support							0.011	
							(0.009)	
Gender wage differential, M-F, thousand								0.007*
								(0.003)
Sex-ratio effect								
for first-son families ($\beta_2 + \beta_3$)	0.123***	0.121**	0.124***	0.084**	0.095***	0.114***	0.119***	0.086*
Percentage increase: sex ratio+0.1	25.7	25.3	25.9	17.4	19.9	23.9	24.8	18.0
Son effect at sex ratio=1 ($\beta_1 + \beta_3$)	-0.012	-0.010	-0.014	-0.012	-0.014	-0.013	-0.012	-0.012
F-statistic ($H_0 : \beta_1 = \beta_2 = \beta_3 = 0$)	16.067	5.699	17.332	8.823	55.916	13.625	16.092	9.849
Observations	5,092	5,092	5,092	5,092	5,092	5,092	5,092	5,092
R-squared	0.181	0.134	0.182	0.184	0.218	0.192	0.182	0.183
Hausman's test p-value	–	0.111	0.053	0.780	0.105	0.062	0.510	0.257
Model	LPM	LPM	LPM	LPM	LPM	LPM	LPM	LPM
Other controls?	YES	YES	YES	YES	YES	YES	YES	YES
Prefecture fixed effects?	YES	NO	YES	YES	YES	YES	YES	YES

Notes: Data are from the 2013 CHFS. The sex-ratio effect for son families is reported in both percentage points and percentages. Standard errors given in parentheses are clustered by the county category.

***Significant at the 1 percent level.

**Significant at the 5 percent level.

*Significant at the 10 percent level.

TABLE 6
Heterogeneity of Sex-Ratio Effect by Wealth

Wealth level (percentile)	Sex-ratio effect for first-son families ($\beta_2 + \beta_3$) (1)	Percentage increase: sex ratio+0.1 (2)	F-statistic ($H_0 : \beta_1 = \beta_2 = \beta_3 = 0$) (3)	Observations (4)	Dependent variable mean (5)
	Dependent variable: stock market participation				
0–50th	-0.013	-5.4	10.424	2,241	0.025
50–90th	0.252***	48.5	224.445	2,201	0.052
50–70th	0.300*	69.5	6.277	1,020	0.043
70–90th	0.214	35.4	3.237	1,181	0.060
90–100th	0.111	8.9	0.887	650	0.125
	Dependent variable: stock share				
0–50th	0.017	19.2	8.389	2,241	0.009
50–90th	0.098	63.5	1.574	2,201	0.015
50–70th	0.118*	106.9	1.831	1,020	0.011
70–90th	0.117	59.3	0.815	1,181	0.020
90–100th	0.031	6.5	2.485	650	0.047
	Dependent variable: risky market participation				
0–50th	0.044	11.1	9.996	2,241	0.039
50–90th	0.272***	32.7	18.578	2,201	0.083
50–70th	0.387	54.3	16.118	1,020	0.071
70–90th	0.261	27.6	2.658	1,181	0.095
90–100th	0.171	9.1	0.462	650	0.188
	Dependent variable: risky share				
0–50th	-0.002	-1.2	7.541	2,241	0.015
50–90th	0.138**	48.0	46.753	2,201	0.029
50–70th	0.223**	96.0	11.904	1,020	0.023
70–90th	0.142	41.8	7.437	1,181	0.034
90–100th	-0.030	-3.6	14.513	650	0.082

Notes: Data are from the 2013 CHFS. The sex-ratio effect for son families is reported in both percentage points and percentages. Standard errors are clustered by the county category.
***Significant at the 1 percent level.
**Significant at the 5 percent level.
*Significant at the 10 percent level.

TABLE 7
Housing and Parental Migration

	Housing		Parental migration		
Dependent variable	Ownership	Mortgage, log	Father	Mother	For children's marriage
	(1)	(2)	(3)	(4)	(5)
First son (β_1)	-0.253**	-3.098*	-0.273***	-0.108**	-0.249**
	(0.129)	(1.618)	(0.096)	(0.051)	(0.101)
Sex ratio (β_2)	-0.034	-0.127	-0.109	0.006	0.053
	(0.084)	(1.167)	(0.073)	(0.039)	(0.073)
First son * Sex ratio (β_3)	0.245**	2.864*	0.269***	0.106**	0.234**
	(0.119)	(1.513)	(0.088)	(0.047)	(0.093)
Sex-ratio effect					
for first-son families ($\beta_2 + \beta_3$)	0.211**	2.737***	0.159**	0.112***	0.287***
Percentage increase: sex ratio+0.1	2.5	27.4	16.0	43.0	248.0
Son effect at sex ratio=1 ($\beta_1 + \beta_3$)	-0.008	-0.234	-0.004	-0.002	-0.015
F-statistic ($H_0 : \beta_1 = \beta_2 = \beta_3 = 0$)	2.406	2.753	4.239	3.380	4.994
Observations	4,210	478	4,210	4,210	922
R-squared	0.191	0.318	0.048	0.033	0.245
Dependent variable mean	0.829	10.017	0.099	0.026	0.012
Model	LPM	OLS	LPM	LPM	LPM
Other controls?	YES	YES	YES	YES	YES
Province fixed effects?	YES	YES	YES	YES	YES

Notes: Data are from the 2010 CFPS survey. The sex-ratio effect for son families is reported in both percentage points and percentages. Standard errors are given in parentheses. In column (3), the sample includes households with at least one migrant worker.

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

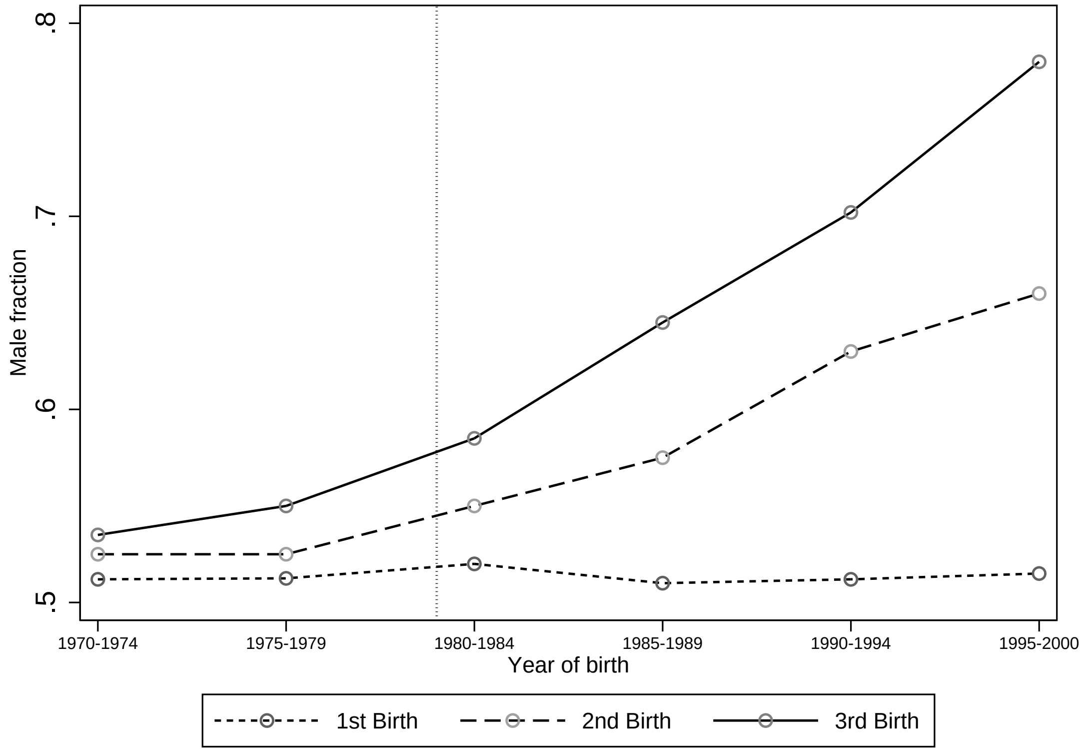

FIGURE 1
Male Fraction of Births by Birth Order in China

Notes: Data are from Ebenstein (2010). The figure shows a steep rise in the sex ratio over the past decades, and the skewness comes from gender selection among second- and higher-order births, rather than among first-order births.

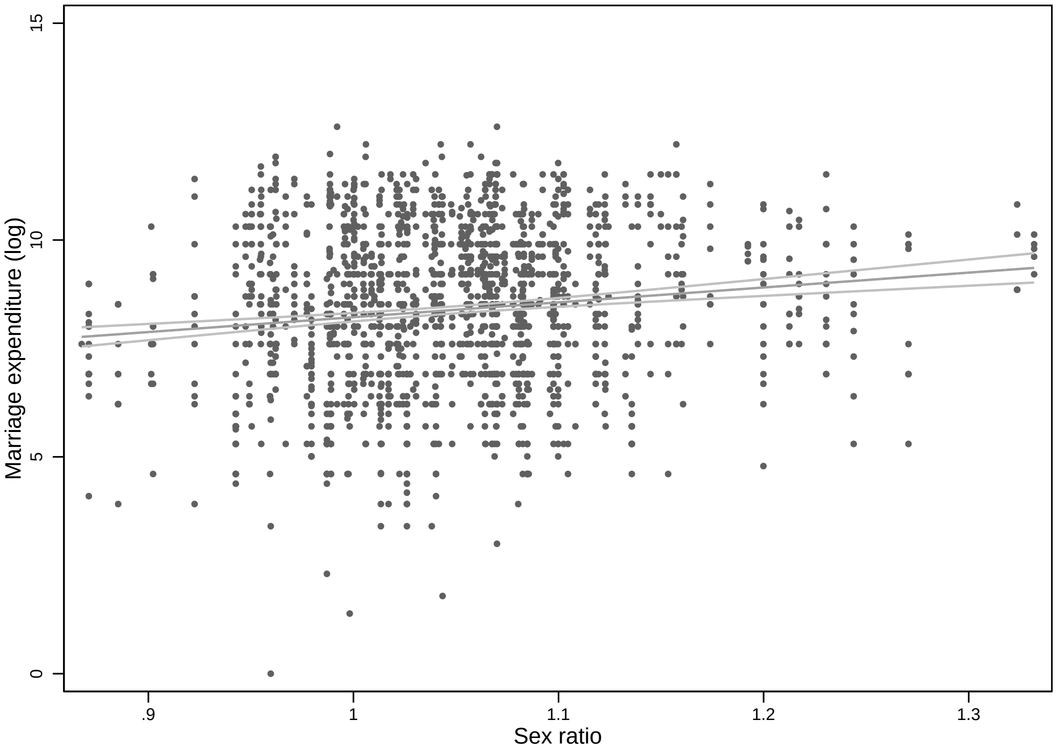

FIGURE 2

Local Sex Ratio and Marriage Expenditure in China

Notes: The sex ratio is calculated for the year 2010 for the cohort between the ages of 10 and 34 years. Data on sex ratios are from the 2010 China Population Census. Data on marriage expenditures are from the 2010 CFPS survey. The figure shows that marriage expenditure is generally larger in regions with higher sex ratios.

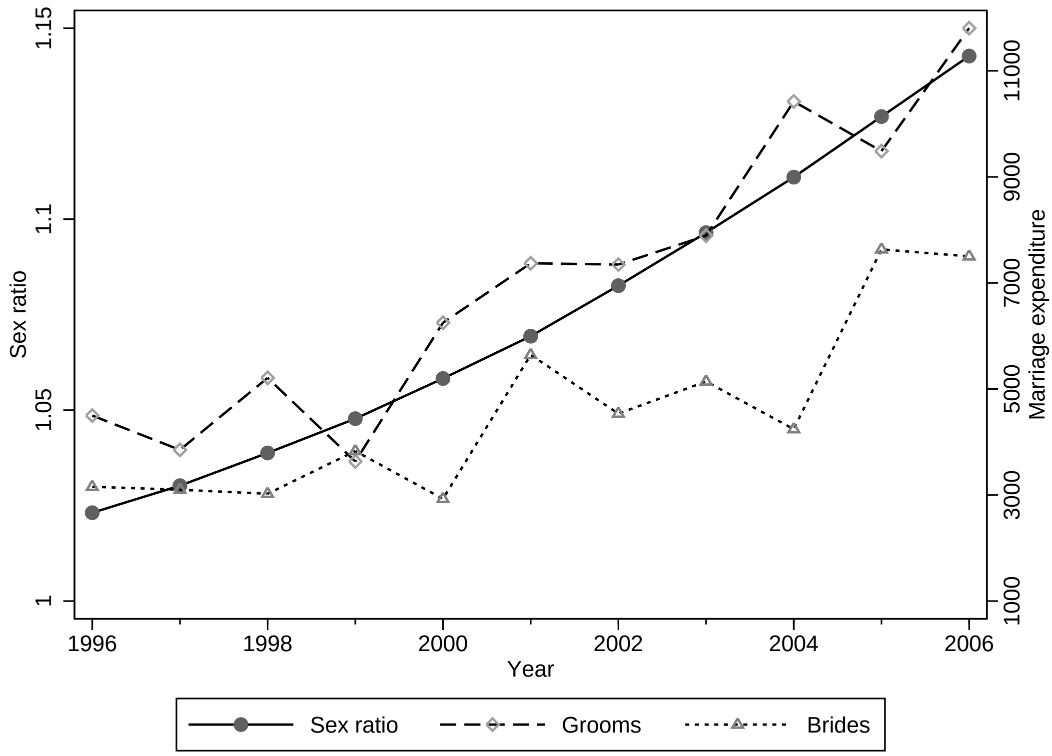

FIGURE 3

Trends in Sex Ratio and Marriage Expenditure in China

Notes: The sex ratio is calculated for each year for the cohort between the ages of 0 and 15 years. Data on sex ratios are projected from the 2010 China Population Census. For example, the sex ratio for the cohort between the ages of 0 and 15 years in 2006 is calculated using the cohort between the ages of 4 and 19 years in 2010, because these two groups are supposed to be the same. Data on marriage expenditure are from Brown et al. (2011). The figure shows that as the sex ratio increases, grooms' families are spending more on marriage over time, whereas brides' families are less so.

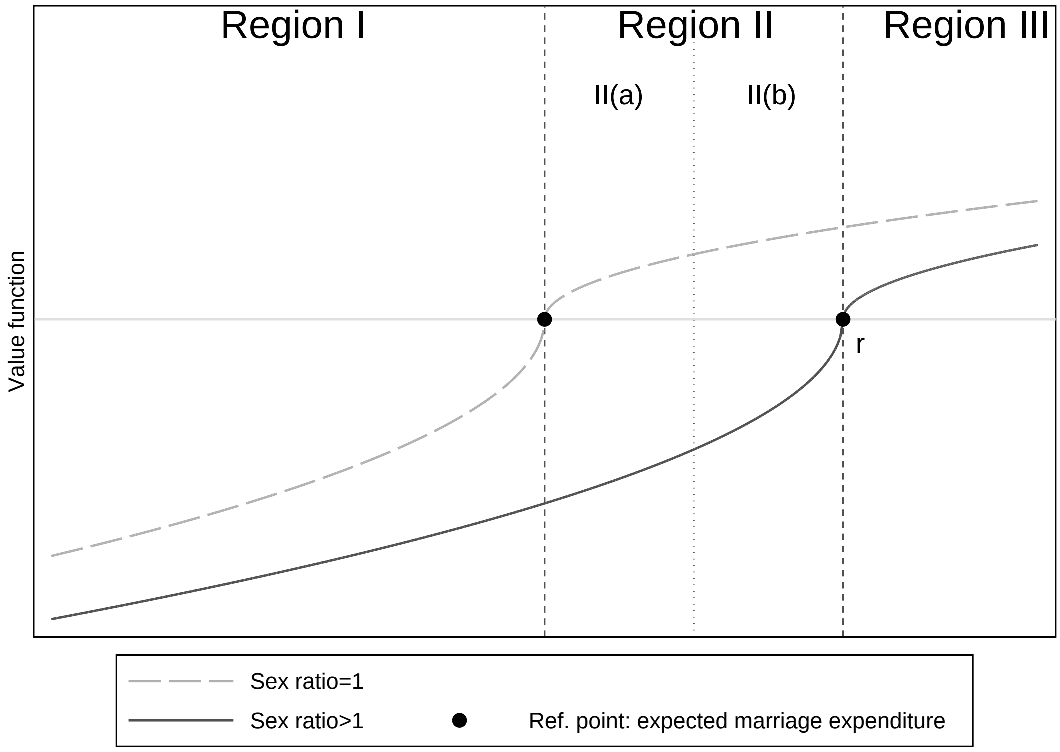

FIGURE 4

Reference-Dependent Value Function for Families with a Son

Notes: The solid S-shaped curve plots the value function at a male-biased sex ratio, and the dashed S-shaped curve plots the value function at a balanced sex ratio: $v(x|r) = (x - r)^{\alpha}$ for $x \geq r$ and $v(x|r) = -\lambda \cdot (r - x)^{\alpha}$ for $x < r$. Utility is derived from marriage expenditure for sons, x, relative to the reference level, r; α and $\lambda > 1$ are parameters. The reference level, r, according to Kőszegi and Rabin (2006, 2007, 2009), is defined as the reference level of marriage expenditure. Parental assessment of gains or losses is determined by this level. Regions I, II, and III respectively include son families who can afford a low, modest, and high marriage expense, whose actual expenditure on sons' marriage would be far below, close to, and above parental expectation.

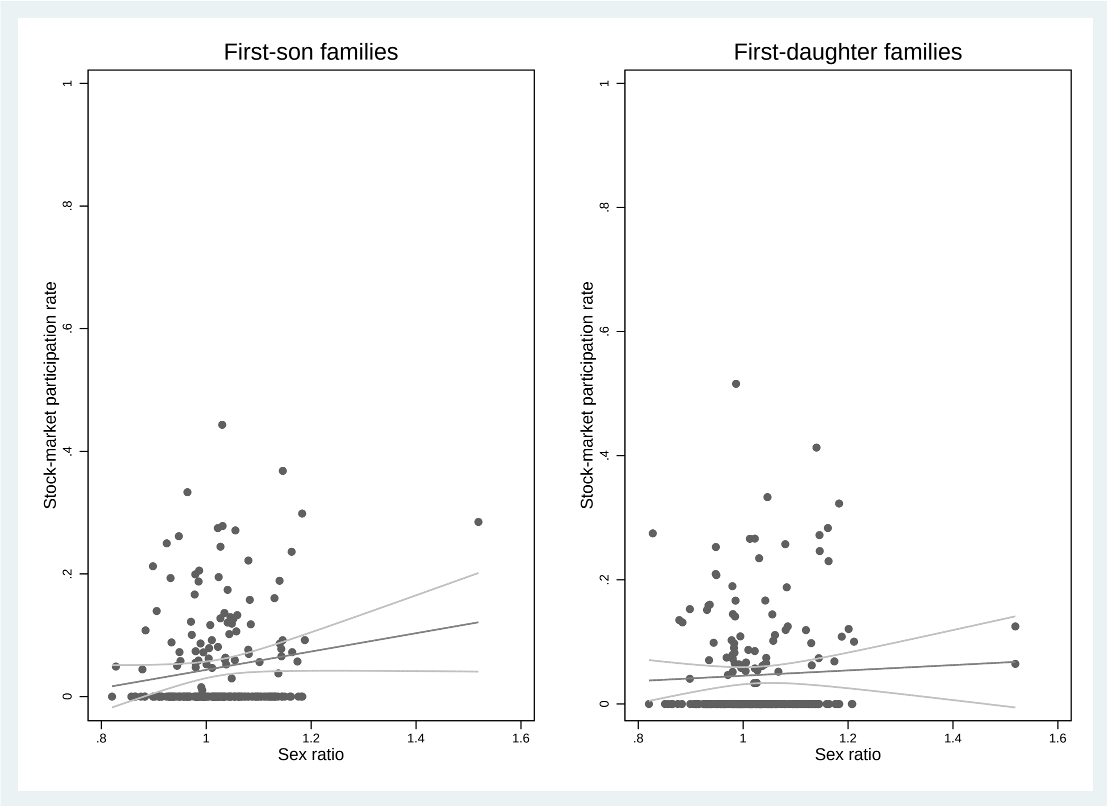

FIGURE 5

Local Sex Ratio and Stock Market Participation in China

Notes: The sex ratio is calculated for the year 2013 for the marriage-age cohort between the ages of 18 and 32 years. Data on sex ratios are projected from the 2010 China Population Census. Data on stock market participation rates are from the 2013 CHFS. The figure shows that the fraction of families with a first-born son who participate in the stock market increases with the sex ratio; the fraction of families with a first-born daughter does not change with the sex ratio.

Appendix

TABLE A1
Additional Robustness Checks

Dependent variable	Stock-market participation			
	One-child families		Other child-gender measures	
	No child-age limit	Child $\geq$ 2 years old	Having any son	Share of sons
	(1)	(2)	(3)	(4)
First son (β_1)	-0.214***	-0.208***		
	(0.044)	(0.045)		
Sex ratio (β_2)	-0.040	-0.026	-0.048	-0.045
	(0.035)	(0.041)	(0.047)	(0.061)
First son * Sex ratio (β_3)	0.193***	0.188***		
	(0.044)	(0.047)		
Having any son (β_1)			-0.179***	
			(0.041)	
Having any son * Sex ratio (β_3)			0.166***	
			(0.033)	
Share of sons (β_1)				-0.200**
				(0.060)
Share of sons * Sex ratio (β_3)				0.187**
				(0.050)
Sex-ratio effect				
for son families ($\beta_2 + \beta_3$)	0.153**	0.162**	0.118**	0.142***
Percentage increase: sex ratio+0.1	23.4	24.2	24.6	29.5
Son effect at sex ratio=1 ($\beta_1 + \beta_3$)	-0.021*	-0.020	-0.013	-0.013
F-statistic ($H_0 : \beta_1 = \beta_2 = \beta_3 = 0$)	10.470	10.371	55.216	16.236
Observations	3,274	3,047	5,092	5,092
R-squared	0.209	0.224	0.181	0.182
Dependent variable mean	0.066	0.067	0.048	0.048
Model	LPM	LPM	LPM	LPM
Other controls?	YES	YES	YES	YES
Prefecture fixed effects?	YES	YES	YES	YES

Notes: Data are from the 2013 CHFS. The sex-ratio effect for son families is reported in both percentage points and percentages. Standard errors given in parentheses are clustered by the county category.
***Significant at the 1 percent level.
**Significant at the 5 percent level.
*Significant at the 10 percent level.

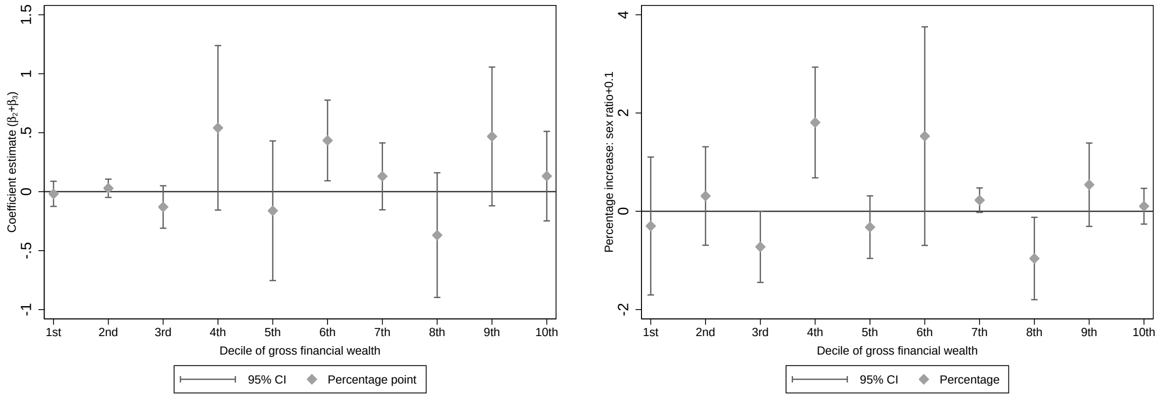

FIGURE A1

Heterogeneity of Sex-Ratio Effect on Stock Market Participation

Notes: Data are from the 2013 CHFS. The sex-ratio effect is for families with a son, and is presented in both percentage points and percentages. The figure shows considerable heterogeneity in such sex-ratio effect.