
High Sex Ratios and Household Portfolio Choice in China

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ABSTRACT


This work examines how high sex ratios (more men than women) affect household portfolio choice. Using data from a nationally representative Chinese household finance survey, we find that a one standard deviation increase in the sex ratio would raise the stock market participation rate by 2.9 percentage points, or 52.2 percent, for families with a son relative to families with a daughter. Our estimates imply that rising sex ratios explain around 10 percent of the significant growth in China's stock market size in recent decades.


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I. Introduction

How households allocate wealth to risky assets—particularly stocks—is of great financial and socioeconomic significance and has profound implications for welfare (Campbell 2006, 2016). This work establishes a link between household portfolio allocations and high sex ratios (that is, more men than women) in China, one of the fastest-growing Asian economies. In recent decades, China has experienced both tremendous growth in the size of its stock market and an unprecedented challenge of high sex ratios. The market value of China's stock markets has increased 15-fold and now accounts for more than 10 percent of the world's total value—second only to the United States (Iskryan 2016). The sex ratio at birth—with a biologically normal range of 1.03–1.08—reached 1.12 in 1990, rose to 1.20 in 2000, and has remained at that (Ebenstein 2010).¹

Rising sex ratios lead to an excess supply of men in the marriage market—more than 30 million, according to Sharygin, Ebenstein, and Das Gupta (2013). Therefore, a man faces unfavorable marriage market conditions and has a large expected marriage expenditure. In Chinese culture and tradition, such financial pressures arising from the marriage market fall largely on his family.² Children's marriage—especially sons'—is an important part of the economic picture for Chinese families. We hypothesize that as sex ratios increase, families with a son expect to spend more to appear financially attractive in order to find the son a marital partner, and they invest more in risky assets to realize it.

Empirical results are in line with our hypothesis. Data are from the 2013 China Household Finance Survey (CHFS), which provides rich information on household financial decisions, plus various demographic and socioeconomic characteristics of family members. We explore cross-sectional variation in sex ratios across counties and examine how this accounts for variation in portfolio choice among families with a firstborn son. While the gender of the first child is random (Ebenstein 2010), the major concern for our identification is that the sex ratio is not randomly assigned across counties; counties with high sex ratios might have other unobserved characteristics that affect household portfolio choice. To address this issue, we use families with a first daughter in the same county as a comparison group. As long as some of such characteristics affect portfolio choice of first-son and first-daughter families in a similar manner within the county, the difference-in-differences strategy reduces the confounding effects of such unobservable characteristics. Regressions also include a variety of controls for parental and household characteristics and prefecture fixed effects to deal with unobserved cross-prefecture heterogeneity (such as average risk tolerance or financial literacy), and allow all controls—including fixed effects—to have differential effects between the two types of families.³

1. Another example is India, in which the stock market value has increased sixfold—more than any other country except China; at the same time, its sex ratio is drastically rising (Amaral and Bhalotra 2017).

2. While sons ultimately provide financial support for aged parents after marriage, the increasing financial pressures on parents from marriage market competition occur before their son's marriage. Before they can relax, parents must accumulate a large amount of money for their son's marriage.

3. The administrative divisions of China consist of four practical levels: province, prefecture, county, and community. There are about nine counties per prefecture. Unobserved heterogeneity is better handled by prefecture dummies than province dummies.

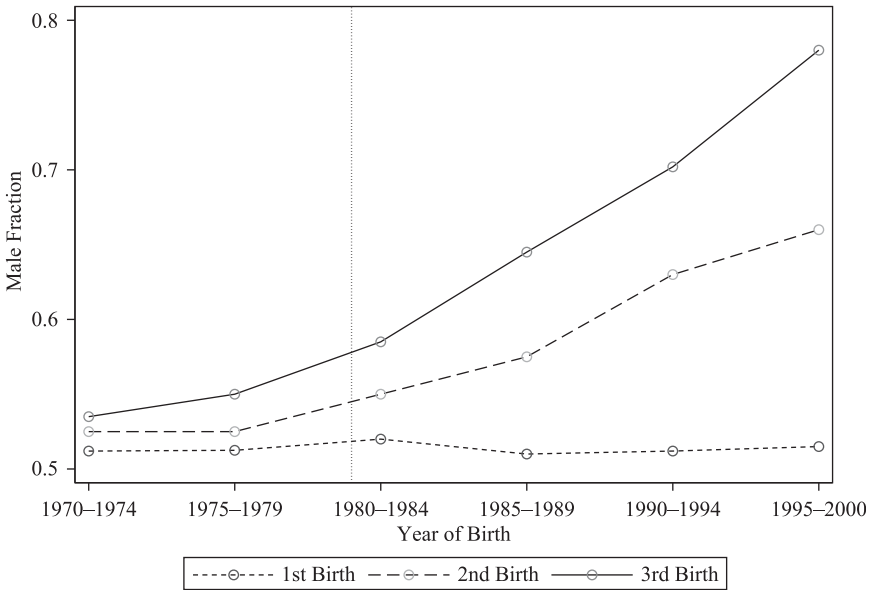


Figure 1
Male Fraction of Births by Birth Order in China

Notes: Data are from Ebenstein (2010). China introduced its family planning policy in 1979, which is represented by the vertical line. The figure shows that more boys than girls have been born in China over the past decades, but first births are exceptions.

We are particularly interested in whether and how high sex ratios affect the portfolio choices of the two types of families differently. This is captured by the coefficient on the interaction between a dummy indicating the first child being a son and the sex ratio. Our identification relies on the assumption that this interaction term is exogenous (even if the sex ratio itself may not be). In China, the first child in a family being a boy or a girl is plausibly random (Ebenstein 2010, 2011; Li and Wu 2011; Wei and Zhang 2011). Indeed, Figure 1 shows that about one-half of first births are boys and the other half are girls. In addition, comparing the effect of high sex ratios for first-son families with the effect for first-daughter families within the county removes unobserved confounders that affect portfolio holdings in a way that is independent of the first-child gender.

The baseline results show that high sex ratios have a significantly stronger impact on stock holdings for first-son families relative to first-daughter families. Specifically, for son families, a one standard deviation increase in the sex ratio (0.09) would increase the stock market participation rate by 2.9 percentage points, or 52.2 percent relative to the mean, and the share of liquid wealth held in stocks by 1.5 percentage points, or 81.7 percent. Using all categories of risky assets yields similar conclusions. We then present a

broad range of robustness checks to demonstrate that our results are confounded by neither potential issues related to son-preferring fertility-stopping rules nor identification concerns related to potential endogeneity of local sex ratios (partly due to gender selection in higher-order births).

We then show that high sex ratios affect parental risk-taking in a manner similar to how they affect portfolio choice. Our risk-taking measure is constructed based on survey questions, as in Malmendier and Nagel (2011). Results show that high sex ratios have a strongly positive effect on financial risk-taking for first-son families relative to first-daughter families. This is consistent with Cameron, Meng, and Zhang (2017), who find that high sex ratios are associated with greater risk-taking and impatience among men. We then examine the sex-ratio effect on risk-taking for three subgroups: low-income families (bottom 25 percent), middle-income families, and high-income families (top 25 percent). We find that the sex-ratio effect on risk-taking is the largest for low-income families.

This work contributes to the expanding literature on the determinants of household portfolio choice. Previous studies focus mostly on developed countries and document the role of participation or borrowing costs, nonstandard preferences, wealth and labor income, personal experience, housing, etc.⁴ We are the first to examine the significant role of unfavorable marriage market conditions that arise from high sex ratios, in the context of a large and fast-developing country. A back-of-the-envelope calculation indicates that rising sex ratios account for about 10 percent of the growth in China's stock market size in recent decades.⁵

This work also adds to the extensive literature on high sex ratios in China and other Asian economies. Numerous studies have shown that high sex ratios have large-scale and wide-ranging implications for households and society.⁶ In particular, our study builds on Wei and Zhang (2011), who find that parents tend to decrease consumption to enhance their sons' competitiveness in the marriage market when sex ratios are high. Using a similar empirical strategy, we augment their work by showing that parents with a son allocate more wealth to risky assets. In addition, our results are in line with the finding of Cameron, Meng, and Zhang (2017) that risk-taking and impatience are greater among men when sex ratios are high.

The next section provides background information. Section III describes the data used in our empirical analyses. Section IV presents the empirical results. Section V discusses potential explanations for our results. Section VI concludes with a policy discussion.

4. See, for example, Benzoni, Collin-Dufresne, and Goldstein (2007); Campbell et al. (2001); Chetty, Sándor, and Szeidl (2017); Cocco (2005); Cocco, Gomes, and Maenhout (2005); Davis, Kubler, and Willen (2006); Gomes and Michaelides (2005); Guiso, Sapienza, and Zingales (2008); Malmendier and Nagel (2011); Polkovnichenko (2007); and Grinblatt and Keloharju (2001).

5. The stock market value increased from 0.42 to 6.6 USD trillion, while GDP increased from 1.2 to 11.2. The sex ratio increased by 0.2. Baseline stock share of wealth was 2 percent.

6. See, for example, Das Gupta (2005); Das Gupta, Ebenstein, and Sharygin (2013); Du and Wei (2016); Ebenstein (2010, 2011); Edlund et al. (2013); Huang and Zhou (2016); Sharygin, Ebenstein, and Das Gupta (2013); and Bulte, Tu, and List (2015).

II. Background: High Sex Ratios in China and the Resulting High Marriage Expenditure

In China, people have historically preferred sons to daughters, which is one of the reasons for high sex ratios. The demand for sons may come from the tradition in a patriarchal society by which men tend to have higher labor productivity, anticipate higher earnings, and shoulder the main responsibility for aging parents (Chew et al. 2018; Das Gupta, Ebenstein, and Sharygin 2013; Ebenstein and Leung 2010; Pitt, Rosenzweig, and Hassan 2012). In 1979, China introduced a family planning policy that stipulates the number of children a couple is allowed to have, which further raises the sex ratio (Li, Yi, and Zhang 2011). People may undertake gender selection to satisfy their dual interests to have a son and comply with the policy. This is made possible by improved access to gender-screening technology (Ebenstein 2010).

A direct consequence of high sex ratios is intensified competition in the marriage market among oversupplied men. In the near future, more than 30 million marriage-age men may not be able to find a partner, which foreshadows a sizable bride shortage (Sharygin, Ebenstein, and Das Gupta 2013). Therefore, men must spend more to appear financially attractive to compete for prospective brides, which bids up marriage expenditure. Marriage expenditure typically consists of a one-time transfer to the bride and her parents, in the form of a bride price. It is usually accompanied by gifts, such as major durable goods. The negotiation of a bride price is traditionally required and continues to be the norm in most areas.⁷ Also, wedding ceremonies call for a large cash outlay. A fundamental component of marriage expenditure is housing, which is also the single largest lifetime financial commitment for most Chinese families. In social norms and marriage customs, a man would provide an apartment for himself and his bride or at least contribute most of the cost. In rural areas, men who have a relatively higher-quality house are much less likely to remain unmarried. In urban areas, men who are homeowners—as opposed to renters—are also less likely to remain single (Wei and Zhang 2011). More generally, a larger marriage expenditure enhances a man's marriage prospects in a competitive marriage market.

It is noteworthy that in China, the child's marriage is an important part of the economic picture for the whole family, rather than only for the child. This is especially the case when the child is a son, who is expected to get married to pass on the family name. In addition, given that today's bid-up marriage expenditure represents a substantial amount of wealth, a man of marriageable age is often unable to afford it by himself. In 2006, for example, the average wedding cost was about 5.5 times per capita income. Therefore, marriage expenditure is largely borne by the man's parents (Brown, Bulte, and Zhang 2011; Wei and Zhang 2011). Parents—who want to help improve their son's marital competitiveness—may increase their expectation of marriage expenditure for children based on local marriage market conditions. Data show that if the local sex ratio is one standard deviation higher, household-level expenditure on wedding ceremonies increases by more than one-fifth.⁸ Online Appendix Figures A1 and A2 show the

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

8. We use data from the 2010 China Family Panel Studies (CFPS) survey, which is regarded as nationally representative of Chinese families. Here sex ratios are at the county level. Regressions control for average household income and population composition, as well as province fixed effects.

positive correlation between sex ratios and marriage expenditure using cross-sectional and time-series data, respectively, especially for grooms' families.

With a larger expected marriage expenditure for children, families might adjust their financial investment decisions. We empirically examine how high sex ratios in China, which signal large marriage expenditure for sons, affect household portfolio choice.

III. Data

We first introduce our data source, the CHFS. We then describe our measures of household portfolio choice. We also provide preliminary evidence on how these measures respond to high sex ratios differently across first-son and first-daughter families.

A. The China Household Finance Survey

The CHFS is a large-scale, nationally representative, random sample of Chinese households. Our empirical analyses mainly rely on data from the 2013 nationwide baseline survey. Its data are consistent with China population census data on the distributions of a variety of demographic and socioeconomic variables (Gan et al. 2013).

The representativeness and high quality of the data are ensured by the large sample size, scientific sampling design, and low refusal rate. The 2013 CHFS covers 28,228 households in 29 (out of 34) provinces.⁹ It implements a three-stage probability-proportional-to-size sampling design. In the first stage, sampling units are counties that are sorted by local GDP per capita, and sampling weights are local populations. The second stage follows virtually the same principle, with sampling units being communities that are sorted by the proportion of nonagricultural population. The third stage is a systematic selection of households with an equal probability. The overall refusal rate is about 11 percent, which is much lower than internationally comparable surveys. For example, in the most recent Survey of Consumer Finances (SCF), the refusal rate is about 30 percent in the area-probability sample, two-thirds in the list sample, and even higher in the high-wealth sample.¹⁰ For the 2010 Eurosystem Household Finance and Consumption Survey, the refusal rate can be as high as 70 percent.¹¹

The CHFS provides detailed data on household financial decisions and activities, such as participation in the stock market and other financial asset markets, housing value and mortgage, income, expenditures, and wealth. Such information is obtained from a primary respondent who has the best knowledge of the household's financial status and, in most cases, is the household head. The CHFS also provides data on all family members' demographic, socioeconomic, and labor-market characteristics, such as age, gender, education, and occupation. Such information is obtained from both the respondent and other family members. As in the SCF, face-to-face interviews are aided by

9. Hong Kong, Macao, Taiwan, Xinjiang, and Tibet are not included.

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

11. The lowest refusal rates are slightly above 10 percent. For more information on this data set and its refusal rates, see Eurosystem Household Finance and Consumption Network (2013a,b).

computer-assisted-interviewing technology, which yields better-quality information than traditional methods (Caeyers, Chalmers, and De Weerd 2012). In addition, we can group individuals by living unit and identify parent–child relationships using a set of identification numbers. These features enable systematic empirical analyses.

Our empirical analyses are based on a cross-sectional sample drawn from the baseline CHFS survey, which provides the most comprehensive set of information on household portfolio choice compared with other wave, and exploit cross-county variation in the sex ratio. This does not render our analyses less strong, partly because variation in sex ratios across time is not that large within a county.

Specifically, we extract a sample of households from the 2013 CHFS, which includes households with at least one child, the eldest child being aged 3–17, and both parents being aged 22–51. As children are under the age of 18 and therefore not yet legal adults, they are less likely to play an important role in family financial decisions. In addition, by placing age limits on both parents and children, we maximize comparability across households. The sample contains 4,363 observations.

B. Measures of Household Portfolio Choice

We construct four measures of household portfolio choice from the CHFS data. The first measure is a dummy indicating stock market participation, which is equal to one if the household has a stock account on the survey date and zero otherwise. The second measure is stock share, defined as the fraction of liquid wealth held in stocks. Liquid wealth is the sum of holdings in risky assets plus holdings in riskless assets. Risky assets include stocks, risky bonds, mutual funds, financial derivatives, and other risky financial products; riskless assets consist of cash and savings accounts. Stock share is set to zero for nonparticipants and those with zero reported liquid wealth by our definition.¹² The other two measures—a dummy for risky asset market participation and risky asset share of liquid wealth—are analogously defined for all categories of risky assets.

1. Summary statistics

Table 1 presents summary statistics for household portfolio choice measures, which are weighted by the CHFS sampling weights.¹³ As Panel A shows, the share of households that participate in the stock market and the risky asset market is 5.6 and 8.4 percent, respectively. On average, households hold 1.8 percent of liquid wealth in stocks and 3.2 percent in all risky assets. These patterns are consistent with the common finding of generally limited household investments in risky financial assets. Of stock market participants—families that hold a stock account—the stock share is 40 percent and the risky asset share is 48.9 percent. This implies that while a large fraction of households avoid participation, for those who participate, their holdings comprise a substantial fraction of family wealth.

12. Setting stock share for nonparticipants to zero yields unconditional estimates, which give the population average causal effect without raising any selection bias (Angrist and Pischke 2008).

13. High-income households are oversampled, as in the SCF. No weighting may generate unrepresentative results; this concern outweighs the loss of efficiency associated with weighting (Malmendier and Nagel 2011).

Table 1
Summary Statistics of Household Portfolio Choice Measures

	All			Stock Market Participants		
	Mean (1)	Min. (2)	Max. (3)	Mean (4)	Min. (5)	Max. (6)
Panel A: Overall						
Stock market participation	0.056 (0.230)	0	1	1 (0)	1	1
Stock share	0.018 (0.115)	0	1	0.400 (0.363)	0	1
Risky-asset market participation	0.084 (0.278)	0	1	1 (0)	1	1
Risky-asset share	0.032 (0.153)	0	1	0.489 (0.367)	0	1
Observations	4,363			245		
	Sex Ratio < Q1			Sex Ratio > Q3		
	Mean		<i>p</i> -Value	Mean		<i>p</i> -Value
	First Son (1)	First Daughter (2)	(H1: SD ≠ 0) (3)	First Son (4)	First Daughter (5)	(H1: SD > 0) (6)
Panel B: First-Son versus First-Daughter Families						
Stock market participation	0.043 (0.202)	0.046 (0.210)	0.807	0.059 (0.236)	0.032 (0.177)	0.018
Stock share	0.020 (0.117)	0.020 (0.126)	0.933	0.023 (0.125)	0.013 (0.104)	0.092
Risky-asset market participation	0.064 (0.245)	0.068 (0.252)	0.813	0.097 (0.296)	0.050 (0.218)	<0.01
Risky-asset share	0.030 (0.149)	0.036 (0.166)	0.566	0.046 (0.182)	0.022 (0.129)	<0.01
Observations	469	457		507	560	

Notes: Data on county-level sex ratios are from the 2010 China population census. Data on other variables are from the 2013 CHFS. Panel A is based on our sample of CHFS households. Stock-market participants are families that hold a stock account. Panel B is based on four subgroups: first-son and first-daughter families in counties with a balanced sex ratio (smaller than the first quartile Q1, 1.09) or a high sex ratio (larger than the third quartile Q3, 1.22). Statistics are weighted by CHFS sampling weights. Standard deviations are given in parentheses.

2. Preliminary evidence

Panel B in Table 1 reports summary statistics for four subsamples: first-son and first-daughter families in counties with a balanced sex ratio (smaller than the first quartile, 1.09) or a high sex ratio (larger than the third quartile, 1.22). We use sex ratios for the child cohort in each county, data on which are from the 2010 population census (more details are given in the next section). We observe that with balanced sex ratios, there is no significant difference in portfolios between the two types of families. But when sex ratios are high, first-son families appear significantly more likely to hold—and hold more—risky assets than first-daughter families. [Online Appendix Table A1](#) conveys virtually the same message. After controlling for potential covariates and prefecture dummies, having a first son increases risky asset holdings for families in high-sex-ratio regions, but this does not affect families in balanced-sex-ratio regions. These patterns suggest different responses to high sex ratios between first-son and first-daughter families in their portfolios and also suggest that, in the absence of pressures arising from the marriage market, child gender itself does not affect household portfolios.

IV. Empirical Results

We have seen that first-son and first-daughter families seem to react differently to high sex ratios in managing their portfolios. In this section, we turn to rigorous empirical analyses. We first present the regression specification. We then report the baseline results and demonstrate the robustness. We also report the results using a survey-based measure of risk-taking and conduct a subgroup analysis.

A. Regression Specification

We estimate the following regression equation:

$$(1) \quad y_{icp} = \beta_0 + \beta_1 son_{icp} + \beta_2 sex_ratio_{cp} + X_{icp} \Gamma_1 + \lambda_p \\ + \beta_3 son_{icp} * sex_ratio_{cp} + son_{icp} * X_{icp} \Gamma_2 + son_{icp} * \lambda_p + \varepsilon_{icp},$$

where the dependent variable, y_{icp} , is a measure of portfolio holdings of household i in county c , prefecture p . We use the four measures defined above as our main dependent variables. The key control variable, son_{icp} , is a dummy equal to one if the firstborn child in household i in county c , prefecture p , is a boy. Another key control variable, sex_ratio_{cp} , is the sex ratio for the child cohort in county c , prefecture p (there are about nine counties per prefecture in China). A vector of additional control variables, X_{icp} , includes various parental and household characteristics: both parents' age, education, hukou (a household registration record that officially identifies a person as a resident of an area), political status, and occupational dummies, plus age of the first child, region of residence, and ethnicity.¹⁴ Regressions also control for prefecture fixed effects, λ_p . The error term is denoted by ε_{icp} .

14. In robustness checks, we include different sets of controls. For example, we add the number of children, average household income, etc. The pattern of results is preserved; see Section V.

In Equation 1, we also control for interactions of the first-son dummy, son_{icp} , with all other control variables, including prefecture fixed effects. Our results therefore are not subject to the restriction that the effects of these controls are the same for first-son and first-daughter families. We are particularly interested in the interaction term coefficient β_3 , which measures the effect of high sex ratios on portfolio holdings for first-son families relative to first-daughter families—it is expected to be significantly positive. This is equivalent to estimating separate models with the sex ratio, other controls, and prefecture dummies for the two types of families and then looking at the difference in the estimated coefficient on the sex ratio between the two models.

We assume that parents infer the local sex ratio relevant for their children, for instance, by observing the gender composition of students in local schools, and react accordingly. We use sex ratios for the cohort that corresponds to the children in our sample (3–17 in 2013).¹⁵ These sex ratios are less likely to be altered by migration, which captures other local conditions, such as job and investment opportunities, and are therefore less prone to measurement error. They are projected using the 0–14 age group in the 2010 population census. Sex ratios are at the county level, as each county can be treated as a local marriage market. China's hukou system presents a formidable obstacle to marriage migration (Davin 2005; Wei and Zhang 2011). The census shows that more than 90 percent of rural residents and 62 percent of urban residents live in their county of birth, and 89 percent of couples are from the same county. Of migrant couples in cities, 82 percent are from the same place, suggesting that migrants often get married before leaving their hometowns. In our regressions, we use the deviation of the sex ratio from the mean in our sample.

Column 1 of Table 2 reports summary statistics for control variables. It shows that in our sample, 51.8 percent of families have a first son. On average, fathers are 40 years old and have 9.2 years of schooling, and mothers are 38 years old and have 8.3 years of schooling; 22.7 percent of fathers and 21.5 percent of mothers are registered as urban residents in the hukou system; and 11.4 percent of fathers and 3.1 percent of mothers are members of the Chinese Communist Party (CCP). In addition, an average first child is 12 years old, 48.5 percent of households reside in urban areas, and 12.8 percent belong to ethnic minorities.

Obtaining unbiased ordinary least squares (OLS) estimates of the sex-ratio effect (β_3) requires that in Equation 1 the error term is not substantially correlated with the interaction between the first-son dummy and the local sex ratio.

The identification partly relies on a well-recognized demographic regularity, that the first child in a family being a boy or a girl is plausibly random.¹⁶ Data from China population censuses (1982, 1990, 2000, and 2010) reveal that high sex ratios in China are driven by imbalances between second- and higher-order births, while the sex ratio for first births is rather stable and falls in the biologically normal range (Figure 1). Parents are least likely to practice gender selection on the first birth, despite their son

15. We find that results are not very sensitive to using sex ratios for different age brackets. This is perhaps due to the persistence of the local level of sex-ratio distortion over time. In addition, as any potential measurement error in sex ratios tends to produce attenuated coefficient estimates, our results can at least be interpreted as lower bounds of the true sex-ratio effect.

16. The interaction term is exogenous as long as the gender of the first child is random, even if the sex ratio is not.

Table 2
Balance Test: First-Son versus First-Daughter Families

	Mean				
	All (1)	First Son (2)	First Daughter (3)	Difference, SD (4)	<i>p</i> -Value (H_1 : Diff. \neq 0) (5)
First son	0.518 (0.500)				
Father's age	39.52 (5.719)	39.54 (5.769)	39.51 (5.666)	0.030 [0.173]	0.655
Father's schooling years	9.170 (3.313)	9.149 (3.263)	9.193 (3.367)	-0.044 [0.104]	0.530
Father's hukou (urban = 1)	0.227 (0.419)	0.226 (0.418)	0.229 (0.420)	-0.003 [0.014]	0.845
Father's political status (CCP = 1)	0.114 (0.318)	0.104 (0.305)	0.126 (0.332)	-0.022 [0.010]	0.129
Mother's age	37.62 (5.815)	37.74 (5.852)	37.50 (5.775)	0.240 [0.176]	0.331
Mother's schooling years	8.347 (3.659)	8.328 (3.548)	8.367 (3.776)	-0.039 [0.115]	0.820
Mother's hukou (urban = 1)	0.215 (0.411)	0.217 (0.412)	0.213 (0.409)	0.004 [0.013]	0.132
Mother's political status (CCP = 1)	0.031 (0.173)	0.031 (0.173)	0.031 (0.174)	-0.000 [0.006]	0.882
First child's age	11.68 (4.105)	11.69 (4.057)	11.67 (4.158)	0.020 [0.127]	0.856
Region of residence (urban = 1)	0.485 (0.500)	0.476 (0.500)	0.496 (0.500)	-0.020 [0.015]	0.319
Ethnicity (minority = 1)	0.128 (0.334)	0.126 (0.332)	0.130 (0.336)	-0.004 [0.010]	0.640
Observations	4,363	2,246	2,117		

Notes: Data are from the 2013 CHFS. Statistics are based on our sample of CHFS households. Statistics are weighted by CHFS sampling weights. Standard deviations are given in parentheses. Standard errors are given in square brackets.

preferences. Before 2015, a second child was officially permitted if the first one was a girl for households in most rural areas, where son preferences appear stronger. This “1.5 children” policy was applicable to residents who accounted for more than 60 percent of the Chinese population and markedly alleviated their motivation to abort a first daughter.¹⁷

Statistical evidence from our sample also validates the randomness of first-child gender. An average family in our sample has 1.5 children, consistent with the above-mentioned policy. Nearly half of the families have a first son, and the other half have a first daughter. Specifically, the mean of the first-son dummy is 0.518, which implies a sex ratio well within the normal range; the standard deviation is 0.5, which suggests that first-child gender is like a random Bernoulli trial in which having a boy or a girl has an equal probability (Table 2).¹⁸ In addition, we regress the first-son dummy on the full set of control variables used in our analyses and find no significant effect of these variables. Most importantly, first-son and first-daughter families have similar predetermined parental and household characteristics, as the balance test in Table 2 shows. For example, 47.6 percent of first-son families and 49.6 percent of first-daughter families are in urban areas. The difference is -0.02 , which is statistically not different from zero at the 10 percent level or below.

In later sections, we present a broad range of robustness analyses to demonstrate that our findings are not driven by identification issues.

B. Baseline Results

Results from estimating Equation 1 based on our sample are reported in Table 3. We mainly use the OLS method in this table and the following tables.¹⁹ Estimations are weighted by the CHFS sampling weights. Standard errors are clustered at the county level and given in parentheses.

In the first column of Table 3, the dependent variable is a dummy indicating stock market participation. The regression is therefore a linear probability model (LPM). At the top of the table we present estimated coefficients on the first-son dummy, the local sex ratio, and their interaction term— β_1 , β_2 , and β_3 .²⁰ The estimated β_3 is 0.326 (standard error 0.160), which is positive and statistically significant at the 5 percent level. This estimate suggests that high sex ratios increase stock market participation among first-son families relative to first-daughter families. Specifically, a one standard deviation (0.09) increase in the sex ratio would, on average, raise the probability that a first-son family participates in the stock market (or overall participation rate) by roughly 2.9 percentage points. This represents a 52.2 percent increase relative to the mean stock ownership probability.

17. The policy was replaced by a nationwide two-child policy in 2015, which further alleviates the motivation to abort the first daughter. Ebenstein (2010, 2011) shows that most Chinese families prefer one boy and one girl to two boys.

18. A Bernoulli random variable with a mean of 0.5 has a standard deviation of 0.5.

19. Standard OLS estimates are consistent with marginal effects from Probit models in most empirical practices (Angrist and Pischke 2008). We have verified that this is also the case in our study.

20. Unreported estimates of coefficients on other controls have the sign and magnitude that one would expect given prior studies' findings.

Table 3
Baseline Results: Sex Ratios and Household Portfolio Choice

Dependent Variable	Stocks		All Risky Assets	
	Participation (1)	Share (2)	Participation (3)	Share (4)
First son (β_1)	-0.229*** (0.062)	-0.069** (0.032)	-0.089 (0.080)	-0.011 (0.045)
Sex ratio (β_2)	-0.027 (0.179)	-0.001 (0.075)	-0.085 (0.199)	0.038 (0.097)
First son \times sex ratio (β_3)	0.326** (0.160)	0.168* (0.101)	0.555*** (0.206)	0.227** (0.115)
Percentage increase: sex ratio+1 SD	52.2	81.7	59.4	64.2
Observations	4,363	4,363	4,363	4,363
R-squared	0.271	0.194	0.304	0.256
Dependent variable mean	0.056	0.018	0.084	0.032
Model	LPM	OLS	LPM	OLS
Other controls	Yes	Yes	Yes	Yes
First son \times other controls	Yes	Yes	Yes	Yes
Prefecture fixed effects	Yes	Yes	Yes	Yes
First son \times prefecture fixed effects	Yes	Yes	Yes	Yes

Notes: Data on county-level sex ratios are from the 2010 China population census. Data on other variables are from the 2013 CHFS. Results are estimated using Equation 1 based on our sample of CHFS households. Percentage increase with a one standard deviation increase in the sex ratio is $0.09\beta_3/\text{sample mean}$. Other controls include various parental and household characteristics—both parents' age, education, hukou, political status, and occupational dummies, plus age of the first child, region of residence, and ethnicity. Interactions of the first-son dummy with these variables as well as prefecture fixed effects are also controlled for. Regressions are weighted by CHFS sampling weights. Standard errors clustered at the county level are given in parentheses. * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$.

We note that the coefficient on the first-son dummy, β_1 , is difficult to interpret, as our regressions include its interactions with all other controls. As discussed in Section IV.A, the empirical approach is equivalent to estimating separate models by subsamples of first-son and first-daughter families; see [Online Appendix Table A2](#) for the estimated sex-ratio effect from separate models and the difference in the effect. Results from generalized Hausman tests formally show that the sex-ratio effect on portfolios is significantly different between the two models.

In Column 2, the dependent variable is the stock share of liquid wealth. Again, the estimated β_3 is positive and statistically significant. This suggests that high sex ratios significantly increase the share of liquid wealth held in stocks by first-son families relative to first-daughter families. Based on the estimate, as the sex ratio increases by one standard deviation, the stock share in first-son families would increase by 1.5 percentage points, or 81.7 percent relative to the mean.

In these results, while the estimated sex-ratio effect appears modest in absolute terms, it becomes substantial when translated to percentages because of the small denominator. As noted in Section III.B, only a small fraction of families in our sample participate in the stock market (5.6 percent), and as a result, the average stock share is also small (1.8 percent). Among those who participate, however, the share is large (40 percent). Prior studies have similar findings. In our hypothesis, families invest in stocks in order to increase expenditure on sons' marriage to increase their chance of finding a partner. It is therefore helpful to quantify how large such an increase is. Our estimates imply that, roughly, a family could spend ¥3,600 more on a child's marriage if it invested in stocks, which represents a 20 percent increase in the average household-level marriage expenditure in China.²¹

The remaining two columns in Table 3 report results from all risky assets. They report evidence that is similar to the first two columns. The evidence is strong, both in terms of statistical significance and economic magnitude, and indicates that high sex ratios induce first-son families to hold more risky assets relative to first-daughter families. Specifically, the probability of holding risky assets (or overall market participation rate) in first-son families would be about five percentage points, or 59.4 percent higher, with a one standard deviation increase in the sex ratio; the share invested in risky assets would be about two percentage points, or 64.2 percent higher.

Household investment in stocks and investment in other risky financial products may not be determined in a similar manner per se. Whether, and how much, households allocate wealth to these assets may also be distinct investment decisions. Therefore, the consistency of our results across these dimensions is impressive. This suggests that marriage market conditions—as measured by child gender and the local level of sex-ratio imbalance—are important for explaining household portfolio choice.

C. Robustness: Potential Issues Related to Son-Preferring Fertility-Stopping Rules

While we have verified that the gender of the first child can be viewed as random, such an argument for identification is potentially not quite complete. Due to son preferences, subsequent fertility decisions may be different between first-son and first-daughter families. The latter typically are more likely to have a second or third child in order to have a boy, while the former are more likely to stop childbearing and therefore have a smaller family (Ebenstein 2011).

One thus might worry that our findings on portfolio allocations may simply reflect the effect of family size. To net out such an effect, we additionally control for the number of children in a family and its interaction with the first-son dummy. Column 1 of Table 4 reports the results using the dummy for stock market participation as the dependent variable. We observe that the results after controlling for family size are not significantly different from the baseline (the first column of Table 3). The coefficient on the interaction

21. We assume that the expected return in the stock market is 7 percent. Our estimation sample shows an average liquid wealth of ¥32,200, and on average, a family invested 40 percent in stocks. We also assume that an average family invests 40 percent of liquid wealth (¥32,200) for four years. Such investments yield ¥3,600 ($¥32,200 \times 0.4 \times 0.07 \times 4$; we use simple interest for simplicity), which we assume the family spends on child (ren)'s marriage. This amount represents a 20 percent increase, given that the average household-level marriage expenditure in China is about ¥18,000, based on 2010 CFPS survey data.

Table 4
Robustness: Addressing Issues Related to Son-Preferring Fertility Stopping Rules

Dependent variable	Stock-Market Participation						
	(1)	(2)	(3)	(4)	(5)	(6)	(7)
First son (β_1)	-0.235*** (0.068)	-0.387*** (0.104)	-0.515*** (0.154)				
Proportion of sons (β_1)				-0.310*** (0.076)	-0.330*** (0.082)		
Having any son (β_1)						-0.199*** (0.073)	-0.228*** (0.077)
Sex ratio (β_2)	-0.021 (0.180)	-0.188 (0.390)	-0.099 (0.309)	-0.098 (0.223)	-0.092 (0.222)	-0.176 (0.290)	-0.174 (0.288)
First son \times sex ratio (β_3)	0.321** (0.161)	0.657* (0.386)	0.705* (0.403)				
Proportion of sons \times sex ratio (β_3)				0.460** (0.232)	0.455* (0.231)		
Having any son \times sex ratio (β_3)						0.464* (0.280)	0.462* (0.279)
# children	-0.004 (0.006)				-0.007 (0.006)		-0.013 (0.009)

(continued)

Table 4 (continued)

Dependent variable	Stock-Market Participation						
	(1)	(2)	(3)	(4)	(5)	(6)	(7)
First son × # children	0.002 (0.009)						
Proportion of sons * # children					0.008 (0.012)		
Having any son × # children							0.013 (0.011)
Observations	4,363	2,726	2,216	4,363	4,363	4,363	4,363
R-squared	0.271	0.236	0.246	0.264	0.264	0.210	0.210
Other controls	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Son × other controls	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Prefecture fixed effects	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Son × prefecture fixed effects	Yes	Yes	Yes	Yes	Yes	Yes	Yes

Notes: Data on county-level sex ratios are from the 2010 China population census. Data on other variables are from the 2013 CHFS. Results are estimated using Equation 1 based on our sample of CHFS households. Column 2 includes one-child families. Column 3 includes one-child families with a child above the age of six. Other controls include various parental and household characteristics—both parents' age, education, hukou, political status, and occupational dummies, plus age of the first child, region of residence, and ethnicity. Interactions of the first-son dummy with these variables as well as prefecture fixed effects are also controlled for. Regressions are weighted by CHFS sampling weights. Standard errors clustered at the county level are given in parentheses. * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$.

term, β_3 , is estimated to be positive and statistically significant, as in the baseline; the magnitude is also similar. While Table 4 reports stock market participation results, using other portfolio choice measures yields similar patterns.

An alternative strategy to deal with the potentially confounding family-size effect is to restrict the sample to families with only one child. This yields 2,726 observations. Results are reported in Column 2 of Table 4 and show a similar pattern to the baseline in terms of the sign and statistical significance of the sex-ratio effect. The pattern remains similar in Column 3, in which we further restrict the sample to families that are less likely to have a second child. Specifically, the sample includes one-child families in which the child is above the age of six, or 2,216 observations. We note that the magnitude of the sex-ratio effect appears larger in these two columns than the baseline.

Another issue raised by son-preferring fertility-stopping rules is that the gender of the first child may not well represent actual child-gender composition and therefore may not adequately capture the expected marriage expenditure of parents. As we have discussed, it is common in China that a second son follows a first daughter. These families may have great expected marriage expenditure as well. To address this concern, we replace the first-son dummy in Equation 1 with the proportion of sons and a dummy for having any son. This yields qualitatively similar results as before.²² We then control for the number of children (as well as its interaction with the child-gender measure) in these empirical exercises, and again we obtain similar results (Table 4, Columns 4–7). These patterns confirm that how families allocate their portfolios in response to high sex ratios depends on child gender.

In summary, robustness results show that concerns related to son-preferring fertility-stopping rules are not likely to be the main driver of our findings.

D. Robustness: Potential Endogeneity of Local Sex Ratios

We have followed the practice of focusing on OLS estimates for the sex-ratio effect (Edlund et al. 2013; Wei and Zhang 2011). But strictly speaking, county-level sex ratios may not be exogenous, even within prefectures. Below we provide evidence that alleviates this concern.

Counties with higher sex ratios perhaps have unobserved characteristics that make families more or less likely to invest in risky assets. This is not an important problem in our data, however, because the correlation between sex ratios and portfolio choice turns out to be small and statistically insignificant; see [Online Appendix Table A3](#). More importantly, using daughter families as a comparison group reduces the confounding effects of such characteristics, as long as they affect the portfolio choices of first-son and first-daughter families in a similar manner within a county. If certain factors affect the two types of families differently, our estimates may be biased. Section II discussed many such factors.

The first possible factor is historical son preferences. For example, some wealthier areas of China retain stronger demand for sons. Also, wealth has a positive effect on risky investments (Calvet, Campbell, and Sodini 2009). To partially evaluate the importance

22. When we use the proportion of sons, the estimate is for the sex-ratio effect for all-son families. When we use a dummy for having any son, the estimate is for the sex-ratio effect for families with at least one son.

of such confounding effects, we control for county-level average household financial wealth—defined as the sum of liquid and illiquid assets—and household income. The second factor is gender difference in earnings. We therefore control for a county-level gender wage differential. The third factor is social old-age support, a lack of which increases the demand for sons. Accordingly we control for the proportion of adults covered by social insurances at the county level. The fourth factor is the implementation of China's family planning policy, which varies from place to place. As fertility is the direct target of this policy, and thus can be regarded as a proxy for its implementation, we control for the average number of children per family in the county. The last factor is technological development, particularly gender-selection technology. This again can be proxied by local average household wealth or income. In these robustness regressions, we also control for the interaction between the first-son dummy and the newly added variable.

Table 5 reports robustness results using the stock market participation dummy as the dependent variable. It shows that controlling for these potentially confounding factors—either individually or collectively—causes a very small difference in our results. In each robustness regression, the estimated coefficient on the interaction term β_3 is similar to the baseline estimate. In particular, we perform a generalized Hausman test to show formally that the estimated sex-ratio effect is not significantly different from the baseline. The lack of sensitivity of our results to the inclusion of various possible sex-ratio confounders implies that omitted variable bias is not likely an issue (Altonji, Elder, and Taber 2005). Therefore, the potential endogeneity of sex ratios may not be a primary concern in our estimation results.

Another reason we focus on OLS estimates for the sex-ratio effect is to avoid problems with common candidates for instruments, such as variation in the financial penalty levied for unauthorized births. China's family planning policy is passed down the administrative chain of command until it is interpreted and adapted to suit local needs (Short and Zhai 1998). Therefore, birth quotas are endogenously stipulated by local governments based on local conditions and may be correlated with household portfolio choice independent of the sex ratio.

E. A Survey-Based Measure of Risk-Taking and Subgroup Analysis

We construct a measure of financial risk-taking based on a 2013 CHFS survey question, in a manner similar to Malmendier and Nagel (2011). In this question, the respondent is asked which of the following statements most precisely describes their degree of risk-taking in investment decisions under financial risk: (i) taking substantial risks and expecting to earn substantial returns; (ii) taking above-average risks and expecting to earn above-average returns; (iii) taking average risks and expecting to earn average returns; (iv) taking below-average risks and expecting to earn below-average returns; and (v) not willing to take any risk. Prior studies have shown that answers to such questions about hypothetical financial decisions are correlated with actual decisions, and they largely reflect risk attitude (Dohmen et al. 2011). We code the answers as a dummy variable that is equal to one if the first or second option is chosen by the respondent and zero otherwise. Its mean is 0.131 in our estimation sample. We use this measure as the dependent variable and estimate Equation 1.

Table 5
Robustness: Addressing Potential Endogeneity of Local Sex Ratios

Dependent Variable	Stock-Market Participation					
	(1)	(2)	(3)	(4)	(5)	(6)
First son (β_1)	-0.228*** (0.063)	-0.221*** (0.072)	-0.251*** (0.064)	-0.251*** (0.093)	-0.314*** (0.089)	-0.170 (0.221)
Sex ratio (β_2)	-0.022 (0.176)	-0.034 (0.168)	-0.032 (0.181)	-0.023 (0.172)	0.052 (0.175)	-0.068 (0.155)
First son \times sex ratio (β_3)	0.337** (0.160)	0.344** (0.167)	0.320* (0.187)	0.339** (0.170)	0.273* (0.157)	0.334* (0.179)
Average household wealth, k	0.022 (0.028)					-0.007 (0.031)
First son \times average household wealth	-0.015 (0.048)					-0.060 (0.039)
Average household income, k		0.002* (0.001)				0.003*** (0.001)
First son \times average household income		-0.000 (0.001)				-0.003*** (0.001)
Gender wage differential, m-f, k			0.001 (0.005)			-0.015*** (0.005)
First son \times gender wage differential			0.011** (0.005)			0.031*** (0.006)

(continued)

Table 5 (continued)

Dependent Variable	Stock-Market Participation					
	(1)	(2)	(3)	(4)	(5)	(6)
Proportion with insurance				0.212 (0.155)		0.091 (0.187)
First son × proportion with insurance				0.067 (0.174)		0.088 (0.193)
Average # children					-0.083** (0.035)	-0.039 (0.042)
First son × average # children					0.066 (0.043)	0.035 (0.047)
Hausman test <i>p</i> -value	0.780	0.993	0.653	0.877	0.568	0.671
Observations	4,363	4,363	4,363	4,363	4,363	4,363
<i>R</i> -squared	0.270	0.273	0.273	0.272	0.271	0.328
Other controls	Yes	Yes	Yes	Yes	Yes	Yes
First son × other controls	Yes	Yes	Yes	Yes	Yes	Yes
Prefecture fixed effects	Yes	Yes	Yes	Yes	Yes	Yes
First son × prefecture fixed effects	Yes	Yes	Yes	Yes	Yes	Yes

Notes: Data on county-level sex ratios are from the 2010 China population census. Data on other variables are from the 2013 CHFS. Results are estimated using Equation 1 based on our sample of CHFS households. The null hypothesis of the Hausman test is that the effect of high sex ratios is equal to that in the baseline—Column 1 of Table 3. Other controls include various parental and household characteristics—both parents' age, education, hukou, political status, and occupational dummies, plus age of the first child, region of residence, and ethnicity. Interactions of the first-son dummy with these variables, as well as prefecture fixed effects, are also controlled for. Regressions are weighted by CHFS sampling weights. Standard errors clustered at the county level are given in parentheses. **p* < 0.1, ***p* < 0.05, ****p* < 0.01.

Table 6
A Survey-Based Measure of Risk-Taking and Subgroup Analysis

Dependent Variable	Risk-Taking Measure									
	All	Bottom 25%	Middle	Top 25%	(1)	(2)	(3)	(4)	(5)	(6)
First son × sex ratio	0.666** (0.267)	1.633** (0.754)	1.767** (0.797)	0.743* (0.435)	0.548 (0.433)	-0.217 (0.836)	-0.650 (0.938)			
First son × sex ratio × skewness			0.093 (0.085)		-0.050 (0.047)		-0.134 (0.124)			
Observations	4,363	1,140	1,140	2,113	2,113	1,110	1,110			
R-squared	0.307	0.372	0.377	0.236	0.238	0.399	0.401			
Dependent variable mean	0.131	0.142	0.142	0.128	0.128	0.123	0.123			
Model	LPM	LPM	LPM	LPM	LPM	LPM	LPM			
First son	Yes	Yes	Yes	Yes	Yes	Yes	Yes			
Sex ratio	Yes	Yes	Yes	Yes	Yes	Yes	Yes			
First son × skewness										
Sex ratio × skewness										
Other controls	Yes	Yes	Yes	Yes	Yes	Yes	Yes			
First son × other controls	Yes	Yes	Yes	Yes	Yes	Yes	Yes			
Prefecture fixed effects	Yes	Yes	Yes	Yes	Yes	Yes	Yes			
First son × prefecture fixed effects	Yes	Yes	Yes	Yes	Yes	Yes	Yes			

Notes: Data on county-level sex ratios are from the 2010 China population census. Data on other variables are from the 2013 CHFS. Results are estimated using Equation 1 based on our sample of CHFS households and income subgroups. Other controls include various parental and household characteristics—both parents' age, education, hukou, political status, and occupational dummies, plus age of the first child, region of residence, and ethnicity. Interactions of the first-son dummy with these variables, as well as prefecture fixed effects, are also controlled for. Regressions are weighted by CHFS sampling weights. Standard errors clustered at the county level are given in parentheses. * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$.

Column 1 of Table 6 reports results using the full sample. Similar to the baseline results on portfolio allocations in Table 3, the estimated interaction term coefficient β_3 is positive and statistically significant. This implies that high sex ratios have a strongly positive effect on financial risk-taking for first-son families relative to first-daughter families. Specifically, the number of first-son families who are willing to take great risk in financial investments would increase by 46 percent as the sex ratio becomes one standard deviation higher. This finding is in line with and supplements our results on household portfolio allocations. (See [Online Appendix Table A4](#) for correlations between the risk-taking measure and household portfolio choice measures.) It is also consistent with Cameron, Meng, and Zhang (2017), who, as noted previously, show that high sex ratios are related to greater risk-taking and impatience among men.

We then divide our sample into three groups—low-income families (bottom 25 percent at the county level), middle-income families, and high-income families (top 25 percent at the county level)—and examine the sex-ratio effect on risk-taking for these groups. Columns 2, 4, and 6 of Table 6 report the results. Comparing the three columns, we find that the sex-ratio effect on risk-taking is the largest for low-income families. For middle-income families, the sex-ratio effect is still positive and statistically significant. But for high-income families, the effect is not significantly different from zero.

For each subgroup of families, we further control for an interaction between skewness of income distribution at the county level and the variable that measures the sex-ratio effect (the first-son dummy interacted with the sex ratio) to account for the potential effect of income distribution on how risk attitude responds to marriage market competition. Results are reported in Columns 3, 5, and 7 of Table 6. We observe that this does not change our conclusion that the sex-ratio effect on risk-taking for son families is the largest for the low-income group. We also observe that the interaction with skewness is not significantly different from zero. These findings again supplement the results on household portfolios, for which a similar subgroup analysis is not feasible due to the negligible fraction of stock market participants in low-income families.

V. Discussion

We have shown that high sex ratios have a strong impact on risky asset holdings for son families relative to daughter families. We have also shown a similar effect on financial risk-taking that is consistent with our risky asset holdings results. In this section, we explore potential explanations for these empirical findings.

Specifically, we discuss theories that explain (i) why expected marriage expenditure increases for families with a son as marriage market competition becomes stronger (due to rising sex ratios) and (ii) why these families become more risk-taking and therefore adopt a more aggressively risky asset-holding strategy, as their expected marriage expenditure increases.

That expected marriage expenditure increases with sex ratios for families with a son can be explained by the marriage matching theory of Becker (2009) and Browning, Chiappori, and Weiss (2014). In this theory, expected marriage expenditure represents the shadow price of finding a marital partner. The shadow price is an equilibrium outcome determined by supply and demand in the marriage market, which in turn depends on the local sex ratio and child gender. When the sex ratio rises, there is an oversupply of

men and a shortage of women in the marriage market; thus, expected marriage expenditure increases for son families. This is also in line with the implications of the status-competition theory. This theory implies that large marriage expenditure may serve to signal attractiveness as a partner in a competitive marriage market because it may indicate wealth that is otherwise unobservable. Prior studies find that expenditure on marriage-related status goods, such as a grandiose wedding and housing, increases as the sex ratio becomes higher (Brown, Bulte, and Zhang 2011; Wei, Zhang, and Liu 2017).

In addition to explaining why expected marriage expenditure responds to changes in marriage market competition, the status-competition theory explains why risk attitude in turn responds to changes in expected marriage expenditure. For example, Robson (1992) define an individual utility that not only depends on one's own wealth but also on the relative standing this wealth induces, implying that concerns about status shape people's risk attitudes. Roussanov and Savor (2014) show that the unmarried allocate more wealth to risky assets, as this can change their relative position in a competitive marriage market. More specifically, the theory with status-based utility predicts that under increasing marriage market competition, those at the lower spectrum of income distribution face greater pressure and therefore are more likely to assume risks (Hopkins and Kornienko 2004). This is broadly consistent with our results in Table 6, which suggest that the sex-ratio effect on risk-taking is the largest for low-income son families.

Another framework that may explain why higher sex ratios render families with a son (but not families with a daughter) more risk-taking is the reference-dependent theory (Kőszegi and Rabin 2007). We assume that altruistic parents who care about their child's marriage have reference-dependent preferences. Under such preferences, parental utility not only depends on the absolute level of wealth but also on a reference level. In the spirit of Kőszegi and Rabin (2006, 2007, 2009), we define the reference level as the parental expectation for their children's marriage expenditure, which is formed based on the marriage market conditions they face. When the sex ratio becomes higher, parents with a son would expect to spend more in order to enhance their son's marriage prospects. This implies that they have a larger expected marriage expenditure and a higher reference level. As a result, they might move to the loss domain from the gain domain. The reference-dependent theory predicts the change in risk attitude when moving from the gain domain to the loss domain. Specifically, loss aversion suggests that the marginal utility of additional marriage expenditure is much higher in losses than in gains, leading to excessive risk-taking (Rabin 2000).²³ Intuitively, families who moved into the loss domain caused by the increase in sex ratio try to move back to the gain domain by taking on financial risk.

Unlike families with a son, those with a daughter may have a reference level of marriage expenditure that is independent of sex ratios. When the sex ratio becomes higher, the marriage market is favorable for parents with a daughter, and they therefore may potentially lower their reference level. On the other hand, parents with a daughter may follow parents with a son to spend more on a child's marriage, in the sense that spouses' bargaining position is determined by their relative contribution to the marriage (Zhang and Chan 1999). These two conflicting motives may result in an insignificant net impact on the reference level of marriage expenditure for families with a daughter.

23. Diminishing sensitivity also implies more risk-taking, since these families go to the convex segment of the value function from the concave segment. However, model implications still hold without this condition.

The reference-dependent theory and status-competition theory are related to each other in the sense that social status is a potential candidate for the reference level, and therefore status seeking corresponds to an increase in the reference level (that is, a larger marriage expenditure). Both theories have the same prediction on changes in risk attitude in response to increasing expected marriage expenditure. Yet the reference-dependent theory remains silent on why expected marriage expenditure increases with marriage market competition.

VI. Conclusion

Using nationally representative Chinese household data, this study shows that high sex ratios render families with a son more likely to hold—and hold more—risky assets in their portfolios relative to families with a daughter. Our estimate implies that rising sex ratios explain around 10 percent of the significant growth in China's stock market in recent decades. These results are relevant to several policy issues. For example, our findings suggest that rising sex ratios may account for some part of the significant growth in the stock market value in China and other Asian economies with similar historical, cultural, and societal settings. In addition, in the sense that high sex ratios affect household participation in the financial market, they may have further implications for financial market outcomes (like asset pricing). Moreover, the sex-ratio effect on financial investment decisions varies across households, depending on the marriage market conditions they face, as well as their income level. This may contribute meaningfully to the evolution of wealth inequality, as highlighted by Campbell (2016).

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