



The glorified mothers of sons: Evidence from child sex composition and parental time allocation in rural China



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ABSTRACT

We study the effects of sons versus daughters on parental joint time allocation between the labor market and the household. Using data from the China Health and Nutrition Survey from 1989 to 2006, we apply a fixed-effects model to control for cross-household heterogeneity in son preference. We find that the birth of sons rather than daughters significantly reduces maternal time spent on household chores, which we argue represents an increase in maternal intra-household bargaining power. However, the effects of sex composition of children on paternal time allocation and maternal time on labor-market activities are weak or mixed. Results are robust to a series of sensitive analyses.

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1. Introduction

The sex ratio, measured as the number of males per 100 females, has been rising drastically in China in recent decades. The normal sex ratio is 103–106; China's rose rapidly from 111.9 to 119.9 between 1990 and 2000 (NBS, 2002), and persisted till 2010.¹ China's biased sex ratio has raised global concerns among both academic researchers and policy makers. Extensive literature has documented the socioeconomic consequences of the rising sex ratio, such as for old-age support (Ebenstein and Leung, 2010), household saving rate (Wei and Zhang, 2011), and crime (Edlund et al., 2013). We estimate the effects of child gender on parental time allocation between the household and the labor market in the context of the rising sex ratio in China. Previous literature has studied the effect of an increased sex ratio on household labor supply (Chiappori et al., 2002; Cruces and Galiani, 2007; Lundberg and Rose, 2002), but few have studied the effect on parental joint time allocation between the household and the labor market. In China, the effect of child gender on parental joint time allocation has received little attention, yet is especially important due to the biased sex ratio.

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¹ The most prevalent explanation for the increasing sex ratio is the implementation of the one-child policy, exacerbated by the traditional son preference and the spread of ultrasound machines (Ebenstein, 2011; Li et al., 2011). Almond et al. (2015) propose that China's rural land reform might also explain this phenomenon.

Estimation of the effects of child gender on parental time allocation is subject to potential endogeneity caused by the unobserved time-invariant heterogeneity in son preference. On the one hand, couples with stronger son preference are more likely to exercise gender-selective abortion under the one-child policy and with the access to ultrasound technology. On the other hand, son preference can affect intra-household time allocation. [Cain et al. \(1979\)](#) demonstrate that in patriarchal societies with stronger son preference, women tend to specialize in household chores while men specialize in market work. Thus, the unobservable cross-household heterogeneity in son preference may bias ordinary least squares (OLS) estimates of the effects of child gender on parental time allocation.

We employ a fixed-effects (FE) estimator to address unobservable cross-household heterogeneity in son preference. Based on Chinese Health and Nutrition Survey (CHNS) data from 1989 to 2006, our main FE estimates consistently show that the presence of sons significantly reduces maternal time spent on household chores. Specifically, mothers with a first-born son spend 37.9% less time on household chores—or 6.8 h per week—than mothers with a first-born daughter. Nonetheless, the estimated effect of child gender on paternal time allocation is weak or mixed, as is the effect on maternal time spent on labor-market activities.

Interpretation of our empirical finding is challenging: How do we distinguish between the preference- and constraint-driven effects of child gender? We investigate three theoretical channels proposed by the literature: the bargaining-power effect, the demonstration effect, and the specialization effect. Our results present clear and strong evidence for the bargaining-power effect, which is driven by preference: Mothers of sons spent significantly less time spent on household chores than mothers of daughters. Moreover, we use leisure time and private consumption as alternative measures of bargaining outcomes and present evidence that sons raise maternal bargaining power. Our findings are consistent with [Li and Wu \(2011\)](#), who find that mothers with a first-born son have a greater role in household decision-making than mothers with a first-born daughter. In contrast, we find weak or mixed evidence for specialization and demonstration effects, which are mainly driven by budget constraints.

We also estimate the effect of child gender on grandparents' co-residence. Multigenerational co-residence is common in rural China, and the interaction between co-residing grandparents and grandchildren is frequent ([Zeng and Xie, 2014](#); [Wu and Li, 2014](#)). We present evidence that paternal grandparents are significantly more likely to co-reside after the birth of first-born sons than daughters, possibly to compensate for the reduced maternal time on household chores. However, no similar response is observed from maternal grandparents. This finding echoes the son preference in rural China: Paternal grandparents reward mothers who give birth to sons by co-residing and sharing household chores. This finding also bolsters the bargaining-power explanation.

We further implement a series of sensitivity analyses. First, we show that there exist no prior observable household heterogeneities that predict the gender of the first child. Second, we show that birth spacing between the first and second births does not differ by the first-child gender, and male-biased abortion is unlikely to bias our main results. Third, we show that FE results are robust to the geographical differences in the access to sex-determination technology and the stringency of the one-child policy. Last, we employ an instrumental variable (IV) estimation using the number of paternal brothers as an instrument for child gender composition. The IV estimates are consistent with our main results. Overall, these robustness analyses support our finding that the birth of a son significantly increases the mother's intra-household bargaining power.

Our study contributes to the literature in three ways. First, we distinguish between preference- and constraint-driven effects of child gender on parental household-workplace time allocation. We present strong evidence for the bargaining-power effect, which is preference-driven, and show weak or mixed evidence for the demonstration or specialization effect. Second, to the best of our knowledge, our study is the first to systematically investigate the effect of child gender composition on parental joint time allocation in a developing country. In addition, we address the endogeneity due to cross-household heterogeneity in son preference by applying an FE estimation. Lastly, we propose and test a new mechanism by which grandparents co-reside and share household chores to compensate for the reduced maternal chore time after the birth of sons. This mechanism sheds light on our understanding of the relationship between son preference and family behavior, especially in developing countries where extended families are prevalent. The rest of the paper is organized as follows. Section 2 discusses the conceptual framework. Section 3 describes the data. Section 4 specifies the econometric model. Section 5 presents and interprets the main results. Section 6 conducts robustness checks, and Section 7 concludes. Additional figures, tables, and discussions can be found in the online Appendix.

2. Conceptual framework

In light of the literature, this section discusses three conceptual channels by which child gender affects parental household-workplace time allocation.

2.1. Bargaining-power effect

This channel is associated with son preference. The birth of sons brings higher utility to fathers who prefer sons to daughters. In a standard divorce-threat bargaining model ([McElroy and Horney, 1981](#)), the birth of sons, relative to daughters,

Table 1
Effects of sons versus daughters on parental time allocation.

	Paternal time		Maternal time	
	Employment	Household chores	Employment	Household chores
Bargaining-power effect	↑	↑	↓	↓
Demonstration effect	↓	↑	↑	↓
Specialization effect	↑	↓	↓	↑
Total effects (sons vs. daughters)	↑↓	↑↓	↑↓	↑↓

Notes: Up/down arrows represent the positive/negative impacts of each effect on parental time allocation between the labor market and household.

changes the threat point of divorce, especially in societies with prevalent son preference. In this case, the birth of sons raises mothers' welfare and increases their intra-household bargaining power.²

In this paper, we use parental time spent on household chores as a proxy for bargaining outcome.³ Furthermore, we investigate the effect of child gender on a full set of measures of parental household-workplace time allocation. Through the household bargaining channel, maternal time on household chores is expected to decrease after the birth of sons relative to the birth of daughters. We expect that maternal time on labor-market activities and paternal time allocation will shift correspondingly.

2.2. Demonstration effect

The demonstration effect works through the comparative advantage of fathers/mothers in raising sons/daughters. The production function of raising sons differs from that of daughters, and the optimal paternal and maternal time investment in sons and daughters also differs. If fathers are more productive in raising the character of their sons, they are more likely to increase the time shared with sons. [Lundberg and Rose \(2002\)](#) claim that increased paternal time with sons represents a demonstration effect, which could be attributed to the more important role of fathers in modeling the traditional social role for sons than for daughters. With this effect, we expect that after the birth of sons, fathers shift their time from the labor market to the household to be with their sons. We anticipate that maternal time allocation will adjust accordingly, such as shifting from the household into the labor market.

2.3. Specialization effect

The third channel is the specialization effect, by which family stability enhances intra-household labor specialization ([Becker, 1985](#)). Since childcare and household work are effort-intensive and normally the responsibility of women, an increase in fertility induces women to reduce labor supply and increase time on home production. In contrast, men are expected to increase their labor in the marketplace. If the birth of sons increases marital stability more than daughters in societies with strong son preference, parents are more likely to follow the specialization rule.⁴

With the specialization effect, fathers of sons are expected to devote more time to labor-market work and mothers of sons to household chores. [Rose and Bias \(2000\)](#) finds consistent reduction in female working hours following the birth of sons compared to that of daughters in both the poorest and less poor households in rural India. [Lundberg and Rose \(2002\)](#) find that men's labor supply and wage rates increase more in response to the birth of sons than to the birth of daughters.

In summary, parents adjust their time allocation differently after the birth of sons and after the birth of daughters. [Table 1](#) displays how the three different channels predict the change in parental time allocation following the birth of sons compared to daughters. The overall effect is theoretically indeterminate and thus requires empirical investigation.

3. China Health and Nutrition Survey

We obtain longitudinal data for rural households from the China Health and Nutrition Survey (CHNS). Data are collected by the Carolina Population Center of the University of North Carolina at Chapel Hill, the National Institute of Nutrition and

² Conceptually, two kinds of bargaining-power effects co-exist. One is a direct effect: The presence of sons increases maternal bargaining power and directly decreases maternal time spent on household chores. The other is an indirect effect: The presence of sons increases maternal bargaining power, which increases marital stability and enhances labor specialization. This indirect effect increases maternal time spent on household chores. Since the direct effect is negative and the indirect effect is positive, reduced-form estimates could be interpreted as a lower bound of the pure direct effect. The specialization effect is discussed in Section 2.3.

³ Researchers have used various indicators to measure bargaining outcomes. For example, [Zhang and Chan \(1999\)](#) use "husband's help with chores" and find that the dowry enhances the bride's bargaining power in the allocation of intra-household output.

⁴ The theoretical mechanism is as follows. Both labor-market and household activities are regarded as learning-by-doing processes. When people spend more time on labor-market work or household chores, their productivity increases. Therefore, husbands and wives have strong incentives to specialize in the market or household. However, the skill of doing household chores is regarded as a special human capital. If women are divorced, their skill in doing household chores in this specific household will be depreciated if they remarry. Therefore, [Becker \(1985\)](#) predicts that when families are more stable, husbands are more likely to specialize in the labor market, whereas wives specialize in household activities.

Food Safety, and the Chinese Center for Disease Control and Prevention. The survey tracks individuals in 1989, 1991, 1993, 1997, 2000, 2004, and 2006 in nine provinces, as shown in Fig. A1 in Appendix A.

We choose CHNS data for three reasons. First, the data record the time usage of each adult household member for both labor-market activities and household chores. Second, the survey collects data on birth history and spousal characteristics for each married woman, independent of co-residence with children. Thus, the data overcome the conventional co-residence bias in studies using household surveys. Finally, high follow-up rates across seven waves facilitate our FE estimation.

We impose the following sample restrictions. First, we restrict the sample to rural households with nonmissing data on time usage for household chores and labor-market activities. Second, we exclude households that appear in only one survey wave.⁵ Third, we exclude households with mothers exceeding 60 years old. Fourth, we exclude households with twins. Compared to single births, twinning affects parents' time allocation not only through child gender effect but also through the unexpected increase in the number of children (Rosenzweig and Wolpin, 2000). After the sample restrictions, our main sample contains 3684 observations from 1346 households in rural China with complete information on parental time allocation, child gender, and a set of household characteristics.

The main dependent variables are parental time spent on household chores and labor-market activities every week. Time spent on household chores is measured as hours spent purchasing, preparing, and cooking food; washing and ironing clothes; and taking care of children under six years old. Time spent on labor-market activities is measured as hours spent on first and second jobs with regular wages, self-employed home gardening, and collective and household farming.⁶ The main independent variables are indicators of child gender, which are defined in the next section.

Table 2 Panels A and B present summary statistics for parental time allocation and demographic characteristics. On average, mothers spend 25 h per week on labor-market activities, and 25 h per week on household chores. Fathers spend 44 h per week on labor-market activities, and only 5 h on household chores. The average age for fathers and mothers is 42 and 40, respectively. Fathers and mothers obtain 7.8 years and 6.1 years of schooling, respectively. More than 84% of parents are Han, which is the ethnic majority.

Table 2 Panels C and D present summary statistics for fertility and household characteristics. In our sample, 49.6% of households have a first-born son and 48.3% a first-born daughter.⁷ This shows that the sex ratio is almost balanced at first birth parity, which conforms to the finding of Ebenstein (2010) based on the 1982–2000 China Census (Appendix Fig. 2A). In our sample, if the first child is a son, approximately 19% of households stop fertility. If the first child is a daughter, 10% of households stop fertility. Table 3 tabulates the distribution of households by fertility and child gender. The table shows that 30% of the households have one child, 35% have two, and 28% have more than two. Appendix Table B4 reports summary statistics for variables that appear only in Appendix tables (Panels A–C) and community characteristics (Panel D). Moreover, Appendix Table B5 tabulates observable household characteristics of families with first-born sons and those with first-born daughters in all available waves. Except for fertility outcomes and grandparents' co-residence status, differences between the two subsamples are economically small and statistically insignificant.

4. Econometric specification

To address unobserved time-invariant heterogeneity in son preference, we use a fixed-effect estimator to examine the effects of first-born child gender on parental time allocation:

$$\ln y_{it} = \alpha_0 + \alpha_1 * fson_{it} + \alpha_2 * fdaughter_{it} + \delta_i + \tau_t + \varepsilon_{it}, \quad (1)$$

where y_{it} represents parental time allocation for household i at survey wave t , which includes four variables: maternal time on household chores, maternal time on labor-market activities, paternal time on household chores, and paternal time on labor-market activities. We measure time usage as hours per week in logarithmic terms. $fson_{it}$ is a binary variable equal to one if household i has a first-born son at wave t and zero otherwise. $fdaughter_{it}$ is a binary variable equal to one if household i has a first-born daughter at wave t and zero otherwise. δ_i and τ_t are household fixed effects and survey-wave fixed effects, respectively. ε is the error term. Standard errors are clustered at the household level in all regressions.

We add a rich set of time-variant controls to Eq. (1). Some omitted time-variant variables may correlate with child gender and parental time allocation simultaneously, leading to a spurious correlation between child gender and parental time allocation (Jayachandran and Kuziemko, 2011). We comprehensively control for individual-, household- and community-level variables to mitigate the concern of omitted variables. Specifically, individual-level time-variant controls include parental age, years of schooling, and migration status. Household-level time-variant controls include the number of children, the age of the first child, whether the household has an infant (age < 6), total household income, and dummies for whether the household has time-saving home appliances. Community-level time-variant controls include whether the community

⁵ Appendix Table B1 compares the main sample with the excluded sample. Except for being surveyed in later waves, being younger, and having fewer children, the excluded sample is similar to the main sample in all observable household characteristics. Moreover, Appendix Table B2 presents summary statistics for the main sample with data from the first available survey wave for each household.

⁶ Based on the CHNS questionnaire, labor-market activities also include work raising livestock/poultry, fishing, small household businesses, and other. Household chores also include cleaning the house. However, data from these categories contain a considerable amount of missing values. In Appendix Table B3, we check the robustness of our main results by including all available time categories and recoding missing values as zero. Results remain robust.

⁷ The two numbers do not add to unity, as some households have not given birth in certain waves.

Table 2
Summary statistics for the main sample.

Variables	Mean	S.D.
Panel A: mother		
Time on household chores (hours/week) ^a	25.392	18.195
Time on the labor market (hours/week) ^b	38.014	29.000
Age	40.406	8.943
Schooling years	6.097	4.183
Ethnicity (Han = 1, minority = 0)	0.841	0.366
Migration status (migrated = 1)	0.001	0.023
Panel B: father		
Time on household chores (hours/week)	5.061	10.434
Time on the labor market (hours/week)	43.689	27.278
Age	42.405	9.678
Schooling years	7.814	3.679
Ethnicity (Han = 1, minority = 0)	0.846	0.361
Migration status (migrated = 1)	0.014	0.118
Panel C: fertility information		
Total number of children	2.002	1.164
First born is son	0.496	0.500
First born is daughter	0.483	0.497
One child, a son	0.193	0.395
One child, a daughter	0.106	0.308
Two children, 1st son 2nd daughter	0.092	0.289
Two children, 1st son 2nd son	0.080	0.271
Two children, 1st daughter 2nd daughter	0.050	0.217
Two children, 1st daughter 2nd son	0.133	0.339
Having infant (age < 6)	0.224	0.417
Age of oldest child	15.526	8.227
Age of youngest child	11.901	7.200
Age of in-sample births ^c	6.070	4.675
Panel D: household information		
Whether have a washing machine	0.467	0.499
Whether have a refrigerator	0.240	0.427
Whether have an electronic cooker	0.513	0.500
Gross household income	17.626	19.608
Paternal grandparents' co-residence	0.127	0.333
Maternal grandparents' co-residence	0.017	0.131
Total number of observations	3684	

^a Household chores include purchasing food, preparing and cooking food, washing and ironing clothes, and taking care of children under 6 years old.

^b Labor market work includes regular employment at first and second jobs, self-employment in home gardening, and collective and household farming.

^c In-sample births are births that occurred during the sample period from 1989 to 2006.

Table 3
Distribution of households by fertility and child gender.

	Households		Households		Households
No children	224 (6.08)	No sons	880 (23.89)	No daughters	1345 (36.51)
One child	1104 (29.97)	One son	1955 (53.07)	One daughter	1559 (42.32)
Two children	1303 (35.37)	Two sons	691 (18.76)	Two daughters	572 (15.52)
Three children or more	1053 (28.58)	Three sons or more	158 (4.29)	Three daughters or more	208 (5.65)
Total	3684 (100)	Total	3684 (100)	Total	3684 (100)

Notes: Percentages are in parentheses.

allows a second child, whether it provides subsidies for one-child families, whether it has ultrasound machines, and whether it has any public or private childcare facilities.⁸

We now interpret the key coefficients of interest, α_1 and α_2 . By controlling for household fixed effects, we essentially classify families into 3 types. Type 1 families have given birth to their first-born son between waves $t - 1$ and t . Type 2

⁸ One concern is that some of these time-variant controls may be endogenous. In Appendix Tables B6 to B9, we present estimation results with and without these controls. Results are robust. See our detailed discussion in Section 5.1.

families have given birth to their first-born daughter between waves $t - 1$ and t . Type 3 families are our reference group: those that have no new birth at first parity between waves $t - 1$ and t .⁹ Therefore, α_1 measures the change in parental time allocation for Type 1 families from wave $t - 1$ to t , relative to the change for Type 3 families. Similarly, α_2 measures the change in parental time allocation for Type 2 families from wave $t - 1$ to t , relative to the change for Type 3 families. We are interested in $\alpha_1 - \alpha_2$, which measures the difference in change in parental time allocation from wave $t - 1$ to t between Type 1 families and Type 2 families. Furthermore, we estimate the effects of child gender on parental time allocation not only at the first birth parity, but also at the second birth parity conditioning on first-child gender¹⁰:

$$\ln y_{it} = \beta_0 + \beta_1 * s1_{it} + \beta_2 * d1_{it} + \beta_3 * s1s2_{it} + \beta_4 * s1d2_{it} + \beta_5 * d1s2_{it} + \beta_6 * d1d2_{it} + \beta_D * D_{it} + \delta_i + \tau_t + \varepsilon_{it}, \quad (2)$$

where s denotes son and d denotes daughter, and the number following s or d represents the birth parity. For instance, $s1d2_{it}$ is a binary variable equal to one if household i has two children in wave t : a first-born son and a second-born daughter. D is the dummy for having three or more children. The definitions of y_{it} , δ_i and τ_t are the same as in Eq. (1). We also include the same set of time-variant controls. Standard errors are clustered at the household level.

The interpretation of regression coefficients in Eq. (2) is similar to that in Eq. (1). By controlling for the household fixed effects, we classify families into 8 types according to the change of child gender composition at first and second birth parities. Types 1 and 2 are families that give birth to the first-born son and first-born daughter, respectively, between waves $t - 1$ and t . Types 3 and 4 families both have a first-born son before wave $t - 1$, but give birth to a second-born son and a second-born daughter, respectively, between waves $t - 1$ and t . Similarly, Types 5 and 6 families both have a first-born daughter before wave $t - 1$, but give birth to a second-born son and a second-born daughter, respectively, between waves $t - 1$ and t . Type 7 are families that give birth to a third or higher-parity child between $t - 1$ and t . Type 8 are families that have no new birth between $t - 1$ and t , which is the reference group.

For Type x ($x = 1, 2, \dots, 6$) families, β_x measures the change in parental time allocation from wave $t - 1$ to t for Type x families, relative to the change for the reference group. Therefore, $\beta_1 - \beta_2$ measures the difference in change in parental time allocation from wave $t - 1$ to t between Type 1 and Type 2 families. Similarly, $\beta_3 - \beta_4$ measures the difference between Type 3 and Type 4 families, and $\beta_5 - \beta_6$ measures the difference between Type 5 and Type 6 families.

5. Effects of child sex composition on parental time allocation

5.1. Main results

We present the estimated effect of child sex composition on parental time allocation in rural China in Table 4. Panel A presents estimates of the effect of first-born child gender on parental time allocation, $\alpha_1 - \alpha_2$, from Eq. (1). Panel B presents estimates of $\beta_1 - \beta_2$, $\beta_3 - \beta_4$, and $\beta_5 - \beta_6$ from Eq. (2). Coefficient interpretations are discussed in Section 4, and detailed point estimates are reported in Appendix B.¹¹

Table 4 Panel A shows that the first-born son significantly reduces maternal chore time, but has a weak and mixed effect on maternal time on labor-market activities and paternal time allocation. A first-born son reduces maternal chore time by 37.9%, or 6.8 h per week (Column (1)). In the labor market, mothers with a first-born son work fewer hours (Column (2)), though the estimate is not statistically significant. Paternal time allocation between labor-market activities and household chores does not respond significantly to the first-born child gender (Columns (3) and (4)).¹²

We adjust the p -values under a multiple-hypotheses setting, as we have four dependent variables of parental time usage (Anderson, 2008). The adjusted p -values still reject the null hypothesis that first-child gender has no effect on maternal time on chores at the 10% level of significance, but do not reject the other null hypotheses at any conventional level of significance.

Table 4 Panel B presents the effect of child gender at the first and second birth parity, conditional on the first-child gender. The estimated effect of child gender at first birth parity is consistent with that in Panel A: Conditional on having one child, a first-born son reduces maternal chore time by 33.6% (6 h per week) relative to a first-born daughter. However, the gender effect dissipates at second birth parity, as estimates of $\beta_3 - \beta_4$ and $\beta_5 - \beta_6$ are small and statistically insignificant.

One possible explanation for the dissipating gender effects at higher birth parities is the primogeniture culture in rural China. Because the oldest son is expected to inherit the family estate in preference to daughters and younger sons, mothers

⁹ The reference group includes two kinds of families: families that have no new birth at any parity between waves $t - 1$ and t , and families that have new birth at the second or higher birth parities between waves $t - 1$ and t .

¹⁰ We thank an anonymous referee for proposing this specification.

¹¹ The full estimation results of Eq. (1) are reported in Table B6 (maternal time allocation) and Table B7 (paternal time allocation). The full estimation results of Eq. (2) are reported in Table B8 (maternal time allocation) and Table B9 (paternal time allocation). For each dependent variable in Eq. (1), we estimate three specifications with different sets of time-variant controls. The first specification includes only individual-level controls; the second additionally includes household-level controls; and the third further includes community-level controls. In Table B6 and Table B7, we show, using Durbin–Wu–Hausman tests, that estimates of $\alpha_1 - \alpha_2$ are the same across the three specifications.

¹² The lack of statistical significance in estimates of $\alpha_1 - \alpha_2$ in Columns (2) and (3) is partially due to the data distribution of dependent variables. Data on maternal labor-market time and paternal chore time have many zero entries as well as very large values. In contrast, the distribution of maternal chore time is more centered.

Table 4
FE estimates of the effects of sex composition of children on parental time allocation in rural China.

Dependent variable	(1)	(2)	(3)	(4)
	Mother		Father	
	Household chores (hours/week)	Labor market (hours/week)	Household chores (hours/week)	Labor market (hours/week)
Panel A: first-child gender as independent variable				
1st son – 1st daughter ($\alpha_1 - \alpha_2$)	–0.379** (0.192)	–0.232 (0.426)	0.056 (0.212)	–0.185 (0.331)
Observations	3684	3684	3684	3684
R-squared	0.183	0.219	0.073	0.160
Number of Households	1346	1346	1346	1346
Additional controls	Yes	Yes	Yes	Yes
Panel B: detailed children gender composition as independent variable				
1st son – 1st daughter ($\beta_1 - \beta_2$)	–0.336** (0.165)	0.010 (0.347)	–0.002 (0.197)	–0.172 (0.280)
Cond. on only 1 child			0.057 (0.221)	0.124 (0.342)
2nd son – 2nd daughter ($\beta_3 - \beta_4$)	–0.074 (0.149)	0.141 (0.330)	0.030 (0.221)	0.553 (0.342)
Cond. on 1st is son				
2nd son – 2nd daughter ($\beta_5 - \beta_6$)	0.079 (0.185)	0.463 (0.313)	0.030 (0.197)	0.553 (0.342)
Cond. on 1st is daughter				
Observations	3684	3684	3684	3684
R-squared	0.189	0.221	0.075	0.163
Number of households	1346	1346	1346	1346
Additional controls	Yes	Yes	Yes	Yes

Notes: This table estimates the effect of child gender on maternal and paternal time allocation between household chores and labor-market activities. The sample includes households that appeared in at least two survey waves, and have non-missing data on parental time allocation. Panel A reports estimates from FE specification in Eq. (1) and Panel B reports those in Eq. (2). Household chores include buying and cooking food, washing and ironing clothes, and taking care of children under 6 years old. Labor market work includes regular job, farming, and gardening. Dependent variables are measured in hours per week, and take the logarithmic terms. Additional controls include parental age, years of schooling, and migration status; number of children, age of the first child, and whether there is an infant (age < 6); total household income; whether the household has time-saving home appliances; whether a second child is allowed if the first is a daughter; whether subsidies are available for one-child families; whether the local community has ultrasound machines; whether the local community has any childcare facilities; and household and survey-wave fixed effects. Standard errors are clustered at the household level.

*Significant at the 10% level.

** Significant at the 5% level.

***Significant at the 1% level.

Detailed regression results for Columns (1) and (2) of Panel A are reported in Table B6, Columns (3) and (4) of Panel A in Table B7, Columns (1) and (2) of Panel B in Table B8, and Columns (3) and (4) of Panel B in Table B9.

are likely to be most “glorified” if giving birth to a first-born son relative to sons at higher parities. Nevertheless, we treat this explanation with caution, as the sample does not contain enough “in-sample” births at higher parities to precisely pin down the gender effects at higher parities.¹³ Moreover, the gender of a higher-parity child is more likely to be endogenous than the first child (Ebenstein, 2010). Overall, we acknowledge this limitation of our data and focus on the gender effect at first birth parity.¹⁴

5.2. Interpreting the main results

5.2.1. The change in intra-household bargaining power

In this section, we discuss the channels through which child gender affects parental time allocation in light of the three conceptual frameworks—bargaining power, demonstration, and specialization—discussed in Section 2.

Empirical evidence supports our prediction of the household bargaining-power effect, in which mothers of first-born sons are glorified and obtain higher intra-household bargaining power. We consider three outcomes of rising bargaining power: lower labor inputs, more leisure, and more private consumption. We show that the birth of a first-born son, relative to a first-born daughter, reduces the mother’s labor inputs in both household chores and labor-market activities (Table 4); raises the mother’s leisure time (Appendix Table B11); and increases the mother’s private consumption (Appendix Table B12). Overall, we find empirical evidence that the birth of a first-born son raises the mother’s intra-household bargaining power, thereby reducing her labor inputs and increasing her leisure time and private consumption.

Evidence for the demonstration effect is weak. The demonstration effect predicts that fathers of first-born sons spend more time in the household to teach their sons the social role of males (Lundberg and Rose, 2002). However, Table 4 shows that paternal time for labor-market activities decreases after the birth of a first-born son, but hardly increases for household

¹³ In-sample births are births that occurred during the sample period, from 1989 to 2006.

¹⁴ We drop households with two or more “in-sample” births and re-estimate Eq. (1) in Appendix Table B10. Results are robust.

chores, notwithstanding that neither estimate is statistically significant. Moreover, we also separately look at parental time spent taking care of children under six years old. We test the hypothesis that fathers spend more time with their first-born sons than first-born daughters, and mothers spend more time with their first-born daughters than first-born sons. Results in Appendix Table B13 show that both parents spend more time with their first-born sons than first-born daughters, which does not support the demonstration effect as a dominant explanation.

Empirical evidence does not support the specialization effect either. Table 4 shows that following the birth of first-born sons, mothers spend significantly less time on household chores and fathers spend less time on labor-market work. Both pieces of evidence contradict the prediction from the specialization effect.

Although all three effects may be in play, the empirical evidence for the bargaining-power effect is the clearest and most robust. The demonstration and specialization effects are weak or dominated, and are estimated imprecisely in the reduced-form framework. Linking to the discussion of preference- or constraint-driven effects of child gender, our evidence suggests that the child-gender effect on parental time allocation in rural China is more likely to be driven by preference rather than budget constraints.

5.2.2. The sex-specific fertility-stopping rule

Besides the three conceptual channels discussed above, we consider whether the sex-specific fertility-stopping rule could drive our main results. It is well documented that in rural China, mothers with sons are more likely to stop their fertility than mothers of daughters (Ebenstein and Leung, 2010; Ebenstein, 2011). Since smaller family size leads to lighter household chores, the negative effect of a first-born son on maternal chore time relative to a first-born daughter may simply reflect the effect of family size.

The sex-specific fertility-stopping rule is unlikely to drive the FE results in Table 4. First, to lessen the potential influence of the family-size effect, we have controlled for the total number of children and included a dummy that indicates the presence of infants (age < 6) in Eq. (1). Therefore, estimates of $\alpha_1 - \alpha_2$ capture the gender effect of first-born child on parental time allocation, netting out the family-size effect.

Second, with household fixed effects, we are estimating the effect of period-to-period changes in child sex composition on parental time allocation (see Section 4). In Eq. (1), α_1 measures the change in time allocation from wave $t - 1$ to t for families that gave birth to first-born son between the two waves (Type 1 families), relative to families with no first-parity birth between the two waves (reference families). Similarly, α_2 measures the change in time use from wave $t - 1$ to t for families that gave birth to first-born daughter between the two waves (Type 2 families), relative to the reference families. Therefore, $\alpha_1 - \alpha_2$ measures the difference in change in parental time allocation from wave $t - 1$ to t between Type 1 families and Type 2 families. Appendix Table B14 summarizes average birth spacing between the first and second births in our sample, and reports an average birth spacing of 3.1 years, regardless of first-child gender. This implies that, from wave $t - 1$ to t , total fertility is unlikely to diverge between Type 1 and Type 2 families. Thus, the difference in changes in parental time allocation captured by $\alpha_1 - \alpha_2$ is unlikely to be driven by the fertility effect.

Finally, we present additional regression evidence. In Appendix Table B15 Column (1), we regress the total number of children on first-child gender dummies, using the same specification as Eq. (1). In Column (2), we regress the total number of children on lagged first-child gender dummies. Regression results confirm that fertility does not differ between Type 1 and Type 2 families. In the long run, first-born-daughter families on average have more children than first-born-son families in rural China. However, in the short run, fertility is unlikely to diverge between these two types of families in our sample.

5.3. Who compensates for the reduced maternal chore time?

As mothers of first-born sons reduce their chore time relative to mothers of first-born daughters, who compensates for the lost chore time? Considering the high rate of multigenerational co-residence (Zeng and Xie, 2014) and the low rate of hired domestic help in rural China, the grandparents' co-residence offers an explanation.

We show that paternal grandparents are more likely to co-reside after the birth of grandsons than granddaughters. In Table 5 Panel A, we regress the dummy for grandparents' co-residence on first-child gender dummies using the same specification as Eq. (1). The birth of a first-born grandson, relative to a first-born granddaughter, raises the probability of paternal grandparents' co-residence by 22.9 percentage points. In Table 5 Panel B, we regress the dummy for grandparents' co-residence on first and second child gender dummies using the same specification as Eq. (2). At first birth parity, having a son rather than a daughter raises the probability of paternal grandparents' co-residence by 14.8 percentage points. At second birth parity, having a son rather than a daughter raises the co-residence probability by 12.1 percentage points if the first-born is a son, and 2.7 percentage points if the first-born is a daughter.¹⁵

However, maternal grandparents show no differential response of co-residence after the birth of grandsons relative to granddaughters. This is not surprising, considering that grandsons carry the family name of the father's lineage, not mother's. Overall, Table 5 shows that paternal grandparents are more likely to co-reside after the birth of grandsons than granddaughters, but maternal grandparents' probability of co-residence is not responsive to grandchild's gender.

¹⁵ Regression details are reported in Appendix Table B16 and Table B17.

Table 5

FE estimates of the effects of sex composition of children on the probability of grandparents' co-residence in rural China.

Dependent variable	(1) Co-residence of paternal parents	(2) Co-residence of maternal parents
Panel A: first-child gender as independent variable		
1st son – 1st daughter ($\alpha_1 - \alpha_2$)	0.229* (0.129)	0.004 (0.006)
Observations	3684	3684
R-squared	0.887	0.945
Number of households	1346	1346
Additional controls	Yes	Yes
Panel B: detailed children gender composition as independent variable		
1st son – 1st daughter ($\beta_1 - \beta_2$)	0.148* (0.085)	-0.014 (0.023)
Cond. on only 1 child		
2nd son – 2nd daughter ($\beta_3 - \beta_4$)	0.121** (0.059)	0.006 (0.052)
Cond. on 1st is son		
2nd son – 2nd daughter ($\beta_5 - \beta_6$)	0.027 (0.060)	0.051 (0.050)
Cond. on 1st is daughter		
Observations	3684	3684
R-squared	0.887	0.945
Number of households	1346	1346
Additional controls	Yes	Yes

Notes: Dependent variable is a dummy variable indicating whether grandparents from paternal/maternal side are co-residing in the household. Additional controls include parental age, years of schooling, and migration status; number of children, age of the first child, and whether there is an infant (age < 6); total household income; whether the household has time-saving home appliances; whether a second child is allowed if the first is a daughter; whether subsidies are available for one-child families; whether the local community has ultrasound machines; whether the local community has any childcare facilities; and household and survey-wave fixed effects. Standard errors are clustered at the household level.

* Significant at the 10% level.

** Significant at the 5% level.

*** Significant at the 1% level.

Regression details of Panel A are reported in Table B16. Regression details of Panel B are reported in Table B17.

The patriarchal culture in rural China and the bargaining-power effect explain the empirical finding in Table 5. Since grandsons carry the father's family name, the mother with sons is glorified by the paternal family and achieves higher intra-household bargaining power. As a result, paternal grandparents reward the mother by co-residing in the household and helping with the household chores.¹⁶ In contrast, since grandchildren do not carry the mother's family name, maternal grandparents have lower incentive to co-reside.

We further consider whether results in Table 5 could be explained by an alternative story. Presumably, son preference may persist across generations in a family. Fathers from more son-preferring families are more likely to have sons; meanwhile, they are also more likely to live with their parents because they have fewer sisters. Therefore, the grandchild being a son is positively correlated with grandparents' co-residence. We argue that this story does not hold in our sample. In the 1960s (average birth year of fathers in our sample), the sex ratio at birth was balanced in China.¹⁷ Moreover, Appendix Table B18 shows that the number of paternal sisters does not correlate with the number of paternal brothers.

Overall, Table 5 shows that paternal grandparents are more likely to co-reside after the birth of grandsons than granddaughters. The purpose of co-residence is most likely to care for the baby and help with household chores, which suggests a substitution between maternal and grandparental time on household chores in rural China—a finding not yet documented in the literature. We believe this mechanism is important for understanding family behavior, especially in developing economies in which extended families are prevalent.

6. Robustness

This section carries out four strands of sensitivity analyses. First, we examine heterogeneity between families with sons and families with daughters. Second, we check the correlation between child gender and birth spacing, and the correlation between child gender and abortion. Third, we investigate cross-region variations in the availability of sex-determination technology and the implementation of the one-child policy. Last, we conduct an IV estimation. Overall, we show that the estimated effects of child gender composition on parental time allocation are not contaminated by confounding factors.

¹⁶ The literature has documented that co-residing grandparents have strong interactions with children (Wu and Li, 2014), and have a positive effect on children's education (Zeng and Xie, 2014). This suggests that paternal grandparents' motivation to better nurture and educate their grandsons may be another reason for their higher co-residence rate.

¹⁷ See Table 1 in World Bank (2006).

6.1. Cross-household heterogeneity

FE identification relies on a common trend assumption: In the absence of treatment—the birth of a first-born son rather than a first-born daughter—treatment and control households would experience similar trends in parental time allocation. Although it is infeasible to check the common trend directly, we check pre-trends in household time allocation and other observable household characteristics. Appendix Table B19 presents summary statistics of parental time allocation and household characteristics by first-child gender for households that gave birth to the first child during the sample period. All variables are measured one period prior to the first birth. We find no significant differences in household characteristics or parental time allocation between the two groups of households, implying no observably different pre-trends that predict the gender of the first-born child.¹⁸

6.2. Birth spacing and abortion

We check whether birth spacing, defined as years between the first and second births, differs between first-born-daughter families and first-born-son families.¹⁹ Appendix Table B14 summarizes birth spacing by child gender. Panel A shows that the birth spacing of first-born-daughter families is 3.10 years, while birth spacing of first-born-son families is 3.12 years (difference 0.02, p -value 0.88). Panel B compares the birth spacing of four groups of families, defined by different combinations of the first- and second-child gender. Cross-group differences in birth spacing are small. For instance, birth spacing between first son and second daughter differs from that between first daughter and second son by only 0.05 years (p -value 0.72). OLS estimation results in Appendix Table B21 confirm that cross-group differences in birth spacing are small and statistically insignificant.

We also examine whether male-biased selective abortion may be correlated with child sex composition. We define the abortion dummy as whether the woman has ever had an abortion, and regress it on three measures of child sex composition: whether the first-born child is a son, whether the couple has at least one son, and the fraction of sons. Appendix Table B22 shows that the OLS coefficients of measures of child sex composition are all small and statistically insignificant, and R^2 are close to zero. However, we exercise caution in interpreting this result due to the extremely low response rate (only 179 non-missing answers).

6.3. Sex-determination technology and the one-child policy

The literature attributes the rise of sex ratio in China to the sex-determination technology and the one-child policy (Ebenstein, 2010; Li et al., 2011). In this section, we test whether these two factors drive our main results.

We show that the sex-determination technology—namely, ultrasound machines—is unlikely to drive our main results. In Eqs. (1) and (2), we have controlled for the binary variable that indicates whether the community is equipped with ultrasound machines. Moreover, we rerun FE regressions separately in the ultrasound-available and unavailable subsamples. Appendix Table B23 shows that estimates of $\alpha_1 - \alpha_2$ are negative in both subsamples and are quantitatively similar (p -value 0.740, Durbin–Wu–Hausman test).²⁰ We conduct a similar analysis to show that the one-child policy is unlikely to drive our main results. In Eqs. (1 and 2), we have controlled for the binary variable that indicates whether the community allows a second child. Moreover, we rerun FE regressions separately on two subsamples: In one subsample, a second child is allowed; in the other, a second child is not allowed. Appendix Table B24 reports that estimates of $\alpha_1 - \alpha_2$ are negative and statistically significant in both subsamples, and quantitatively similar (p -value 0.565, Durbin–Wu–Hausman test).²¹ Furthermore, considering that the stringency of the one-child policy is correlated with the share of ethnic minorities in the population (Li and Zhang, 2007), we rerun FE regressions separately on the ethnic-majority subsample and ethnic-minority subsample. Appendix Table B25 shows that estimates of $\alpha_1 - \alpha_2$ are similar across subsamples.

6.4. Instrumental variable estimates

We conduct an IV estimation as an additional robustness check. We focus on two measures of child sex composition: the number of sons and the fraction of sons. Details of the IV specification are discussed in Appendix C1. We include the same set

¹⁸ Similarly, in Appendix Table B20 we summarize observable household characteristics by new-birth gender for households that gave birth to a new child during the sample period. All variables are measured one period prior to the new birth. For each observable household characteristic, we present the mean by new-birth gender, the difference of the means, and the p -value of t -test on the difference. Additionally, we regress the new-birth gender on observable household characteristics and report the joint F -statistic. Both the t -tests and the F -test fail to reject the null that one-period-prior characteristics do not predict the gender of the new child.

¹⁹ The interpretation of the difference in birth spacing between these two types of families is ambiguous (Almond et al., 2013). On the one hand, birth spacing of first-born-daughter families is shorter because these families are more eager to have a second child. On the other hand, if first-born-daughter families were to seek gender-selective abortion at second birth, their birth spacing would be longer.

²⁰ We conduct Durbin–Wu–Hausman tests on all four types of parental time allocation. Detailed p -values are reported at the bottom of Appendix Table B23. All tests fail to reject the null that estimates of $\alpha_1 - \alpha_2$ are the same across the two subsamples.

²¹ Detailed p -values of the Durbin–Wu–Hausman tests are reported at the bottom of Appendix Table B25. All tests fail to reject the null that estimates of $\alpha_1 - \alpha_2$ are the same across the two subsamples.

of controls as in Eq. (1), and cluster standard errors at the household level. The IV estimation sample excludes households with no children, and includes those with observations in only one survey wave. Appendix Table B26 presents summary statistics for the IV sample.

We use the number of paternal brothers and its interaction with the household's ethnic minority dummy as instruments. The number of paternal brothers is correlated with child sex composition for ethnic minorities, because the tribal culture fosters strong interhousehold ties and strong substitution between a man's own son and his brother's sons to carry the family name. As a result, minority males with more brothers face less pressure to pursue having a son (Chu, 2001). On the other hand, intra-household time allocation is not directly correlated with the number of paternal brothers, because male adults establish separate households and separate their finances from their brothers' soon after marriage (Parish and Whyte, 1980). Therefore, the number of paternal brothers is correlated with child sex composition, but is arguably exogenous to daily household decisions on time allocation.

Appendix Table B28 reports IV results, which are consistent with the FE results.²² Panel A shows the effect of number of sons on parental time allocation: Conditional on having the same number of children, mothers with one more son reduce their time spent on household chores by 23%. Similar to FE results in Table 4, effects of child gender on maternal time on labor-market activities and paternal time allocation are mixed and statistically insignificant. Panel B presents the results of regressing the parental time allocation on the fraction of sons, and shows a consistent pattern.

We implement the overidentification test to examine the validity of instruments.²³ In all four specifications, *p*-values of Hansen *J* test are larger than 0.3 and fail to reject the null hypothesis that the overidentification restriction is valid. Moreover, we report the Kleibergen–Paap *F*-statistics for weak identification and Kleibergen–Paap LM statistics for underidentification (Stock and Yogo, 2005; Kleibergen and Paap, 2006). No evidence of weak identification or underidentification is found.

We conduct a series of sensitivity checks by varying the IV estimation sample: restricting to the minority subsample (Appendix Tables B29 and B30); collapsing multiple observations for each household to the mean (Appendix Table B31); taking one random observation for each household (Appendix Table B32); and taking the first observation for each household (Appendix Table B33). Results are robust. Discussions of these sensitivity analyses are presented in Appendix C2.

7. Conclusion

This paper empirically examines the effect of sons versus daughters on parental joint time allocation between the household and the labor market. We employ a fixed-effects estimation method and find that the presence of sons exerts a negative and statistically significant effect on maternal time on household chores. Specifically, mothers with a first-born son, rather than a first-born daughter, reduce their chore time by 37.9% (6.8 h per week). Estimated response of maternal time on labor-market activities and paternal time allocation are mixed and statistically insignificant.

Our empirical findings are consistent with the prediction of the household bargaining model. Due to son preference in rural China, mothers with sons, especially first-born sons, are glorified by their families and gain higher bargaining power. The key outcome of this higher bargaining power is a reduction in maternal chore time. Additional empirical evidence includes increased maternal leisure time, increased maternal private consumption, and a higher likelihood of paternal grandparents' co-residence. These patterns are consistent with those previously documented by studies of the household bargaining model (Zhang and Chan, 1999). We do not find strong empirical support for alternative interpretations such as the demonstration effect, specialization effect, or sex-specific fertility-stopping rule.

Our findings have two implications. First, the study corroborates the view that child sex composition can directly affect parental time allocation between the labor market and the household (Rosenzweig and Wolpin, 2000). Our results cast doubt on the validity of using child sex composition as an instrument for other endogenous variables such as fertility. Second, our results improve understanding of the consequences of China's rising sex ratio; in particular, the rise in sex ratio intergenerationally benefits females by enhancing mothers' intra-household bargaining power. Nevertheless, the general equilibrium effect deserves further exploration.

Appendix A. Supplementary data

Supplementary data associated with this article can be found, in the online version, at <https://doi.org/10.1016/j.jebo.2017.11.012>.

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²² First-stage results are reported in Appendix Table B27 and discussed in Appendix C.

²³ Specifically, we regress residuals from the second-stage regression of the IV estimation on all of the explanatory variables plus the IVs, and obtain the unadjusted R^2 . We then compute $N \cdot R^2$ as the Hansen *J* statistic, where *N* is the number of observations. Under the joint null hypothesis in which the instruments are valid and correctly excluded from the estimated equation, the test statistic is distributed as χ^2 with degrees of freedom equal to the number of over-identifying restrictions (Hayashi, 2000).

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