Boys, Girls, and Grandparents:  
The Impact of the Sex of Preschool-Age Children  
on Family Living Arrangements and Maternal Labour Supply

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Abstract

This paper considers how the sex of preschool-aged children can affect their extended families’ living arrangements and maternal labor supply. Using China Health and Nutrition Survey (CHNS) data, we find that among less-educated mothers, the incidence of co-residing with the paternal grandmother is at least 8.6 percentage points higher if the firstborn is a boy. At the same time, maternal labor supply increases by 2.892 days per month. By contrast, for educated mothers who presumably have high opportunity costs of missing work, the propensity for co-residence is higher, and the working hours are longer relative to less-educated mothers, regardless of the child’s sex. Given son preference, the heterogeneous effects of child gender are consistent with our premise that sharing boys’ companionship and the labor division between generations are the two main forces determining both co-residence and household time allocation. This paper not only provides an improved understanding of the factors determining co-residence, but also lends empirical support to policies aiming to improve the well-being of girls.

Key Words: Living arrangement, gender inequality, childcare, maternal labor supply.
**Introduction**

A large and growing volume of empirical analyses use household-level surveys and, almost without fail, take household composition as exogenous. However, there is increasing recognition that household resource allocations are generally tied to living arrangements. For example, Foster and Rosenzweig (2002) note that co-production is tied to co-residence; Eggleston, Sun and Zhan (2016) find that the increasing use of medical services by the elderly is accompanied by a decline in co-residence with their adult children; Edmonds, Mammen and Miller (2005) show that the receipt of a pension triggers rearrangement of the family, while also changing child labor supply and school attendance (Edmonds 2006). An improved understanding of living arrangements is not only important with regard to empirical methodology; it also alerts researchers to the importance of internalizing living arrangements in their analytical frameworks to get a comprehensive understanding of household decision making.

This paper focuses on investigating the effect of child gender on co-residence and maternal time use. In most of the existing literature, the gender difference in investment in children and the family living arrangements are not reconciled in one framework. On the one hand, boys’ advantages are found along the dimensions of childcare time (Barcellos et al., 2014) and purchased inputs such as nutrition (Chen et al., 1981; Das Gupta, 1987), healthcare (Basu, 1989; Ganatra and Hirve, 1994) and vaccinations (Borooah, 2004). On the other hand, some recent studies show that under son preference, boys’ grandparents are more likely to co-reside and contribute to child rearing, especially to childcare and chores (Chen, Liu and Mair, 2011; Liu and Chen, 2016). In fact, investment in children is highly likely to be tied to co-residence, to the extent that a child is plausibly believed to be a household public good and that childcare time is determined in view of the household labor division.

In this paper, we formulate and test how child gender affects co-residence and maternal time use simultaneously, through these two channels. First, in many developing countries, under prevalent son preference, the companionship of a boy is likely to be deemed as more valuable than that of a girl. Therefore, a boy—and his parents—are more likely to co-reside with the grandparents. With the grandmother involved in domestic activities, the maternal
labor supply of households with boys is likely to increase, and the maternal time spent on domestic work decreases under the assumption of some extent of substitutability between the mother’s and grandmother’s time in the domestic realm.

Second, fast economic growth and the widening wage gap between generations result in the possibility of a scale economy from labor division. Mothers with a high wage rate are more likely to co-reside with grandparents, counting on their auxiliaries in the domestic realm and, thus, working longer hours on the job. This second channel implies a heterogeneous effect of child gender among mothers with different potential wage rates: For high-wage women, household decision making is driven not only by the incentive to share boys’ companionship, but also by labor division. Thus, the impact of the child’s sex is relatively weak. By contrast, for low-wage women, the surplus from labor division is limited, and, therefore, the child’s sex has a stronger impact.

In light of the simple model described above, our empirical analysis compares the likelihood of co-residence across different child gender and different maternal schooling, which is used as a proxy for potential wage in the labor market. The analysis uses China Health and Nutrition Survey (CHNS) data. In the household module of the CHNS, married women under age 52 were asked about their birth history and intergenerational linkages to their parents and in-laws, including living arrangements. To estimate the effect of child gender on co-residence, our analysis focuses on families with preschool-aged firstborn children because, according to the prior research, the sex of the firstborn child is arguably exogenous (Li et al., 2005; Li, 2007; Ebenstein, 2010; Sun and Zhao, 2016).

We find that among mothers with potentially low wage rates—i.e., mothers with only a primary school or no education—those with a firstborn boy are 8.6 percentage points more likely than those with a firstborn girl to live with their mother-in-law during the child’s preschool years. At the same time, among these less-educated mothers, those with a male firstborn spend 2.9 more days per month on paid work and significantly less time on family

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1 We want to focus on families with preschool-aged child(ren)—i.e., children between 0 and 6 years old—because for most families, this is the critical period when help with domestic work is most needed.
chores, such as cooking food and washing and ironing clothes. These findings are consistent with our premise that boys are more likely to have a co-residing grandmother who helps perform domestic work and enables more labor supply from the mother. There is no evidence of a reduction in maternal childcare if the preschool child is a boy, which is possibly because the labor of the mother and grandmother in terms of feeding, washing and supervising children is not fully substitutable.\(^2\) Consistent with the implications of heterogeneity of the impact of the child’s sex among women with different job opportunities, we find that educated women tend to co-reside with their in-laws and work longer hours generally, regardless of child gender.

The evidence from the reduced-form regression supports our hypothesis that the incentives of public good sharing and labor division drive household allocations and living arrangements. In addition, although it does not speak directly to the effect of the grandmother’s assistance on maternal time use, the reduced-form estimates suggest that co-residence could have a “childcare effect” or “family-chore-alleviation effect” on maternal labor supply, if we have an apriori expectation that under son preference, women should spend less time in the labor force participation.

Along these lines, this paper also builds on the literature on maternal labor supply and childcare arrangements. One strand of this literature proposes structural models to estimate the effect of paid care costs on maternal labor supply (Michalopoulos et al., 1992; Kimmel and Connelly, 2007). Another strand estimates reduced-form models of household choices, exploiting variations in childcare prices (Blau and Robins, 1988; Connelly, 1992) or natural experiments of subsidies for childcare (Anderson and Levine, 1999; Blau, 2003; Baker et al., 2005; Lefevre and Merrigan, 2008). The existing research focuses mainly on the United States and Canada, where the effect of lower childcare costs on maternal labor supply is modest. There is comparatively little research on the developing world, partly because of its lack of childcare institutions and partly because relatives generally provide childcare at little or no direct cost. This paper is based

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\(^2\) In CHNS, childcare includes only physical labor devoted to feeding, washing and supervising children. Activities such as washing children’s clothes and cooking children’s meals are regarded as family chores instead of childcare.
on the observation that childcare is tied to intergenerational living arrangements in developing countries. The empirical evidence supports the conjecture that the level of maternal labor supply is responsive to extended-family living arrangements. Similar evidence was found by Maurer-Fazio et al. (2011), who show that the presence of older people in the household increases prime-age urban women’s labor force participation in China.

The remainder of the paper is organized as follows: Section 2 provides background information on household divisions and childcare arrangements and the rising intergenerational income gap in China. In Section 3, we provide a theoretical framework and derive testable predictions. Section 4 presents summary statistics and graphical evidence. Section 5 describes reduced-form specifications, reports empirical results, and provides interpretations. In Section 6, we offer concluding remarks.

**Family Division, Childcare Arrangements, and the “Leaning in” of Grandparents**

Extended families dominate Chinese society. Family and kinship ties are guided by a strong patriarchal and patrilocal tradition. Sons are regarded as permanent members of the family line, while daughters leave their natal family behind and move to live with their husbands’ families at the time of marriage.\(^3\) According to Lang (1946), the extended family is generally composed of parents, their unmarried children, and one or more married sons with their wives and children.

The extended family later splits, in a process referred to as family division. “Family division” is generally defined as one married son and his wife and children moving out from their parents’ house when the extended family has multiple married sons (Cohen 1992).\(^4\) Several studies find that economic factors in large part determine when adult children move out of their parents’ house, and that demographic factors and social norms also play important roles.\(^5\)

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\(^3\) This patrilocality norm is not unique to China, but is also prevalent in other Asian countries like India and Vietnam (Ebenstein 2014).

\(^4\) It is rare for parents to live alone because the only son is obliged to take care of his elderly parents and provide for their funerals (Silverstein et al. 2006).

\(^5\) Intergenerational co-residence is of two main types. Grown-up sons and their wives and children may live in their parents’ place, or move their parents into their own houses to take care of them. The first type is the most prevalent.
Figure 1 depicts the proportion of co-residence against the age of the firstborn child in China. The ratio of co-residence has been declining, a finding consistent with the patrilocal tradition that women move to their in-laws’ house after getting married and later live in their own place. The greatest drop-off in the co-residence ratio is during the period when the firstborn child is between 0 and 6 years of age, when the proportion of co-residence declines from 0.75 to 0.42. China’s Law of Compulsory Education stipulates the school-entrance age as 6 years. Given the lack of formal institutions, childcare by household members is most important before the child attains school age, and the intensiveness of childcare demand declines with the child’s age.

Table 1 presents the percentage of each living arrangement across the age of the firstborn child. If we compare households in which the firstborn child is below school age with households where the firstborn child is 6–12 years of age, the ratio of co-residing with the paternal grandmother drops from 50.4 to 30.03 percentage points, and the ratio of living in an adjacent dwelling or the same courtyard, in another house in the same village or community, or in another village or community in the same county or city, increases by 11.94, 6.51, and 3.86 percentage points, respectively. By contrast, only a small portion of children live with their maternal grandmothers, and the ratio of co-residence does not change with the child’s age. This is not surprising considering the patrilocal tradition.

The lack of formal institutions providing domestic services, including childcare, is well documented in the literature. Paternal grandmothers play important roles in helping with childcare and other family chores. Figure 2 provides direct evidence of the time contributed toward domestic activities by paternal grandmothers. According to Fig. 2, a co-residing paternal grandmother on average spends 10.5 hours per week on childcare, 19.4 minutes per day on cleaning the house, 30.4 minutes on cooking, and 53.3 minutes on washing and ironing clothes.

The role that paternal grandmothers play in assisting domestic activities gains much more

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According to CHNS data, when focusing on co-residing households, 82.04% of the households report the household head to be one of the grown-up sons’ parents, and only 15.47% report the household head to be a grown-up son. The lack of formal childcare institutions in China is well documented (Jacka 1997; Kilburn and Datar 2002). Moreover, the scarce facilities are targeted more toward toddlers and children than infants. In our sample, more than 60% of the women had never used any kind of childcare service provided by a nonhousehold member.
importance in the context of China’s rapid economic growth: its gross domestic product (GDP) grew by nearly 10%, and household income in urban and rural areas increased by 8% between 1978 and 2012. As a consequence of this rapid growth, Li and Ding (2003) and Zhang et al. (2005) indicate that the rate of private returns to education increased from less than 2% in the early 1990s to around 10% at the beginning of the new century. This has led to a wider income gap between workers with different education levels.

The stock of human capital shows a sharp gradient across generations (Cai 2015): Among the working-age population in China, the education level gets lower with age. For instance, people aged 20 have 9 years of education, on average, while those aged 60 have only 6. Their higher level of education enables younger generations to benefit more from China’s rapid economic growth.

Considering the big wage gap between generations, grandparents’ decision to help raise their grandchildren makes clear economic sense, and the “leaning in” observed among grandparents is documented by researchers (Hermalin et al. 1998; Maurer-Fazio et al. 2011; Silverstein and Cong 2013) and reported by many journalists.⁷

Model

Assume that the extended family is composed of a preschool child, a mother, and a grandmother. Suppose the wages for the mother and the grandmother in family \( i \) are \( w_i^P \) and \( w_i^G \), respectively. In the context of China’s fast economic growth, without loss of generality, we assume \( w_i^G = 0 \) to simplify our analysis. To ensure a positive consumption for the grandmother, we assume she has unearned income \( y > 0 \). Both parents and grandparents have a time endowment of \( T \) hours.

Suppose the utility function of the parent (\( P \)) is:

\[
U^P(x_i^P, z_i) = x_i^P + z_i
\]

⁷ As reported in the U.S. periodical The Atlantic, a 68-year-old interviewee, Ida Lang, said, “My role staying home with the kids allows for my family’s success.” To her, the decision to provide childcare for her grandchildren is not difficult. “Just look at how much money she can make!” she exclaims, referring to her daughter-in-law. See http://www.theatlantic.com/china/archive/2013/09/in-china-its-the-grandparents-who-lean-in/280097/.
where $x_i$ is private consumption and $z_i$ is the child’s quality.

The utility of the grandparent ($G$) is:

$$U^G(x^G_i, s_i) = \begin{cases} 
  x^G_i + s_i & \text{if } C \\
  x^G_i & \text{if } NC
\end{cases}$$

Assume that the grandparent derives utility from a child’s companionship only when co-residing with the child. To sharpen the idea of the model, we illustrate the grandparent’s son preference by assuming that the grandparent only cares about the sex of the child, that is, $s_i = 1$ if the child is a boy and $s_i = 0$ if it is a girl.8

A child’s quality $z_i$ is produced by the time input of mother ($t^P_i$) and grandmother ($t^G_i$), and maternal time is partially substitutable by the grandmother’s time, that is, the effective childcare input is $t_i = t^P_i + \beta t^G_i$, $0 < \beta < 1$. Specifically, the production function of $z_i$ is given by

$$z_i = f(t^P_i, t^G_i) = \begin{cases} 
  -\infty & \text{if } t^P_i + \beta t^G_i < t \\
  \rho(t^P_i + \beta t^G_i) & \text{if } t^P_i + \beta t^G_i \in [t, \bar{t}] \\
  \rho \bar{t} & \text{if } t^P_i + \beta t^G_i = \bar{t}
\end{cases}$$

where $\rho$ is the marginal product of time input, which is a constant for simplicity. This production function assumes lower and upper bounds of time input, respectively, that is, (1) raising a child requires a minimal effective input of $t$; and (2) any more input beyond $\bar{t}$ cannot further increase the child’s quality.

Regarding maternal time spent on domestic activities, the mother involves herself in child rearing and family chores. We assume $\beta T < t$, implying that a minimum amount of maternal time ($t - \beta T$) is required in child rearing, for example, for breast feeding. By contrast, we suppose running a household requires $d$ hours on chores, which can all be assumed by the grandmother. Let $\bar{t} + d < T$, meaning if not at a job outside the home, 

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8 The model can be easily generalized to the scenario that the grandmother cares about both the sex and quality of the child.
the mother is always able to provide all family chores.

If the mother and grandmother live separately ($t^G = 0$), the grandmother gets utility $y$ and the mother solves the following problem:

$$\max_{t^P_i \in [0,T]} \quad x^P_i + f(t^P_i, 0) \quad \text{(NC)}$$

$$s.t. \quad x^P_i = (T - d - t^P_i)w_i$$

$NC$ represents for non-co-residence. The maternal time on childcare is

$$t^P_{i, NC} = \begin{cases} \bar{t} & \text{if } w_i < \rho \\ t & \text{if } w_i > \rho \end{cases}$$

The sum of the mother and grandmother’s utility with non-co-residence is

$$V_{i, NC}(w_i) = \begin{cases} (T - d - \bar{t})w_i + \rho \bar{t} + y & \text{if } w_i < \rho \\ (T - d - t)w_i + \rho t + y & \text{if } w_i > \rho \end{cases}$$

Consider the scenario under the arrangement of co-residence ($C$). Assume that there exists a disutility $\varepsilon_i$ for a family $i$ if two generations live together, $\varepsilon \in \mathcal{N}(-\infty, +\infty)$. $\varepsilon$ reflects disutility such as the loss of privacy, etc., and the cumulative distribution function is $\Phi(\varepsilon)$.

The joint family solves the following problem:

$$\max_{t^P_i, d^P_i} \quad \left[ x^P_i + f(t^P_i, t^G_i) \right] + \left[ x^G_i + s_i \right] - \varepsilon_i \quad \text{(C)}$$

$$s.t. \quad x^P_i + x^G_i = (T - d^P_i - t^P_i)w_i + y$$

$$t^G_i + (d - d^P_i) = T$$

where $d^P_i$ is the mother’s time on chores. We then derive

$$d^P_{i,C} = 0$$
\[ t_{i,c}^P = \begin{cases} T - \beta T - d & \text{if } w_i < \rho \\ t - \beta T - d & \text{if } w_i > \rho \end{cases} \]

We can denote the total utility of co-residence as \( V_{i,c}(w_i, s_i) - \varepsilon_i \)

\[ V_{i,c}(w_i, s_i) = \begin{cases} \left[ T + \beta(T - d) - \bar{t} \right] w_i + \rho \bar{t} + s_i + y & \text{if } w_i < \rho \\ \left[ T + \beta(T - d) - t \right] w_i + \rho t + s_i + y & \text{if } w_i > \rho \end{cases} \]

See all the proofs to derive \((d_{i,c}^P, t_{i,c}^P)\) in the Appendix.

**Lemma 1.** All else equal, upon co-residence, the maternal labor supply increases by \( d + \beta(T - d) \) compared with non-co-residence.

The intuition of the lemma is as follows. Since the grandmother’s and mother’s time are fully substitutable with regard to family chores and the grandmother is less productive in childcare, upon co-residence, the grandmother will then assume all the family chores \( d \) and spend the rest of her time \((T - d)\) on childcare. Correspondingly, compared with separate living, the maternal labor supply will increase by \( d + \beta(T - d)\), where \( \beta(T - d) \) is the effective time input on childcare by the grandmother.

The household members compare the utility of co-residence \((V_{i,c}(w_i, s_i) - \varepsilon_i)\) with that of non-co-residence \((V_{i,NC}(w_i))\). They co-reside if

\[ \varepsilon_i < \Delta_i \equiv V_{i,c}(w_i, s_i) - V_{i,NC}(w_i) = (d + \beta(T - d))w_i + s_i \]

where \( \Delta_i \) is the surplus from co-residence. The first term \((d + \beta(T - d))w_i\) represents the surplus from labor specialization between the mother and grandmother. The second term \( s_i \) represents the surplus from the mother and grandmother sharing the companionship of the child if it is a boy.

We can denote the likelihood of co-residence as \( \Psi(w_i, s_i) = \Phi[(d + \beta(T - d))w_i + s_i] \).

It is easy to have the propositions as follows. Please see all the proofs of Propositions and Lemma in the Appendix.
**Proposition 1:** All else equal, co-residence is more likely when the child is a boy.

**Proposition 2:** All else equal, co-residence is more likely when the mother’s wage is greater.

**Proposition 3:** When the child is a boy, as the mother’s wages increase, the probability of co-residence does not increase as much as when the child is a girl.

Proposition 3 implies a greater effect of the sex of the child on co-residence for low-wage mothers. Among these women \( w_i < \rho \), the surplus from labor specialization \((d + \beta(T - d)w_i)\) is small, hence son preference plays a more important role in the decision making on co-residence. By contrast, for high-wage women \( w_i < \rho \), the surplus from labor specialization \((d + \beta(T - d)w_i)\) is large, so co-residence relies less on the sex of the child.

Together with Lemma 1, Proposition 3 also implies greater effects of the sex of the child on maternal labor supply for low-wage mothers. This is because the sex of the child affects co-residence more effectively among low-wage women and at the same time, labor specialization upon co-residence predicts an increase in maternal labor supply by \( d + \beta(T - d) \).

**Data and Graphical Evidence**

**Data Descriptions**

The analyses use CHNS data. Using a random cluster process, the CHNS draws a sample of approximately 4,400 households with a total of 26,000 individuals in 9 provinces. Among villages and townships within counties and urban and suburban neighborhoods within cities were selected randomly. Using a fixed frame, the survey has been conducted every 3–4 years since 1989. This paper uses data from all waves except for 1989, in which

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9. Guangxi, Guizhou, Henan, Hubei, Hunan, Jiangsu, Liaoning, and Shandong are represented in all the waves, except for 1997. In 1997, Liaoning was replaced by Heilongjiang. For more details, please see the website of CHNS: [http://www.cpc.unc.edu/projects/china](http://www.cpc.unc.edu/projects/china).

10. CHNS allows for the tracking of respondents from intact households or from households that have separated from the original family but still live in the same sample area. The survey design does not follow changes to individuals who separated from their families between survey waves.
year key variables are unavailable. Appendix Table A1 summarizes the factors behind increases in household size. Most such increases occur when brides move in with their in-laws.

In the CHNS’s household frame, married women under age 52 were asked about their birth history and intergenerational linkages. These linkages include living arrangements, defined by the location of the parents’ and parents-in-law’s dwellings relative to that of the married women, and the basic demographics of the parents and in-laws. This detailed information on intergenerational linkages is important, considering that household members in most panel data do not necessarily live jointly.

Survey data on married women can then be linked to information on how they allocate their time, including for jobs on the market, childcare, and household chores (such as preparing and cooking food, washing and ironing clothes, and housecleaning). The employment module includes both labor supply and the wage rate. However, in the CHNS’s adult survey, only 42% of women in rural areas have off-farm employment. Therefore, we use education as a predictor of the opportunity cost of caregiving.

**Mean Comparison and Graphical Evidence**

In line with prior literature (Li et al. 2005; Li 2007; Ebenstein 2010; Sun and Zhao 2016), this study considers the sex of the firstborn child as exogenous. The sample is confined to women whose firstborn child is below school age.

Figure 3 presents the structure of living arrangements against that of women’s education levels by the sex of their firstborn. It is shown that for the less-educated women (the bars to the left), a firstborn son predicts a higher propensity of co-residence. By contrast, for educated women (the bars to the right), no such gender difference is exhibited.\(^{11}\)

Figure 4 shows maternal labor supply by gender and by schooling: less-educated women work more if they have firstborn sons, while the work levels of educated women do not vary much by their firstborn’s gender. Figures 3 and 4 are consistent with the model

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\(^{11}\) We also conduct a survival/hazard analysis of co-residence for male and female children and yield very similar results to those shown in Fig. 3. See also Appendix Fig 1.
propositions.

To see the significance levels of the mean comparisons, we conduct a $t$-test on the likelihood of co-residence, maternal labor supply (days per month), and time spent on different family chores. Table 2, columns (1) and (4) show the combined mean of co-residency status or maternal time use by maternal education levels. Columns (2) and (3), (5) and (6) show the mean comparison between the sex of the child within the same maternal education set.

As shown in Table 2, columns (2) and (3), among those with only primary education or no education, mothers with firstborn sons are 10.6 percentage points more likely to co-reside with the paternal grandmother than mothers with firstborn daughters. The difference is significant at the 1% level. At the same time, among less-educated mothers, those with firstborn sons work 3.2 more days per month on the job and spend significantly less time on family chores than those with firstborn daughters. The difference is the least significant, at the 10% level. Columns (5) and (6) show that the differences according to the sex of the child are mitigated as maternal education increases.

**Regression Evidence and Interpretations**

In this section, to obtain more precise estimations, we control for individual heterogeneities and conduct regression analyses. In addition, we discuss identification particulars. We also provide interpretations and further auxiliary evidence.

**Living Arrangements**

The main regression analysis uses the sample of households in which the firstborn child is below school age. This sample is used because the sex of the firstborn is arguably exogenous. Dahl and Morretti (2008) use it in a reduced-form specification to detect the demand for sons, using data from the United States. Moreover, Li et al. (2005), Li (2007), Ebenstein (2010), and Sun and Zhao (2016) examine China’s population census data and find that the sex of the firstborn is quite random.
The specification is:

(1) \( coriside_i = \alpha + FBS_i \cdot \beta_1 + SCH_i \cdot \beta_2 + FBS_i \times SCH_i \cdot \beta_3 + X_i \delta + \lambda_p + \tau_t + \epsilon_i \)

In Eq. (1), \( coriside_i \) is a dummy variable indicating whether mother \( i \) and the paternal grandmother live jointly. \( FBS_i \) is a dummy indicator of whether mother \( i \)'s firstborn child is a boy. \( SCH_i \) is a dummy indicator of whether mother \( i \) finishes middle school education, \( \lambda_p \) is the provincial fixed effect, and \( \tau_t \) is the calendar-year fixed effect. \( X_i \) includes the woman’s age and age squared to control the life-cycle pattern of co-residence. We also include covariates reflecting matching quality, which will be discussed in more detail in Section 5. The standard errors are clustered at the province level and adjusted using the Moulton factor.\(^{12}\)

The coefficient of interest is \( \beta_1 \). It captures the gender differential in co-residence for households with less-educated mothers. \( \beta_2 \) reflects the difference in co-residence between mothers with different education levels, if the firstborn is a girl. The effect of the mother’s opportunity cost of missing work is likely to be absorbed in \( \beta_2 \). \( \beta_3 \) reflects the mitigation of child gender disparity by maternal education. Thus, \( \beta_3 \) is expected to be negative.

Using the sample in which the firstborn is below school age, Table 3, columns (1) and (2), report the gender difference in co-residence among households. Column (1) presents the estimands, controlling for calendar-year fixed effects and community fixed effects. Column (2) further controls for parental age and the father’s education level. The estimands are consistent with Fig. 3, showing a higher likelihood of co-residence if the firstborn is a boy among households with a less-educated mother. In Panel A, the linear probability model in columns (1) and (2) shows a more-conservative estimand compared with the Probit model in Panel B: among less-educated mothers, a firstborn son leads to an 8.6 percentage points increase in the likelihood of co-residence and the difference is significant at the 5% level.

\(^{12}\) Since the variable of interest, the sex of the firstborn, is time invariant, we use the longitudinal data as repeated cross-sections. The estimation is weighed by the frequency in which a respondent appears in the sample.
Using multinomial Logit as an alternative model, we analyze multiple living arrangements. Since very few women live in a different village than their in-laws, we use women living in the same village but a different neighborhood as the reference group and examine the relative occurrence of co-residence and of living in an adjacent dwelling or the same courtyard. Appendix Table A2, column (1), reports the ratio of the probability of choosing co-residence over the probability of choosing the baseline category, and column (2) reports the z-score. The results show a very similar pattern to those presented in Table 3. Table A2 columns (3) and (4), show that the risk of living in an adjacent dwelling relative to the base group is not affected by the sex of the child or maternal education. This is consistent with the conjecture that economic connections are most intensive among co-residing members and weaken once the married son moves out from the parents’ house.\textsuperscript{13}

How differences in co-residence align with differences in the sex of firstborn grandchildren should not be easily attributed to a preference for sons (and grandsons) without considering other channels. For example, the sex of the firstborn could affect family size. Jensen (2003) analyzes the son-preferring fertility-stopping rules in India, showing that girls on average have more siblings than boys. In China, during the period of our sample, the family planning policy for rural residents allowed only one child if the firstborn was a boy, but allowed for a second birth when the firstborn was a girl. More children demand more childcare time and presumably an extra caregiver. But the effect of family size on co-residence is unclear. On the one hand, co-residence could be more likely as family size increases because of the higher productivity of the grandmother; on the other, the mother’s domestic productivity also rises, which decreases her labor supply and therefore decreases the co-residing surpluses from labor specialization.

As a robustness check on the estimands in Table 3, columns (1) and (2), we confine the sample to households with only one child. Columns (3) and (4) present the results using this subsample. The magnitudes of the coefficients are very similar to those presented in columns (1) and (2) in both tables, and the significant level drops a little due to a smaller...
sample size. The consistency of the results indicates that the firstborn gender is unlikely to affect co-residence through affecting the number of preschool children.

We conduct another robustness check to test the premise that decisions on family living arrangements hinge on the demand for childcare and related domestic work. We fit Eq. (1) using the placebo composed of households in which the oldest child is 16 or above, the age at which most children finish compulsory education and start to be regarded as adults in most legal and social circumstances. Appendix Table A3 shows that among these households, either the sex of the child or the maternal education level is significantly related to co-residence. In addition, the magnitudes of the coefficients are much smaller than those in Table 3.14

Finally, we address the possibility that the gender difference could reflect a different demand for care, that is, boys could be genetically different from girls and need more care, or their caregivers could believe this to be so. However, similar to what Barcellos et al. (2014) have found in India, researchers in China find that infant mortality is consistently higher among girls than boys. According to China’s 2000 population census, female infant mortality is 33.75 per thousand, while male infant mortality is only 23.92 per thousand. This difference does not support the thesis that boys need more care than girls.

Maternal Participation in the Labor Force and Domestic Realm

In this subsection, we investigate the model’s implication for maternal time use. As predicted by lemma 1, upon co-residence, the mother will adjust her time use. In particular, since the time of the mother and grandmother are to a large extent substitutable as regards family chores (such as cooking and washing clothes), it is likely that the mother will reduce her time on these activities and spend more time on the job. The empirical work focuses on households with less-educated mothers, because, as shown in Table 3, within this subset, the sex of the firstborn exhibits a stronger impact on co-

14 The insignificant coefficients reported in Table A3 are not because of the higher mortality of the grandmother; compared with the 14.8 percentage points mortality rate of the paternal grandmother in households with preschool children, the mortality in the placebo group only increased to 20.5 percentage points.
residing with the paternal grandmother.\textsuperscript{15}

This implication is clearly illustrated by the mean comparison in Table 2. Compared with mothers with firstborn girls, those with firstborn sons are 10.6 percentage points more likely to co-reside with a mother-in-law, work 3.2 days more per month, and spend significantly less time on family chores.

To get a more precise estimation, we fit the following regression:

\( y_i = \alpha + FBS_i \cdot \gamma + X_i \theta + \lambda_p + \tau_t + \epsilon_i \)

where \( y_i \) is the outcome variable. In particular, we examine maternal labor force participation, time spent on work (days per week), and each domestic activity, respectively. \( FBS_i \) is the dummy indicator of having a firstborn son, \( X_i \) is maternal age and age squared, \( \lambda_p \) is the provincial fixed effect, and \( \tau_t \) is the calendar-year fixed effect. Standard errors are clustered at the provincial level and corrected using the Moulton factor.

Table 4 reports the results of estimation. Panel A and B column (1) show the results of the ordinary least squares (OLS) and Probit estimation, respectively. The more conservative estimand is that among less-educated mothers, a firstborn son increases the likelihood of co-residence by 8.9 percentage points. This difference is significant at the 5\% level. In columns (2) to (6), Panel A and B show OLS and Tobit estimands, respectively. Column (2) shows that having a firstborn son increases labor supply by at least 2.9 days per month. The estimand is significant at the 10\% level. Columns (3) to (5) show that if the firstborn is male, the mother spends less time on domestic chores (such as cleaning the house, cooking, and washing clothes). This is consistent with the interpretation that since the time of the mother and grandmother is highly substitutable on these chores, co-residence shortens the mother’s time spent on these activities. The estimands of the child’s gender effect are not as significant compared with the mean comparison shown in

\textsuperscript{15} Educated mothers are likely to co-reside with their mothers-in-law regardless of the sex of the child, presumably because of the potentially big surpluses from labor division between generations. It could also be the case that educated mothers have more resources, which render the grandmother’s role less critical. We discuss this heterogeneity in more detail in the next subsection.
Table 2, presumably because controlling provincial and survey-year fixed effects reduces the variations. The OLS and Tobit estimands on the effects on maternal time spent on washing clothes are significant at the 10% level. The effect on the length of time spent cooking is only significant in the Tobit estimation, while that on the length of time spent on cleaning the house is not statistically significant in either estimation when controlling for provincial and survey-year fixed effects.

Column (6) presents the effect of the sex of the firstborn on maternal time spent on childcare. Neither the mean comparisons using raw data in Table 2 nor the regression results in Table 3 are significant. The $t$-value in Table 2 is -0.53. Therefore, there is no evidence showing that help from the grandmother “crowds out” the mother’s time on childcare.\(^{16}\) This is presumably because in the CHNS the physical care of children is defined as washing, dressing, feeding, and supervising children; it does not include cooking food, and washing the children’s clothes or cleaning the children’s room. The former cannot be fully substituted by the grandmother (breastfeeding offers an extreme example), while the latter is likely to be the margins that grandmothers are more likely to contribute to.

Similar to the robustness check conducted when estimating the effect of the sex of the child on co-residence, we confine the sample to households with only one child and fit Eq. (2) and present the results in Appendix Table A4. The magnitudes and significance levels of the coefficients are similar to that in Table 4. The robustness alleviates the concern that maternal time allocation could be mostly affected by fertility.

As regards time use, together with lemma 1, proposition 3 predicts a mitigation of the impact of child sex among educated women. The mean comparison in columns (1) and (4) in Table 2 are in favor of the prediction. To probe this heterogeneity in the use of time more precisely, we fit Eq. (2) using the subsample of women with a middle school or higher education, and the regression evidence is shown in Table 5. None of the estimands on the impact of the sex of the firstborn child is significant in Table 5, and the

\(^{16}\) If grandmothers are also involved in childcare (as shown in Fig. 2, panel 2d), it means that boys are likely to receive more hours of physical care in the presence of a grandmother.
magnitudes are much smaller compared with those in Table 4.

Finally, we conduct a placebo test to show that the sex of the child affects maternal time allocation through labor division between generations. In Appendix Table A5, we show that for extended families in which the paternal grandmothers have passed away, the sex of the child has no impact on time use, regardless of maternal education.

**Validity of Maternal Education as the Proxy for Job Opportunities**

In this paper, we use education as a proxy for maternal job opportunities. We address the validity from the following few perspectives. *First*, to address the concern that the heterogeneity could be caused by differences in more- and less-educated women’s choices of spouse,\(^\text{17}\) when estimating Eq. (1) using the whole sample, we include spouses’ characteristics to control matching quality, including the schooling years and age of the husband. As shown in Table 3, columns (2) and (4), the estimands remain very similar to that in the benchmark analysis. The consistency indicates that the heterogeneity is less likely driven by marriage matching to the extent that the pattern detected remains when controlling such important matching dimensions as education and age.

*Second*, maternal education could reflect the mother’s attitude toward childcare and toward co-residence and time use. Evidence suggests that educated mothers do not regard the caretaking activities of other caregivers as substitutes for maternal care. For example, Bertini et al. (2003) find that breastfeeding is associated with high maternal education. In addition, it is likely that educated mothers appreciate privacy more than the less educated. Therefore, the cost of co-residence is higher. In both cases, we would expect to see that more-educated mothers are less likely to co-reside with their in-laws, which is contrary to the results in Table 3.

*Third*, maternal earnings have an income effect. On the one hand, if richer mothers are better able to afford durable goods (such as housing, television, etc.), the income effect predicts that there is a greater possibility of them moving out of their in-laws’ house. On the other hand, richer mothers are more capable of providing instrumental help to elderly

\(^{17}\) For example, more-educated women could want to marry a husband from a family with a weaker son preference.
in-laws and could therefore be more likely to stay with them. This channel also predicts that educated mothers are more likely to live with a mother-in-law. However, if co-residence is driven by the incentive of providing instrumental help to elderly parents, the correlation between maternal schooling and co-residence should be stronger among older women, which is disproved by the results shown in Table A3. If we focus on households where grandchildren are 16 or older, maternal schooling has no impact on living arrangements.

**Concluding Remarks**

Gender inequality in childhood has been scrutinized in numerous studies. This paper focuses on the outcome of co-residing with a grandmother. We find that boys are on average more likely to live with their grandmothers, and this co-residence does not crowd out maternal time on childcare. Grandmothers are most likely to assist in household chores such as cooking and washing clothes. This finding implies that boys are likely to receive more childcare time than girls. Also, the mothers of boys work longer hours on the job than the mothers of girls, which could mean higher labor income for boys’ families, given the rising intergenerational wage gap in China.

Exploring the heterogeneity of the child-gender effect, we find that gender inequality in co-residence is greatly mitigated by an increase in maternal education. This is consistent with the conjecture that co-residence is also determined by the mother’s opportunity cost of missing work. The finding that gender inequality is more prevalent among households with less-educated mothers is disconcerting, because these households have fewer resources and are thus more dependent on a grandmother’s help. Therefore, (grand)son-biased co-residence potentially can do more harm to children who are already in a disadvantaged economic situation.

China has launched a series of programs aiming to improve the well-being of girls. For example, the “Care for Girls” program involves lectures to grandparents on gender equality, health examinations of girls to ensure they receive good care from their families, schooling subsidies for girls, and small loans to families with only daughters to help them develop an income-generating household economy. This paper lends empirical support to
such policies, especially on the policy component of changing grandparents’ bias against girls.

In a more general picture, this paper carries important implications for assessing the effects of public policies aimed at improving the well-being of children and women in the developing world, especially those whose target is alleviating gender inequality and improving girls’ living conditions. Since the sex of a child is very likely to affect (extended) household living arrangements, caution should be used when interpreting the effects of such policies.

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18 Such as DWCRA in India, PROGRESA in Mexico, and Care for Girls in China.

19 In addition, the analysis of this paper demonstrates that if panel data surveys drop entire households (or portions of them) that did not remain intact as part of the survey design, inferences about the policies’ effects from these data would thus be biased to the extent that household formations are nonrandom (Foster and Rosenzweig 2002).
Appendices

Proofs of Propositions and Lemmas

• Derive \((d_{i,C}^p, t_{i,C}^p)\)

If \(w_i < \rho\), then the optimal child input is \(\bar{t}_i\). To see this, note that, given any \(d_i^p\), since \(\bar{t} + d < T\), if the child input is less than \(\bar{t}\), then the total surplus can be greater by raising \(t_i^p\). Thus, \(t_i^p + \beta[T - (d - d_i^p)] = \bar{t}\), i.e., \(t_i^p = \bar{t} - \beta[T - (d - d_i^p)]\). The problem turns out to be

\[
\max_{d_i^p} (T - d_i^p - \bar{t} + \beta[T - (d - d_i^p)])w_i + \rho\bar{t} + s_i
\]

leading to \(d_{i,C}^p = 0\) and \(t_{i,C}^p = \bar{t} - \beta(T - d)\).

If \(w_i > \rho\), the the optimal child input is \(t_i\). To see this, note that, given any \(d_i^p\), if the child input is greater than \(t\), then the total surplus can be greater by reducing \(t_i^p\). Thus, \(t_i^p + \beta[T - (d - d_i^p)] = t_i\), i.e., \(t_i^p = t_i - \beta[T - (d - d_i^p)]\). The problem turns out to be

\[
\max_{d_i^p} (T - d_i^p - t_i + \beta[T - (d - d_i^p)])w_i + \rho t_i + s_i
\]

leading to \(d_{i,C}^p = 0\) and \(t_{i,C}^p = t_i - \beta(T - d)\).

• Proof of Lemma 1

Note that \((T - t_{i,C}^p) - (T - t_{i,NC}^p) = d + \beta(T - d)\).■

• Proof of Proposition 1

Note that \(\Phi[(d + \beta(T - d))w_i + 1] > \Phi[(d + \beta(T - d))w_i]\). ■

• Proof of Proposition 2

\(\Phi[(d + \beta(T - d))w_i]\) increases in \(w_i\). Hence, as mother’s wage increases the probability of co-residence for the girl family increases.

• Proof of Proposition 3

Note that
\[
\frac{\partial}{\partial w_i} \Phi[(d + \beta(T - d))w_i + 1] > \Phi[(d + \beta(T - d))w_i] = \beta(T - d)[\Phi[(d + \beta(T - d))w_i + 1] - \Phi[(d + \beta(T - d))w_i]] < 0,
\]

implying that for high-wage mothers, the impact of wage on the probability of co-residence is less. ■

23
References


Chen, F., Liu, G. and Mair, C.A., 2011. Intergenerational ties in context:


### Table 1 Grandmothers’ living arrangements, by grandchildren’s age

<table>
<thead>
<tr>
<th>Location of grandmothers’ homes, relative to that of a mother with at least one preschool-age child</th>
<th>Paternal grandmother’s dwelling</th>
<th>Maternal grandmother’s dwelling</th>
</tr>
</thead>
<tbody>
<tr>
<td>Age of the firstborn child</td>
<td>[0,6)</td>
<td>[6,12]</td>
</tr>
<tr>
<td>Co-reside</td>
<td>50.40</td>
<td>30.03</td>
</tr>
<tr>
<td>Adjacent dwelling or same courtyard</td>
<td>20.50</td>
<td>32.44</td>
</tr>
<tr>
<td>Another house in same village or community</td>
<td>15.84</td>
<td>22.35</td>
</tr>
<tr>
<td>Another village or community in same county or city</td>
<td>8.82</td>
<td>12.68</td>
</tr>
<tr>
<td>Another county or city in same province</td>
<td>2.65</td>
<td>2.37</td>
</tr>
</tbody>
</table>

**Source:** China Health and Nutrition Survey (CHNS) data.

**Notes:** The table summarizes the proportion of each living arrangement defined by the location of the in-laws’ or parents’ dwelling relative to that of mothers with preschool-age child(ren).
### Table 2 
Mean comparison of mothers’ living arrangements and activities, by their education level and the sex of a firstborn preschool-age child

<table>
<thead>
<tr>
<th></th>
<th>Women with no education or primary school education</th>
<th>Women with middle school education and above</th>
</tr>
</thead>
<tbody>
<tr>
<td></td>
<td>Combined</td>
<td>Firstborn boy</td>
</tr>
<tr>
<td>Co-reside with paternal</td>
<td>0.477</td>
<td>0.526**</td>
</tr>
<tr>
<td>grandmother (1=Yes; 0=No)</td>
<td>(0.5)</td>
<td>(0.5)</td>
</tr>
<tr>
<td></td>
<td>N=604</td>
<td>N=323</td>
</tr>
<tr>
<td>Labor supply (average number of days/month worked last year)</td>
<td>13.307</td>
<td>14.794*</td>
</tr>
<tr>
<td></td>
<td>N=254</td>
<td>N=136</td>
</tr>
<tr>
<td></td>
<td>(15.561)</td>
<td>(15.278)</td>
</tr>
<tr>
<td></td>
<td>N=384</td>
<td>N=203</td>
</tr>
<tr>
<td></td>
<td>N=668</td>
<td>N=356</td>
</tr>
<tr>
<td>Avg # mins/day spent on washing and ironing clothes</td>
<td>40.893</td>
<td>38.155*</td>
</tr>
<tr>
<td></td>
<td>N=607</td>
<td>N=317</td>
</tr>
<tr>
<td>Avg # hours/week spent on childcare</td>
<td>14.123</td>
<td>13.280</td>
</tr>
<tr>
<td></td>
<td>(21.236)</td>
<td>(21.703)</td>
</tr>
<tr>
<td></td>
<td>N=553</td>
<td>N=289</td>
</tr>
</tbody>
</table>

*Source:* China Health and Nutrition Survey (CHNS) data.

*Notes:* The sample is confined to women whose firstborn child is below school age.

The daggers and stars represent the significance level of the *t*-test comparing the means between women with a firstborn son and a firstborn daughter at each level of education. †p<.10; *p<.05; **p<.01
Table 3 The gender difference in grandchildren co-residing with a paternal grandmother

<table>
<thead>
<tr>
<th></th>
<th>Households in which the oldest child is below 6</th>
<th>Households in which the only child is below 6</th>
</tr>
</thead>
<tbody>
<tr>
<td></td>
<td>(1)</td>
<td>(2)</td>
</tr>
<tr>
<td><strong>Panel A. OLS estimands</strong></td>
<td></td>
<td></td>
</tr>
<tr>
<td>Firstborn boy ($\beta_1$)</td>
<td>0.086*</td>
<td>0.086*</td>
</tr>
<tr>
<td>(Yes=1; 0 otherwise)</td>
<td>(0.038)</td>
<td>(0.038)</td>
</tr>
<tr>
<td>Mother finishes middle school ($\beta_2$)</td>
<td>0.106**</td>
<td>0.116**</td>
</tr>
<tr>
<td>(Yes=1; 0 otherwise)</td>
<td>(0.032)</td>
<td>(0.033)</td>
</tr>
<tr>
<td>Firstborn son* finishes middle school ($\beta_3$)</td>
<td>-0.110*</td>
<td>-0.111**</td>
</tr>
<tr>
<td></td>
<td>(0.043)</td>
<td>(0.043)</td>
</tr>
<tr>
<td><strong>Panel B. Probit estimands (partial derivative at mean level)</strong></td>
<td></td>
<td></td>
</tr>
<tr>
<td>Firstborn boy ($\beta_1$)</td>
<td>0.104**</td>
<td>0.104**</td>
</tr>
<tr>
<td>(Yes=1; 0 otherwise)</td>
<td>(0.028)</td>
<td>(0.029)</td>
</tr>
<tr>
<td>Mother finishes middle school ($\beta_2$)</td>
<td>0.126**</td>
<td>0.137**</td>
</tr>
<tr>
<td>(Yes=1; 0 otherwise)</td>
<td>(0.023)</td>
<td>(0.027)</td>
</tr>
<tr>
<td>Firstborn son* finishes middle school ($\beta_3$)</td>
<td>-0.130**</td>
<td>-0.132**</td>
</tr>
<tr>
<td></td>
<td>(0.043)</td>
<td>(0.044)</td>
</tr>
<tr>
<td><strong>Control variables</strong></td>
<td>×</td>
<td>×</td>
</tr>
<tr>
<td><strong>Calendar-year fixed effects</strong></td>
<td>×</td>
<td>×</td>
</tr>
<tr>
<td><strong>Provincial fixed effects</strong></td>
<td>×</td>
<td>×</td>
</tr>
<tr>
<td><strong>Observations</strong></td>
<td>2,552</td>
<td>2,552</td>
</tr>
</tbody>
</table>

*Source: China Health and Nutrition Survey (CHNS) data.*

*Notes: The sample is confined to women whose firstborn child is below school age.*

Standard errors (robust to correlations of residuals within provinces and heteroscedasticity) are in parentheses. The standard errors are corrected using the Moulton factor.

The daggers and stars represent the significance level †p<.10; *p<.05; **p<.01.

Regressions also include a constant term, urban/rural status, calendar-year fixed effects, and provincial fixed effects. In columns (2) and (4), the control variables include parental age, father’s education level.
Table 4 The impact of the sex of the child on family living arrangements and maternal time allocation among less-educated mothers

<table>
<thead>
<tr>
<th>Variables</th>
<th>Co-reside with the paternal grandmother</th>
<th>Labor supply (days/month)</th>
<th>Clean house (mins/day)</th>
<th>Prepare and cook food (mins/day)</th>
<th>Wash and iron clothes (mins/day)</th>
<th>Take care of children (hrs/week)</th>
</tr>
</thead>
<tbody>
<tr>
<td></td>
<td>(1)</td>
<td>(2)</td>
<td>(3)</td>
<td>(4)</td>
<td>(5)</td>
<td>(6)</td>
</tr>
<tr>
<td><strong>Panel A. OLS estimands</strong></td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>Firstborn boy ($\gamma$)</td>
<td>0.089*</td>
<td>2.892†</td>
<td>-0.926</td>
<td>-2.024</td>
<td>-4.402†</td>
<td>-1.081</td>
</tr>
<tr>
<td>(0.039)</td>
<td>(1.501)</td>
<td>(1.529)</td>
<td>(1.337)</td>
<td>(2.558)</td>
<td>(1.770)</td>
<td></td>
</tr>
<tr>
<td><strong>Panel B. Probit/tobit estimands</strong></td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>Firstborn boy ($\gamma$)</td>
<td>0.107*</td>
<td>4.059†</td>
<td>-1.551</td>
<td>-2.505†</td>
<td>-4.556†</td>
<td>-1.696</td>
</tr>
<tr>
<td>(0.045)</td>
<td>(2.324)</td>
<td>(1.759)</td>
<td>(1.473)</td>
<td>(2.713)</td>
<td>(1.871)</td>
<td></td>
</tr>
<tr>
<td>Observations</td>
<td>604</td>
<td>246</td>
<td>344</td>
<td>619</td>
<td>559</td>
<td>515</td>
</tr>
</tbody>
</table>

Source: China Health and Nutrition Survey (CHNS) data.

Notes: The sample is confined to women whose firstborn child is below school age and who have no or only a primary education. Standard errors (robust to correlations of residuals within provinces and heteroscedasticity) are in parentheses. The standard errors are corrected using the Moulton factor. The daggers and stars represent the significance level †p<.10; *p<.05.

Regressions also include a constant term, maternal age, age squared, urban/rural status, calendar-year fixed effects, and provincial fixed effects.
Table 5 The impact of a child’s sex on living arrangements and maternal time allocation among educated mothers

<table>
<thead>
<tr>
<th>Variables</th>
<th>Co-reside with the paternal grandmother</th>
<th>Labor supply (days/month)</th>
<th>Clean house (mins/day)</th>
<th>Prepare and cook food (mins/day)</th>
<th>Wash and iron clothes (mins/day)</th>
<th>Take care of children (hrs/week)</th>
</tr>
</thead>
<tbody>
<tr>
<td><strong>Panel A. OLS estimands</strong></td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>Firstborn boy (γ)</td>
<td>-0.028</td>
<td>0.593</td>
<td>0.371</td>
<td>0.232</td>
<td>0.606</td>
<td>-1.134</td>
</tr>
<tr>
<td></td>
<td>(0.022)</td>
<td>(0.516)</td>
<td>(0.936)</td>
<td>(1.185)</td>
<td>(1.317)</td>
<td>(1.129)</td>
</tr>
<tr>
<td><strong>Panel B. Probit/tobit estimands</strong></td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>Firstborn boy (γ)</td>
<td>-0.037</td>
<td>0.870</td>
<td>0.639</td>
<td>1.279</td>
<td>1.395</td>
<td>-1.072</td>
</tr>
<tr>
<td></td>
<td>(0.025)</td>
<td>(0.697)</td>
<td>(1.103)</td>
<td>(1.585)</td>
<td>(1.526)</td>
<td>(1.262)</td>
</tr>
<tr>
<td>Observations</td>
<td>1,948</td>
<td>1,595</td>
<td>1,297</td>
<td>1,892</td>
<td>1,785</td>
<td>1,616</td>
</tr>
</tbody>
</table>

Source: China Health and Nutrition Survey (CHNS) data.
Notes: The sample is confined to women whose firstborn child is below school age and who attain middle school education or above. Standard errors (robust to correlations of residuals within provinces and heteroscedasticity) are in parentheses. The standard errors are corrected using the Moulton factor. Regressions also include a constant term, maternal age, age squared, urban/rural status, calendar-year fixed effects, and provincial fixed effects.
Appendix Tables

Table A1. Demystifying the increase in family size

<table>
<thead>
<tr>
<th>CHNS wave</th>
<th>2004</th>
<th>2006</th>
<th>2009</th>
<th>2011</th>
</tr>
</thead>
<tbody>
<tr>
<td>Total number of households</td>
<td>3,014</td>
<td>3,107</td>
<td>3,075</td>
<td>3,593</td>
</tr>
<tr>
<td><strong>Households with new members joined</strong></td>
<td>679</td>
<td>452</td>
<td>679</td>
<td>1,000</td>
</tr>
<tr>
<td>Among which (pt.)</td>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>New baby born</td>
<td>42.56</td>
<td>47.57</td>
<td>41.97</td>
<td>44.70</td>
</tr>
<tr>
<td>Marriages</td>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>Brides moved to in-laws’ house</td>
<td>19.88</td>
<td>20.13</td>
<td>24.01</td>
<td>19.70</td>
</tr>
<tr>
<td>Grooms moved to in-laws’ house</td>
<td>0.15</td>
<td>2.43</td>
<td>1.03</td>
<td>0.50</td>
</tr>
<tr>
<td>Spouse moved into a single-person household</td>
<td>5.15</td>
<td>6.19</td>
<td>7.66</td>
<td>3.80</td>
</tr>
<tr>
<td>Elderly parents moved in with adult children</td>
<td>5.89</td>
<td>0.42</td>
<td>1.77</td>
<td>0.70</td>
</tr>
<tr>
<td>Age of the elderly when moved in (years)</td>
<td>67.1</td>
<td>73.5</td>
<td>78.1</td>
<td>77.9</td>
</tr>
<tr>
<td>Relatives/friends of the household head moving in</td>
<td>26.37</td>
<td>23.26</td>
<td>23.56</td>
<td>30.6</td>
</tr>
</tbody>
</table>

*Source:* China Health and Nutrition Survey (CHNS) data.
Table A2. Multinomial estimation of the living arrangements of women with preschool-age boys

<table>
<thead>
<tr>
<th></th>
<th>Co-reside with parental grandmother</th>
<th>Adjacent dwelling or same courtyard as parental grandmother</th>
</tr>
</thead>
<tbody>
<tr>
<td></td>
<td>Relative risk</td>
<td>Z-score</td>
</tr>
<tr>
<td></td>
<td>(1)</td>
<td>(2)</td>
</tr>
<tr>
<td>Firstborn son ($\beta_1'$)</td>
<td>1.350</td>
<td>1.69</td>
</tr>
<tr>
<td>Finish middle school ($\beta_2'$)</td>
<td>1.746*</td>
<td>2.54</td>
</tr>
<tr>
<td>Firstborn son* finishes middle school ($\beta_3'$)</td>
<td>0.636*</td>
<td>-2.72</td>
</tr>
<tr>
<td>Age</td>
<td>0.784</td>
<td>-1.63</td>
</tr>
<tr>
<td>Age squared</td>
<td>1.003</td>
<td>1.24</td>
</tr>
<tr>
<td>Husband’s age</td>
<td>0.931**</td>
<td>-2.72</td>
</tr>
<tr>
<td>Husband finishes middle school</td>
<td>0.921</td>
<td>-0.50</td>
</tr>
</tbody>
</table>

Observations: 2,564 2,564

Source: China Health and Nutrition Survey (CHNS) data.

Notes: The sample is confined to women whose firstborn child is below school age. The outcome variable is a categorical indicator of living arrangements. 1=co-reside; 2=living in an adjacent dwelling or same courtyard; 3=living in another house in the same village or community. Category 3: Living in another house in the same village or community is set as the base. The daggers and stars represent the significance level $\dagger p<.10$; $* p<.05$; **$p<.01$. Regressions also include a constant term, urban/rural status calendar-year fixed effects, and provincial fixed effects.
**Table A3.** The gender difference in co-residence with a paternal grandmother among households with a firstborn child over 15 years of age

<table>
<thead>
<tr>
<th></th>
<th>Panel A. OLS estimands</th>
<th>Panel B. Probit estimands (partial derivative at mean level)</th>
</tr>
</thead>
<tbody>
<tr>
<td></td>
<td>(1)</td>
<td>(2)</td>
</tr>
<tr>
<td>Firstborn boy ($\beta_1$)</td>
<td>0.017</td>
<td>0.017</td>
</tr>
<tr>
<td>(Yes=1; 0 Otherwise)</td>
<td>(0.021)</td>
<td>(0.021)</td>
</tr>
<tr>
<td>Mother finishes middle school ($\beta_2$)</td>
<td>0.026</td>
<td>0.024</td>
</tr>
<tr>
<td>(Yes=1; 0 Otherwise)</td>
<td>(0.020)</td>
<td>(0.021)</td>
</tr>
<tr>
<td>Firstborn son* finishes middle school ($\beta_3$)</td>
<td>-0.030</td>
<td>-0.035</td>
</tr>
<tr>
<td></td>
<td>(0.027)</td>
<td>(0.028)</td>
</tr>
</tbody>
</table>

Control variables
- Calendar-year fixed effects
- Provincial fixed effects

Observations 3,398 3,031

*Source:* China Health and Nutrition Survey (CHNS) data.

*Notes:* The sample is confined to women whose firstborn child is over 15 years of age. Standard errors (robust to correlations of residuals within provinces and heteroscedasticity) are in parentheses. The standard errors are corrected using the Moulton factor. Regressions also include a constant term, urban/rural status, calendar-year fixed effects, and provincial fixed effects. In column (2), the control variables include parental age and father’s education level.
Table A4. The impact of a child’s sex on family living arrangements and maternal time allocation among less-educated mothers (in households with one child)

<table>
<thead>
<tr>
<th>Variables</th>
<th>Co-reside with the paternal grandmother (1=Yes; 0=No)</th>
<th>Labor supply (days/week)</th>
<th>Clean house (mins/day)</th>
<th>Prepare and cook food (mins/day)</th>
<th>Wash and iron clothes (mins/day)</th>
<th>Take care of children (hrs/week)</th>
</tr>
</thead>
<tbody>
<tr>
<td>Panel A. OLS estimands</td>
<td>Firstborn boy (γ)</td>
<td>0.078†</td>
<td>0.892*</td>
<td>-1.444</td>
<td>-2.439*</td>
<td>-3.256</td>
</tr>
<tr>
<td></td>
<td></td>
<td>(0.047)</td>
<td>(0.417)</td>
<td>(1.756)</td>
<td>(1.436)</td>
<td>(2.912)</td>
</tr>
<tr>
<td>Panel B. Probit/tobit estimands</td>
<td>Firstborn boy (γ)</td>
<td>0.094†</td>
<td>1.219*</td>
<td>-2.404</td>
<td>-2.869*</td>
<td>-3.526</td>
</tr>
<tr>
<td></td>
<td></td>
<td>(0.051)</td>
<td>(0.608)</td>
<td>(1.977)</td>
<td>(1.735)</td>
<td>(3.027)</td>
</tr>
</tbody>
</table>

Observations: 364 183 255 407 374 328

Source: China Health and Nutrition Survey (CHNS) data.

Notes: The sample is confined to women whose firstborn child is below school age and who have primary education or no education. Standard errors (robust to correlations of residuals within provinces and heteroscedasticity) are in parentheses. The standard errors are corrected using the Moulton factor. The daggers and stars represent the significance level †p < .10; *p < .05. Regressions also include a constant term, maternal age, age squared, urban/rural status, calendar-year fixed effects, and provincial fixed effects.
Table A5. The impact of a child’s sex on maternal time allocation among less-educated mothers across households with the grandmother alive and passed away

<table>
<thead>
<tr>
<th></th>
<th>Time spent on domestic activities</th>
<th></th>
<th></th>
<th></th>
<th></th>
</tr>
</thead>
<tbody>
<tr>
<td></td>
<td>Labor supply (days/month)</td>
<td>Clean house (mins/day)</td>
<td>Prepare and cook food (mins/day)</td>
<td>Wash and iron clothes (mins/day)</td>
<td>Take care of children (hrs/week)</td>
</tr>
<tr>
<td></td>
<td>(1)</td>
<td>(2)</td>
<td>(3)</td>
<td>(4)</td>
<td>(5)</td>
</tr>
<tr>
<td><strong>Panel A. Households with the grandmother alive</strong></td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>Firstborn boy (γ)</td>
<td>5.796*</td>
<td>-1.377</td>
<td>-2.685†</td>
<td>-4.924†</td>
<td>-0.579</td>
</tr>
<tr>
<td></td>
<td>(2.828)</td>
<td>(1.700)</td>
<td>(1.505)</td>
<td>(2.752)</td>
<td>(1.814)</td>
</tr>
<tr>
<td>Observations</td>
<td>209</td>
<td>296</td>
<td>526</td>
<td>476</td>
<td>434</td>
</tr>
<tr>
<td><strong>Panel B. Households with the grandmother passed away</strong></td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>Firstborn boy (γ)</td>
<td>-0.910</td>
<td>-0.862</td>
<td>3.415</td>
<td>6.871</td>
<td>-0.677</td>
</tr>
<tr>
<td>Observations</td>
<td>32</td>
<td>41</td>
<td>85</td>
<td>76</td>
<td>72</td>
</tr>
</tbody>
</table>

Source: China Health and Nutrition Survey (CHNS) data.

Notes: The sample is confined to women whose firstborn child is below school age and who have no or only a primary education. Standard errors (robust to correlations of residuals within provinces and heteroscedasticity) are in parentheses. The standard errors are corrected using the Moulton factor. The daggers and stars represent the significance level †p<.10; *p<.05.

Regressions also include a constant term, maternal age, age squared, urban/rural status, calendar-year fixed effects, and provincial fixed effects.
Figures

Fig. 1 Ratio of women co-residing with in-laws to the age of firstborn child

*Source:* China Health and Nutrition Survey (CHNS) data.

*Notes:* This figure shows the ratio of women co-residing with in-laws against the age of the firstborn child. The vertical axis is the ratio of the number of co-residing households to all households.
Fig. 2 Length of time spent on family chores, by family living arrangements

Source: China Health and Nutrition Survey (CHNS) data.

Notes: The figure shows time spent on domestic activities by the mother, paternal grandmother, and other household members in co-residing and non-co-residing households. The light gray bars represent the time contributed by paternal grandmothers, which counts for a large portion of the total time spent on each of the family chores.
Fig. 3 Proportion of women co-residing with their mother-in-law by their education level and the sex of the firstborn child

Source: China Health and Nutrition Survey (CHNS) data.

Notes: The sample is confined to women whose firstborn child is below school age. The bars are the sample means of the dummy indicator of co-residing with the mother-in-law (1=co-reside; 0 otherwise) by women’s education level and the sex of their firstborn child.
Fig. 4 Time spent on the job (days per month) by women’s education level and the sex of the firstborn child

Source: China Health and Nutrition Survey (CHNS) data.

Notes: The sample is confined to women whose firstborn child is below school age. The bars are the sample mean of time spent on the job (days per month) by women’s education level and the sex of the firstborn child.
Appendix Fig. A1. The duration of co-residence by child sex

**Source:** China Health and Nutrition Survey (CHNS) data.

**Notes:** The steps in each figure are survival functions for households with a firstborn son and daughter, respectively. The “failure” in this survival analysis is non-co-residence or extended family division. The sample is composed of women whose firstborn child is below school age at the starting time of the analysis.