Journal of Development Economics 166 (2024) 103191

Contents lists available at ScienceDirect

Journal of Development Economics

journal homepage: www.elsevier.com/locate/devec

Regular Article

Fertility and delayed migration: How son preference protects young girls against mother–child separation

Zibin Huang [a], Xu Jiang [b], Ang Sun [b,*]

[a] College of Business, Shanghai University of Finance and Economics; Shanghai Institute of International Finance and Economics, China
[b] School of Finance, Renmin University, China

ARTICLE INFO

Keywords:
Left-behind children
Fertility
Migration
Son preference
Child development

ABSTRACT

Mother–child separation harms children's development. This concern is particularly relevant in rapidly urbanizing societies with massive migration. However, factors that increase the probability of children becoming separated from their migrating parents are not well understood. In this study, we find that during early childhood, specifically at the age of 2 years, girls are less likely to become separated from their parents than boys are. This is due to son preference, which leads parents to attempt to have another child after a firstborn daughter, mostly when the daughter is 2 years old. Furthermore, migrant women often stay in their hometowns during pregnancy, which prevents mother–daughter separation. Our placebo tests indicate that this gender gap at the age of 2 years does not exist in provinces with birth interval restrictions or strict one-child policies. The gap is wider in regions with stronger son preference or higher historical population outflow rates.

1. Introduction

Early childhood experiences establish the foundation for optimum health, growth, and neurodevelopment throughout the lifespan (Cunha et al., 2006, 2010; Heckman, 2006; Cunha and Heckman, 2007). Parental companionship, especially from one's mother, plays a critical role in building this foundation (Waldfogel et al., 2002; Baum, 2003; Luby et al., 2012; Zhao et al., 2014; Luby et al., 2016). However, knowledge of who is more likely to experience this separation and why is rather lacking.

This question is particularly relevant for countries that have undergone rapid urbanization. According to the United Nations Development Programme (2009), over 900 million people worldwide migrate away from home for work. For instance, China has witnessed a massive wave of migration in the past few decades. As of 2021, the total number of migrant workers in China was approximately 300 million (The State Council of PRC, 2022). When workers migrate to more developed regions, they often leave their children behind in their hometowns owing to high childcare costs or limited access to public schools and kindergartens in urban areas (Cortes, 2008, 2015; Graham and Jordan, 2011; Zhang et al., 2014). In China, in 2015, the number of left-behind children was estimated to be as high as 70 million (Huang, 2022).

Using a 0.35% sample from China's 2010 population census, we examine gender differences in parent–child separation for pre-school-aged firstborn children in rural China.[1] We find that parents are less likely to leave girls behind than boys in early childhood; this difference is most prominent at age 2. In line with this result, we find the likelihood of parents migrating for work to be significantly higher for boys than for girls at age 2. Such gender differences persist across a battery of robustness checks.

We argue that these outcomes are driven by the fact that in the context of son preference, parents with firstborn girls are more likely to try to have a second child shortly after their first child's birth (Ebenstein, 2010). Additionally, mothers usually remain in their hometowns during pregnancy and childbirth, restricted by the reimbursement policies of health insurance. The reimbursement rate is only 30%–40% for medical costs outside one's registered residential (*hukou*) region, compared with 50%–75% for expenditures that occur within one's locality (Yi and Gu, 2015). According to statistics from the 2012 wave of the China Migration Dynamic Study, over 66.9% of migrant mothers stay in their

* Corresponding author.
E-mail addresses: huangzibin@mail.shufe.edu.cn (Z. Huang), xu.jiang@ruc.edu.cn (X. Jiang), ang.sun@gmail.com (A. Sun).

[1] We select only firstborn children because the genders of high-order births in China are not random. Parents may select their children's gender through abortion or even abandonment (Chen et al., 2013; Bao et al., 2023). However, the sex of the first child in areas with the 1.5 or two-child policy can be considered random. We discuss this in more detail in Section 5.1.

https://doi.org/10.1016/j.jdeveco.2023.103191
Received 30 November 2022; Received in revised form 4 September 2023; Accepted 10 September 2023
Available online 16 September 2023
0304-3878/

hometowns for most or all of their pregnancies, and 80.9% give birth in their hometowns. The fact that pregnant women postpone migration or return to their hometowns effectively prevents mother–daughter separation until a mother gives birth to her second child, which is around the time when her firstborn turns 2 years old. No such factors appear to restrict separation for boys in early childhood. Thus, son preference paradoxically protects girls from becoming separated from their mothers in early childhood.

We investigate the abovementioned mechanism of fertility-caused migration delay by conducting two placebo tests. First, we compare provinces with different requirements for the interval between a woman's first and second births. Some provincial governments mandate a waiting period of at least four years or require the mother to reach a certain age before having a second child to control population growth. Our analysis reveals that gender differences in early childhood separation do not exist in regions with such restrictions, as parents cannot freely respond to the birth of a firstborn girl by immediately having a second child. Second, we compare provinces with strict one-child policies to those with more relaxed policies. In response to protests from rural communities, some provinces implemented policies allowing families to have two children (the two-child policy) or to have a second child when the first child is a girl (the 1.5-child policy). We note that the observed gender gap in early childhood separation only occurs in these provinces with the 1.5 or two-child policy. By contrast, we do not observe this gender gap in provinces with strict one-child policies, as families are not permitted by law to respond to the birth of a firstborn girl by having a second child.

To further examine this mechanism, we conduct two heterogeneity analyses. First, we categorize the sample into groups based on different strengths of son preference. We use the regional sex ratio of the cohort of interest as a proxy for the strength of son preference in the locality. We detect a significant gender difference in being left behind only in regions with skewed sex ratios. In addition, according to the past findings that indicate education weakens son preference (Chung and Gupta, 2007), we examine heterogeneity across parental education and find that gender differences in mother–child separation exist only in families with lower levels of parental education. Second, we explore gender differences across regions with different historical ratios of population outflow, revealing that gender differences in parent–child separation are more prominent in regions with high historical migration rates. These results align with our premise that significant gender differences in parent–child separation at 2 years of age are likely to be driven by parents' motivation to attempt to have a son shortly after the birth of a firstborn girl.

The separation of mothers from their sons can explain another phenomenon: that of girls largely outperforming boys in both cognitive and non-cognitive domains (Kimball, 1989; Szatmari et al., 1989; Campbell et al., 2000; Gilliam, 2005; Ready et al., 2005; Beaman et al., 2006; Entwisle, 2018). Numerous studies have shown the detrimental effect of mother–child separation on children's development (Waldfogel et al., 2002; Baum, 2003; Luby et al., 2012; Luby et al., 2016). Accordingly, we briefly explore the correlation between children's experiences of being left behind and their human capital outcomes. We find that becoming separated from their parents in early childhood is associated with a lower class ranking after attaining 10 years old. Regarding gender gap in academic performance, according to a simple back-of-the-envelope calculation, we find that the experience of being left behind at age 2 accounts for 5.0%–20.2% of the total gender difference in class ranking. These results suggest that more companionship from mothers in early childhood benefits girls' long-term development and that the gender gap in mother–child separation is associated with the gender gap in human capital.

Our study extends the current literature from several angles. First, various studies have illustrated that left-behind children experience huge human capital losses (Sun et al., 2010; He et al., 2012; Wen and Lin, 2012; Chen, 2013; Mou et al., 2013; Zhang et al., 2014; Gao et al., 2015; Mu and De Brauw, 2015; Ren and Treiman, 2016; Meng and Yamauchi, 2017; Luo, 2020; Huang, 2022). Specifically, Yue et al. (2016) note that maternal migration during early childhood has significant negative effects on the cognitive development of children from rural China. Separation from mothers between 6 months and 2 years of age negatively affects children's cognitive scores, dietary quality, height growth, and weight. The present study adds to the existing literature by investigating gender differences in the likelihood of being left behind, providing important implications for addressing gender inequality in parental care in the early childhood.

Second, many studies have discussed the consequences of son preference from various perspectives since Sen (1992) proposes the "missing girls" phenomenon. Son preference influences both fertility and investment in children (Filmer et al., 2008; Liu, 2014; Wang, 2019). In China and South Asia, son preference causes gender-selective abortions of female fetuses, resulting in a skewed sex ratio at birth (Park and Cho, 1995; Das Gupta et al., 2003; Ebenstein, 2010; Chen et al., 2013). In India, Das (1987) describes the "stop-after-a-son" rule for fertility. As a consequence of this rule, Jayachandran and Kuziemko (2011) observe a significantly lower breastfeeding duration for daughters than for sons. Moreover, son preference leads to unequal resource distribution between boys and girls in early childhood, causing gender inequality in dimensions such as educational opportunities, nutritional intake, and health conditions (Pitt and Rosenzweig, 1990; Rose, 1999; Borooah, 2004; Wang, 2005; Gong et al., 2005; Jayachandran and Pande, 2017; Iqbal et al., 2018). However, to our knowledge, this study is the first to highlight that son preference can paradoxically protect girls by reducing mother–child separation.

Third, this study sheds light on the puzzling phenomenon in which, on average, girls perform better academically than boys (Kimball, 1989; Campbell et al., 2000; Entwisle, 2018). Other than biological influences, economists have focused on the role of social influences. For example, Bertrand and Pan (2013), Fan et al. (2015), Aucejo and James (2019) and Autor et al. (2019) indicate that boys are more vulnerable than girls in families with single parents, employed mothers, or low socioeconomic status. Boys in such families are more likely to experience externalizing problems, school suspensions, and eventually lower educational attainment. In this study, we identify a novel channel as a parallel mechanism that can be regarded as a good example of the fact that, compared to girls, boys are not only more vulnerable to separation from their parents but are also more likely to experience such unfavorable changes.

In general, our study makes two main contributions to the literature. First, it is the first study to causally examine and discuss the gender disparity in children being left behind by their parents. Second, we uncover a novel finding that son preference can paradoxically act as a protective factor for girls by reducing the likelihood of separation between mothers and their daughters.

The remainder of the paper is organized as follows. In Section 2, we describe the background of the widespread son preference and parent–child separation in China. In Section 3, we give an overview of our dataset and present some preliminary graphical evidence. In Section 4, we present difference-in-differences (DID) estimates of gender disparities in parent–child separation and explain fertility motives as the channel driving such differences. In Section 5, we address the details of identification and provide interpretations. Section 6 explores the heterogeneities of gender differences in parent–child separation to better elucidate this mechanism. In Section 7, we report the robustness checks and provide supplementary evidence. Section 8 discusses the importance of early childhood separation using a back-of-the-envelope calculation, and Section 9 concludes the paper.

2. Background

2.1. Son preference and gender disparities in children's development

Son preference is a long-standing social problem in many countries (Sen, 1992; Dahl and Moretti, 2008; Blau et al., 2020). It is uniquely severe in Asia, especially in regions with a long history of Confucian culture (Ebenstein, 2021). Aside from the most salient manifestation—namely the "missing girls" issue (Sen, 1992), reflected by acutely skewed sex ratios in China (Ebenstein, 2010), South Korea (Edlund and Lee, 2013), and India (Clark, 2000; Mitra, 2014)—girls usually occupy an inferior position in terms of family resource allocation during childhood. Parents often invest less in their daughters regarding both education and nutrition (Gong et al., 2005; Jayachandran and Pande, 2017). Scholars and policymakers are particularly concerned about such gender inequality because early childhood is the most important stage for human capital formation (Cunha and Heckman, 2007; Cunha et al., 2010). Owing to the imbalanced resource allocation of nutritional resources and medical care, compared with boys, girls are more likely to be underweight and have higher morbidity and mortality rates (Chen et al., 1981; Dancer et al., 2008).

However, despite such disadvantages, on average, girls perform better in reading and writing, and even in subjects considered more stereotypically "masculine" such as math in primary and secondary school (Kimball, 1989; Campbell et al., 2000; Entwisle, 2018). Data from the 2010–2018 China Family Panel Study (CFPS) shows that, among children aged 10–18 years in patriarchal rural China, girls consistently outperform boys by 5–10 percentiles (10%–20%) of class ranking. These differences are displayed in Appendix Figure A1. This has led to a reverse gender gap in educational attainment in a large part of the world (Murnane, 2013; Treiman, 2013). Both biological and social factors contribute to girls' better performance.

2.2. Parent–child separation

2.2.1. Left-behind children: the cause and status quo

In rapidly urbanizing China, the massive migration from rural to urban regions exceeded 300 million people in 2021. A huge number of migrant laborers, who are in their 20s and 30s, exhaust themselves during long shifts, live frugally in dormitories provided by their employers, and leave their young children behind in their hometowns. Such separation can be heartbreaking for both parents and children; however, it seems inevitable given the high living costs in big cities and institutional barriers such as the *hukou* system.

Hukou is a registered residency status that directly connects to the public services one can access. The modern *hukou* system dates back to 1958 when China completed its socialist transformation; it served as part of the central planning economic system to restrict rural–urban migration as urban workers received much higher social welfare support, and the prices of agricultural products were deliberately undervalued (Lin, 2011). This system originally registered each individual in their home location, and any migration without the government's permission was prohibited. The *hukou* system was relaxed but remained in effect after China's economic reform in the 1980s. People were allowed to migrate; however, in destination cities, non-registered residents were restricted from receiving public services, including public medical insurance and public education (Song, 2014; Sieg et al., 2021; Huang, 2022). Such services from private providers are rather expensive and it is extremely difficult for migrants to pay the premiums.

Aside from a lack of educational opportunities, children's cost of living is much higher in urban compared to rural areas. According to a report by the leading business media *Caixin* (YuWa Population Research, 2022), as of 2019, the average cost of rearing a child from age 0–17 years was 300,000 RMB (44,000 USD) in rural China but doubled to 630, 000 RMB (93,000 USD) in urban regions, accounting for about 43% of 18 years of total disposable income for a family of three.[2] Thus, a lack of educational opportunities and high living costs in urban areas result in many parents leaving their children behind in their hometowns when they migrate to cities for work. This produces a large group of disadvantaged children who are separated from their parents. According to a report by the United Nations International Children's Emergency Fund (UNICEF), the total number of left-behind children in China numbered around 70 million in 2015 (Huang, 2022).

2.2.2. Disadvantages for left-behind children

Left-behind children grow up in an unfriendly environment without parental companionship or supervision, which causes problems such as deficits in academic performance (Zhang et al., 2014; Meng and Yamauchi, 2017), nutritional intake (Wen and Lin, 2012; Mu and De Brauw, 2015), and mental health (He et al., 2012; Gao et al., 2015). Specifically, relevant to our study, Yue et al. (2016) highlight the extreme detrimental impact of separation during early childhood.

Using CFPS data, we observe a similar pattern of disadvantages for left-behind children. Table 1 shows that preschool-aged children separated from their mothers in 2010 developed worse cognitive and non-cognitive abilities by 2018 compared with those who remained with their mothers. Specifically, they had a lower class rank, lower scores on math and word tests administered by investigators, and a stronger tendency toward depression. The effects of separation from one's father are more limited, which is consistent with the fact that mothers are generally the primary caregivers.

3. Data, summary statistics, and graphical evidence

We use two datasets to explore gender differences during early childhood in the likelihood of parent–child separation and parental fertility behaviors, and one dataset for the back-of-the-envelope calculation. In this section, we describe our working sample, present summary statistics, and use raw data to show the graphical evidence of age-specific hazards of separation from one's parents.

3.1. Data and summary statistics

Our analysis of gender disparities in parent–child separation, particularly mother–child separation, uses a 0.35% sample from the 2010 population census, implemented by the National Bureau of Statistics. All individuals in China are expected to be interviewed regarding their basic demographic information and socioeconomic status, including jobs, fertility history, current residence, and registered residence (*hukou*). With respect to our research question, the main advantage of this dataset is its large sample size. We define families with either father or mother having agricultural *hukou* as rural families. We focus on children under school age: The 2010 sample includes 91,941 rural children who were under school age at the time of the survey.

The census surveys individuals who either reside at their registered address or in other places but are registered as a household member in the residency document (the *hukou* book). For those not living at their registered household address, the census inquires as to whether they reside in other counties, cities, or provinces. The census records each registered household member's relationship with the household head; thus, we can match children with their parents. Utilizing such

[2] The median disposable income per capita in China in 2019 was 26,523 RMB (National Bureau of Statistics, 2020).

Table 1
Comparison between children with and without the experience of parental separation.

	Separated from the mother			Separated from the father		
	Yes	No	Diff.	Yes	No	Diff.
	(1)	(2)	(3)	(4)	(5)	(6)
Class ranking	0.524	0.608	−0.084***	0.588	0.601	−0.013
	(0.301)	(0.280)	(0.005)	(0.289)	(0.282)	(0.579)
Math test score	6.278	8.045	−1.767***	6.715	8.135	−1.420***
	(7.433)	(6.324)	(0.002)	(7.330)	(6.207)	(0.003)
Word test score	15.986	18.161	−2.175**	16.557	18.262	−1.705**
	(12.694)	(10.865)	(0.030)	(12.301)	(10.752)	(0.036)
Depression	30.833	29.640	1.193*	30.362	29.633	0.729
	(7.874)	(6.813)	(0.056)	(6.801)	(7.023)	(0.151)
Self-control	62.901	62.397	0.504	63.287	62.211	1.075
	(13.771)	(18.185)	(0.758)	(13.703)	(18.694)	(0.417)
Locus of control	57.307	58.395	−1.088	58.102	58.290	−0.188
	(10.386)	(9.938)	(0.226)	(9.284)	(10.217)	(0.797)
Self-esteem	34.913	36.399	−1.485	37.470	35.806	1.663
	(37.386)	(37.558)	(0.659)	(37.486)	(37.545)	(0.543)
Trust	57.025	56.779	0.247	56.728	56.838	−0.111
	(13.178)	(14.594)	(0.849)	(13.328)	(14.723)	(0.916)

SOURCE—Longitudinal data from the 2010 and 2018 waves of the CFPS.
NOTE—The sample comprises children aged 0–5 years in rural families. The sample is divided into two groups according to whether children were left behind by their mothers or fathers in 2010. The variables capture cognitive and non-cognitive skills and were collected in 2018. The entries are the means. Standard deviations or *p*-values of *t*-tests (diff. columns) are reported in parentheses.
* significant at the 10% level.
** significant at the 5% level.
*** significant at the 1% level.

information, we identify parent–child separation; that is, we deem a situation to involve parent–child separation if the child resides at the registered address while the parent does not.[3] To focus on the "true essence" of parent-child separation, we exclude the circumstances of easy commuting or frequent visits, and regard children as being separated from a parent if the parent live in a different county.

To investigate gender differences in parental behaviors regarding fertility and migration, we restrict our sample in the main regression in the following manner. First, to avoid the potential issues of gender-selective abortions against female fetuses, we restrict the sample to firstborn children as several studies have documented the randomness of firstborn gender (Li, 2007; Ebenstein, 2010; Sun and Zhao, 2016). Second, in our main regression, we exclude six provinces that are under a restrictive one-child policy (Beijing, Tianjin, Shanghai, Jiangsu, Sichuan and Chongqing). Although, as the data indicates, approximately 40.5% of people in these provinces actually have more than one child, this rate is substantially lower than the 55.0% in other areas. We also exclude ethnic minorities and provinces with a large proportion (at least 20%) of minorities (Tibet, Xinjiang, Qinghai, Guangxi, Guizhou, Ningxia, Yunnan, and Inner Mongolia), which are usually subject to very different fertility policies. Third, because the return on human capital investment is highest in early childhood, and because decisions regarding migration and living arrangements are most likely associated with fertility behavior in the years following a woman's first birth, we focus on the preschool age (0–5 years old) of the firstborn. Fourth, as we are interested in mothers with the limited option of either remaining at home with their child or migrating without the child, in the main regression, we exclude households that are able to migrate with their children. Children in these households comprise only 2.8% of the 91,941 rural children under school age. We incorporate these households into the multi-logit analysis as part of the robustness check (see Section 7.2).[4]

Table 2 presents the summary statistics of our sample. As shown in Panel A, 15.1% of rural children aged 0–5 years are separated from at least one parent; 7.3% and 14.3% are separated from their mothers and fathers, respectively. Among all cases of parent–child separation, 97.3% of the parents are migrant workers who have left their children behind at their registered residential/*hukou* address.[5]

Table 2
Summary statistics of Children's living arrangement and (family) characteristics.

Variable	Mean	SD
	(1)	(2)
Panel A. Living arrangement		
Coresident with both parents	0.849	0.358
Separated from mother only	0.008	0.088
Separated from father only	0.078	0.268
Separated from both parents	0.065	0.246
Observations	18,435	
Panel B. Children's (family) characteristics		
Have sibling(s)	0.175	0.380
Mother's age	28.219	3.784
Father's age	29.460	4.317
Mother registered as a rural resident in the *hukou* system	0.976	0.153
Father registered as a rural resident in the *hukou* system	0.980	0.140
Mother's years of schooling	8.994	1.946
Father's years of schooling	9.344	1.826
Observations	18,435	

SOURCE— 0.35% sample from the 2010 China Population Census.
NOTE—The sample comprises 0–5-year-old firstborn children from rural Han families. Provinces under a restrictive one-child policy or with large minority groups are excluded. Children migrating with parents are also excluded. Panel A includes variables related to the current residency status of children (whether they are separated from their parents or not). Panel B presents individual and household characteristics.

[3] In a small portion of households, the parents reside at the registered address, but the children do not. In this circumstance, some children board with or are adopted by relatives. Such households amount to only 0.1% of the sample.

[4] We also exclude children whose parents are remarried, divorced or widowed in our working sample.

[5] If the parent does not live in the registered residency address, census further asks the reason on such separation. We specify a parent as a migrant worker if the answer is "*Going to another place for business or work*".

Table 2 Panel B displays the summary statistics of some individual and household characteristics. In our sample, 17.5% of the firstborn children aged 0–5 years have at least one sibling. In our defined rural families, the parents are predominantly registered as rural residents in the *hukou* system, with an average of around nine years of education; the mothers and fathers have mean ages of 28.2 and 29.5 years, respectively.

In addition to the 2010 population census, we use two supplementary datasets. First, we use the China Migrant Dynamic Survey (CMDS) to explore the mechanism driving gender differences in separation. The CMDS is a repeated cross-sectional survey on migrant households implemented by the National Health Commission from 2009 to 2021. Second, we employ the CFPS to investigate the relationship between parent–child separation and human capital outcomes. The CFPS is a nationally representative, longitudinal survey of Chinese communities, families, and individuals (Xie and Hu, 2014); it contains several measurements of human capital for children above age 10. We describe these two datasets in more detail in Sections 4 and 8, respectively.

3.2. Graphical evidence

For this subsection, we compare the age-specific means of the likelihood of mother–child separation and that of mothers being migrant laborers for boys versus girls in rural Han families in China. We calculate the likelihood of mother–child separation as the ratio of the number of children whose mothers live in a different county (from the child's place of residence), to the total number of children. The likelihood of a mother being a migrant laborer is the ratio of the number of children whose mothers work or run a business in a different county (from the child's place of residence), to the total number of children.

Panels (a) and (b) of Fig. 1 show the age-specific likelihood of mother–child separation and a mother being a migrant laborer, respectively, in early childhood, after controlling for household characteristics, including parents' ages (and the quadratic terms), the dummy indicator of rural residency in the *hukou* system, and parents' years of schooling. In both figures, the dashed and solid lines represent girls and boys, respectively.

Fig. 1 suggests three empirical patterns. First, Panel (a) presents an extremely low propensity of separation or maternal migration in the year of childbirth. There is no gender difference in separation at this point. Second, the likelihood in both figures increases substantially for both boys and girls at the age of 1, which is generally when the lactation period ends. At age 2, the likelihood of separation and parental migration for boys apparently exceeds that for girls. Third, the gender difference in separation gradually disappears as the children grow older. Thus, we can see an interesting pattern where the gender gap in mother–child separation and a mother being a migrant worker seems exceedingly pronounced at around age 2.

These findings are consistent with our premise that, in the context of son preference, parents will want a second child if their firstborn child is a girl. During the pregnancy and delivery of the second child, the mother will stay in her hometown with her firstborn child. Specifically, we observe the gender gap appearing at around age 2 of her firstborn child, and this is mainly attributed to higher levels of separation among households with boys at the age of 2 compared to other ages.

4. Gender differences in parent–child separation

To carefully quantify the magnitude and significance levels, we use regression analysis to examine the details of gender differences (see Section 4.1). We explore the mechanism of fertility-caused migration delay (see Section 4.2) and derive supporting evidence (see Section 4.3).

4.1. Empirical strategy and regression evidence

For the benchmark analysis, we run the following DID regression to estimate the gender differences of interest across age groups:

$$y_{icl} = \alpha + \sum_{\tau=1}^{5} \beta_\tau I(l_{icl} = \tau) \times Boy_{icl} + \sum_{\tau=1}^{5} \delta_\tau I(l_{icl} = \tau) + Boy_{icl} \bullet \mu + X_{icl} \cdot \varphi + \gamma_c + \varepsilon_{icl} \quad (1)$$

The subscripts *i, c,* and *l* denote the individual child, prefecture (city), and child's age in 2010, respectively. The outcome variable y_{icl} indicates whether child i in city c of age l lives separately from a parent. Specifically, for child i, $y_{icl} = 1$ if a parent lives in a different county from the child's residence place; $y_{icl} = 0$ if the parent and the child live at the same place. Boy_{icl} is a dummy indicator for whether the firstborn child i is a boy. The coefficient of it is the gender fixed effect and represents the gender difference at age 0. In other words, if gender results in varying risks of becoming left-behind children at all ages, including even in the year of birth, then this age-invariant difference should be captured by μ. $I(l_{icl} = \tau)$ denotes whether the firstborn child is at age τ in 2010.

We also interact the boy dummy with each age dummy for $\tau = 1$–5, respectively. δ_τ is the age-specific separation hazard for girls. Under the assumption that the firstborn gender is random, coefficient β_τ captures the age-variant gender differences regarding the hazard of parent–child separation for the birth cohort aged τ in 2010 for $\tau = 1, \ldots, 5$, respectively. Vector X_{icl} includes control variables such as parents' ages (and their quadratic terms), the dummy indicator of rural residency in the *hukou* system and years of schooling. γ_c is the prefecture fixed effect. Standard errors are clustered at the county level.[6]

We discuss three particulars regarding the abovementioned specification in Section 5: the randomness of gender, differentiating between age effects and birth cohort effects, and the interpretation of the coefficients of interactions as the age-varying gender difference.

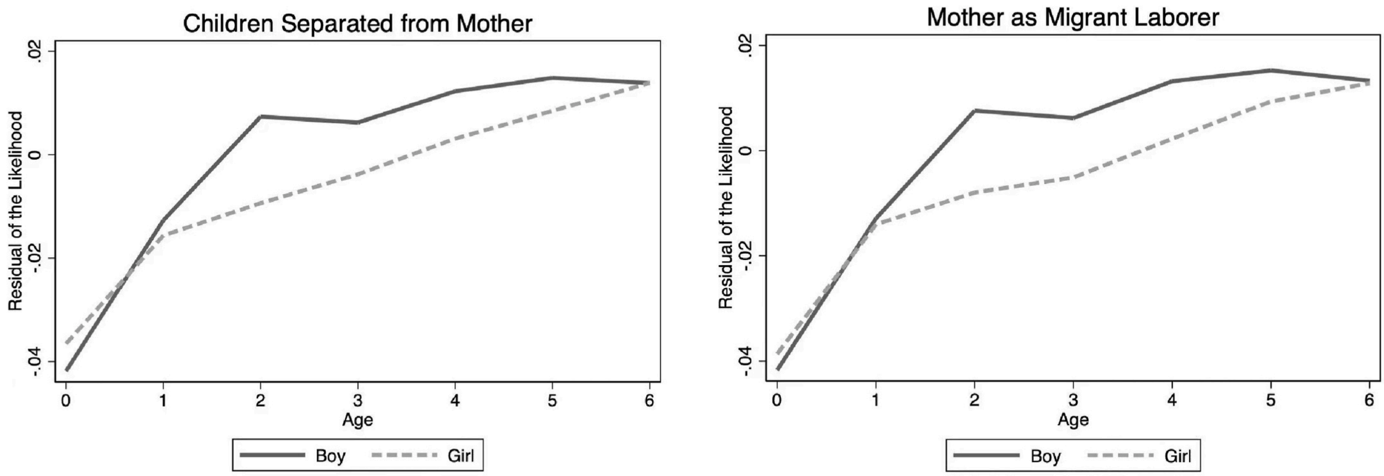
4.1.1. The hazard of parent–child separation

Table 3 displays the results for age-specific gender differences in terms of the parent–child separation hazard. Columns (1)–(3) show the age-variant gender differences in the likelihood of a child separating from their mother, father, and both parents, respectively. The gender difference of mother–child separation is close to zero at baseline (age 0). Given that $\mu \approx 0$, as seen in Table 3, we use the "age-variant gender differences" and "the total gender gap at each age" interchangeably as the interpretation of β_τ. We will discuss and differentiate the interpretations in detail in Section 5.3.

The gender difference remains very small at age 1, which is in line with the conjecture that the mother, having given birth in her hometown, would spend some time recovering and breastfeeding before migrating for work, regardless of the child's gender. This is also in line with the pattern in Fig. 1(a), which shows that the propensities for mother–child separation for boys and girls closely track each other from 0 to 1 years of age.

However, at age 2, boys are more likely to live separately from both parents. Specifically, the estimates in Column (1) of Table 3 indicate that boys are 3.1 percentage points more likely to be separated from their mothers than girls at age 2. This is equivalent to a 44.9 percent increase in the probability of separation given that the average probability of girls being left behind at this age is 6.9%. Columns (2) and (3) demonstrate that 2-year-old boys are 3.5 and 2.8 percentage points more likely to be separated from their father and from both parents, respectively, compared to girls of the same age. In Column (2), while the magnitude of

[6] The population census covers all residents in mainland China and does not have a stratified sampling design. Thus, based on the recommendations of Abadie et al. (2023), it is not necessary to cluster the standard error at a level that is too high. We choose to cluster it at the county level to be more conservative. There were approximately 2856 counties in China in 2010. Their average area is 4000 km^2 and their average population is 480,000.



(a) Children separated from the mother (b) The mother is a migrant laborer

Fig. 1. Residual of the likelihood of mother-child separation and that of mother being migrant laborer against children's age. SOURCE—0.35% sample from the 2010 China Population Census. The sample comprises 0–6-year-old children from rural Han families. Children migrating with parents are also excluded. Provinces under a restrictive one-child policy or with large minority groups are excluded. The horizontal axis is the children's age. In Subfigure (a), the vertical axis represents the residual of the likelihood of mothers currently residing in a different county. In Subfigure (b), the vertical axis denotes the residual of the likelihood of mothers currently migrating for work for children of certain age. The residuals are obtained from a regression of the dummy indicator of separation (migration) on household characteristics, including parents' ages (and the quadratic terms), the dummy indicator of rural residency in the *hukou* system, and parents' years of schooling.

Table 3
Age-specific gender differences in the probability of parental separation.

	Separated from the mother	Separated from the father	Separated from both parents
	(1)	(2)	(3)
Boy	−0.005	−0.007	−0.003
	(0.008)	(0.013)	(0.007)
Boy × Age 1	−0.006	−0.004	−0.007
	(0.012)	(0.017)	(0.011)
Boy × Age 2	0.031***	0.035**	0.028**
	(0.012)	(0.017)	(0.011)
Boy × Age 3	0.010	0.007	0.005
	(0.012)	(0.017)	(0.011)
Boy × Age 4	0.005	0.000	0.000
	(0.012)	(0.017)	(0.011)
Boy × Age 5	0.011	0.009	0.009
	(0.012)	(0.017)	(0.011)
Observations	18,435	18,435	18,435
R^2	0.122	0.153	0.116

SOURCE—0.35% sample from the 2010 China Population Census.
NOTE—The sample comprises 0–5-year-old firstborn children from rural Han families. Provinces under a restrictive one-child policy or with large minority groups are excluded. Children migrating with parents are also excluded. The outcomes are the dummy indicators of whether the child is currently separated from the mother (Column (1)), the father (Column (2)), or both parents (Column (3)). The control variables are parents' ages (and the quadratic terms), the dummy indicator of rural residency in the *hukou* system, parents' years of schooling, and city-level fixed effects. Standard errors are clustered at the county level.
* significant at the 10% level.
** significant at the 5% level.
*** significant at the 1% level.

the point estimate seems slightly greater than that in Column (1), it amounts to a much smaller percent (25.9 percent) increase, as the sample mean of father–child separation at age 2 is 13.5%. The positive gender difference persists at ages 3 and 4 but fades with no statistical significance.

The age-specific likelihood of separation is visualized in Fig. 2. Panels (a) and (b) depict the point estimates and the 90% confidence intervals of the regression coefficients presented in columns (1) and (2) of Table 3. As shown in the two figures, there is a notable increase in the estimates at age 2. Conversely, we do not observe a statistically significant gender gap at any other age.

4.1.2. *Parental migration for work*

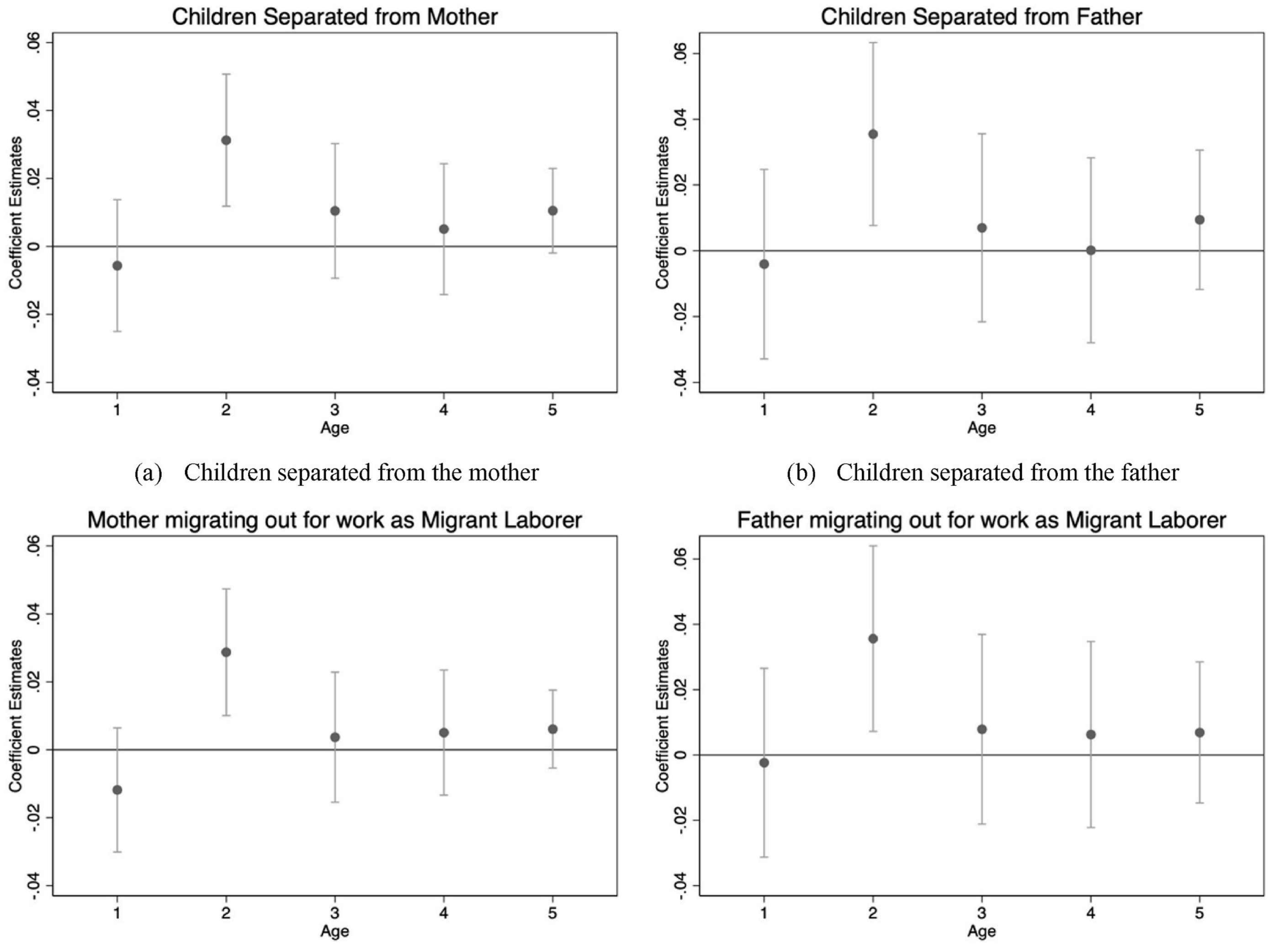
We redefine $y_{icl} = 1$ if a child's parent does not live at the registered residence address because of migration for work or business, and $y_{icl} = 0$ if otherwise. All other features are the same as in regression (1). In our sample, 97.3% of separations result from parents migrating for work and leaving their children behind; thus, we expect the pattern of the age-specific gender differences in the likelihood of parents' work-induced migration to be consistent with what is presented in Table 3.

The results are reported in Table 4. For boys of 2 years of age, mothers and fathers are 2.9 and 3.6 percentage points more likely to migrate for work, respectively. The signs of the coefficients remain consistently positive between ages of 3 and 5 years; however, the magnitude declines and becomes statistically insignificant. This pattern is similar to what is shown in Table 3, which is in line with the fact that most separations during childhood are driven by parents' work-related migration. Together with Fig. 1, the diminished gender difference actually indicates a rather bleak picture of children's living environment in rural China: Regardless of the child's gender, after the firstborn child attains schooling age, over 9% of mothers and 15% of fathers eventually leave their children behind to seek better work opportunities.

Panels (c) and (d) of Fig. 2 depict the point estimates and 90% confidence intervals of the regression coefficients displayed in Table 4. A salient feature is the upward jump of the point estimates at age 2 shown in the two figures, whereas we detect no statistically significant gender gap at any other age. This pattern aligns with that presented in panels (a) and (b) of Fig. 2, implying a close relationship between the parent-child separation and the parents' migration.

4.2. *Mechanism: fertility-caused migration delay*

We use the CMDS to examine the mechanism underlying the abovementioned patterns; it covers multiple dimensions of migrant families, especially birth history. The birth history module includes the time of each birth and the child's gender. In addition, the 2012 wave particularly provides specific information on the locations where



(a) Children separated from the mother (b) Children separated from the father

(c) The mother is a migrant laborer (d) The father is a migrant laborer

Fig. 2. The estimate of the age-specific gender differences in the likelihood of parent-child separation (parent being migrant labor). SOURCE—0.35% sample from the 2010 China Population Census. The sample comprises 0–5-year-old firstborn children from rural Han families. Provinces under a restrictive one-child policy or with large minority groups are excluded. Children migrating with parents are also excluded. The horizontal axis indicates the children's age. The vertical axis is the estimate of the gender differences in the likelihood of child-mother separation and child-father separation for Subfigure (a) and Subfigure (b), and is that of the gender differences in the likelihood of mother and father being migrating for work or business for Subfigure (c) and Subfigure (d). Age 0 is the reference group.

mothers stayed during their pregnancies and the times of giving birth. Therefore, we utilize the 2012 wave data for the purpose of our analysis.

Using the 2012 CMDS data, we depict the ratio of the number of families who have second children to the total number of families that already have one child. We plot the ratio by birth interval and firstborn gender. Two features are quite salient in Fig. 3(a). First, Chinese parents are, on average, 11.9 percentage points (110.7%) more likely to give birth to a second child when their firstborn is a girl. This finding is consistent with much of the prior research documenting the impact of son preference on fertility (Ebenstein, 2010). Second, the peak timing for migrant parents to give birth to a second child is 2 years after the first birth. This peak of a 24-month interval is not a new concept in China and has been well-documented among different populations worldwide (United Nations, 1976; Wenlock, 1977; Liu and Zou, 2011; Buckles and Munnich, 2012; Zhang et al., 2019). The biological explanation was first proposed by Potter (1963), who dissects birth spacing into five components[7] and concludes that the average birth interval ranges between 18 and 27 months. Fig. 3(a) displays an obvious peak at 24 months for households with firstborn boys and those with girls.[8]

For the results presented in Fig. 3(b), we perform the same exercises using the sample of families with migrant parents from the 2010 population census. We find very similar patterns to those reflected in Fig. 3 (a).

Migrant laborers with registered residences in rural areas generally give birth in their hometowns. The high medical costs in cities are likely to play a critical role in this decision. As described in Section 2.2.1, the dual *hukou* system makes rural residents ineligible for urban health insurance. Medical expenses incurred in cities, including prenatal care and

[7] The components are: pregnancy, post-partum amenorrhea, anovulatory cycles, "the time required for conception after ovulation is established," and an additional period to account for pregnancy wastage.

[8] Families with firstborn girls may want to rush to have a son as soon as possible to promote their reputation in their family and clan (Zhao and Li, 2018).

Table 4
Age-specific gender differences in labor-oriented parental migration.

	The mother is a migrant laborer	The father is a migrant laborer	Both parents are migrant laborers
	(1)	(2)	(3)
Boy	−0.001	−0.008	0.001
	(0.007)	(0.013)	(0.006)
Boy × Age 1	−0.012	−0.002	−0.012
	(0.012)	(0.017)	(0.011)
Boy × Age 2	0.029***	0.036**	0.025**
	(0.011)	(0.018)	(0.011)
Boy × Age 3	0.004	0.008	−0.002
	(0.011)	(0.017)	(0.010)
Boy × Age 4	0.005	0.006	0.000
	(0.012)	(0.018)	(0.011)
Boy × Age 5	0.006	0.007	0.005
	(0.011)	(0.017)	(0.010)
Observations	18,435	18,435	18,435
R^2	0.124	0.145	0.116

SOURCE—0.35% sample from 2010 the China Population Census.
NOTE—The sample comprises 0–5-year-old firstborn children from rural Han families. Provinces under a restrictive one-child policy or with large minority groups are excluded. Children migrating with parents are also excluded. The outcomes are the dummy indicators of whether the mother (Column (1)), the father (Column (2)), or both parents (Column (3)) currently migrate for work. The control variables are parents' ages (and the quadratic terms), the dummy indicator of rural residency in the *hukou* system, parents' years of schooling, and city-level fixed effects. Standard errors are clustered at the county level.
* significant at the 10% level.
** significant at the 5% level.
*** significant at the 1% level.

childbirth, can only be reimbursed in registered residential counties through rural residents' insurance (i.e., rural cooperative medical services). The reimbursement rate is quite low for expenses incurred in regions outside of one's registered residence. To illustrate the discriminatory medical reimbursement system, Yi and Gu (2015) surveyed the reimbursement rate in multiple provinces for inpatient costs in the registered residential county, outside the county but within the registered province, and outside the registered province. Taking Henan (which is one of the largest family provinces of migrant workers) as an example, from 2007 to 2011, the inpatient reimbursement rate in the locality was raised from 50% to 75%, while the rate for expenditures outside the *hukou* province remained between 30% and 35%. In addition to the low reimbursement rate, the reimbursement process itself is quite inconvenient for migrants. Usually, not only must they personally pay all costs in the city first but they are also required to go through a complicated bureaucratic procedure to receive reimbursement from the government in their hometowns. This significantly increases the cost of giving birth in locations other than one's hometown (Huang and Wu, 2020).

Using CMDS data, Table 5 suggests that 64.2% of migrant mothers stay in their hometowns throughout their pregnancy, and another 2.7% stay for most of their pregnancy; 80.9% give birth in their hometowns. This means that pregnant women postpone migration or return to their hometowns when they are pregnant and preparing for childbirth.

The fertility behavior of migrant parents, especially mothers, has clear implications for their migration and residential arrangements. Combining the evidence on the likelihood, timing, and location of a mother's second experience of childbirth, we argue that girls are more likely than boys to have a younger sibling before reaching the age of 2, which means they experience more companionship from their parents, especially their mothers, in early childhood. Thus, we uncover an intriguing paradox in which son preference increases the likelihood that

Table 5
Birth and pregnancy locations for migrant mothers.

Variable	Frequency in the sample	Proportion
	(1)	(2)
Panel A. Pregnancy location		
Mainly the place of migration	13,482	13.7%
Place of migration	19,025	19.4%
Mainly hometown	2678	2.7%
Hometown	63,130	64.2%
Panel B. Birth place		
Place of migration	15,933	16.2%
Hometown	79,557	80.9%
Other places	2825	2.9%

SOURCE—The 2012 wave of the CMDS.
NOTE—The sample comprises rural Han families. Provinces under a restrictive one-child policy or with large minority groups are excluded. Panel A shows the frequency and proportion of each location where mothers resided during pregnancy. Panel B presents the frequency and proportion of each location where mothers resided at the time of giving birth.

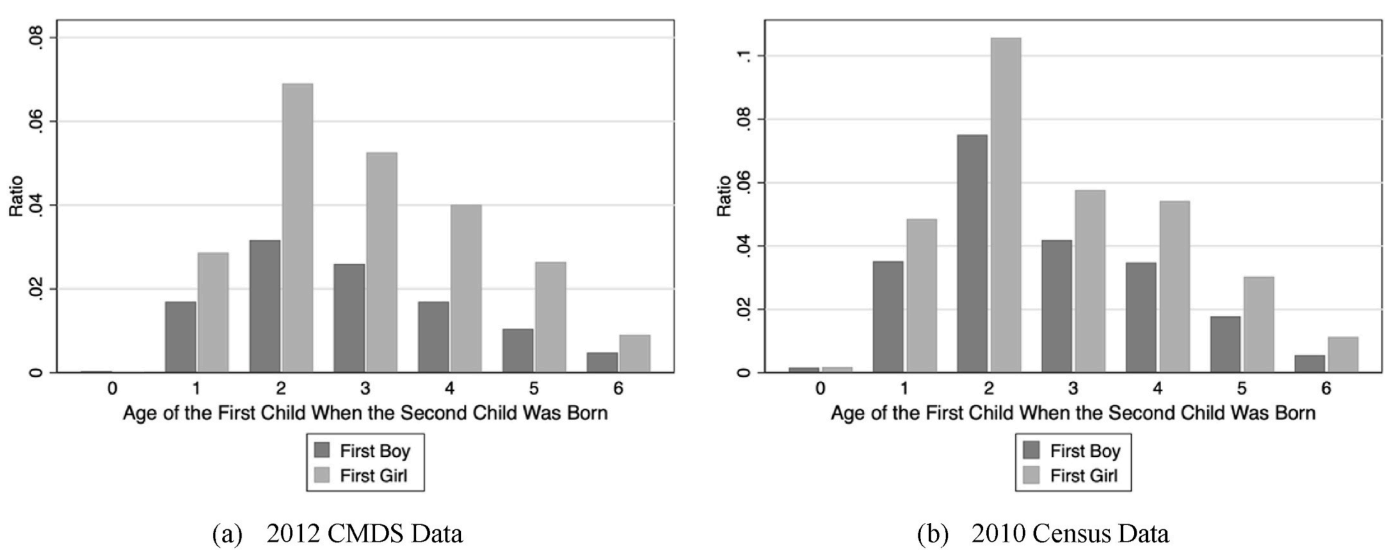


Fig. 3. Ratio of the number of households giving birth to a second child to the total number of households with at least one child. Subfigure (a) uses data from the 2012 wave of the CMDS, and Subfigure (b) uses the 0.35% sample from the 2010 China Population Census. Both samples consist of rural Han families with migrant parents, excluding province of origin under a restrictive one-child policy or with large minority groups. The dark bars indicate households with firstborn boys, while the light bars represent households with firstborn girls.

firstborn girls will live in a better environment owing to increased interaction with and companionship from their parents.

4.3. More supporting evidence of the mechanism

4.3.1. The "Age-2 effect"

Migration decisions may differ between the parents of boys and girls in general, which we discuss in detail in Section 7.3. However, the existing explanations in previous studies do not suggest a peak in migration and separation, particularly at age 2, of firstborn children. In this subsection, we further establish the mechanism of fertility-driven delay in migration by highlighting the fertility-caused "age-2 effect." Essentially, we conduct two placebo tests. We find that when the timing of a second childbirth is restricted, or when the second birth is disallowed by China's family planning policy, the "age-2 effect" of residential separation and parental migration disappears.

Provinces with Birth Interval Restrictions. The peak of the 2-year birth interval depicted in Fig. 3 aligns perfectly with the high mother–son separation rate at 2 years of age found in the benchmark regression. However, if parents cannot plan births at will, no significant gender differences in mother–child separation will be observed at any specific age below that specified as the minimum required birth spacing.

We refer to the variations in the restrictions on birth spacing in China's family planning policies. In the 1980s and 1990s, the stipulations on the interval between the first and second birth were generally imposed in three ways. First, in some provinces, only the minimum interval was stipulated and had to be at least 4 years (e.g., Gansu, Hubei, and Jiangxi). Second, some provinces stated that the minimum age at which women could have a second child should be at least 28 (e.g., Jiangsu, Shandong, and Liaoning). Third, in provinces such as Hebei, Henan, and Hunan, both requirements were applied. Such stringent requirements have since been relaxed in some provinces with the promulgation of the Population and Family Planning Law in 2002. Prior restrictions on birth spacing have been changed or entirely abolished in a new round of revisions to local population and family planning regulations. By 2010, 14 provinces (i.e., Jilin, Hainan, Shanghai, Gansu, Xinjiang, Zhejiang, Hunan, Inner Mongolia, Hubei, Guangdong, Shanxi, Shaanxi, Jiangxi, and Guizhou) had abolished any requirements on birth spacing. We regard provinces where the requirements have been completely abolished as the group with no birth interval restrictions. The remaining sample is the reference group, which is subjected to at least some requirements.

As shown in Fig. 4(a), in regions where parents can freely decide on the timing of the first two births, the birth interval distribution exhibits a striking peak at 2 years. In contrast, as shown in Fig. 4(b), the distribution appears much flatter in provinces with birth interval restrictions. The regression results using the two subsamples are reported in Table 6. As expected, Column (1) indicates that the statistically significant higher rate of mother–son separation at age 2 only appears in regions without birth interval constraints, and the magnitude is larger than that of the benchmark regression using the whole sample. Likewise, the gender differences in the propensity of father–child separation and of separation from both parents at age 2 are significant when no birth interval constraints are imposed, as seen in Columns (3) and (5), respectively. In contrast, in provinces with constraints on the timing of the second birth, the birth interval peak is flattened, and gender differences in separation at age 2 are weakened, as presented in Columns (2), (4), and (6).

Provinces with a Strict One-Child Policy. Provinces under a strict one-child policy will not be affected by the channel of fertility-caused migration delay, because a second birth is not allowed. We exclude such provinces from the benchmark regressions. However, these provinces could serve as a natural placebo group in our analysis.

Table 7 illustrates the age-specific disparities between genders in the likelihood of parental separation in provinces that strictly enforce the one-child policy. The results in Column (1) do not reveal any gender gap for mother–child separation at age 2. Similarly, we do not detect any significant gender differences in father–child separation or separation from both parents in Columns (2) and (3), respectively.

4.3.2. Beyond age 2

In this section, we use variations in the birth interval in addition to the peak of 2 years. Although having a younger sibling may be endogenous, the exact timing is more likely to be random. We leverage the exogenous component of the birth interval to deepen our understanding of the delay in migration caused by fertility. Fig. 5 utilizes data from the CMDS. Each dot in Fig. 5 represents the mean starting time of mothers' most recent migration for each birth interval. We observe a clear, positive correlation between the starting time of maternal migration (age of the first child at the time of her most recent migration experience) and the interval between the first and second birth (age difference between the first and second child). This demonstrates that mothers who experience longer birth intervals prior to the arrival of their second child defer migration, thus supporting the argument that birth spacing plays a crucial role in determining the timing of mother–child separation.

5. Identification and interpretation particulars

In this section, we discuss the randomness of gender, differentiation between age effects and birth cohort effects, and the interpretation of the coefficients of interactions as the age-specific gender difference in separation.

5.1. The randomness of gender

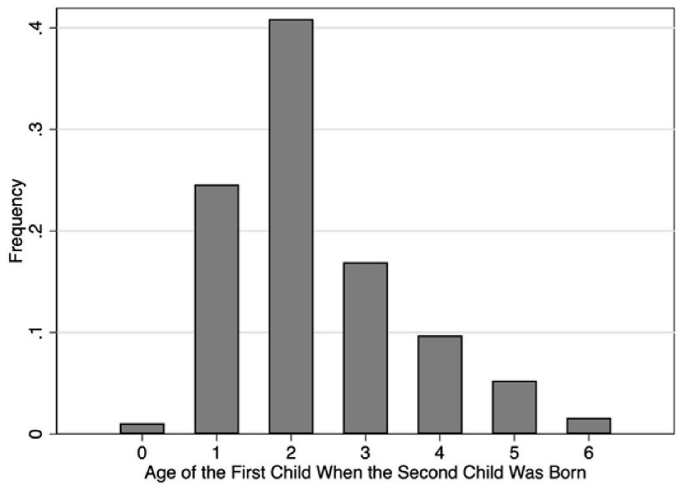
If gender is not random, the results reported in Tables 3 and 4 could be confounded by factors other than the firstborn child's gender. For example, parents with a stronger son preference could select against female fetuses (Chen et al., 2013). If such parents are less educated (Chung and Gupta, 2007), they will likely incur higher costs and presumably gain fewer benefits from migration. Although this particular possibility only biases the difference downward, we hope that gender is close to being randomly assigned because other potential unobservable factors associated with son preference could be erroneously omitted, causing ambiguous bias. We probe this random gender assignment from two perspectives.

First, well-established evidence shows that the firstborn child's gender is random, and gender selections are generally performed at higher parities. Ebenstein (2010) investigates samples from China's population census in 1982, 1990, and 2000, and finds the sex ratio of firstborn children in China to be perfectly balanced. Examining China's 2010 census data, we note that the firstborn sex ratio (i.e., the ratio of the number of firstborn boys to the number of firstborn girls), is 105.2. This ratio falls within the range of a natural sex ratio at birth, which is 105.0–107.0 (Banister, 2004). Additionally, we exclude regions subject to a stringent one-child policy. By restricting the sample in this manner, we focus on the sample without a strong incentive to prenatally select the fetus's gender because parents can always have a second child if the firstborn is a girl.

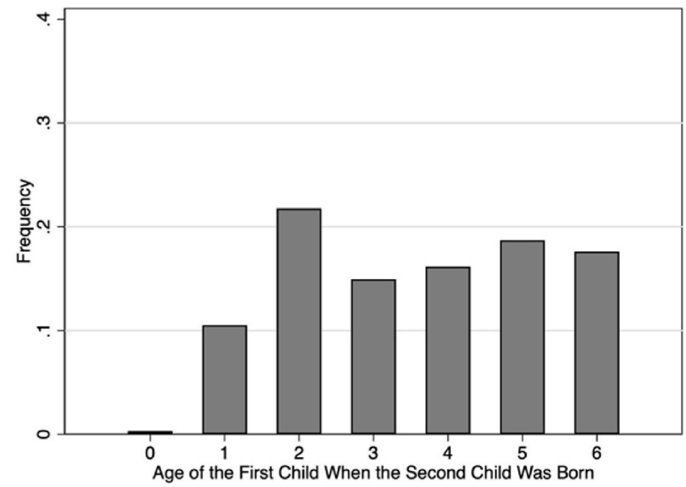
Second, we implement a balance test by regressing the gender of firstborn children on family demographic factors, including parents' ages, parental education level, and whether the family members are registered as rural residents in the *hukou* book. We also control for the fixed effects of city and children's age. As Table 8 indicates, the firstborn gender is uncorrelated to any household characteristics when fitting the data with both the ordinary least squares (OLS) and probit models. Thus, we infer that gender is close to randomly assigned, at least to the extent that all observable factors available in the census are balanced between subsamples of different genders.

5.2. Age versus birth cohort effects

When using cross-sectional data to identify age-specific effects, an



(a) Without birth interval restrictions



(b) With birth interval restrictions

Fig. 4. Histogram of spacing (years) between first and second births. SOURCE—0.35% sample from the 2010 China Population Census. The sample includes rural Han families, excluding provinces under a restrictive one-child policy or with large minority groups. Subfigure (a) employs subsamples from provinces with no birth interval restrictions, while Subfigure (b) uses subsamples from provinces with some birth interval restrictions.

Table 6
Placebo test 1: Age-specific gender differences across regions with and without birth interval restrictions.

	Separated from the mother		Separated from the father		Separated from both parents	
	Without restrictions	With restrictions	Without restrictions	With restrictions	Without restrictions	With restrictions
	(1)	(2)	(3)	(4)	(5)	(6)
Boy	−0.011	0.001	−0.020	0.000	−0.006	−0.000
	(0.015)	(0.008)	(0.022)	(0.016)	(0.013)	(0.007)
Boy × Age 1	−0.003	−0.009	−0.017	0.007	−0.011	−0.006
	(0.025)	(0.012)	(0.030)	(0.021)	(0.022)	(0.012)
Boy × Age 2	0.060**	0.011	0.070**	0.015	0.053**	0.011
	(0.024)	(0.012)	(0.031)	(0.021)	(0.023)	(0.012)
Boy × Age 3	0.047	−0.000	0.045	−0.004	0.029	−0.001
	(0.031)	(0.012)	(0.036)	(0.020)	(0.028)	(0.011)
Boy × Age 4	−0.010	0.003	0.006	−0.005	−0.008	−0.001
	(0.030)	(0.012)	(0.038)	(0.020)	(0.028)	(0.012)
Boy × Age 5	−0.028	0.008	0.025	0.003	−0.031	0.008
	(0.033)	(0.012)	(0.046)	(0.020)	(0.029)	(0.011)
Observations	4551	13,883	4551	13,883	4551	13,883
R^2	0.101	0.141	0.133	0.167	0.098	0.135

SOURCE—0.35% sample from the 2010 China Population Census.
NOTE—The sample comprises 0–5-year-old firstborn children from rural Han families. Provinces under a restrictive one-child policy or with large minority groups are excluded. Children migrating with parents are also excluded. The outcomes are the dummy indicators of whether the child is currently separated from the mother (Columns (1) and (2)), the father (Columns (3) and (4)), or both parents (Columns (5) and (6)). For each outcome, the regression uses subsamples without and with restrictions on the spacing between the first and second birth, respectively. The control variables are parents' ages (and the quadratic terms), the dummy indicator of rural residency in the *hukou* system, parents' years of schooling, and city-level fixed effects. Standard errors are clustered at the county level.
* significant at the 10% level.
** significant at the 5% level.
*** significant at the 1% level.

assumption of homogeneous birth cohorts must be imposed. Without tracking an individual over time, we cannot differentiate between the effects of a specific age or birth year. We utilize the longitudinal data of the CFPS from 2010 to 2018 to fit Equation (1). The panel structure enables us to control for the gender-specific birth cohort effect by including birth year dummies and the interactions of such dummies with the indicator of the child's gender. In addition, we control for city-level fixed effects. The standard errors are clustered at the county level.

The results are reported in Table 9. Boys are more likely to be separated from their parents at age 2; the estimates are significant at the 1% and 5% levels, after controlling for the birth year fixed effects. The point estimates are larger than those in the benchmark regression using the census data. This may be because separation is more generally defined in the CFPS: The CFPS only asks if parents currently reside at the registered address; if not, it does not further ask whether they live in a different county or simply in a different house on the same street. Thus, separation could actually include cases in which commuting and frequent visits are feasible. As such, despite the larger magnitude of the point estimates, the results in Table 9 effectively differentiate between age and birth cohort effects, indicating that parental migration is more likely to occur when the child reaches a certain age instead of being driven by birth year-specific effects.

5.3. *Interpretation of the gender differences*

In the DID design specified as Equation (1), we interpret β_τ as the age-varying gender difference at age τ if assuming little gender difference in separation during the year of childbirth. Tables 3 and 4 both

Table 7
Placebo test 2: Age-specific gender differences in the probability of parental separation under a restrictive one-child policy.

	Separated from the mother	Separated from the father	Separated from both parents
	(1)	(2)	(3)
Boy	0.010	0.033	0.005
	(0.031)	(0.045)	(0.029)
Boy × Age 1	−0.018	−0.023	−0.017
	(0.041)	(0.060)	(0.037)
Boy × Age 2	−0.002	0.014	−0.003
	(0.048)	(0.060)	(0.044)
Boy × Age 3	−0.031	−0.066	−0.014
	(0.044)	(0.059)	(0.043)
Boy × Age 4	0.025	−0.052	0.022
	(0.046)	(0.057)	(0.045)
Boy × Age 5	−0.029	−0.069	−0.035
	(0.042)	(0.055)	(0.040)
Observations	2286	2286	2286
R^2	0.113	0.164	0.109

SOURCE—0.35% sample from the 2010 China Population Census.
NOTE—The sample comprises 0–5-year-old firstborn children from rural Han families and is confined to six provinces with restrictive one child policy (Beijing, Tianjin, Shanghai, Jiangsu, Sichuan and Chongqing). Children migrating with parents are excluded. The outcomes are the dummy indicators of whether the child is currently separated from the mother (Column (1)), the father (Column (2)), or both parents (Column (3)). The control variables are parents' ages (and the quadratic terms), the dummy indicator of rural residency in the *hukou* system, parents' years of schooling, and city-level fixed effects. Standard errors are clustered at the county level.
* significant at the 10% level.
** significant at the 5% level.
*** significant at the 1% level.

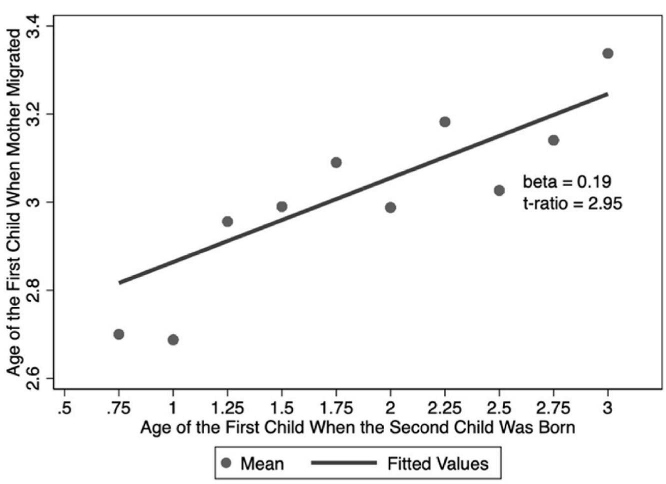


Fig. 5. Association between the age of the first child when the mother migrated and when the second child was born. SOURCE—The 2012 wave of the CMDS. The sample includes rural Han families, excluding provinces under a restrictive one-child policy or with large minority groups. The dots indicate the mean start time of mothers' most recent migration for each birth interval, and the line is a simple linear fit of the raw data.

report a slightly and insignificantly higher propensity for parents to separate from girls at the baseline age of zero.

Some readers may doubt the significance of the age-varying gender difference. After all, the general difference between 2-year-old boys and girls is the sum of the gender effect (μ) and the interaction term (β_2). If the fixed effect (μ) is negative, a potential concern would be whether the gender difference at age 2 is sustained as statistically significant after totaling the difference at age 0. Hence, we use a more general specification as a robustness check, as described in the following:

Table 8
Balance tests: Regression of the firstborn Child's gender on household characteristics.

	Dependent variable: Having a firstborn son	
	Linear probability model	Probit model
	(1)	(2)
Mother's age	−0.000	−0.000
	(0.002)	(0.002)
Father's age	−0.001	−0.001
	(0.001)	(0.001)
Mother registered as a rural resident in the *hukou* system	0.000	0.000
	(0.002)	(0.002)
Father registered as a rural resident in the *hukou* system	−0.002	−0.002
	(0.002)	(0.002)
Mother's years of schooling	−0.000	−0.000
	(0.002)	(0.002)
Father's years of schooling	−0.001	−0.001
	(0.001)	(0.001)
P-value of the F-test	0.701	0.687
Observations	18,435	18,435
R^2/PseudoR^2	0.014	0.010

SOURCE—0.35% sample from the 2010 China Population Census.
NOTE—The sample comprises 0–5-year-old firstborn children from rural Han families. Provinces under a restrictive one-child policy or with large minority groups are excluded. Children migrating with parents are also excluded. Controls include age and city-level fixed effects, respectively. Columns (1) and (2) apply a linear probability model and a probit model, respectively. F-tests are conducted to test whether these characteristics are jointly significant, and the *p*-values of the F-tests are reported.
* significant at the 10% level.
** significant at the 5% level.
*** significant at the 1% level.

$$y_{icl} = \alpha + \sum_{\tau=0}^{5} \theta_\tau I(l_{icl} = \tau) \times Boy_{icl} + \sum_{\tau=1}^{5} \delta_\tau I(l_{icl} = \tau) + X_{icl} \cdot \varphi + \gamma_c + \varepsilon_{icl} \quad (2)$$

In Equation (2), we replace the dummy indicator of Boy_{icl} with the interaction of $I(l = 0) \times Boy_{icl}$. All other features are the same as in Equation (1).

The advantage of this equation is that it allows the coefficients of the interaction terms to directly reflect the overall gender difference in the probability of separation from one's parents at each age from 0 to 5 years; however, the drawback is the inability to isolate the baseline age-invariant fixed effect of child gender. Such baseline fixed effect could be sensible in the context of son preference. For example, parents may generally feel more reluctant to leave sons behind than daughters. If we focus on the proposed mechanism (i.e., that child-age-varying household fertility behavior drives the gender difference), the coefficient β_τ in Equation (1) would be more informative. However, if we are interested in the magnitude of the overall gender difference, θ_τ in Equation (2) would be clearer and more direct.

The results of estimating Equation (2) are reported in Tables A1 and A2 of the Appendix. The magnitude of the age-specific gender difference at age τ ($\tau = 1, \ldots 5$) is $\mu + \beta_\tau$ in Equation (1); that is, θ_τ in Equation (2). θ_2 is significant at the 1% and 5% levels in terms of mother–child and father–child separation, respectively.

6. Heterogeneity analyses

To better understand the mechanism we propose, we present two heterogeneity analyses in this section to examine the mechanism from different angles. First, we compare regions or families with different degrees of son preference. The pattern of mother–son separation should be more prominent in places or families with a stronger son preference. Second, as separations are mostly due to work-oriented parental

Table 9
Controlling for birth year cohort fixed effects: Age-specific gender differences in the probability of parental separation.

	Separated from the mother	Separated from the father	Separated from both parents
	(1)	(2)	(3)
Boy	−0.007	−0.052	−0.000
	(0.027)	(0.054)	(0.027)
Boy × Age 1	0.013	0.010	0.006
	(0.031)	(0.061)	(0.029)
Boy × Age 2	0.092***	0.116**	0.072**
	(0.029)	(0.058)	(0.029)
Boy × Age 3	0.022	0.018	0.004
	(0.035)	(0.060)	(0.032)
Boy × Age 4	0.065	0.036	0.055
	(0.042)	(0.063)	(0.037)
Boy × Age 5	0.016	−0.051	0.005
	(0.032)	(0.055)	(0.032)
Boy × Year 2012	0.025	0.026	0.009
	(0.033)	(0.041)	(0.031)
Boy × Year 2014	−0.020	−0.004	−0.028
	(0.032)	(0.039)	(0.028)
Boy × Year 2016	−0.021	0.030	−0.020
	(0.037)	(0.043)	(0.035)
Boy × Year 2018	0.011	0.070*	0.024
	(0.034)	(0.038)	(0.032)
Observations	3218	3218	3218
R^2	0.151	0.166	0.135

SOURCE—Longitudinal data of the CFPS, including the 2010, 2012, 2014, 2016, and 2018 waves.
NOTE—The sample comprises 0–5-year-old firstborn children from rural families. Provinces under a restrictive one-child policy or with large minority groups are excluded. Children migrating with parents are also excluded. We do not include the 2020 wave in this regression as this period coincided with the height of the COVID-19 pandemic in China. The imposition of nationwide lockdowns for multiple months profoundly altered the economic landscape and significantly disrupted population migration. The outcomes are the dummy indicators of whether the child is currently separated from the mother (Column (1)), the father (Column (2)), or both parents (Column (3)). The control variables are parents' ages (and the quadratic terms), the dummy indicator of rural residency in the *hukou* system, parents' years of schooling, and city-level fixed effects. Standard errors are clustered at the county level.
* significant at the 10% level.
** significant at the 5% level.
*** significant at the 1% level.

migration, we compare regions with different historical rates of working migrants. We expect to see more mother–son separation during early childhood in regions with higher rates of migrant laborers.

6.1. Heterogeneity across subjects with different degrees of son preference

Following the introduction of ultrasound B-scans in the early 1980s, Chinese families gradually performed prenatal gender selection by aborting female fetuses, particularly for higher-order births following daughters, which dramatically skewed the sex ratio at birth (Poston et al., 1997; Chen et al., 2013). Using 2000 population census data, we calculate the province-specific sex ratio of rural preschool-aged children (0–5 years) and employ it as a proxy for the degree of son preference. In this calculation, we do not exclude children of higher parities, who are prone to gender selection and drive most regional variations in sex ratio. We categorize the sample into two groups based on the median provincial sex ratio and conduct benchmark regressions for each group separately.

Table 10 shows the regression outcomes in provinces with skewed and balanced sex ratios for children. For each outcome, the regression involves subsamples above (skewed) and below (balanced) the mean sex ratio for rural children aged 0–5 years old. In provinces with high sex ratios (i.e., Column (1)), the likelihood of a gender gap in mother–child separation at age 2 is 5.7 percentage points, which is significant at the 1% level. The 5.7-percentage-point difference amounts to nearly 50% of the sample mean of the separation proportion. In sharp contrast, the difference in counties with low sex ratios is of a much smaller magnitude and is not statistically significant, as seen in Column (2). Similar contrasts are observable when comparing Column (3) with Column (4) and Column (5) with Column (6).

A similar argument can be made in terms of parental education level. Higher levels of parental (especially maternal) education are associated with lower son preference (Chung and Gupta, 2007). Hence, we consider parents' years of schooling as an alternative proxy for the degree of son preference. We divide the sample into two groups: children whose parents' education level does not exceed high school (referred to as less educated), and those with at least one parent who graduated from high school or above (referred to as educated).

Table 11 reports the regression results. In Column (1), among children with less educated parents, the estimate for gender differences at age 2 is similar to that in the benchmark regression. In contrast, the point estimate in Column (2) is much smaller among children with more educated parents, and the effect is not statistically significant. The gender differences in being separated from one's father or from both parents are also larger for children with less educated parents, as presented in Columns (3)–(6), although the gender differences in father–-child separation are not significant in either of the subsamples. These findings are consistent with the conjecture that son preference is the main driving force behind the higher propensity toward mother–child separation in early childhood. Of course, we should not neglect the fact that more educated parents may better understand the important role of parental companionship in building the foundation of child development and therefore stay with their children longer (Luo et al., 2019).

6.2. Heterogeneity across regions with different historical migrant rates

If parent–child separation is mostly driven by parental migration, the patterns documented in the benchmark regressions should be more prominent in regions with traditionally high rates of job-oriented migration. The predetermined rate of migrant labor means comparatively high potential wages (Tombe and Zhu, 2019), better amenities in destination cities (Roback, 1982; Khanna et al., 2021), and established job referral connections or networks with relatives and friends from the same hometown (Card, 2009; Colas, 2019).

Using 2000 population census data, we calculate the county-specific net rate of population outflow as the ratio of the difference between the registered population outflow and resident population inflow in relation to the population with a registered residence. Equivalently, it can be defined as the ratio of the difference between the population with registered residency and the actual resident population to the population with registered residency. Using this county-specific ratio, we divide the sample into two groups by the median of net outflow rates.[9]

The results are shown in Table 12. For each outcome variable, we compare the age-specific gender differences in traditionally high outflow regions with those in low outflow regions. As expected, the gender differences in parent–child separation are of a larger magnitude and higher significance for traditionally high outflow regions. This finding supports the view that gender differences are related to parental work-related migration.

7. Robustness checks

7.1. City-level migration

The previous analysis supports the premise that son preference leads to consecutive childbirth for parents with a firstborn girl. This would

[9] As we are focusing on regions with net population outflows, we take the median within counties where the net population outflow rate is positive.

Table 10
Heterogeneous effects by sex ratio of birth cohorts.

Sex ratio group	Separated from the mother		Separated from the father		Separated from both parents	
	Skewed	Balanced	Skewed	Balanced	Skewed	Balanced
	(1)	(2)	(3)	(4)	(5)	(6)
Boy	−0.008	−0.003	−0.004	−0.011	−0.005	−0.002
	(0.013)	(0.006)	(0.022)	(0.013)	(0.012)	(0.005)
Boy × Age 1	−0.015	0.009	−0.013	0.009	−0.016	0.007
	(0.021)	(0.010)	(0.028)	(0.018)	(0.019)	(0.008)
Boy × Age 2	0.057***	0.004	0.055*	0.014	0.049**	0.007
	(0.020)	(0.010)	(0.029)	(0.017)	(0.020)	(0.009)
Boy × Age 3	0.018	0.006	0.008	0.010	0.006	0.008
	(0.021)	(0.010)	(0.028)	(0.016)	(0.019)	(0.008)
Boy × Age 4	0.010	0.001	−0.010	0.014	0.003	−0.001
	(0.021)	(0.009)	(0.029)	(0.017)	(0.019)	(0.008)
Boy × Age 5	0.020	0.003	−0.001	0.025	0.017	0.003
	(0.021)	(0.008)	(0.029)	(0.017)	(0.019)	(0.007)
Observations	9907	8528	9907	8528	9907	8528
R^2	0.093	0.051	0.115	0.058	0.089	0.054

SOURCE—0.35% sample from the 2010 China Population Census, 1% sample from the 2000 China Population Census.
NOTE—The sample comprises 0–5-year-old firstborn children from rural Han families. Provinces under a restrictive one-child policy or with large minority groups are excluded. Children migrating with parents are also excluded. The outcomes are the dummy indicators of whether the child is currently separated from the mother (Columns (1) and (2)), the father (Columns (3) and (4)), or both parents (Columns (5) and (6)). For each outcome, the regression uses subsamples above (Skewed) and below (Balanced) the median of the provincial sex ratio, respectively. The sex ratio is calculated among rural children who were 0–5 years old using 2000 population census. The control variables are parents' ages (and the quadratic terms), the dummy indicator of rural residency in the *hukou* system, parents' years of schooling, and city-level fixed effects. Standard errors are clustered at the county level.
* significant at the 10% level.
** significant at the 5% level.
*** significant at the 1% level.

Table 11
Heterogeneous effects by parents' years of schooling.

Education group	Separated from the mother		Separated from the father		Separated from both parents	
	Less educated	Educated	Less educated	Educated	Less educated	Educated
	(1)	(2)	(3)	(4)	(5)	(6)
Boy	−0.003	−0.006	−0.001	−0.019	0.004	−0.019
	(0.009)	(0.015)	(0.016)	(0.024)	(0.008)	(0.014)
Boy × Age 1	−0.009	−0.002	−0.020	0.040	−0.016	0.017
	(0.014)	(0.026)	(0.020)	(0.035)	(0.013)	(0.024)
Boy × Age 2	0.032**	0.020	0.034	0.026	0.026**	0.027
	(0.013)	(0.026)	(0.021)	(0.035)	(0.013)	(0.026)
Boy × Age 3	0.006	0.025	−0.002	0.039	−0.004	0.038
	(0.014)	(0.024)	(0.020)	(0.034)	(0.013)	(0.023)
Boy × Age 4	0.002	0.018	−0.003	0.001	−0.009	0.035
	(0.014)	(0.024)	(0.020)	(0.034)	(0.013)	(0.023)
Boy × Age 5	0.015	−0.016	0.012	−0.013	0.008	0.002
	(0.013)	(0.025)	(0.020)	(0.034)	(0.012)	(0.023)
Observations	14,617	3803	14,617	3803	14,617	3803
R^2	0.129	0.147	0.157	0.178	0.122	0.146

SOURCE—0.35% sample from the 2010 China Population Census.
NOTE—The sample comprises 0–5-year-old firstborn children from rural Han families. Provinces under a restrictive one-child policy or with large minority groups are excluded. Children migrating with parents are also excluded. The outcomes are the dummy indicators of whether the child is currently separated from the mother (Columns (1) and (2)), the father (Columns (3) and (4)), or both parents (Columns (5) and (6)). For each outcome, the regression uses subsamples with less educated and more educated parents, respectively. The subsample referred to as "less educated" comprises children whose parents' highest education level is no greater than junior high school, and the subsample referred to as "more educated" comprises households with at least one parent who has graduated from high school or above. The control variables are parents' ages (and the quadratic terms), the dummy indicator of rural residency in the *hukou* system, parents' years of schooling, and city-level fixed effects. Standard errors are clustered at the county level.
* significant at the 10% level.
** significant at the 5% level.
*** significant at the 1% level.

prevent or postpone migration and cause gender differences in the probability of children becoming separated from their parents, especially their mothers. In the benchmark regression, separation is defined as parents and children living in different counties. A concern might exist in that, with transportation infrastructure such as expressways, parents living in other districts and counties may often commute to their hometowns to visit their children. Such separation may not result in serious consequences for either cognitive or non-cognitive development.

In this subsection, we consider parent–child separation across cities or prefectures, in which case the left-behind children may be harmed more. In fact, in our sample, parents migrating across cities account for more than 90% of parent–child separations. We replace the outcome in Equation (1) as city-level separation and perform the regressions again, the results of which are presented in Table 13. Similar to the baseline

Table 12
Heterogeneous effects by historical population outflow rate.

Emigration rate group	Separated from the mother		Separated from the father		Separated from both parents	
	High	Low	High	Low	High	Low
	(1)	(2)	(3)	(4)	(5)	(6)
Boy	−0.019	0.004	0.005	−0.011	−0.013	0.004
	(0.018)	(0.008)	(0.028)	(0.014)	(0.017)	(0.007)
Boy × Age 1	−0.014	−0.003	−0.055	0.020	−0.028	0.001
	(0.027)	(0.013)	(0.038)	(0.019)	(0.026)	(0.012)
Boy × Age 2	0.060**	0.013	0.044	0.025	0.051**	0.014
	(0.025)	(0.013)	(0.038)	(0.019)	(0.025)	(0.012)
Boy × Age 3	0.048*	−0.009	0.013	0.005	0.025	−0.006
	(0.029)	(0.012)	(0.038)	(0.018)	(0.026)	(0.010)
Boy × Age 4	0.036	−0.009	−0.005	0.003	0.025	−0.011
	(0.028)	(0.012)	(0.038)	(0.019)	(0.026)	(0.011)
Boy × Age 5	0.032	−0.002	−0.019	0.023	0.021	0.002
	(0.028)	(0.012)	(0.035)	(0.019)	(0.027)	(0.010)
Observations	5771	12,645	5771	12,645	5771	12,645
R^2	0.127	0.116	0.150	0.139	0.118	0.112

SOURCE—0.35% sample from the 2010 China Population Census, 1% sample from the 2000 China Population Census.
NOTE—The sample comprises 0–5-year-old firstborn children from rural Han families. Provinces under a restrictive one-child policy or with large minority groups are excluded. Children migrating with parents are also excluded. The outcomes are the dummy indicators of whether the child is currently separated from the mother (Columns (1) and (2)), the father (Columns (3) and (4)), or both parents (Columns (5) and (6)). For each outcome, the regression uses subsamples above (high) and below (low) the median of the net rates of population outflow, respectively. The county-specific net rate of population outflow is calculated as the ratio of the difference between the registered population outflow and the resident population inflow to the population with the registered residence. As we are focusing on regions with net population outflows, we take the median within counties where the net population outflow rate is positive. The historical migration flow is calculated from using 2000 population census. The control variables are parents' ages (and the quadratic terms), the dummy indicator of rural residency in the *hukou* system, parents' years of schooling, and city-level fixed effects. Standard errors are clustered at the county level.
* significant at the 10% level.
** significant at the 5% level.
*** significant at the 1% level.

outcomes, at the city level, firstborn boys have a significantly higher probability of becoming separated from their parents at age 2 than firstborn girls.

7.2. Migrant children

Parents could potentially arrange their family residence in three ways: (a) live together with their children in their hometown; (b) migrate and leave their children behind in their hometown; and (c) migrate together with their children. Parent–child separation occurs in (b). We focus on the channel how parental fertility behaviors stop them from leaving hometowns. Thus, we wish to compare families who separate and those who remain together in their hometowns (i.e., (b) and (a)). Hence, we drop migrant-children families (i.e., (c)) who are likely to be more resourceful and can afford to give birth in big cities.

However, excluding migrant children could potentially bias our estimation on gender parity in parent-child separation. On the one hand, if providing companionship to boys brings higher utility to parents, boys are more likely to migrate with parents and therefore dropped from the sample. Since the non-separated boy families are less represented in the sample mother-boy separation will be overestimated. On the other hand, if parents are more likely to take girls (who are generally more mature and docile compared with boys) with them, then the mother-boy separation will be underestimated.

As explained in Section 3.1, migrant children (0–5 years) only account for 2.8% of the entire sample. This small portion is unlikely to lead

Table 13
Alternative definition of separation: Age-specific gender differences in the probability of parent–child separation.

	Separated from the mother	Separated from the father	Separated from both parents
	(1)	(2)	(3)
Boy	−0.005	−0.009	−0.003
	(0.007)	(0.012)	(0.007)
Boy × Age 1	−0.003	0.005	−0.005
	(0.012)	(0.017)	(0.011)
Boy × Age 2	0.031***	0.036**	0.029***
	(0.011)	(0.017)	(0.011)
Boy × Age 3	0.011	0.011	0.007
	(0.011)	(0.016)	(0.010)
Boy × Age 4	0.004	−0.002	0.001
	(0.012)	(0.017)	(0.011)
Boy × Age 5	0.009	0.013	0.009
	(0.011)	(0.016)	(0.010)
Observations	18,435	18,435	18,435
R^2	0.125	0.161	0.119

SOURCE—0.35% sample from the 2010 China Population Census.
NOTE—A separation is identified when the parent currently resides in a city different from the child's place of residence. The sample comprises 0–5-year-old firstborn children from rural Han families. Provinces under a restrictive one-child policy or with large minority groups are excluded. Children migrating with parents are also excluded. The outcomes are the dummy indicators of whether the child is currently separated from the mother (Column (1)), the father (Column (2)), or both parents (Column (3)). The control variables are parents' ages (and the quadratic terms), the dummy indicator of rural residency in the *hukou* system, parents' years of schooling, and city-level fixed effects. Standard errors are clustered at the county level.
* significant at the 10% level.
** significant at the 5% level.
*** significant at the 1% level.

to any substantial change in the interpretation of the gender differences in parental separation. However, to address this issue more rigorously, we reintroduce migrant children into the sample and adopt a multinomial logit model. We estimate the gender differences in children separated from and migrating with their parents, respectively, with families staying together in their hometowns as the reference group. Table 14 reports the logit coefficients in Panel A and the average marginal effects in Panel B. The estimates display a similar pattern for children separated from their parents, compared with the main regression. In contrast, we do not observe any significant gender differences for migrant children, either economically or statistically.

7.3. Supplementary explanations

This study links to prior research on the influence of children's gender on parents' labor supply, occupational choices, and extended household living arrangements. Previous studies propose channels such as imposed pressures from preparing for a son's future competition in the marriage market and attracting help from grandparents. However, we argue that these studies do not undermine our contribution, as they do not offer an explanation for why the gender gap in mother–child separation specifically occurs at age 2.

First, using the China Health and Nutrition Survey, Wang (2019) finds that a firstborn boy decreases the maternal labor supply and increases paternal labor-related migration. Fan et al. (2018) find a decline in maternal family chores but weak or mixed evidence regarding the paternal and maternal labor supply. Both studies consider changes in household time allocation as consequences of the shift in bargaining power caused by the child's gender. Wei and Zhang (2011) demonstrate that fathers of firstborn boys face pressure to save money to purchase a house in order to gain an advantage for their sons in a potential future marriage. Hence, parents of sons are more likely to migrate for work, as well as to work longer and in less safe environments (Tan et al., 2021).

Table 14
Multinomial logit analysis of age-specific gender differences.

	Separated from the mother	Migrated with the mother	Separated from the father	Migrated with the father	Separated from both parents	Migrated with both parents
	(1)	(2)	(3)	(4)	(5)	(6)
Panel A. Coefficient estimates						
Boy	−0.131	0.026	−0.041	−0.089	−0.116	−0.085
	(0.214)	(0.274)	(0.124)	(0.276)	(0.229)	(0.282)
Boy × Age 1	−0.073	−0.114	−0.082	0.020	−0.080	0.051
	(0.267)	(0.355)	(0.161)	(0.360)	(0.278)	(0.364)
Boy × Age 2	0.549**	0.230	0.295*	0.372	0.533*	0.362
	(0.258)	(0.363)	(0.166)	(0.381)	(0.275)	(0.382)
Boy × Age 3	0.215	0.148	0.042	0.287	0.163	0.320
	(0.262)	(0.363)	(0.165)	(0.368)	(0.275)	(0.375)
Boy × Age 4	0.134	0.186	−0.013	0.268	0.077	0.280
	(0.248)	(0.348)	(0.164)	(0.349)	(0.262)	(0.356)
Boy × Age 5	0.271	0.424	0.097	0.587	0.245	0.514
	(0.263)	(0.358)	(0.166)	(0.363)	(0.281)	(0.372)
Panel B. Average marginal effects						
Boy	−0.008	0.001	−0.004	−0.002	−0.006	−0.002
Boy × Age 1	−0.004	−0.002	−0.009	0.001	−0.004	0.001
Boy × Age 2	0.031	0.004	0.029	0.007	0.027	0.007
Boy × Age 3	0.012	0.003	0.003	0.006	0.008	0.007
Boy × Age 4	0.007	0.004	−0.003	0.006	0.003	0.006
Boy × Age 5	0.015	0.009	0.007	0.013	0.011	0.011
Observations	18,932	18,932	18,932	18,932	18,932	18,932

SOURCE—0.35% sample from 2010 China Population Census.
NOTE—The sample comprises 0–5-year-old firstborn children from rural Han families. Provinces under a restrictive one-child policy or with large minority groups are excluded. The outcome is the categorical variable with each value indicating whether the child is currently at home with the parents (reference category), separated from the parent(s), or migrated with the parent(s). The control variables are parents' ages (and the quadratic terms), the dummy indicator of rural residency in the *hukou* system, parents' years of schooling, and city-level fixed effects. Standard errors are clustered at the county level. Panel A reports coefficient estimates, and Panel B presents average marginal effects.
* significant at the 10% level.
** significant at the 5% level.
*** significant at the 1% level.

Second, Sun et al. (2019) find that families with firstborn sons are more likely to live with the son's paternal grandparents.[10] Paternal grandmothers help with family chores and attend to children, which allows mothers to join the labor market and work outside the home. Their findings support our explanation. If son preference leads to grandparents being more willing to take in a grandson, the option of migrating for better work opportunities would seem more feasible.

We do not intend to rule out any of these channels; however, we argue that no prior findings weaken the novelty of our research. Our findings offer an original explanation of the higher propensity toward parent–son separation because we provide concrete, detailed evidence on age-specific gender differences in separation and show that the timing of consecutive fertility matches that of mother–child separation. No past research has highlighted any specific implications regarding the timing of parental migration.

8. Discussions: the importance of "Age-2" separation and the gender gap in school performance

Due to the lack of knowledge, information, and resources, many migrant parents live separately from their children (Yue et al., 2016; Zhou and Liu, 2016; Yue et al., 2017), and the negative relationship between being left behind and children's cognitive and non-cognitive skills has been found by a big volume of literature (Sun et al., 2010; He et al., 2012; Wen and Lin, 2012; Chen, 2013; Mou et al., 2013; Zhang et al., 2014; Gao et al., 2015; Mu and De Brauw, 2015; Ren and Treiman, 2016; Yue et al., 2016; Meng and Yamauchi, 2017; Luo, 2020; Huang, 2022).

In this section, we combine the results from the CFPS data with that yielded from the 2010 population census to assess the contribution of the gender gap in separation at certain age to that in future schooling performance by a back-of-the-envelope calculation. First, we estimate the overall association between separation and later academic ranking. In the 2010 and 2012 waves of the CFPS, the children's module asked whether a child was currently left behind by their migrating parent(s). In the 2018 and 2020 waves, for children above 10 years old, the panel data recorded the academic test score class ranking percentiles. We then estimate the relationship between being left behind in 2010 or 2012 and the academic performance when the same child was observed again in the 2018 or 2020 survey.[11]

Column (1) in Panel An of Appendix Table A3 presents the associations between the experience of being separated at certain age and children's class ranking percentile after attaining 10 years old. It shows that separating from the mother at age 2 is associated with 8.4 percentile decrease in future academic performance. To conduct the back-of-the-envelope calculation, we multiply them with the gender difference in

[10] Specifically, under China's patriarchal and patrilocal traditions, after marrying, young couples live with their elderly parents. The extended family splits when the youngest child gets married. Parents with a firstborn son are less likely to move away from the extended family.

[11] The regression equation is $ClassRanking_{icl,t+8} = \alpha + \sum_{\tau=2}^{5}\beta_\tau I(l=\tau) \times Separation_{iclt} + \sum_{\tau=3}^{5}\delta_\tau I(l_{iclt}=\tau) + X_{iclt}\cdot\varphi + \lambda_t + \varepsilon_{iclt}$. The subscripts *i, c, l* and *t* denote the individual child, prefecture, child's age and the survey year (2010 or 2012), respectively. τ starts from 2 because those who turned 10 in 2018 were 2 in 2010, the earliest wave of the CFPS. Standard errors are clustered at the village level. Under the assumption that different birth cohorts are comparable after controlling for household characteristics (annual household income, parents' ages and the quadratic terms, the dummy indicator of rural residency in the *hukou* system, and parents' years of schooling), school characteristics (the dummy indicators of whether the school is a key school and whether the class is a key class), and year fixed effect, we interpret the coefficient β_τ as the association between separation at age τ and the performance after attaining age 10.

the age-specific likelihood of mother-child separation, i.e., the baseline regression coefficients presented in Column (1) of Table 3. The products are exhibited in Column (2) of Appendix Table A3. These products represent the gender difference in class ranking attributable to early mother-child separation. Finally, we divide the entries in Column (2) by the total gender difference in the class ranking[12] and report the ratio in Column (3). This percentage suggests the contribution of gender difference in mother-child separation at certain age to the gender difference in later class ranking.

Appendix Table A3 Panel A suggests the separation at age 2 contributes most to future class ranking. On the one hand, separation at the ages of 2 and 3 are likely to have most salient impact on future academic performance, as evidenced by the associations in column (1). This is consistent with Heckman's conjecture of the "first 1000 days of life" (Heckman, 2008). On the other hand, the gender disparity in separations is most pronounced at age 2. As a result, the contribution of gender difference in separation at age 2 (5.01%) is 5.7 times higher than at age 3 (0.88%). In summary, the back-of-the-envelope calculation shows the gender difference in separation from the mother at the age of 2 contributes the highest explanatory power to the gender differences in future academic performance.

As a sensitivity check on the estimated contribution, we use an IV estimator to gauge the association between separation and schooling performance presented in Column (1). Parents' decisions regarding childcare arrangements and work-oriented migration are generally made jointly and are potentially correlated with parents' unobserved characteristics. We employ a shift-share instrumental variable method in the vein of Card (2009). Specifically, we calculate a labor demand shock on each potential migration destination province in each year (2010 and 2012). For each county of origin, we calculate the average "pulling shock" as the average of provincial labor demand shock weighted by traditional migration share from the origin county to the destination province (see Appendix B for details on the construction of the IV). The instrument for each interaction term is the product of the average "pulling shock" and the child's age dummy. The coefficients of the IV estimations are reported in Columns (4). The consequent calculations parallel to Columns (2) and (3) are presented in Columns (5) and (6) of Appendix Table A3. We see a similar pattern that the gender disparity in separation at age 2 has the largest explanatory power for the gender difference observed in later school performance.[13] The first stage of the 2SLS estimation is reported in Appendix Table A4.

The number of siblings could influence the performance of the firstborn children.[14] Under son preference, boys are more likely to have fewer siblings than firstborn girls. To address the concern that girls may benefit from the environment of more siblings, Appendix Table A3 Panel B further controls for the number of siblings. The results are very similar to that presented in Panel A.

9. Conclusions

The experience of being separated from one's parents in early childhood is well-known to cause both short- and long-term drawbacks in cognitive and non-cognitive development. However, no prior research has focused on gender differences in the likelihood of being left behind by migrant parents.

Although it appears surprising, in this study, we find an unexpected consequence of son preference in rural China in which firstborn girls are less likely to be left behind during their early childhood than firstborn boys. To continue attempting to have a son, parents—especially mothers of firstborn girls—are less likely to migrate for work. Using population census data, we note a surge in second childbirths and significant gender differences in the incidence of parents migrating for work at the point at which the firstborn girl is 2 years old. Such a pattern is highly consistent with our prediction that parents' consecutive fertility behavior following a firstborn daughter postpones or terminates their migration plans and prevents mother–daughter separation during early childhood. To this extent, son preference actually protects girls by providing them with maternal care and companionship in early childhood.

Such results shed light on controversies in prior literature focusing on unequal family resource allocation which harms girls. Some studies predict that gender inequality will disappear with economic growth (Chung and Gupta, 2007); however, others assert that institutional or cultural factors will hinder the path toward gender equity (Duflo, 2012). In this study, we are not confined by the traditional angle of the distribution of purchased goods between male and female children; instead, we focus on a largely neglected—yet extremely important—parental factor (i.e., parental companionship in early childhood). The unintended consequences of son preference during a time of fast economic growth and urbanization show that the interaction of son preference and growth could lead to more complex effects than those revealed by prior literature.

This result is consistent with the observation that although boys obtain more household resources, they lag behind girls academically. According to the World Bank (2018, 2020), in 2018, 3.4% of boys in primary schools worldwide were repeating a grade, which was greater than the percentage of girls (2.9%). Additionally, boys who persisted to the last grade of primary school accounted for 81.2% of the cohort in 2020, while the ratio was 84.3% for girls. Our back-of-the-envelope calculation indicates that at least 5.0% of the gender difference in school performance can be attributed to that of early childhood separation from one's parents. One caveat regarding the understanding of our findings is that the inequality we observe does not mean girls live in a superior environment compared to that of boys. Postponing or interrupting job-related migration could result in substantial economic costs and difficulties in supporting one's family, especially in providing necessary childcare. Therefore, the calculation can be regarded as a lower bound.

The issue of left-behind children is not unique to China. According to BBC News (2015), although China has the most left-behind children (61 million), many also reside in the Philippines (9 million), Sri Lanka (1 million), and Tajikistan (100,000). The issue also affects European countries such as Romania (350,000) and Moldova (180,000) as well as nations across Latin America, including Mexico (500,000) and Ecuador (218,000). Many organizations provide aid to these millions of children, including non-governmental organizations run by China and UNICEF. They advocate for and offer direct assistance to left-behind children. For example, UNICEF provides services in multiple Chinese provinces, including social and emotional development and health administration. However, a more promising way to address this crisis would be for policymakers to remove barriers to social services in different localities, which would allow parents to be with their children while pursuing better work opportunities.

[12] The gender differences in academic ranking are 5.26, 7.70, 5.93, and 9.42 percentiles for age 10, 11, 12, and 13, respectively. We link separation age 2, 3, 4, and 5 to their academic performance at age 10, 11, 12, and 13, respectively. Thus, in calculating the explanatory power, we divide each separation age in Column (2) of Table A3 by each corresponding gender difference in academic ranking after attaining 10 years old.

[13] We find a larger coefficient size of IV regression compared with the OLS regression. This implies a potential positive selection in terms of parents' migration and separation from their children. Of course, the larger estimate could also be resulted from the nature of LATE estimator.

[14] Siblings compete for limited parental input, which may be detrimental to intellectual development (Blake, 1981; Downey, 1995, 2001; Cáceres-Delpiano, 2006). Conversely, Black et al. (2005) and Conley and Glauber (2006) propose that these findings could be a misinterpretation of the birth order effect as the effect of siblings. In addition, siblings could provide each other more opportunities to socialize, teach, and learn, which enhances non-cognitive skills (Cicirelli, 1974; Weisner, 1989; Azmitia and Hesser, 1993; Downey et al., 2015; Qian, 2018).

Authors' statement

All authors have accepted responsibility for the entire content of this manuscript and approved its submission.

This article is the authors' original work, hasn't received prior publication and isn't under consideration for publication elsewhere.

Authors state no conflict of interest.

Sun acknowledges financing from National Natural Science Foundation of China (No. 72373157), Fundamental Research Funds for the Central Universities, and the Research Funds of Renmin University of China (No. 22XNA003) and fund for building world-class universities (disciplines) of Renmin University of China (No. KYGJC2022006).

Zibin acknowledges partial financial support from the Key Program of the National Natural Science Foundation of China (Grant No. 72034006), and the Shanghai Education Science Research Program (Grant No. C2023077).

Zibin Huang, Xu Jiang, Ang Sun.

Data availability

Data will be made available on request.

Appendix A. Additional Tables and Figures

Table A1
Alternative Specification: Age-Specific Gender Differences in the Probability of Parental Separation

	Separated from the mother	Separated from the father	Separated from both parents
	(1)	(2)	(3)
Boy × Age 0	−0.005	−0.007	−0.003
	(0.008)	(0.013)	(0.007)
Boy × Age 1	−0.010	−0.011	−0.010
	(0.010)	(0.013)	(0.009)
Boy × Age 2	0.027***	0.029**	0.025***
	(0.009)	(0.012)	(0.009)
Boy × Age 3	0.006	0.000	0.002
	(0.009)	(0.011)	(0.008)
Boy × Age 4	0.000	−0.007	−0.003
	(0.009)	(0.011)	(0.009)
Boy × Age 5	0.006	0.003	0.006
	(0.009)	(0.010)	(0.008)
Observations	18,435	18,435	18,435
R^2	0.122	0.153	0.116

SOURCE—0.35% sample from the 2010 China Population Census.
NOTE—The sample comprises 0–5-year-old firstborn children from rural Han families. Provinces under a restrictive one-child policy or with large minority groups are excluded. Children migrating with parents are also excluded. The outcomes are the dummy indicators of whether the child is currently separated from the mother (Column (1)), the father (Column (2)), or both parents (Column (3)). The control variables are parents' ages (and the quadratic terms), the dummy indicator of rural residency in the *hukou* system, parents' years of schooling years, and city-level fixed effects. Standard errors are clustered at the county level.
*significant at the 10% level.
**significant at the 5% level.
***significant at the 1% level.

Table A2
Alternative Specification: Age-Specific Gender Differences in Labor-Oriented Parental Migration

	The mother is a migrant laborer	The father is a migrant laborer	Both parents are migrant laborers
	(1)	(2)	(3)
Boy × Age 0	−0.001	−0.008	0.001
	(0.007)	(0.013)	(0.006)
Boy × Age 1	−0.013	−0.010	−0.012
	(0.009)	(0.013)	(0.009)
Boy × Age 2	0.027***	0.028**	0.025***
	(0.009)	(0.012)	(0.009)
Boy × Age 3	0.002	−0.000	−0.002
	(0.009)	(0.011)	(0.008)
Boy × Age 4	0.004	−0.002	0.001
	(0.009)	(0.011)	(0.009)
Boy × Age 5	0.005	−0.001	0.005
	(0.008)	(0.011)	(0.008)
Observations	18,435	18,435	18,435
R^2	0.124	0.145	0.116

SOURCE—0.35% sample from the 2010 China Population Census.
NOTE—The sample comprises 0–5-year-old firstborn children from rural Han families. Provinces under a restrictive one-child policy or with large minority groups are excluded. Children migrating with parents are also excluded. The outcomes are the dummy indicators of whether the mother (Column (1)), the father (Column (2)), or both parents (Column (3)) currently migrate for work. The control variables are parents' ages (and the quadratic terms), the dummy indicator of rural residency in the *hukou* system, parents' years of schooling, and city-level fixed effects. Standard errors are clustered at the county level.
*significant at the 10% level.
**significant at the 5% level.
***significant at the 1% level.

Table A3
Contribution of Gender Difference in Mother-Child Separation at Each Age to That in Long-Term Schooling Performance

	OLS			IV		
	(1) Assc w. Class rank	(2) (1)*Genderdiff in separation	(3) Percent explained	(4) Assc w. Class rank	(5) (4)*Genderdiff in separation	(6) Percent explained
Panel A. Back-of-the-envelope calculation (benchmark)						
Separated from the mother at Age 2	−8.432*	−0.264	5.007%	−34.084*	−1.065	20.239%
	(4.448)			(20.432)		
Separated from the mother at Age 3	−6.454	−0.068	0.877%	−10.405	−0.109	1.414%
	(4.124)			(15.569)		
Separated from the mother at Age 4	−1.668	−0.009	0.144%	−11.351	−0.058	0.980%
	(3.824)			(25.437)		
Separated from the mother at Age 5	−5.435	−0.057	0.608%	−11.078	−0.117	1.238%
	(4.227)			(16.957)		
Observations	1814			1726		
Panel B. Back-of-the-envelope calculation (controlling for the number of siblings)						
Separated from the mother at Age 2	−8.371*	−0.262	4.971%	−34.190*	−1.069	20.302%
	(4.463)			(20.421)		
Separated from the mother at Age 3	−6.358	−0.067	0.864%	−11.244	−0.118	1.528%
	(4.134)			(15.279)		
Separated from the mother at Age 4	−1.712	−0.009	0.148%	−12.356	−0.063	1.067%
	(3.857)			(25.521)		
Separated from the mother at Age 5	−5.533	−0.058	0.619%	−12.894	−0.136	1.441%
	(4.248)			(16.981)		
Observations	1814			1726		

SOURCE—Longitudinal data from the 2010, 2012, 2018, and 2020 waves of the CFPS.

NOTE—The sample consists of individuals aged 2–5 years old in either the 2020 or 2012 wave, who were also observed in the 2018 and 2020 waves. The entries in Column (1) are associations between the percentile class ranking in either 2018 or 2020 and whether the child was separated from the mother at certain age. In Column (4), the indicators of separation are instrumented by the shift-share Bartik instruments interacted with each age. In Panel A, the control variables are household characteristics (annual household income, parents' ages and the quadratic terms, the dummy indicator of rural residency in the *hukou* system, and parents' years of schooling), school characteristics (the dummy indicators of whether the school is a key school and whether the class is a key class), and year fixed effect. Panel B further includes the number of siblings to control for the effect of siblings. Standard errors are clustered at the village level. Columns (2) and (5) show the product of the likelihood of the gender gap in age-specific mother–child separation from Table 3 and the entries in Columns (1) and (4), respectively (the unit is percentile). Columns (3) and (6) divide (2) and (5) by the gender difference in schooling performance. The ratio is interpreted as the proportion of the gender differences in class ranking that can be attributed to the gender differences in the probability of mother–child separation at each specific age.

* significant at the 10% level.
** significant at the 5% level.
*** significant at the 1% level.

Table A4
The First Stage of IV Regression

	(1)	(2)	(3)	(4)
	Separated from mother at Age 2	Separated from mother at Age 3	Separated from mother at Age 4	Separated from mother at Age 5
Shift share × Age 2	2.495***			
	(0.623)			
Shift share × Age 3		2.401***		
		(0.535)		
Shift share × Age 4			1.482***	
			(0.453)	
Shift share × Age 5				2.799***
				(0.518)
F-Statistic	16.06	16.97	11.67	17.17
Observations	1726	1726	1726	1726
R^2	0.171	0.165	0.129	0.144

SOURCE—Longitudinal data from the 2010 and 2012 waves of the CFPS.

NOTE—The sample comprises 2-5-year-old children from the 2010 and 2012 waves of the CFPS. The potential endogenous variable is whether the child was separated from the mother at certain age. The instruments are county-year level Bartik shift-share interacted with age dummies. The IV captures the average labor demand shock in destination provinces, weighted by the immigration rate for the destination provinces from each original county. The control variables are household characteristics (annual household income, parents' ages and the quadratic terms, the dummy indicator of rural residency in the *hukou* system, and parents' years of schooling), school characteristics (the dummy indicators of whether the school is a key school and whether the class is a key class), and year fixed effect. Standard errors are clustered at the village level. The coefficient estimates of the Bartik IVs and the partial *F*-statistics are reported.

* significant at the 10% level.
** significant at the 5% level.
*** significant at the 1% level.

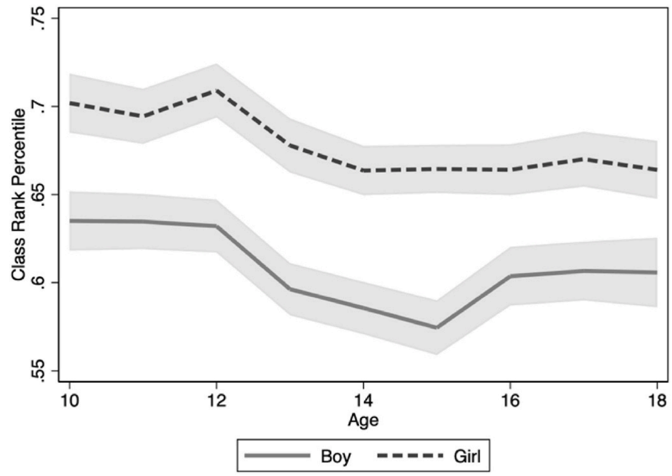


Fig. A1. Gender differences in the percentile of class ranking against age.

SOURCE—2010, 2012, 2014, 2016, 2018 and 2020 waves of the CFPS. The gray band is the 90% confidence interval for the mean.

Appendix B. Construction of Bartik IV

It is important to note that decisions on migration and living arrangements are made jointly and the determinants may be correlated with parents' unobservable characteristics. To gauge the association estimated in Appendix Table A3 Column (1), we construct a shift-share Bartik instrument to implement a 2SLS estimation. This instrument is similar to the traditional Card instrument (Card, 2009); however, our method focuses on instrumenting for migration from one's original (home) county. Let R denote the set of destination provinces. For individual from home county j in year t, the instrument is defined as follows:

$$z_{jt} = \sum_{r \in R} (L_{rt} - L_{rt-1}) \frac{Mig_{jr,2010}}{Mig_{j,2010}},$$

where L_{rt} is the number of employed workers in destination province r during that year, and L_{rt-1} refers to the number of employed workers in that province during the previous year. The shift component is defined as the difference, i.e., $(L_{rt} - L_{rt-1})$, which indicates the labor demand shock in destination province r in year t. In the last fraction term, $Mig_{jr,2010}$ indicates the number of migrants from home county j to destination province r in year 2010.[15] $Mig_{j,2010}$ refers to the total number of migrants from j to any place other than j in 2010. Thus, $\frac{Mig_{jr,2010}}{Mig_{j,2010}}$ is the fraction of people migrating from home county j to destination province r out of all migrants from j in 2010. The employment data come from the China Labor Statistical Yearbook and the migration data come from the Population Census in 2010.

The above procedure assigns greater weights to provinces with higher traditional migration sizes. These provinces are usually closer to the destination r. For instance, migrants from *Zhengding* County in *Hebei* Province are more likely to be affected by the labor market shocks in *Beijing*, which is adjacent to *Heibei* Province, than those in *Guangdong* Province, which is more than one thousand miles away. Short geographical distance generally means lower migration costs and stronger networks of job referral.

The maintaining assumption of this IV method is that the changes of local migration behaviors in all destination j during period t-1 and t would be the same if there is no difference in labor demand shocks $(L_{rt} - L_{rt-1})$ across any of the destination provinces ($r \in R$) (Goldsmith-Pinkham et al., 2020).

We construct the instrument of each separation variable as the product of z_{jt} and the according child's age dummy. The results of the first stage are presented in Appendix Table A4. The IV regression coefficients are presented in Appendix Table A3 Column (4), which suggest that separation from one's mother in early childhood has a persistent and substantial negative impact on the later academic performance of children, with the most significant negative effect observed at age 2. This is consistent with the findings using OLS regression.

References

Abadie, Alberto, Athey, Susan, Imbens, Guido W., Wooldridge, Jeffrey, 2023. When should you adjust standard errors for clustering? Q. J. Econ. 138 (1), 1–35.
Aucejo, Esteban M., James, Jonathan, 2019. Catching up to girls: understanding the gender imbalance in educational attainment within race. J. Appl. Econom. 34 (4), 502–525.
Autor, David, Figlio, David, Karbownik, Krzysztof, Roth, Jeffrey, Wasserman, Melanie, 2019. Family disadvantage and the gender gap in behavioral and educational outcomes. Am. Econ. J. Appl. Econ. 11 (3), 338–381.
Azmitia, Margarita, Hesser, Joanne, 1993. Why siblings are important agents of cognitive development: a comparison of siblings and peers. Child Dev. 64 (2), 430–444.
Banister, Judith, 2004. Shortage of girls in China today. J. Popul. Res. 21 (1), 19–45.
Bao, Xiaojia, Galiani, Sebastian, Li, Kai, Long, Cheryl, 2023. Where have all the children gone? An empirical study of child abandonment and abduction in China. J. Econ. Behav. Organ. 208, 95–119.
Baum, Charles L., 2003. Does early maternal employment harm child development? An analysis of the potential benefits of leave taking. J. Labor Econ. 21 (2), 409–448.
BBC News, 2015. Who Are the "Left behind Children" Around the World? https://www.bbc.com/news/av/world-33996631. (Accessed 27 August 2023).
Beaman, Robyn, Wheldall, Kevin, Kemp, Coral, 2006. Differential teacher attention to boys and girls in the classroom. Educ. Rev. 58 (3), 339–366.
Bertrand, Marianne, Pan, Jessica, 2013. The trouble with boys: social influences and the gender gap in disruptive behavior. Am. Econ. J. Appl. Econ. 5 (1), 32–64.
Black, Sandra E., Devereux, Paul J., Kjell, G., 2005. The more the merrier? The effect of family size and birth order on children's education. Q. J. Econ. 120 (2), 669–700.
Blake, Judith, 1981. Family size and the quality of children. Demography 18, 421–442.

[15] Including the home province when migrating from j to other counties within the home province.

Blau, Francine D., Kahn, Lawrence M., Brummund, Peter, Cook, Jason, Larson-Koester, Miriam, 2020. Is there still son preference in the United States? J. Popul. Econ. 33 (3), 709–750.
Borooah, Vani K., 2004. Gender bias among children in India in their diet and immunisation against disease. Soc. Sci. Med. 58 (9), 1719–1731.
Buckles, Kasey S., Munnich, Elizabeth L., 2012. Birth spacing and sibling outcomes. J. Hum. Resour. 47 (3), 613–642.
Cáceres-Delpiano, Julio, 2006. The impacts of family size on investment in child quality. J. Hum. Resour. 41 (4), 738–754.
Campbell, Jay R., Hombo, Catherine M., Mazzeo, John, 2000. NAEP 1999 Trends in Academic Progress: Three Decades of Student Performance. ED Pubs, Jessup, MD, pp. 20794–21398. PO Box 1398.
Card, David, 2009. Immigration and inequality. Am. Econ. Rev. 99 (2), 1–21.
Chen, Joyce J., 2013. Identifying non-cooperative behavior among spouses: child outcomes in migrant-sending households. J. Dev. Econ. 100 (1), 1–18.
Chen, Lincoln C., Huq, Emdadul, Stan, d'Souza, 1981. Sex bias in the family allocation of food and health care in rural Bangladesh. Popul. Dev. Rev. 55–70.
Chen, Yuyu, Li, Hongbin, Meng, Lingsheng, 2013. Prenatal sex selection and missing girls in China: evidence from the diffusion of diagnostic ultrasound. J. Hum. Resour. 48 (1), 36–70.
Chung, Woojin, Gupta, Monica Das, 2007. The decline of son preference in South Korea: the roles of development and public policy. Popul. Dev. Rev. 33 (4), 757–783.
Cicirelli, Victor G., 1974. Relationship of sibling structure and interaction to younger sib's conceptual style. J. Genet. Psychol. 125 (2), 37–49.
Clark, Shelley, 2000. Son preference and sex composition of children: evidence from India. Demography 37 (1), 95–108.
Colas, Mark, 2019. Dynamic Responses to Immigration. Opportunity and Inclusive Growth Institute.
Conley, Dalton, Glauber, Rebecca, 2006. Parental educational investment and children's academic risk estimates of the impact of sibship size and birth order from exogenous variation in fertility. J. Hum. Resour. 41 (4), 722–737.
Cortes, Rosalia, 2008. Children and Women Left behind in Labour Sending Countries: an Appraisal of Social Risks. United Nations Children's Fund (UNICEF), Policy, Advocacy and Knowledge Management Section, Division of Policy and Practice, New York.
Cortes, Patricia, 2015. The feminization of international migration and its effects on the children left behind: evidence from the Philippines. World Dev. 65, 62–78.
Cunha, Flavio, Heckman, James, 2007. The technology of skill formation. Am. Econ. Rev. 97 (2), 31–47.
Cunha, F., Heckman, J.J., Lochner, L., Masterov, D.V., 2006. Interpreting the evidence on life cycle skill formation. Handbook of the Econ. Educ. 1, 697–812.
Cunha, Flavio, Heckman, James J., Schennach, Susanne M., 2010. Estimating the technology of cognitive and noncognitive skill formation. Econometrica 78 (3), 883–931.
Dahl, Gordon B., Moretti, Enrico, 2008. The demand for sons. Rev. Econ. Stud. 75 (4), 1085–1120.
Dancer, Diane, Rammohan, Anu, Smith, Murray D., 2008. Infant mortality and child nutrition in Bangladesh. Health Econ. 17 (9), 1015–1035.
Das, Narayan, 1987. Sex preference and fertility behavior: a study of recent Indian data. Demography 517–530.
Das Gupta, Monica, Jiang, Zhenghua, Li, Bohua, Xie, Zhenming, Chung, Woojin, Bae, Hwa- Ok, 2003. Why is son preference so persistent in East and South Asia? A cross-country study of China, India and the Republic of Korea. J. Dev. Stud. 40 (2), 153–187.
Downey, Douglas B., 1995. When bigger is not better: family size, parental resources, and children's educational performance. Am. Socio. Rev. 746–761.
Downey, Douglas B., Condron, Dennis J., Deniz, Yucel, 2015. Number of siblings and social skills revisited among American fifth graders. J. Fam. Issues 36 (2), 273–296.
Duflo, Esther, 2012. Women empowerment and economic development. J. Econ. Lit. 50 (4), 1051–1079.
Ebenstein, Avraham, 2010. The "missing girls" of China and the unintended consequences of the One Child Policy. J. Hum. Resour. 45 (1), 87–115.
Ebenstein, Avraham, 2021. The historical origins of son preference: Patrilocality and missing women. Available at SSRN 3829406.
Edlund, Lena, Lee, Chulhee, 2013. Son Preference, Sex Selection and Economic Development: the Case of South Korea. National Bureau of Economic Research. NBER Working Paper 18679.
Entwisle, Doris R., 2018. Children, Schools, and Inequality. Routledge.
Fan, Xiaodong, Fang, Hanming, Markussen, Simen, 2015. Mothers' Employment and Children's Educational Gender Gap. National Bureau of Economic Research. NBER Working Paper 21183.
Fan, Yi, Yi, Junjian, Yuan, Ye, Zhang, Junsen, 2018. The glorified mothers of sons: evidence from child sex composition and parental time allocation in rural China. J. Econ. Behav. Organ. 145, 249–260.
Filmer, Deon, Friedman, Jed, Schady, Norbert, 2008. Development, Modernization, and Son Preference in Fertility Decisions. World Bank Policy Research Working. *Paper* 4716.
Gao, Qin, Hong, Li, Zou, Hong, Cross, Wendi, Bian, Ran, Liu, Yan, 2015. The mental health of children of migrant workers in Beijing: the protective role of public school attendance. Scand. J. Psychol. 56 (4), 384–390.
Gilliam, Walter S., 2005. Prekindergarteners Left behind: Expulsion Rates in State Prekindergarten Systems. Foundation for Child Development, New York, NY.
Goldsmith-Pinkham, Paul, Isaac, Sorkin, Henry, Swift, 2020. Bartik instruments: what, when, why, and how. Am. Econ. Rev. 110 (8), 2586–2624.
Gong, Xiaodong, Arthur Van Soest, Zhang, Ping, 2005. The effects of the gender of children on expenditure patterns in rural China: a semiparametric analysis. J. Appl. Econom. 20 (4), 509–527.
Graham, Elspeth, Jordan, Lucy P., 2011. Migrant parents and the psychological well-being of left-behind children in Southeast Asia. J. Marriage Fam. 73 (4), 763–787.
He, Bingyan, Fan, Jingyi, Liu, Ni, Li, Huijuan, Wang, Yanjun, Williams, Joshua, Wong, Kaishing, 2012. Depression risk of 'left-behind children' in rural China. Psychiatr. Res. 200 (2–3), 306–312.
Heckman, James J., 2006. Skill formation and the economics of investing in disadvantaged children. Science 312 (5782), 1900–1902.
Heckman, James J., 2008. The case for investing in disadvantaged young children. CESifo DICE Report 6 (2), 3–8.
Huang, Zibin, 2022. Peer Effects, Parental Migration and Children's Human Capital: A Spatial Equilibrium Analysis in China. Working Paper.
Huang, Xian, Wu, Bingxiao, 2020. Impact of urban-rural health insurance integration on health care: evidence from rural China. China Econ. Rev. 64, 101543.
Iqbal, Neelam, Anna, Gkiouleka, Milner, Adrienne, Montag, Doreen, Gallo, Valentina, 2018. Girls' hidden penalty: analysis of gender inequality in child mortality with data from 195 countries. BMJ Glob. Health 3 (5), e001028.
Jayachandran, Seema, Kuziemko, Ilyana, 2011. Why do mothers breastfeed girls less than boys? Evidence and implications for child health in India. Q. J. Econ. 126 (3), 1485–1538.
Jayachandran, Seema, Pande, Rohini, 2017. Why are Indian children so short? The role of birth order and son preference. Am. Econ. Rev. 107 (9), 2600–2629.
Khanna, Gaurav, Liang, Wenquan, Mobarak, Ahmed Mushfiq, Song, Ran, 2021. The Productivity Consequences of Pollution-Induced Migration in China. National Bureau of Economic Research. NBER Working Paper 28401.
Kimball, Meredith M., 1989. A new perspective on women's math achievement. Psychol. Bull. 105 (2), 198.
Li, Shuzhuo, 2007. Imbalanced sex ratio at birth and comprehensive intervention in China. In: Presented at 4th Asia Pacific Conference on Reproductive and Sexual Health and Rights, Hyderabad, India.
Lin, Justin Yifu, 2011. Demystifying the Chinese Economy. Cambridge University Press, Cambridge.
Liu, Haoming, 2014. The quality-quantity trade-off: evidence from the relaxation of China's one-child policy. J. Popul. Econ. 27 (2), 565–602.
Liu, Shuang, Zou, Mingru, 2011. Birth interval between first and second child and its policy implications. Popul. Res. 35 (2), 83–93 (*in Chinese*).
Luby, Joan L., Barch, Deanna M., Belden, Andy, Gaffrey, Michael S., Tillman, Rebecca, Babb, Casey, Nishino, Tomoyuki, Suzuki, Hideo, Botteron, Kelly N., 2012. Maternal support in early childhood predicts larger hippocampal volumes at school age. Proc. Natl. Acad. Sci. USA 109 (8), 2854–2859.
Luby, Joan L., Belden, Andy, Harms, Michael P., Tillman, Rebecca, M Barch, Deanna, 2016. Preschool is a sensitive period for the influence of maternal support on the trajectory of hippocampal development. Proc. Natl. Acad. Sci. USA 113 (20), 5742–5747.
Luo, Xiaoman, 2020. How Does Parental Out-Migration Affect Left-Behind Children's Schooling Outcomes? Working Paper.
Luo, Renfu, Jia, Fang, Yue, Ai, Zhang, Linxiu, Lyu, Qijia, Shi, Yaojiang, Yang, Meredith, Kotb, Sarah, Scott, Rozelle, 2019. Passive parenting and its association with early child development. Early Child. Dev. Care 189 (10), 1709–1723.
Meng, Xin, Yamauchi, Chikako, 2017. Children of migrants: the cumulative impact of parental migration on children's education and health outcomes in China. Demography 54 (5), 1677–1714.
Mitra, Aparna, 2014. Son preference in India: implications for gender development. J. Econ. Issues 48 (4), 1021–1037.
Mou, Jin, Griffiths, Sian M., Fong, Hildy, Dawes, Martin G., 2013. Health of China's rural– urban migrants and their families: a review of literature from 2000 to 2012. Br. Med. Bull. 106 (1), 19–43.
Mu, Ren, De Brauw, Alan, 2015. Migration and young child nutrition: evidence from rural China. J. Popul. Econ. 28 (3), 631–657.
Murnane, Richard J., 2013. US high school graduation rates: patterns and explanations. J. Econ. Lit. 51 (2), 370–422.
National Bureau of Statistics, 2020. Income and Consumption of Residents in 2019. http://www.stats.gov.cn/sj/zxfb/202302/t20230203_1900600.html. (Accessed 27 August 2023).
Park, Chai Bin, Cho, Nam-Hoon, 1995. Consequences of son preference in a low-fertility society: imbalance of the sex ratio at birth in Korea. Popul. Dev. Rev. 59–84.
Pitt, Mark M., Rosenzweig, Mark R., 1990. Estimating the intrahousehold incidence of illness: child health and gender-inequality in the allocation of time. Int. Econ. Rev. 969–989.
Poston Jr., Dudley, L., Gu, Baochang, Liu, Peihang Peggy, McDaniel, Terra, 1997. Son preference and the sex ratio at birth in China: a provincial level analysis. Soc. Biol. 44 (1–2), 55–76.
Potter, Robert G., 1963. Birth intervals: structure and change. Popul. Stud. 17 (2), 155–166.
Qian, Nancy, 2018. The Effect of China's One Child Policy on Sex Selection, Family Size, and the School Enrolment of Daughters, vol. 296. Towards Gender Equity in Development.
Ready, Douglas D., LoGerfo, Laura F., Burkam, David T., Lee, Valerie E., 2005. Explaining girls' advantage in kindergarten literacy learning: do classroom behaviors make a difference? Elem. Sch. J. 106 (1), 21–38.
Ren, Qiang, Treiman, Donald J., 2016. The consequences of parental labor migration in China for children's emotional wellbeing. Soc. Sci. Res. 58, 46–67.
Roback, Jennifer, 1982. Wages, rents, and the quality of life. J. Polit. Econ. 90 (6), 1257–1278.

Rose, Elaina, 1999. Consumption smoothing and excess female mortality in rural India. Rev. Econ. Stat. 81 (1), 41–49.
Sen, Amartya, 1992. Missing women. BMJ Br. Med. J. (Clin. Res. Ed.) 304 (6827), 587.
Sieg, Holger, Yoon, Chamna, Zhang, Jipeng, 2021. Migration Controls, Fiscal Externalities, and Access to Educational Opportunities in China. Working Paper.
Song, Yang, 2014. What should economists know about the current Chinese hukou system? China Econ. Rev. 29, 200–212.
Sun, Ang, Zhao, Yaohui, 2016. Divorce, abortion, and the child sex ratio: the impact of divorce reform in China. J. Dev. Econ. 120, 53–69.
Sun, Meiping, Ma, Rui, Zeng, Yang, Luo, Fengji, Zhang, Jing, Hou, Wenjun, 2010. Immunization status and risk factors of migrant children in densely populated areas of Beijing, China. Vaccine 28 (5), 1264–1274.
Sun, Ang, Zhang, Chuanchuan, Hu, Xiangting, 2019. Boys, girls, and grandparents: the impact of the sex of preschool-aged children on family living arrangements and maternal labor supply. Demography 56 (3), 813–833.
Szatmari, Peter, Offord, David R., Boyle, Michael H., 1989. Ontario child health study: prevalence of attention deficit disorder with hyperactivity. JCPP (J. Child Psychol. Psychiatry) 30 (2), 219–223.
Tan, Zhibo, Wei, Shang-Jin, Zhang, Xiaobo, 2021. Deadly discrimination: implications of "missing girls" for workplace safety. J. Dev. Econ. 152, 102678.
The State Council of PRC, 2022. The Number of Migrant Workers Increases by 6.91 Million in 2021. https://www.gov.cn/xinwen/2022-04/29/content_5688045.htm. (Accessed 27 August 2023).
The World Bank, 2018. 2020. Data Bank: World Development Indicators. https://databank.worldbank.org/source/world-development-indicators. (Accessed 27 August 2023).
Tombe, Trevor, Zhu, Xiaodong, 2019. Trade, migration, and productivity: a quantitative analysis of China. Am. Econ. Rev. 109 (5), 1843–1872.
Treiman, Donald J., 2013. Trends in educational attainment in China. Chinese Sociol. Revi. 45 (3), 3–25.
United Nations Development Programme, 2009. Human Development Report 2009. https://hdr.undp.org/system/files/documents/human-development-report-2009-english.human-development-report-2009-english. (Accessed 27 August 2023).
United Nations. Economic Commission for Europe, 1976. Fertility and Family Planning in Europe Around 1970: A Comparative Study of Twelve National Surveys, vol. 58. New York: United Nations.
Waldfogel, Jane, Han, Wen-Jui, Brooks-Gunn, Jeanne, 2002. The effects of early maternal employment on child cognitive development. Demography 39 (2), 369–392.
Wang, Wendy, 2005. Son preference and educational opportunities of children in China—"I wish you were a boy!". Gend. Issues 22 (2), 3–30.
Wang, Qingfeng, 2019. Son Preference and Human Capital. Working Paper.
Wang, Shing-Yi, 2019. The Labor Supply Consequences of Having a Boy in China. National Bureau of Economic Research. NBER Working Paper 26185.
Wei, Shang-Jin, Zhang, Xiaobo, 2011. The competitive saving motive: evidence from rising sex ratios and savings rates in China. J. Polit. Econ. 119 (3), 511–564.
Weisner, Thomas S., 1989. Comparing sibling relationships across cultures. In: Sibling Interaction across Cultures: Theoretical and Methodological Issues. Springer New York, New York, NY, pp. 11–25.
Wen, Ming, Lin, Danhua, 2012. Child development in rural China: children left behind by their migrant parents and children of nonmigrant families. Child Dev. 83 (1), 120–136.
Wenlock, R.W., 1977. Birth spacing and prolonged lactation in rural Zambia. J. Biosoc. Sci. 9 (4), 481–485.
Xie, Yu, Hu, Jingwei, 2014. An introduction to the China family panel studies (CFPS). Chinese Sociological Review 47 (1), 3–29.
Yi, Fujin, Gu, Huangqian, 2015. The influence of discriminatory reimbursement rate of New Cooperative Medical Scheme to rural labor mobility. China Rural Survey 3, 2–15 (*in Chinese*).
Yue, A., Sylvia, S., Bai, Y., Shi, Y., Luo, R., Rozelle, S., 2016. The Effect of Maternal Migration on Early Childhood Development in Rural China.
Yue, A., Shi, Y., Luo, R., Chen, J., Garth, J., Zhang, J., et al., 2017. China's invisible crisis: cognitive delays among rural toddlers and the absence of modern parenting. China J. 78 (1), 50–80.
YuWa Population Research, 2022. A Report on the Child Raising Cost in China. http://download.caixin.com/upload/shengyubaogao2022.pdf. (Accessed 27 August 2023).
Zhang, Hongliang, Behrman, Jere R., Fan, C. Simon, Wei, Xiangdong, Zhang, Junsen, 2014. Does parental absence reduce cognitive achievements? Evidence from rural China. J. Dev. Econ. 111, 181–195.
Zhang, Cuiling, Liu, Hongyan, Wang, Ying'an, 2019. Shifts in second birth level and spacing following changes in fertility-limiting policies in Chinese provinces, 1984–2016. Popul. Dev. 25 (4), 11–22 (*in Chinese*).
Zhao, Xindong, Li, Xiang, 2018. Research on influence factors of migrant female population's birth interval: based on the data of migrant population dynamic monitoring survey in China 2016. Statist. Res. 35, 69–80 (in Chinese).
Zhao, Qiran, Yu, Xiaohua, Wang, Xiaobing, Glauben, Thomas, 2014. The impact of parental migration on children's school performance in rural China. China Econ. Rev. 31, 43–54.
Zhou, Qin, Liu, Guoen, 2016. The difference of health insurance: based on the study of the local population and migrants. Nankai Econ. Stud. 1 (*in Chinese*).