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Birth order and the gender gap in educational attainment

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Abstract While studies of birth order effects on human capital formation for developed countries abound, less is known about these effects in a developing country context. Harnessing rich childbearing history data on senior parents in China, I provide within-family estimates of the impact of birth order on adult children's completed schooling, emphasizing heterogenous effects across gender. I find evidence that, holding the size of the family fixed, a daughter's schooling decreases with the number of younger siblings, while a son's schooling increases with the number of younger siblings. Birth order differences in age at marriage and provision of intergenerational support to parents are possible explanations for the observed patterns in schooling. My findings suggest that the one-child policy, despite having contributed to worsening the sex-ratio imbalance in China, could have helped reduce the gender gap in educational attainment.

Keywords Human Capital · Gender Gap · Birth Order · Family Planning Policies

JEL classification J10 · J13 · J26 · O12 · O15

1 Introduction

A central problem in various fields of economics and other social sciences is to understand what drives intra-familial disparities in parental investment and,

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consequently, economic success across siblings.[1] Within this research agenda, an important question concerns the role of birth order in explaining differences in children's outcomes. Traditionally, studies addressing this question focused on developed countries and, with a few exceptions, concluded that being the earliest born confers many benefits.[2] More recently, however, attention has shifted to developing countries (e.g., Ejrnæs and Pörtner (2004), Edmonds (2006), Emerson and Souza (2008), Dammert (2010), Tenikue and Verheyden (2010), and De Haan et al. (2014)) and, interestingly, the findings suggest a reverse pattern: earlier-born children fare worse than later borns.[3]

In this paper, I contribute to this expanding literature by (i) estimating birth order effects on completed years of schooling in China, (ii) examining how the effects differ across gender, and (iii) investigating novel mechanisms that might give rise to these effects. Gender heterogeneity in birth order effects are plausible, especially in the Chinese context where traditional gender roles and other social norms are likely to shape parents' perception of the returns to resource investment in their children. Different from previous studies in developing countries I focus on adult children outcomes.[4] I harness rich retrospective data on childbearing history collected for a nationally-representative sample of senior Chinese women and their adult children. The data enables me to produce estimates rid of bias resulting from correlation between a child's birth order and any other unobservable attributes shared by siblings. I also utilize information on adult children's marriage market outcomes, their financial and instrumental support to old-age parents, and early-life parental time

[1] The seminal studies by Becker and Tomes (1976) Behrman et al. (1982) posit that children's outcomes are determined by differences in birth endowments and parental resource investments, such as schooling. Parental resource allocation can reinforce endowment differences across siblings or compensate for these differences. Other studies have explored how birthweight (Behrman et al. (1994), Rosenzweig and Zhang (2009)), birth order (Behrman and Taubman (1986)) and gender (Rosenzweig and Schultz (1982)) influence intra-household resource allocation by affecting the returns to resource investment.

[2] See Behrman and Taubman (1986), Hanushek (1992), Black et al. (2005), Conley and Glauber (2006), Kantarevic and Mechoulan (2006), Bjerkedal et al. (2007), Black et al. (2007), Zajonc and Sulloway (2007), Price (2008) Booth and Kee (2009), De Haan (2010), Monfardini and See (2010), Hotz and Pantano (2015). The findings indicate that first borns have higher levels of schooling, earnings, IQ, cognitive ability, lower incidence of teenage pregnancy, and receive more parental supervision. While the empirical literature appears to have reached some consensus regarding the sign and magnitude of birth order effects, less is known about the underlying mechanisms driving these effects. Previous studies have put forth a wide range of explanations (see Behrman and Taubman (1986)): financial and parental time constraints create unequal distribution of resources across birth order; lower-quality intellectual environment available to higher-order children because they are raised when the average IQ in the family is lower; later borns may have inferior genetic endowment, as egg and sperm quality deteriorates with parents' age; parents stopping childbearing after a problematic child could mean lower quality of later-born children; parents could engage in strategic parenting to make an example out of the first borns, driving them to monitor the oldest child more closely than the others (Hotz and Pantano (2015)).

[3] Possible explanations for positive birth order effects in poor countries offered by the literature are differences in comparative advantage in child labor across age (Edmonds (2006)), and limited parental resources pushing older children to paid work at the expenses of their schooling (Dammert (2010), De Haan et al. (2014)).

[4] Most studies examine birth order effects on young children's schooling and child labor (the exception is Ejrnæs and Pörtner (2004), which examine completed education). De Haan et al. (2014) also examines birth order effects in parental time inputs and child cognition.

inputs to elucidate some of the possible mechanisms underlying the birth order patterns by gender in China.

The data for this study come from the 2013 and 2014 waves of the China Health and Retirement Longitudinal Study (CHARLS). The 2013 CHARLS provides information on the date of birth of every child ever born, which allows me to accurately measure birth order among adult siblings. For every living child the survey collects data on the educational attainment, income level, living arrangement, and financial support given and received, among others. The 2014 CHARLS Life History survey gathers additional information on early-life outcomes such as early-childhood care, age at first marriage, and marriage partner characteristics. The CHARLS, therefore, provides a unique opportunity to explore channels leading to birth order effects by gender in the Chinese context.

Gender determination technology, coupled with stringent caps on number of births due to the one-child policy, has contributed to the sex-ratio imbalance in China. Sex-selective abortion also poses a threat to identification as gender is no longer randomly distributed across birth rank. To circumvent this problem, I carry out the empirical analysis on different samples depending on the senior mother's exposure to the policy and run a series of robustness checks on alternative samples and using different specifications. My preferred estimates are obtained using the adult children born to women above childbearing age when the family planning policy came into effect and before ultrasound technology was popularized.

My within-family estimates show that, holding family size constant, lowering the birth order of a child by one is associated with a 0.31 year reduction in schooling if the child is a female but a 0.26 year *increase* in the schooling if the child is a male. The results are similar when estimates are produced after splitting the sample between rural and urban residents. Results obtained using an alternative sample of younger children drawn from the 1982 Population Census of China confirm my main findings: birth order effects on literacy rates, primary school completion, and accumulated years of schooling are positive for girls and negative for boys. Additionally, birth order reduces the likelihood girls do housework but increases the probability for boys. The exception is paid work, where positive birth order effects are observed for boys and girls.

Next, I turn to possible mechanisms underlying gender differences in birth order effects. First, birth order effects are not significant among daughters from a subsample of educated urban parents, but negative birth order effects among sons are stronger in this group. This might suggest that resource constraint could lead to worse outcomes for older daughters but cannot account for the reverse pattern for sons. Second, older children marry at an earlier age and are more likely to marry young, which could explain the lower educational attainment of older daughters if marriage migration interrupts schooling. Finally, older daughters are less likely to help their senior parents financially than younger siblings and older sons transfer larger sums of money. Therefore, from an efficiency standpoint, parents might perceive the returns to investing in the schooling of older sons as larger and of older daughters as smaller. There is still the possibility that my findings are explained by preferential treatment of children according to birth order and gender. If that were the case, I should see the same patterns in other early-life outcomes. I find no evidence, however, of birth order differences in the likelihood of being cared for by both biological parents at the ages 0 to 5, nor on the likelihood that the mother was working at the time the child was born.

Assessing the differential role birth order plays in the human capital formation of boys and girls is key to evaluating the impact of family planning policies implemented throughout the developing world, as it introduces another margin by which family size influences outcomes. This is particularly true of China, where the ramifications of the one-child policy are yet to be fully understood. My findings suggest that, by curbing family size, the policy had a stronger positive effect on human capital formation of females relative to males; despite having contributed to the sex-ratio imbalance in China, it might have helped to reduce the gender gap in educational attainment among surviving children.

The next sections are organized as follows. "The Chinese context" section highlights the characteristics of the Chinese society and economy that are relevant to the study of birth order effects. "Data and empirical strategy" section presents the empirical strategy and provides a detailed description of the data used in the empirical analysis. "Effect of birth order on schooling" section shows the main findings and discusses possible threats to identification and choice of functional forms. "Channels" section sheds light on possible mechanisms explaining birth order effects by gender. "Concluding remarks" section concludes.

2 The Chinese context

In this section, I describe some important features of the Chinese economy and social norms that are essential for understanding how family size and birth order play a role in human capital formation.

The Chinese economy underwent profound transformations in the past decades, in part propelled by the reforms of the late 70s. In agriculture production, the organization of economic activity shifted from a communal system towards an individual responsibility system, which led to large gains in agricultural productivity (Mcmillan et al. 1989). China's GDP more than quadrupled during the 1978–1996 period, real per capita disposable income more than tripled in the urban areas and almost quadrupled in the rural areas (Yao 1999). Urbanization rose in the post-reform period. In 1978, <20% of the population lived in cities; now, it is >50% World Bank (2014). According to the World Bank, China is currently the third economy in the world, with its GDP a little over 10 trillion of US$.

With nearly 1.3 billion people, China is also the most populous country in the world. Its rapid population growth was one of the motives underlying the country's family planning policies. Measures to curb fertility were initiated around 1972, when couples were offered economic incentives to have fewer children with a birth spacing of at least 4 years apart (Qian 2009). In 1979, its more draconian form known as one-child policy was put in place. Under the one-child policy, individuals of Han ethnicity were restricted to having one child. Violators faced punishment in the form of large fines. It was not uncommon for women to be subjected to forced abortions and sterilizations (Banister 1987). Children born irregularly were often denied access to public health care and education. Under this policy, first-born children have a clear advantage over the later borns in families with more than one child.

In 1984, the Central government gave local governments the power to determine their own limits on fertility, generating regional differences in the policy mandates. While the one-child restriction was strongly enforced on urban couples, many rural parents were allowed to have a second child when the first-born child was a girl. In some local communities, only-child couples and couples working in risky activities were also granted permission to have a second child. Recently, it is estimated that the family planning policy imposes a one-child restriction on 35% of the population (urban residents), a 1.5-child limit on 54% (rural residents), and two or three child on the remaining 11% (residents living in remote areas) (Ebenstein 2010).

It is a well-known feature of the Chinese society that parents favor sons at the expenses of daughters, especially in rural areas. Traditional gender roles could be one explanation for the prevalence of son bias. In rural areas, male children provide labor to the family farm. Social norms also dictate that children should be the providers of financial support and other types of care to parents during old age. It is believed that this role is fulfilled mainly by sons, as daughters are expected to care for her husband's family after marriage. Son preferences have been identified as one of the causes underlying the phenomenon called "missing girls", which refers to the highly imbalanced sex-ratio at birth (Sen 1990). Studies have found that the one-child policy has caused an increase in the sex-ratio imbalance in China, particularly after ultrasound technology was introduced in the 1980s (Ebenstein 2010; Qian 2008; Johansson and Nygren 1991). Son bias influences fertility behavior and alters the gender distribution across birth orders. Because parents have strong preferences for sons, they are more likely to stop having children after having a boy, which leads to a higher sex-ratio among higher birth orders.

These unique features of the Chinese society create challenges to any attempt to estimate birth order effects by gender. By design, the one-child policy constrains the human capital development of later-born children due to the sanctions imposed on births occurring outside of the one-child rule. Additionally, the introduction of gender determination technology in the early 1980s fostered the practice of sex-selective abortion, rendering the assumption that gender is randomly distributed across birth order unrealistic. The availability of data on women above the child-bearing age at the time the policy was implemented allows me to overcome some of these challenges. I discuss my approach in more detail in the next section.

3 Data and empirical strategy

3.1 Longitudinal data on older households

I draw data from the 2013 and 2014 waves of the China Health and Retirement Longitudinal Study (CHARLS), which is a nationally-representative sample of senior Chinese residents. The CHARLS surveyed 10,257 households with members aged 45 years or older, residing in 150 counties or districts and 450 villages or urban communities across China. Senior respondents from sampled households provided detailed information on childbearing history, including the number of children alive and deceased, adopted, and biological. They also answered questions regarding the gender, the year, and the month of birth of each child. The 2013 CHARLS asks parents for a rich set of information on basic demographic characteristics of surviving children, including the highest level of completed schooling, income brackets, migration status, and marital status. Data on financial and time transfers seniors respondents received from each child are also available. The 2014 CHARLS Life History provides retrospective data on how the senior respondents raised their children (i.e., whether the child was cared for by grandparents at a young age, whether the child went to kindergarten at the age of 5, and mother's working status and number of maternity leave days when the child was born). In the childrearing history module, senior respondents informed the year in which their children were married and the age and educational attainment of their spouses. The data provide a unique opportunity to investigate birth order effects by gender on a wide range of outcomes, as well as to assess possible mechanisms underlying these birth order effects on schooling.

I restrict my sample to the surviving biological children born to senior respondents whose youngest child was born on or after 1979 when the one-child policy started. The children in the sample were then 34 years or older in 2013 and had already completed their schooling. Birth order is calculated using information on the year and month of birth of children reported alive in 2013. Because the survey collects data on both resident and non-resident children, it allows me to accurately measure birth order and avoid sample selection bias.[5] When information on gender or date of birth of a child is missing, I drop all the children born to that child's parents. I also exclude families with multiple births from the estimating sample.

Table 1 presents summary statistics of parents and children's characteristics. The sample is split according to the child's gender and a measure of the child's mother's exposure to the one-child policy. To construct this measure, I use the year of birth of the senior female respondents and compute the fraction of their childbearing years—ages 15 to 40—that falls under the policy.[6] I use 1979 as the starting date of the one-child policy.[7] Columns (1)–(9) show mean and standard deviations for the "before-one-child-policy" (BOCP) samples. [8] Columns labeled "Diff" report the difference in means between daughters and sons and the standard error of the difference in parentheses.

[5] Because the CHARLS collects data on birth date of alive and deceased children, it also allows me to measure the birth order of surviving children after factoring in the birth order of deceased siblings. The birth order effects and their gender patterns are robust to using alternative measures of birth order.

[6] Setting the childbearing years to 15 to 49 years reduces the sample size by nearly 80%. In "Effect of birth order on schooling" section, however, I present a robustness check to setting the childbearing years to 15 to 49 using an alternative sample.

[7] According to Qian (2009), China began providing couples with economic incentives to have fewer children around 1972 ("Later, longer, fewer" policy). In "Effect of birth order on schooling" section, I present estimates using an alternative sample of mothers past their childbearing years in 1972.

[8] Columns (1)–(2) summarize the sample of adult children born to women that were not exposed to the policy (that is, past the age of 40 in 1979). Columns (4)–(5) report numbers for children born to women who had <10% of their childbearing years exposed to the policy, and columns (7)–(8) for children born to women who had less than 25% exposed. Columns (10)–(11) present summary statistics for the "after-one-child-policy" sample (AOCP), which are children born to women with 25% or more of their childbearing years exposed to the policy.

Table 1 Summary statistics of sample of adult children by gender and test of difference in means

	BOCP (<1%)			BOCP (<10%)			BOCP (<25%)			AOCP (≥25%)		
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)	(11)	(12)
Variable	Sons	Daughters	Difference	Sons	Daughters	Difference	Sons	Daughters	Difference	Sons	Daughters	Difference
Parents' characteristics												
Mother's age	78.9 (4.36)	79.1 (4.67)	0.11 (0.34)	77.4 (4.72)	77.3 (4.97)	−0.063 (0.31)	74.6 (5.42)	74.5 (5.57)	−0.16 (0.29)	63.0 (3.24)	62.8 (3.31)	−0.18 (0.18)
Father's age	80.8 (4.27)	80.8 (4.98)	0.061 (0.35)	79.2 (4.82)	79.2 (5.68)	−0.058 (0.34)	76.7 (5.69)	76.6 (6.10)	−0.032 (0.31)	66.0 (4.24)	65.8 (4.10)	−0.24 (0.22)
Mother's education	1.58 (3.32)	1.61 (3.23)	0.032 (0.25)	1.73 (3.43)	1.93 (3.49)	0.21 (0.22)	2.23 (3.64)	2.40 (3.82)	0.17 (0.20)	2.70 (3.72)	2.48 (3.57)	−0.23 (0.20)
Father's education	3.83 (4.45)	3.99 (4.63)	0.16 (0.34)	4.12 (4.57)	4.15 (4.61)	0.028 (0.30)	4.59 (4.60)	4.65 (4.68)	0.050 (0.24)	5.33 (4.07)	5.56 (4.23)	0.24 (0.22)
Mother is Han	0.96 (0.20)	0.94 (0.24)	−0.021 (0.018)	0.96 (0.20)	0.94 (0.24)	−0.020 (0.016)	0.95 (0.21)	0.94 (0.24)	−0.014 (0.013)	0.94 (0.24)	0.95 (0.23)	0.0084 (0.013)
Father is Han	0.93 (0.26)	0.92 (0.27)	−0.0081 (0.029)	0.93 (0.25)	0.94 (0.25)	0.00065 (0.022)	0.94 (0.23)	0.94 (0.23)	0.0023 (0.015)	0.95 (0.21)	0.95 (0.22)	−0.0035 (0.013)
Agric hukou	0.70 (0.46)	0.70 (0.46)	−0.0049 (0.034)	0.71 (0.46)	0.71 (0.45)	0.0032 (0.030)	0.72 (0.45)	0.71 (0.45)	−0.011 (0.024)	0.80 (0.40)	0.74 (0.44)	−0.056 (0.022)
Child's characteristics												
Age	49.9 (7.02)	50.3 (7.16)	0.36 (0.25)	48.9 (6.88)	49.2 (6.99)	0.29 (0.22)	47.2 (6.69)	47.5 (6.76)	0.24 (0.17)	39.8 (3.85)	39.8 (3.91)	0.030 (0.12)
Years of schooling	8.48 (3.92)	6.58 (4.62)	−1.90 (0.15)	8.45 (3.86)	6.79 (4.50)	−1.66 (0.13)	8.55 (3.78)	7.02 (4.39)	−1.53 (0.11)	8.73 (3.50)	7.84 (4.18)	−0.89 (0.12)
No. of siblings	3.83 (1.47)	3.91 (1.46)	0.081 (0.052)	3.71 (1.47)	3.79 (1.44)	0.082 (0.046)	3.46 (1.45)	3.58 (1.42)	0.11 (0.037)	2.07 (0.98)	2.23 (1.05)	0.16 (0.032)
No. of young bro	1.02 (1.12)	1.09 (1.10)	0.066 (0.040)	0.97 (1.08)	1.08 (1.09)	0.10 (0.034)	0.90 (1.03)	1.01 (1.03)	0.11 (0.027)	0.53 (0.71)	0.69 (0.76)	0.16 (0.023)

Table 1 continued

	BOCP (<1%)			BOCP (<10%)			BOCP (<25%)			AOCP (≥25%)		
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)	(11)	(12)
Variable	Sons	Daughters	Difference	Sons	Daughters	Difference	Sons	Daughters	Difference	Sons	Daughters	Difference
No. of young sis	0.82 (1.00)	0.88 (1.01)	0.059 (0.036)	0.81 (0.99)	0.85 (0.99)	0.040 (0.031)	0.75 (0.95)	0.82 (0.98)	0.063 (0.025)	0.47 (0.68)	0.46 (0.70)	−0.0080 (0.022)
No. of old bro	1.05 (1.14)	1.02 (1.10)	−0.033 (0.040)	1.00 (1.10)	0.99 (1.07)	−0.012 (0.034)	0.92 (1.04)	0.90 (1.02)	−0.015 (0.027)	0.53 (0.71)	0.61 (0.75)	0.077 (0.023)
No. of old sis	0.93 (1.07)	0.92 (1.04)	−0.011 (0.038)	0.93 (1.05)	0.88 (1.02)	−0.047 (0.032)	0.89 (1.04)	0.84 (1.00)	−0.049 (0.026)	0.54 (0.76)	0.47 (0.70)	−0.072 (0.023)
Birth order	2.98 (1.63)	2.94 (1.63)	−0.043 (0.058)	2.92 (1.59)	2.87 (1.59)	−0.046 (0.050)	2.80 (1.53)	2.74 (1.52)	−0.055 (0.039)	2.07 (1.03)	2.08 (1.06)	0.0027 (0.033)
First born	0.22 (0.41)	0.23 (0.42)	0.011 (0.015)	0.22 (0.42)	0.24 (0.43)	0.017 (0.013)	0.24 (0.43)	0.25 (0.44)	0.018 (0.011)	0.35 (0.48)	0.35 (0.48)	0.0048 (0.015)
Last born	0.24 (0.43)	0.22 (0.42)	−0.019 (0.015)	0.25 (0.44)	0.22 (0.42)	−0.032 (0.013)	0.28 (0.45)	0.23 (0.42)	−0.050 (0.011)	0.38 (0.48)	0.33 (0.47)	−0.049 (0.015)
No. of observations	1704	1464	3168	2205	1906	4111	3175	2791	5966	2267	1758	4026
No. of families	373	337	710	493	458	951	751	710	1461	790	621	1411

Notes: Summary statistics calculated from estimating sample. Estimating sample excludes families with twins, with only one child, and those with children born after 1979. BOCP <1%, <10%, <25%, and AOCP ≥25% are subsamples of adult children born to mothers with <1%, 10%, 25%, and >25% of their childbearing years exposed to the one-child policy, respectively. Columns (1), (2), (4), (5), (7), (8), (10), and (11) present means and standard deviation (in parentheses); columns (3), (6), (9), and (12) present difference in means between daughters and sons and the standard error of the difference (in parentheses). Data source: CHARLS 2013

The average age among the adult children in the BOCP (<1%) sample is 50 years. The mean age among their senior mothers is 79 years.[9] Sons have on average 1.90 more years of schooling than daughters, and the difference is statistically significant. Except for schooling, sons and daughters in the BOCP (<1%) sample have similar characteristics. There are no significant gender differences in parental characteristics such as age, schooling, ethnicity, and hukou status. Additionally, the average age, number of siblings, number of younger and older siblings, and birth order are statistically the same. A look at the AOCP (≥25%) sample, however, reveals gender differences in other outcomes besides schooling: males have a smaller number of siblings, are less likely to be from families with agricultural hukou, and less likely to be the last-born child in the family.

3.2 An empirical model for estimating birth order effects

My estimation approach harnesses within-family variation to estimate birth order effects in schooling, so that the specifications presented below control for family size effects. Following Edmonds (2006) and Hotz and Pantano (2015), I begin my exploration by using a specification that imposes linearity across birth order

$$\begin{aligned}\text{Schooling}_{ijc} = \alpha + \beta\, \text{No.yng sib}_{ijc} + \delta\, \text{No.yng sib}_{ijc} \\ \times \text{Son}_{ijc} + \gamma \text{Son}_{ijc} + \lambda_j + \lambda_c + \epsilon_{ijc},\end{aligned} \tag{1}$$

where Schooling_{ijc} is completed years of schooling of child i in family j born in year c, No.yng sib_{ijc} is the number of younger siblings of child i in family j, Son_{ijc} is the child's gender, λ_j is a family fixed effect, and λ_c is a birth cohort fixed effect. Note that because I am exploiting within-family variation in the number of younger siblings, the size of the family (that is, the total number of siblings) is held constant.[10] With family fixed effects included in the model, birth order effects are identified from differences in schooling between a child and the child's next youngest sibling. The next youngest is a sibling with one fewer younger sibling but the same total number of siblings. So in a family of two children, the schooling of the first born is compared to the schooling of the second born; in a family of three children, the schooling of the second born is compared to the schooling of the third born, and the first born to the second born, and so on. The interaction with child's gender allows birth order differences to vary between sons and daughters of the same family.

Within-family estimates of β and δ do not suffer from bias due to family specific heterogeneity that correlates with the choice of family size, such as taste for quantity

[9] The advanced maternal age could raise concerns about sample selection if mortality rates are systematically correlated with mother's unobservable heterogeneity. The other samples feature younger mothers and are less susceptible to selective mortality. In the empirical analysis, I present the main empirical results for each of these samples separately. Furthermore, I address this issue in the robustness check section by reporting estimates obtained using data on young children born to younger mothers drawn from the 1982 Population Census of China.

[10] For studies on the effects of family size on children's outcomes, see Rosenzweig and Wolpin (1980), Rosenzweig and Schultz (1987), Hanushek (1992), Roy and Foster (1997), Miller (2009), Sinha (2005), Black et al. (2005), Cáceres-Delpiano (2006), Li et al. (2008), Qian (2009), Rosenzweig and Zhang (2009), Angrist et al. (2010), Black et al. (2010).

and quality of children. Estimating gender differences in birth order effects using Chinese data, however, may be proven challenging due to several reasons. First, in 1979 the Chinese government instituted the one-child policy which, in combination with prenatal sex determination technologies, some studies have shown led to an increase in the sex-ratio imbalance because of sex-selective abortion (Chen et al. 2013; Ebenstein 2010). As a consequence, one cannot expect child's gender to be randomly distributed across birth orders. I attempt to mitigate this problem by disaggregating the estimating samples according to the mother's exposure to the policy. My preferred results are estimated on a sample of children born to women who where past the age of 40 in 1979.[11]

Even if sex-selective abortion is not a concern, some threats to identification remain because of the prevalence of son preferences in the Chinese context. Unequal treatment of sons and daughters might lower survival rates among female children. Parents may also be more likely to stop childbearing after achieving the desired number of sons. In that case, females would be more likely to have a larger number of siblings and number of younger siblings. They would also be less likely to be the last child born in the family. I find no evidence that this is the case among my preferred sample. Table 1 shows that differences in family size and birth order between males and females in the BOCP (<1%) and BOCP (<10%) samples are small and not statistically significant. Additionally, Supplementary Appendix Table A.1 shows OLS and family FE regressions of the number of younger siblings on the child's gender after controlling for family size. Among the BOCP (<1%) sample, both OLS and family FE estimates show no association between gender and number of younger siblings. For all BOCP samples the family FE estimates also show no gender differences. The estimates using the AOCP sample, on the other hand, show strong negative association between gender and the number of younger siblings. This is expected as most parents in the AOCP sample were having children when the ultrasound technology was introduced.[12]

Using within-family variation in birth order by gender could introduce another source of bias because identification relies on families with at least two-girls/two-boys. Households with at least two girls may systematically differ from households with at least two boys (De Haan et al. 2014). To see if that is the case, Supplementary Appendix Table A.5 reports difference in the means of parental characteristics (age, years of schooling, ethnicity, and hukou status) between a subsample of families with at least two sons and another of families with at least two daughters. There are no significant differences between the two subsamples. Nonetheless, if offspring gender composition correlates with how parents treat

[11] Ultrasound technology was only available to the public during the 1980s and was first adopted in the urban areas (Chen et al. 2013). Therefore, I do not expect the mothers in the BOCP (<1%) sample to have had the ability to learn the gender of the child prior to birth. Additionally, I present results estimated on an alternative sample of mothers who where above the childbearing age in 1972 when family planning policies were first introduced in China.

[12] Supplementary Appendix Tables A.2, A.3, and A.4 confirm these findings using birth rank, first-born dummy, and last-born dummy as dependent variables. OLS and FE estimates indicate that females are not more likely to be first borns and males are not more likely to be last borns in the BOCP (<1%) and BOCP (<10%) samples.

some children relative to the others, this source of unobserved heterogeneity is not captured by family fixed effects models (Edmonds 2006). While there is no easy way to tackle this problem, I address this issue by proposing an alternative specification. In the absence of prenatal gender determination technology, the gender of the first-born child is arguably random.[13] An alternative approach to estimating gender differences in birth order effects compares the difference in educational outcomes between first and last borns in families with first-born males and first-born females

$$\begin{aligned} \text{Schooling}_{ijc} = \alpha + \beta \, \text{First born}_{ijc} + \delta \, \text{First born}_{ijc}, \\ \times \text{First} - \text{born son}_j + \gamma \text{Son}_{ijc} + \lambda_j + \lambda_c + \epsilon_{ijc}, \end{aligned} \tag{2}$$

where First born_{ijc} is a dummy that takes on value one if i is a first-born child from family j and born in cohort c, and zero if the child is the last-born child; First-born son_j is an indicator for whether the eldest child from family j is male. λ_j is a family fixed effect and λ_c is a cohort fixed effect. Son_{ijc} is an indicator for the child's gender, and it helps to control for differences between eldest-male and eldest-female families in the gender of the last-born child. Equation (2) is estimated on a sample of first- and last-born children.

4 Effect of birth order on schooling

Table 2 presents OLS and family FE estimates of Eq. (1) which omit the gender interaction. While the OLS estimates point to a strong negative association between the number of younger siblings and schooling after controlling for family size, within-family estimates do not show a relationship in any of the samples. Clearly, comparisons of educational outcomes across different birth orders that do not account for unobserved family heterogeneity will bias the results towards finding positive birth order effects.

All estimates presented subsequently include control for family fixed effects. Table 3 displays estimates of Eq. (1). The results presented in even-numbered columns add an interaction between the total number of siblings and child's gender as a robustness check. Allowing birth order effects to differ by child's gender generates a different set of conclusions. The estimates obtained on the BOCP samples indicate that, holding family size constant, daughters born earlier have fewer years of schooling than later-born siblings. Sons born earlier, however, have more years of schooling than later-born siblings. According to the estimates on the BOCP (<1%) sample, having one additional younger sibling is associated with a 0.31 year reduction in schooling for daughters, but a 0.26 year increase for sons. The estimates are robust to inclusion of an interaction between family size and child's gender, suggesting that the interaction between number of younger siblings and gender is not

[13] Indeed, the results from Supplementary Appendix Table A.3 suggest there is no association between gender and the likelihood of being a first born in the BOCP samples (the coefficient in very small in the BOCP (<1%) sample—0.00029). Furthermore, Supplementary Appendix Table A.6 reports summary statistics of first borns by gender. Among the BOCP (<1%) and BOCP (<10%) samples, first-born males and females do not differ significantly in parental observable characteristics. There is also no evidence that first-born females have more siblings on average than first-born males.

Table 2 The effect of birth order on completed years of schooling

	BOCP (<1%)		BOCP (<10%)		BOCP (<25%)		AOCP (≥25%)	
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
No. yng sib	−0.26*** (0.088)	−0.029 (0.11)	−0.22*** (0.077)	−0.041 (0.094)	−0.16** (0.065)	−0.049 (0.077)	0.15 (0.093)	−0.016 (0.12)
Son	1.87*** (0.15)	1.92*** (0.14)	1.70*** (0.13)	1.78*** (0.12)	1.52*** (0.10)	1.67*** (0.097)	0.80*** (0.11)	1.16*** (0.11)
Family FE	No	Yes	No	Yes	No	Yes	No	Yes
Mean education	7.23	7.23	7.36	7.36	7.53	7.53	8.22	8.22
No. of observations	3168	3168	4111	4111	5966	5966	4025	4025

Notes: Standard errors clustered at the household level in parentheses. Dependent variable is completed years of schooling. No. yng sib is the child's number of younger siblings. Estimating sample excludes families with twins, those with children born after 1979, and those with only one child. BOCP <1%, <10%, <25%, and AOCP ≥25% are subsamples of adult children born to mothers with less than 1%, 10%, 25%, and >25% of their childbearing years exposed to the one-child policy, respectively. Additional controls included but not reported (OLS: child's and mother's year of birth indicators, mother's education and rural/urban status, and family size; FE: child's year of birth indicators). Data source: 2013 CHARLS. Stars indicate statistical significance. ***<0.01, **<0.05, *<0.1

picking up possible family size effects on the gender differential in educational outcomes.[14,15]

In China, there are marked disparities in socio-economic status between urban and rural areas. In particular, the rural setting is characterized by stronger son preferences possibly due to heavier reliance on male children for farm labor and old-age security (Oliveira 2016; Choukhmane et al. 2014; Banerjee et al. 2014). Gender and birth order differentials in parental investment in human capital may be amplified in rural settings if these families face tighter resource constraints. Next, I check for differences in birth order effects between these two populations. Supplementary Appendix Table A.9 presents estimates obtained after splitting the adult children sample by their senior mother's rural/urban status. The gender pattern and magnitudes of birth order effects among the BOCP samples are similar to Table 3. In the AOCP sample, there birth order effects for female children are not present.

My estimates are in contrast to those produced by the literature on birth order effects in a developing country context. Parish and Willis (1993) use data from Taiwanese siblings and find evidence of positive birth order effects for both males and females. Edmonds (2006) finds evidence for Nepal that older daughters receive

[14] I also checked for the differences in birth order effects by gender across families of different sizes by adding interactions with the number of siblings. The estimates are robust to adding these interactions. The results are not reported but are available upon request.

[15] A remaining concern is the imperfect measurement of birth order when using the date of birth of surviving children to determine their ranking in the family. To check if it bias the main results, Supplementary Appendix Table A.7 reports estimates for a subset of families whose children were all alive in the survey year, 2013. The results are robust to excluding families with deceased children from the estimating sample. The estimates are also robust to using families with deceased children in 2013 but factoring in the birth order of dead children when calculating birth rank (Supplementary Appendix Table A.8 displays the results).

Table 3 Gender differences in the effect of birth order on completed years of schooling

	BOCP (<1%)		BOCP (<10%)		BOCP (<25%)		AOCP (≥25%)	
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
No. yng sib	−0.31*** (0.11)	−0.29** (0.12)	−0.30*** (0.10)	−0.28*** (0.10)	−0.29*** (0.082)	−0.27*** (0.085)	−0.17 (0.13)	−0.074 (0.13)
No. yng sib × Son	0.57*** (0.081)	0.54*** (0.084)	0.54*** (0.076)	0.51*** (0.078)	0.50*** (0.064)	0.45*** (0.067)	0.32*** (0.095)	0.14 (0.10)
Son	0.79*** (0.19)	0.48 (0.41)	0.75*** (0.17)	0.46 (0.35)	0.79*** (0.13)	0.50* (0.27)	0.80*** (0.14)	0.14 (0.23)
No. sib × Son		0.096 (0.11)		0.093 (0.096)		0.100 (0.079)		0.38*** (0.11)
$\beta+\delta=0$ (p-values)	0.03	0.03	0.02	0.02	0.02	0.03	0.26	0.62
Mean education	7.23	7.23	7.36	7.36	7.53	7.53	8.22	8.22
No. of observations	3168	3168	4111	4111	5966	5966	4025	4025

Notes: Standard errors clustered at the household level in parentheses. Family fixed-effects estimates. The p-value presented is for testing the statistical significance of $\beta+\delta$ (β is the coef. on No. yng sib; δ is the coef. on No. yng sib × Son). Dependent variable is completed years of schooling. No. yng sib is the child's number of younger siblings. Son is an indicator for whether the child is male. No. sib is the total number of siblings. Estimating sample excludes families with twins, those with children born after 1979, and those with only one child. BOCP <1%, <10%, <25%, and AOCP ≥25% are subsamples of adult children born to mothers with <1%, 10%, 25%, and >25% of their childbearing years exposed to the one-child policy, respectively. Additional controls included but not reported (child's year of birth indicators). Data source: 2013 CHARLS. Stars indicate statistical significance. ***<0.01, **<0.05, *<0.1

less investment in human capital (as they are more likely to do housework compared to younger siblings), but there is no evidence of birth order effects for sons. De Haan et al. (2014) find positive birth order effects for both genders, although the effect is stronger in families with first-born boys. The other studies do not draw a distinction between male and female children.

5 Robustness checks

5.1 Selective mortality

Selecting a sample of children born to mothers above childbearing age when the one-child policy came into effect could introduce a sample selection bias due to selective mortality. The older women that survived and were part of the survey might have different preferences for investment in children across birth orders than those in the population of interest. This is a serious concern that I attempt to address by providing further evidence using the 1982 Chinese Population Census.[16] The sample includes women that had all of their their childbearing years exposed to the one-child policy using 1979 as the starting year of the policy.[17] The young children in the estimating sample were aged 10 to 25 in 1982 and had at least one sibling. Table 4 shows within-family estimates of the effects of the number of younger siblings on several outcomes by child gender. All estimates add child's year of birth indicators. [18]

Columns (1) and (2) present estimates from linear probability models for the likelihood that the child is literate and the probability that a child older than 13 had completed primary school, respectively. The results are consistent with previous findings. An increase in the number of younger siblings, holding family size constant, is associated with a 2 percentage point reduction in the likelihood that a female child is literate, but a 2 percentage point increase in the likelihood that a male child is literate. The estimates are very similar for primary school completion. Column (3) uses the accumulated years of schooling as the outcome variable. The results again confirm previous findings. For daughters, having an additional younger sibling is associated with a 0.17 year reduction in accumulated years of schooling, whereas the same increase is associated with a 0.29 year increase for sons. [19] Columns (4) and (5) present additional estimates of the birth order effects on market work and housework. An increase in the number of younger siblings is consistently associated with worse outcomes for female children. They are more likely to work and do chores and less likely to be enrolled in school. For male children, a larger number of younger siblings

[16] I draw my sample from the IPUMS-International database (Minnesota Population Center 2015).

[17] These are women who were older than 40 in 1979. I get similar results from a sample of women older than 49 in 1979 (See Supplementary Appendix Table A.10).

[18] I also produced estimates, which add an interaction between the total number of siblings and the child's gender as a robustness check. The estimates of birth order effects by gender are virtually the same. The results are not reported but are available upon request.

[19] Note that while in Table 3 the outcome is completed years of schooling of adult children, in Table 4 it is the accumulated years of schooling of young children. This difference can explain why the magnitudes of the effects are different. Nonetheless, the estimates are qualitatively similar.

Table 4 Gender differences in the effect of birth order on younger children's outcomes

	Literate	Primary	Yrs. schooling	Work	Housework
	(1)	(2)	(3)	(4)	(5)
No. yng sib	−0.020*** (0.00074)	−0.021*** (0.00087)	−0.17*** (0.0071)	0.020*** (0.00061)	0.0023*** (0.00027)
No. yng sib × Son	0.040*** (0.00047)	0.043*** (0.00052)	0.46*** (0.0043)	−0.0052*** (0.00038)	−0.0044*** (0.00018)
Son	0.056*** (0.00083)	0.044*** (0.00097)	0.50*** (0.0077)	−0.031*** (0.00083)	−0.0066*** (0.00032)
$\beta + \delta = 0$ (p-value)	0.00	0.00	0.00	0.00	0.00
Mean dep var	0.91	0.91	7.45	0.62	0.01
No. of observations	1,260,791	1,029,946	1,260,791	1,253,095	1,253,095

Notes: Standard errors clustered at the household level in parentheses. Family fixed-effects estimates. The p-value presented is for testing the statistical significance of $\beta + \delta$ (β is the coef. on No. yng sib, δ is the coef. on No. yng sib × Son). Dependent variable in column (1) is an indicator for whether the child is literate; in column (2) it is an indicator for whether the child (older than 13 years) has completed primary school; in column (3) it is the accumulated years of schooling; in column (4) an indicator for whether the child is currently employed; in column (5) an indicator for whether the child is currently doing housework (and not at school). No. yng sib is the child's number of younger siblings. Son is an indicator for whether the child is male. Estimating sample includes young children from women who were older than 40 in 1979. Children were aged 10 to 25 in 1982 and had at least one sibling. Additional controls included but not reported (child's year of birth indicators). Data source: 1982 China Population Census. Stars indicate statistical significance. ***<0.01, **<0.05, *<0.1

is associated with a smaller likelihood of doing housework, although it increases the likelihood of engaging in market work.[20]

5.2 Birth order dummies

In this subsection, I check for the robustness of the main results to an alternative functional form that allows for non-linear birth order effects:

$$\begin{aligned}\text{Schooling}_{ijc} = \alpha + \sum_{r=2}^{6} \beta_r \text{rth} - \text{born}\,\text{dght}_{ijc} + \sum_{r=2}^{6} \delta_r \text{rth} - \text{born}\,\text{son}_{ijc} \\ + \gamma \text{Son}_{ijc} + \lambda_j + \lambda_c + \epsilon_{ijc},\end{aligned} \tag{3}$$

where rth-borndght$_{ijc}$ (rth-born son$_{ijc}$) is an indicator for whether child i is the rth-born daughter (son) from family j born in cohort c, λ_j is a family fixed effect, and λ_c is a cohort fixed effect. Son$_{ijc}$ is an indicator for male child. Table 5 presents the estimate of Eq. (3) when regressing a primary school completion indicator on birth order dummies using a sample of younger children drawn from the 1982 Chinese Population Census. To check for differences in gender patterns in birth order effects across family size, columns (2) to (6) break down the estimates by the total number of children in family. The new results reveal the same patterns I find in the linear specifications: later-born daughters are more likely to have completed primary school than earlier-born daughters while later-born sons are less likely to have finished primary school than earlier-born sons. The exception is families with six children, where the birth order patterns for daughters seem to be non-linear. These families, however, account for only 3.8% of the families in the sample.[21] I then run the same specification on the BOCP (<1%) and BOCP (<10%) samples of adult children drawn from the 2013 CHARLS. Supplementary Appendix Table A.13 reports the estimates.[22] The smaller size yields less precise estimates. Nonetheless, the results point to the same gender patterns.

5.3 Other threats to identification

In the context of a family fixed effects model, identification of birth order effects for sons (daughters) relies on families with at last two sons (two daughters); if birth order effects by gender interact with offspring gender composition effects, and parents can alter the gender composition, controlling for family fixed effects will not be enough to yield unbiased estimates of the former. There are a few reasons why I believe this source of bias might not be so severe. Because the women in my preferred sample (BOCP) were no longer in childbearing age when ultrasound technology became

[20] Because data are only available for children who reside with their parents at the time the census was conducted, birth order may be inaccurately measured. Supplementary Appendix Table A.11 presents estimates on a sample of younger children from mothers who had all their ever-born children co-residing in 1982. The results are qualitatively similar to those reported in Table 4.

[21] Supplementary Appendix Table A.12 presents estimates using the indicator for whether the child is doing housework (and not at school) as outcome. I find again that the likelihood of doing housework decreases with birth order for daughters but it increases for sons.

[22] Because of the small sample size, I do not present estimates separately by family size.

Table 5 Gender differences in the effect of birth order on younger children's primary school completion: birth order indicators by family size

	All fam.	2-child fam.	3-child fam.	4-child fam.	5-child fam	6-child fam
	(1)	(2)	(3)	(4)	(5)	(6)
2nd-born daughter	0.00049 (0.0014)	0.0047* (0.0026)	0.0050** (0.0023)	0.0023 (0.0026)	−0.00048 (0.0037)	−0.025*** (0.0062)
3th-born daughter	0.0085*** (0.0020)		0.016*** (0.0034)	0.020*** (0.0038)	0.022*** (0.0052)	−0.011 (0.0087)
4th-born daughter	0.016*** (0.0029)			0.033*** (0.0053)	0.048*** (0.0071)	0.0085 (0.012)
5th-born daughter	0.034*** (0.0050)				0.066*** (0.0098)	0.045*** (0.015)
6th-born daughter	0.046*** (0.015)					0.059*** (0.022)
2nd-born son	−0.032*** (0.0012)	−0.020*** (0.0024)	−0.027*** (0.0020)	−0.036*** (0.0024)	−0.042*** (0.0033)	−0.047*** (0.0054)
3th-born son	−0.045*** (0.0019)		−0.054*** (0.0033)	−0.064*** (0.0037)	−0.066*** (0.0051)	−0.084*** (0.0083)
4th-born son	−0.050*** (0.0027)			−0.094*** (0.0051)	−0.087*** (0.0070)	−0.12*** (0.011)
5th-born son	−0.062*** (0.0043)				−0.13*** (0.0092)	−0.14*** (0.014)
6th-born son	−0.068*** (0.011)					−0.17*** (0.019)
Son	0.15*** (0.0011)	0.083*** (0.0020)	0.12*** (0.0019)	0.17*** (0.0023)	0.22*** (0.0032)	0.26*** (0.0055)
Mean dep var	0.91	0.91	0.91	0.91	0.91	0.91
No. of observations	1,000,459	179,282	279,216	277,497	185,286	79,178

Notes: Standard errors clustered at the household level in parentheses. Family fixed-effects estimates. Dependent variable is an indicator for whether the child (older than 13 years) has completed primary school. Son is an indicator for whether the child is male. Estimating sample includes young children from women who were older than 40 in 1979. Children were aged 13 to 25 in 1982 and had at least one sibling and at most five siblings. Birth order by gender indicators presented. The omitted categories are 1-born daughter and 1-born son. Additional controls included but not reported (child's year of birth indicators). Data source: 1982 China Population Census. Stars indicate statistical significance. ***<0.01, **<0.05, *<0.1

widely available to the population, parents' ability to choose the gender composition of their offspring was very limited. Nonetheless, when son preferences are strong, parents might stop having children after the birth of boys, which systematically alters the gender composition of their offspring. The discussion from "Data and empirical strategy" section suggests that, at least in the before-one-child policy samples, the child's gender is uncorrelated with birth order and the likelihood that the child is a first- or last-born child. Furthermore, parents with at least two girls are not statistically different from parents who had at least two boys in terms of their education, hukou status, and ethnicity.[23] As an additional robustness check, I estimate Eq. (1) separately on a sample of children from families with at least two sons and a sample of children from families with at least two daughters. As Supplementary Appendix Table A.14 shows, the estimates of birth order effects by gender are very similar in the two samples.

To address additional identification concerns, I run an alternative specification that compares the schooling outcomes of first borns and last borns within the same family, and allows for the effect to depend on the gender of the first-born child, as explained in "Data and empirical strategy" section, Eq. (2). Comparing first born to last borns also avoids possible bias due to over-representing children from larger families in the base group. Table 6 presents the results. According to column (1), compared to the last-born sibling, first borns have 1.4 fewer years of schooling in families with first born females, while first borns have 0.6 more years of schooling in families with first-born males; these estimates control for the gender of the last born.[24] One worry might be that, even if the gender of the first child is arguably random, the family environment in which first-born daughters were raised is different from those of first-born sons. For example, first-born females may have a larger number of siblings, so that the interaction term could be capturing the impact of family size or birth spacing. Adding an interaction between the first-born dummy and family size and another between the first-born dummy and birth spacing does not change the conclusions for the BOCP samples but it does lead to insignificant birth order effects for females in the AOCP sample.

Finally, it is worth noting that couples in China received economic incentives to have fewer children before the OCP came into effect in 1979. Starting in 1972, the policy "Later, longer, fewer" rewarded parents who waited at least 4 years to have another child after a child was born (Qian 2009). Although ultrasound technology was not introduced to the public until the early 1980s, one might still be concerned that the family planning policies of the early 1970s intensified parental preferences for male children, which could bias the birth order by gender estimates presented earlier. Another concern is that women over 40 were still in childbearing age. Restricting my estimating sample to children born to CHARLS senior female respondents who were older than 40 in 1972 or older than 49 in 1979 would result in a very small sample. To circumvent this problem, I use information on siblings of CHARLS senior respondents drawn from the 2014 Life History survey. All senior

[23] See Supplementary Appendix Table A.5.

[24] It is worth noting that these estimates, although large in magnitude, are only statistically significant at 10% in the BOCP (<10%) sample. Supplementary Appendix Table A.15 reports results from the same specification using an alternative (larger) sample. In the larger sample the estimates are highly significant.

Table 6 Gender differences in the effect of birth order on completed years of schooling, sample of first- and last-born children

	BOCP (<1%)		BOCP (<10%)		BOCP (<25%)		AOCP (≥25%)	
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
First born	−1.47*** (0.48)	−1.53*** (0.52)	−1.31*** (0.41)	−1.27*** (0.44)	−0.96*** (0.30)	−0.92*** (0.32)	0.36 (0.27)	0.60** (0.30)
First born × First-born son	2.04*** (0.47)	2.04*** (0.47)	1.94*** (0.40)	1.94*** (0.40)	1.41*** (0.29)	1.41*** (0.29)	0.47* (0.28)	0.42 (0.29)
Son	0.20 (0.32)	0.19 (0.32)	0.23 (0.28)	0.23 (0.28)	0.51** (0.21)	0.50** (0.21)	0.78*** (0.20)	0.80*** (0.20)
First born × No. siblings		0.055 (0.15)		−0.033 (0.13)		−0.038 (0.11)		−0.36** (0.16)
First born × Age gap		0.77 (1.70)		0.51 (1.33)		0.34 (0.85)		0.31 (0.32)
$\beta + \delta = 0$ (p-value)	0.21	0.27	0.08	0.08	0.08	0.08	0.00	0.00
Mean education	7.23	7.23	7.36	7.36	7.53	7.53	8.22	8.22
No. of observations	1447	1447	1934	1934	2967	2967	2837	2830

Notes: Standard errors clustered at the household level in parentheses. Family fixed-effects estimates. The p-value is for testing the statistical significance of $\beta + \delta$ (β is the coef. on First born; δ is the coef. on First born × First-born son). Dependent variable is completed years of schooling. Estimating sample excludes families with twins, those with children born after 1979, and those with only one child. BOCP <1%, <10%, <25%, and AOCP ≥25% are subsamples of adult children born to mothers with <1%, 10%, 25%, and >25% of their childbearing years exposed to the one-child policy, respectively. First born takes on value equal to one if the child is a first born and zero if the child is a last born. First born son is an indicator for whether the child's mother had a male first born. No. sibs is the total number of siblings. Age gap is the age gap between first and last borns. Son is an indicator for male child. Additional controls included but not reported (child's year of birth indicators). Data source: 2013 CHARLS. Stars indicate statistical significance. ***<0.01, **<0.05, *<0.1

CHARLS respondents in the 2013 wave were asked to provide basic demographic information on every sibling who had survived until age 6, including the sibling's gender, year of birth, and highest level of education attained. I then use data on CHARLS senior respondents' education and their siblings' education to provide estimates of birth order effects in Eq. (2) for a sample of adult children born to women who were born on or before 1929 (older than 42 years in 1972). Supplementary Appendix Table A.15 reports the results. The estimates confirm previous findings. First-born daughters have on average 0.7 fewer years of schooling than their youngest sibling, whereas first-born sons have 0.5 more years of schooling than their youngest sibling. Restricting the sample to children born to mothers with no schooling yields a larger birth order effect for daughters. It is still possible that an eldest daughter's education is impacted by the fact that females have a larger number of siblings than males. However, the results are robust to including interactions between the first-born dummy and family size, as well as an interaction with the age gap between the eldest and youngest child in the family, as column (3) indicates.

6 Channels

6.1 Resource constraints

A family facing financial constraints may choose to allocate the household resources in a way that hurts the older daughters and benefits the older sons. The need to complement the household income would lead parents to allocate some of their children's time towards household production or paid labor. If the returns to schooling (or the perceived returns) of daughters are smaller than the returns to schooling of sons, and older children have a comparative advantage in household production, the older daughters' schooling would be negatively affected. At the same time, as the family diverts resources from daughters to sons due to the perceived higher returns to sons' schooling, the older sons would likely benefit more as they face less sibling competition for those resources. It is possible, therefore, that resource constraints create gender differences in birth order effects *within the same family*. In that case, heterogeneity in the magnitudes of within-family birth order differences across households facing different degrees of resource constraints could shed light on whether this is a possible channel at play.[25] With that in mind, birth order effects should be amplified in families with low socio-economic status (SES). I define as low SES the families whose senior mothers have no schooling and live in rural areas, and high SES the families whose senior mothers have some schooling and live in urban areas. Table 7 displays estimates for each group using the BOCP (<1%) sample. The results suggest that, relative to lower SES families, the disadvantage in schooling faced by older daughters in higher SES households disappears, but older sons seem to hold an even larger advantage over their younger siblings. Therefore, while I cannot rule out that financial constraint could explain the

[25] Among existing studies that have resorted to this empirical approach to search for suggestive evidence of a financial constraint channel are Black et al. (2005) and De Haan et al. (2014).

Table 7 Gender differences in the effect of birth order on completed years of schooling, by mother's socio-economic status

	Whole sample		Low SES		High SES	
	(1)	(2)	(3)	(4)	(5)	(6)
No. yng sib	−0.31*** (0.11)	−0.29** (0.12)	−0.38** (0.18)	−0.38** (0.18)	−0.100 (0.33)	−0.14 (0.33)
No. yng sib × Son	0.57*** (0.081)	0.54*** (0.084)	0.59*** (0.12)	0.59*** (0.12)	0.64** (0.25)	0.71*** (0.25)
Son	0.79*** (0.19)	0.48 (0.41)	1.58*** (0.30)	1.58** (0.68)	−0.86* (0.48)	−0.41 (0.82)
No. sib × Son		0.096 (0.11)		−0.0013 (0.17)		−0.18 (0.22)
$\beta+\delta=0$ (p-value)	0.03	0.03	0.23	0.22	0.15	0.13
Mean education	7.23	7.23	5.81	5.81	11.29	11.29
No. of observations	3168	3168	1563	1563	442	442

Notes: Standard errors clustered at the household level in parentheses. Family fixed-effect estimates. The p-value presented is for testing the statistical significance of $\beta+\delta$ (β is the coef. on No. yng sib; δ is the coef. on No. yng sib × Son). Dependent variable is completed years of schooling. No. yng sib is the child's number of younger siblings. Son is an indicator for whether the child is male. Estimating sample excludes families with twins, those with children born after 1979, and those with only one child. Estimates are obtained on a sample of adult children born to mothers with <1% of their childbearing years exposed to the one-child policy. Column (1) uses the entire sample; column (2) restricts sample to children born to mothers with no schooling and living in rural areas; column (3) restricts sample to children born to mothers with some schooling and living in urban areas. Additional controls included but not reported (child's year of birth indicators). Data source: 2013 CHARLS. Stars indicate statistical significance. ***<0.01, **<0.05, *<0.1

presence of positive birth order effects, this channel alone cannot account for the reverse pattern across gender.

6.2 Marriage outcomes

Marriage institutions interacted with traditional gender roles could give rise to positive birth order effects among females if older daughters are pushed to marry earlier than their sisters.[26] Next, I explore the relationship between birth order and age at first marriage.[27] In columns (1) and (2) from Table 8, I present the estimated coefficients on the number of younger siblings and its interactions with the child gender when the outcome variables are the child's age at first marriage and an indicator for marrying before the minimum legal age (22 for women and 20 for men). The results support the hypothesis that birth order impacts the time of marriage. More precisely, within-family comparisons reveal that lowering the birth order by one is associated with a 0.8 year reduction in the age at marriage for sons and daughters. Additionally, the decrease in birth order is associated with a 7.2 percentage points increase in the likelihood of marrying young, which is when most children are at school age. There is no evidence of gender differences in this effect.[28] The resulting impact of early marriage on educational outcomes, however, could be different for sons and daughters if marriage only disrupts schooling for females. In the Chinese context, daughters are expected to move to their husband's house after marriage,

[26] In a study conducted using data from four South Asian countries (Bangladesh, India, Nepal, and Pakistan), Vogl (2013) finds that the older daughters' educational attainment and marriage outcomes are hindered by the presence of younger sisters. His explanation for this effect is that parents will hasten to marry the eldest daughter to allow enough time to find grooms for her younger sisters. The practice of following the birth order when marrying children of same sex, according to the author, can have many explanations. Parents may want to marry all daughters while they are still young and able to command a better groom in the marriage markets. Additionally, given the prevalence of the practice, marrying daughters out of order could send a negative signal about the quality of the older sibling. An earlier study by Parish and Willis (1993) also finds that the presence of younger sisters is associated with worse educational outcomes and lower age at first marriage among Taiwanese women. These two studies underscore the importance of traditional marriage institutions in accounting for positive birth order effects in non-western societies.

[27] Vogl (2013) estimates the effect of younger sisters on older daughter's parental co-residence since he does not possess data on the age at marriage. The idea is that, since a female child moves out of her parents' home upon marriage, the presence of younger sisters should decrease the likelihood that an older daughter is currently living with her parents. I provide a more direct test by using the actual age at marriage, as in Parish and Willis (1993). Additionally, the availability of data on the age at marriage for all surviving children born to the same mother allows me to control for unobserved parental preferences for children's marriage outcomes.

[28] If I break down the number of younger siblings according to gender, I find that an additional younger brother is associated with a 2.3 percentage points increase in the likelihood of marrying early for both sons and daughters, whereas an additional younger sister is associated with a 2.2 percentage points increase for sons and a 3.9 percentage points increase for daughters. Although an additional younger sister has a larger effect on marriage outcomes for females than an additional younger brother, the difference is not statistically significant (p-value = 0.33). Therefore, I find no evidence of gender heterogeneity in the effect of birth order on age at marriage in the Chinese setting. These estimates are not reported but are available upon request; they should be interpreted with caution due to the endogeneity in sibling sex composition in the Chinese context.

Table 8 Gender differences in birth order effects on other outcomes

	Age marriage	Married young	Proximity	Made transf.	Log transf.
	(1)	(2)	(3)	(4)	(5)
No. yng sib	−0.81*** (0.21)	0.072*** (0.019)	0.022* (0.013)	−0.029** (0.014)	0.016 (0.042)
No. yng sib × Son	0.089 (0.11)	−0.014 (0.011)	−0.050*** (0.0086)	0.023** (0.0099)	0.061** (0.027)
Son	0.76** (0.31)	0.16*** (0.030)	0.38*** (0.024)	−0.24*** (0.025)	−0.0024 (0.066)
$\beta+\delta=0$ (p-value)	0.00	0.00	0.04	0.64	0.08
Mean dep var	23.48	0.23	0.17	0.66	6.51
No. of observations	2245	2245	3122	3328	2233

Notes: Standard errors clustered at the household level in parentheses. Family fixed-effects estimates. The p-value presented is for testing the statistical significance of $\beta+\delta$ (β is the coef. on No. yng sib; δ is the coef. on No. yng sib × Son). Dependent variable in column (1) is the child's age at first marriage; in column (2) it is an indicator for whether the child married before the minimum legal age of marriage (22 for women and 20 for men); in column (3) it is an indicator for whether the child lives close to parents (in the same household or in an adjacent dwelling); in column (4) it is an indicator for whether the child helped their parents financially in the previous year; in column (5) it is the log of transfer amount, conditional on positive transfers. No. yng sib is the child's number of younger siblings. Estimating sample excludes families with twins, those with children born after 1979, and those with only one child. Coefficients in columns (1)–(4) are estimated on the BOCP (<1%) sample; estimates in columns (4)–(5) obtained on the BOCP (<20%) to increase sample size due to loss of observations after controlling for child's income. Additional controls included but not reported (columns (1)–(3): child's year of birth indicators; columns (4)–(5): child's year of birth indicators and income). Data source: CHARLS 2013. Stars indicate statistical significance. ***<0.01, **<0.05, *<0.1

while sons continue to live with their parents. Marriage migration can contribute to disrupting the female's pursuit of education.[29]

6.3 Child-provided old-age support

Until not long ago many Chinese residents lacked access to a pension system. According to the 2000 Chinese Census, only 8.2% of the non-working rural population aged 60 and older had pension benefits (Wang 2006). Despite the coverage expansion that took place in the past 5 years, benefit levels are still low and account for a small percentage of the average income of rural residents (Oliveira 2016). As a result, senior parents rely heavily on their adult children to provide economic support throughout old age. Indeed, 82.5% of the senior respondents in the BOCP (>1%) sample received financial support from a child in 2012. If there are social norms in place which dictate that the responsibility for providing for parents during old age depends on the child's birth order and gender (or their interaction), parents may invest in their children's human capital along those dimensions in order to maximize old-age support.

To shed some light on this possible mechanism, columns (3), (4) and (5) from Table 8 present estimates of the relationship between birth order and old-age support.[30] Birth order patterns in proximity do not conform with the idea that Chinese parents are more likely to live close to the eldest son. Indeed, an increase in the number of younger siblings is associated with a 2.8 percentage points reduction in the likelihood a male child will live close to parents.[31] On the other hand, there is some weak evidence that older daughters are more likely to live near her parents than their younger siblings (although the coefficient in only significant at a 10%). The latter finding might seem at odds with the finding that older daughters marry earlier than their younger siblings. However, it is worth mentioning that the measure of parental proximity refers to the survey year, when the majority of the adult children (94%) are already married.

Turning to the likelihood of providing financial support, I find that non-resident older daughters are less likely to make cash or in-kind transfers to their parents (the likelihood decreases 2.9 percentage points with the number of younger siblings). While there is no indication of an association between birth order and the likelihood of providing support for sons, the numbers from column (5) suggest that, conditional

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[29] Indeed, female children are more likely to live outside of their parents' village/town than male children.

[30] The dependent variable in column (3) is an indicator for living close to parents (either in the same household or in an adjacent dwelling). The outcome in column (4) is an indicator for whether a non-resident child made cash or in-kind transfers to parents in the previous year and in column (5) it is the (log) amount of transfers, conditional on making a transfer. The estimates presented in columns (4) and (5) include control for the child's household income brackets. The child's income bracket is reported by the child's parents. I run the regressions in columns (4) and (5) on the BOCP (<20%) sample to increase the sample size after losing observations due to missing values for child's income.

[31] This finding agrees with the results in Konrad et al. (2002). The authors find evidence that in Germany a first-born child has a higher probability of living further way from parents, leaving the burden of caring for elderly parents to the second born child; the authors do not allow for gender heterogeneity in birth order differences. By contrast, Wakabayashi and Horioka (2009) conclude that in Japan it is the eldest-male child the one more likely to live with his elderly parents, which they attribute to traditions and social norms.

on making some transfer, older sons transfer larger amounts (the amount increased 6.1% with the number of younger siblings).

In summary, there is evidence that financial old-age support increases in birth order for females and decreases for males. If these findings reflect parents' expectations regarding a child's role in old-age support according to their gender and birth order, it makes sense that parents will invest more in educating the older male children in order to increase their earnings potential and, consequently, their ability to provide financial assistance.[32]

Finally, I cannot rule out the possibility that these birth order patterns in old-age provision result from preferential treatment of children according to birth rank and gender. If the eldest son has the highest status in the family, parents will favor them over the other children. In that case, they might feel more altruistic towards their parents and, consequently, provide more financial assistance.

6.4 Parental preferences for birth order

Another possible explanation for birth order effects on schooling is the differential treatment of children due to parental preferences for birth order and gender. In this case, one would expect these preferences to be reflected not only in their formal schooling but also in other early-childhood investments. I take a further look into the relationship between birth order by gender and other measures of parental inputs by exploring retrospective data from the 2014 CHARLS Life History survey. Table 9 presents the results. In column (1), the outcome is an indicator for whether the child was cared for by both biological parents at the ages 0 to 2; in column (2) it is an indicator for whether the child was cared for by both parents at the ages 3 to 5. Column (3) uses an indicator for whether the mother was working when the child was born, and column (4) an indicator for whether the child went to kindergarten when younger than 5 years old. There is no support for the hypothesis that children received differential early-childhood investments depending on their birth order and gender.

7 Concluding remarks

While a large number of quantitative studies of birth order effects has been produced for developed countries, little is known about the existence and magnitude of these effects in the context of poor countries, particularly in China. In face of the several efforts to curb family size put forth by policy makers in the developing world, more studies of the impact of birth rank on children's outcomes are warranted. Employing rich data on the childbearing history of older Chinese mothers to provide within-family estimates of birth rank on adult children's schooling, I find that birth order effects differ by child's gender. Controlling for family size, having a larger number

[32] Because proximity could be considered as a form of old-age support, the results in column (3) suggest that older daughters (sons) provide more (less) instrumental assistance to their parents. I argue that schooling, however, does not make a child more productive in providing instrumental support. On the contrary, it would raise the child's opportunity cost of time spent in elderly care.

Table 9 Gender differences in birth order effects on parental inputs

	Cared for by biological parents (1)	Mom worked (2)	Kindergarten (3)
No. yng sib	0.0082 (0.0058)	−0.00093 (0.010)	0.0010 (0.0070)
No. yng sib × Son	0.00076 (0.0040)	−0.0026 (0.0055)	0.0043 (0.0035)
Son	0.0035 (0.011)	−0.021 (0.013)	−0.0082 (0.0095)
$\beta+\delta=0$ (p-value)	0.16	0.75	0.45
Mean dep var	0.72	0.47	0.08
No. of observations	2837	2400	2826

Notes: Standard errors clustered at the household level in parentheses. Family fixed-effects estimates. The p-value presented is for testing the statistical significance of $\beta+\delta$ (β is the coef. on No. yng sib; δ is the coef. on No. yng sib × Son). Dependent variable in column (1) is an indicator for whether the child was at the care of both biological parents at the ages 0 to 5; in column (2) it is an indicator for whether the mother worked when the child was born; in column (3) it is in an indicator for whether the child was in kindergarten at the age of 5. No yng sib is the child's number of younger siblings. Son is an indicator for whether the child is male. Estimating sample excludes families with twins, those with children born after 1979, and those with only one child. The results are estimated on the BOCP (<1%) sample. Additional controls included but not reported (child's year of birth indicators). Data source: CHARLS 2013. Stars indicate statistical significance. ***<0.01, **<0.05, *<0.1

of younger siblings is associated with worse educational outcomes for female children, but better outcomes for male children. Birth order differences in age at marriage and provision of financial old-age support are possible culprits.

To illustrate the implications of these findings for evaluating family planning policies, I use my estimates to provide back-of-the-envelope calculations of the impact of the one-child policy on educational attainment for males and females. McElroy and Yang (2000) exploit variation in penalties for above-quota births and conclude that the policy led to a 0.33 reduction in average number of children per family. Another study by Oliveira (2016) uses CHARLS data and finds that an exogenous increase in the number of children results in a 0.64 year decrease in the average schooling of adult children born to senior Chinese parents. Altogether, these estimates suggest that the one-child policy can account for 15% of the difference in the average years of schooling of adult children in the BOC (<1%) and AOC (>25%) samples. When I combine those estimates with my estimates of birth order effects by gender, I conclude that the policy can account for 15% of the schooling difference between the BOCP and AOCP female samples, but only 5% of the difference between the BOCP and AOCP male samples. Overall, my back-of-the-envelope calculations suggest that the modest reduction in family size created by the one-child policy can account for roughly 23% of the observed difference in the male-female gap in educational attainment between the BOCP and AOCP samples.

In light of these findings, I conclude that the one-child policy led to stronger positive effects on human capital formation of females relative to males. The evidence seems to suggest that the policy, despite having contributed to the sex-ratio imbalance in China, helped reduce the gender gap in educational attainment among surviving children.

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Compliance with ethical standards

Conflict of interest The author declares that she have no conflict of interests.

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