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Yi Min,
Robert J.R. Elliott,
Eric Melander

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Patriarchal Norms and Missing Girls: Evidence from the Unintended Consequences of Chinese Reforms

Yi Min Robert J.R. Elliott Eric Melander*

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Abstract

In this paper we document how deep-seated cultural norms result in unintended demographic responses to two reforms in China: (1) the One-Child Policy and (2) land reform under the Household Responsibility System. Constructing and validating an index of historical patriarchal norms based on soil suitability for plough use and combining this with fertility and sex ratio data, we estimate the heterogeneous impacts of the reforms in a difference-in-differences framework. Results show that in counties characterised by strong patriarchy, where son preference is plausibly higher, families with a firstborn girl resisted population control and saw smaller decreases in second-child fertility after the implementation of the One-Child Policy. Moreover, land reform heightened the male-bias in the second-child sex ratio among firstborn-girl families, resulting in approximately 0.89 million missing girls in high-patriarchy counties as a result of norms-induced sex selection. Our results reveal an enduring legacy of historical norms and the importance of cultural forces as drivers of gender (in)equality.

Keywords: Patriarchal norms, son preference, One-Child Policy, land reform.

JEL Codes: J13, J16, N35, O15, Q18, Z13

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1 Introduction

Chinese economic development since the market reform and opening-up in the late 1970s has been remarkable. Equally remarkable is how quickly Chinese sex ratios became male-biased around the same time, appearing to be a counterexample to the generally positive relationship between economic development and gender equality observed elsewhere (Duflo, 2012).[1] The potential role of culture, in particular patriarchal norms that may give rise to a preference for sons, therefore warrants closer examination. A growing body of literature has demonstrated that geographical legacies, historical agricultural technologies, and linguistic structures have had an enduring effect on gender roles across societies (Alesina et al., 2013; Carranza, 2014; Hansen et al., 2015; Fredriksson and Gupta, 2023; Davis et al., 2024). While economic development and policy interventions may attenuate cultural persistence (Xue, 2024; Lippmann et al., 2020), not all cultural traits are easily altered (Giavazzi et al., 2019). This raises an important question: how do patriarchal cultural norms shape gender-based fertility decisions?

Importantly, in a Chinese context, economic development and rising male-biased sex ratios coincided with the introduction of both the One-Child Policy (hereafter, OCP) and land reforms (decollectivisation) under the Household Responsibility System (hereafter, HRS). Each reform was rolled out (at different speeds and locations) across China beginning in 1978.[2] The OCP imposed a legal constraint on fertility (only one child), and the HRS, by returning childbearing costs to farmers after the decollectivisation of agriculture, also restricted feasible fertility due to standard budget constraint arguments (Chen and Xie, 2025). Notably, the decline in fertility after the introduction of both policies has been shown to have unintended effects: sex selection, to have at least one son, became an increasingly preferred and economically viable strategy, resulting in a rising male-biased sex ratio (Ebenstein, 2010; Edlund et al., 2013; Almond et al., 2019). However, the deeper factors shaping such policy side effects, that is, patriarchal norms giving rise to son preference, remain under-explored (Norling, 2018; China Population and Development Research Center, 2021). The purpose of this paper is to investigate whether the behavioural responses to

[1] Appendix Figure A1 shows the evolution of the sex ratio and GDP per capita over the period 1960-2022, with the sex-ratio peaking around 2006 at 1.17 male births per female birth.

[2] One can observe in Figure A1 that in recent years the male sex ratio declined after the progressive relaxation of the OCP, although sex ratios have remained higher than the biological norm of 1.05 even after the Three-Child Policy was introduced in 2021.

the OCP and HRS are driven by deep-rooted patriarchal cultural norms. Answering this question is fundamental to understanding why Chinese sex ratios are so persistently male-biased even in the presence of remarkable economic development.

The first objective of our paper is to construct an index of patriarchy based on soil suitability for historical plough-use, in the spirit of Alesina et al. (2013), but adapted for the Chinese context by additionally considering wetland-rice suitability in robustness analyses (see Section 3 for further discussion). The result is a proxy for historical patriarchal gender norms that varies across disaggregated geographical units within China. Validation exercises, which correlate our index with present-day proxies of gender norms, show that our index of patriarchy captures dimensions of culture that persist to the present day.[3] The cultural nature of our index is further underscored by its persistence even in the face of positive income shocks driven by female productivity in tea cultivation (Qian, 2008). Combining our index with county gazetteer data from Almond et al. (2019) on the rollout of the OCP and HRS, we are able to estimate the separate impact of these reforms on fertility and sex ratios using a difference-in-differences framework. Given the quasi-randomness of the sex of the first child (Ebenstein, 2010; Li et al., 2011; Babiarz et al., 2019), we consider families with a firstborn girl as treated and compare them to families with a firstborn boy.[4] We then trace out this difference before and after each reform, with a focus on the differences between places characterised by high and low historical patriarchy (effectively a triple-difference) in their behavioural responses to the reforms.

We find that patriarchal cultural norms strongly influence the impact of the OCP and the HRS. Our first finding is that high-patriarchy counties have 19.2% higher second-child fertility among families with a firstborn girl after the implementation of the OCP, relative to low-patriarchy counties where deep-rooted son preference is plausibly less prevalent. Back-of-the-envelope calculations suggest that the interaction of the OCP with underlying patriarchal norms resulted in approximately 4 million additional second births. Our second finding is that the HRS increases the male-bias in second-child sex ratios by around 3.9 percentage points in families with a firstborn girl. Again, this effect is driven by high-patriarchy counties, confirming the sex selection mechanisms of Almond

[3] We also cross-validate our suitability-based index against the Patriarchy Index of Szołtysek et al. (2017) and find a strong correlation between our index and the gender dimensions of patriarchy used in the construction of the Patriarchy Index.

[4] We validate this quasi-randomness descriptively in Section 3.

et al. (2019). We calculate that the interaction between the HRS and the underlying son preference in high-patriarchy places resulted in around 0.89 million missing girls.[5]

This paper makes three main contributions. First, we demonstrate how deeply embedded cultural values can shape the impact of large-scale economic reforms such as the OCP and HRS. While previous work has estimated the direct impact of the OCP and HRS on fertility and sex ratios (Almond et al., 2019; Chen and Xie, 2025), there has been little research unpacking the norms that give rise to such policy responses and explaining why these responses were substantially stronger in particular places, especially for the male-skewed sex ratio (Jia and Kung, 2025). Notably, work from a diverse range of contexts shows that women's economic empowerment is not straightforwardly followed by broader empowerment, and may even have the reverse effect (Cheng et al., 2022; Bastagli et al., 2016; Bhalotra et al., 2020; Rosenblum, 2015). Thus, our findings suggest an important role of underlying culture in shaping the efficacy of reforms.

Second, we provide a new perspective on gender norms in contemporary China by examining how norms vary across space by tracing their deep historical roots. Existing studies have explained the variation in son preference with reference to ideal family sizes and gender composition (Norling, 2018; China Population and Development Research Center, 2021). In this paper we examine how the variation in historical patriarchal norms serves as a deep-rooted cultural mechanism influencing son preference even within the same ethnicity, and how cultural heterogeneity manifests itself in the presence of stringent institutional constraints, that is, patriarchal norms continue to shape varying behaviour under the two reforms.

Third, our work speaks to the phenomenon of missing women, first described by Sen (1990). Subsequent studies have emphasised the key role of declining fertility and access to pre-natal sex-selective abortion in driving the rise in the male sex ratio (Bhalotra and Cochrane, 2010; Jayachandran, 2017). The issue is particularly acute in China, where the majority of missing women are lost at birth (missing *girls*), rather than during adulthood, in contrast to patterns observed in countries such as India (Anderson and Ray, 2010). However, few studies have investigated the cultural determinants of the geographical distribution of missing women within a single country, especially in the Chinese context. Examples of studies documenting the spatial patterns of missing women in other

[5] This magnitude is economically meaningful given that the estimated number of missing girls at the second birth in 1990 was approximately 2.72 million (Ebenstein, 2010).

geographies include Mavisakalyan and Minasyan (2023) for Azerbaijan and Fenske et al. (2025) for India. By highlighting the differential responses to economic reform induced by differences in deep-seated cultural norms, we provide further evidence that economic measures alone may prove insufficient to achieve gender equality and that shifting cultural norms may be equally important.

2 Historical and Institutional Background

2.1 Patriarchy in China

The term patriarchy, from the Greek "rule of the father", is a hierarchical system where men are privileged by virtue of their biological sex and use this to control and dominate women (Millett, 2016). This system is typically embedded in traditional family institutions, which manifest patriarchy through dual hierarchies of generations and gender, namely, the power of seniors over juniors and men over women (Santos and Harrell, 2016). In this paper, we focus on the gender dimension.

Patriarchy in China dates back to the Zhou Dynasty (roughly 1046 BCE to 771 BCE), during which family and state were closely linked under ritual propriety (*Li*). This included a system of primogeniture (Lee, 1999), giving rise to a gendered division of labour with female roles historically confined to household production and domestic chores in the private sphere (Walby, 1990). Confucianism, which upheld this system, became the cultural orthodoxy during the Han Dynasty (roughly 206 BCE to 220 CE) and later served as the foundation for the imperial examination system used to select government officials (Chen et al., 2020). Within this framework, the absence of a male heir was regarded as an extinction of the lineage and a violation of filial piety. Women were expected to adhere to their prescribed roles, that is, to subordinate to their fathers before marriage, to their husbands after marriage, and to their sons after their husband's death (Liang et al., 2021). Chastity was regarded as a moral obligation and upheld as a model of virtue, with widow remarriage condemned as a transgression under patriarchal norms (Jia and Kung, 2025).

In the 1950s, family reforms initiated by the Chinese Communist Party (CCP) sought to liberate women from the family and empower them through the introduction of a range of gendered policies, such as the New Marriage Law (right to divorce) and equal access to education (Qian, 2024). During the Cultural Revolution (1966 to 1976), Chairman Mao introduced the slogan "women hold up half the sky". Although this movement was still rooted in patriarchal norms, it allowed women into traditionally male domains based on male-centred standards, and officially brought women into the public sphere, significantly increasing female participation in the labour force and in politics (Stacey, 1975; Lee, 1999). However, this did not lead to full gender equality in China (Evans, 2021) and women were still largely associated with the domestic sphere.

In modern China, traditional gender roles persist, with women bearing a greater share of domestic responsibilities, even when they participate in the labour market alongside their husbands or earn higher relative incomes (Hu, 2018; Evans, 2021; Hancock et al., 2025). Wage gaps, occupational segregation and under-representation in political institutions remain despite women's increased access to education (Benstead, 2021). Further, the emphasis on male lineage has led patriarchal norms to take on more implicit and subtle forms in the private sphere. For example, the practice of naming children with the mother's surname instead of the father's (especially for women from families without sons) ostensibly can be interpreted as a challenge to patriarchal norms, but in essence often serves to preserve the grandfather's lineage aspirations (Qi, 2018). The motivation to preserve the lineage not only encourages higher fertility rates to ensure at least one son (thereby exacerbating imbalanced sex ratios), but also amplifies the gender gap in education and labour force participation (Zhang and Ma, 2017; Zhang, 2022; Huang et al., 2025).

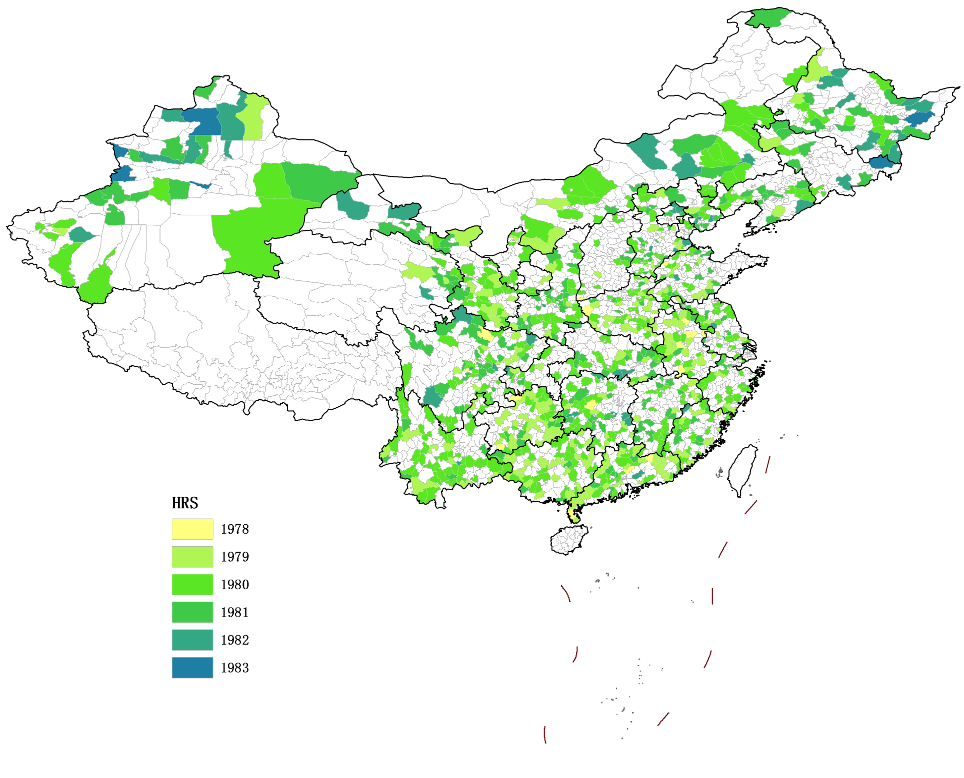
2.2 Land reform under the Household Responsibility System

In 1978, farmers in Xiaogang Village, Anhui Province, were the first to experience land reform under the "Household Responsibility System" (HRS), where farmers became responsible for their own profits and losses by contracting land with the collective (Zhou, 1996). Collective land remained publicly owned, but use rights were contracted to individual households. While this was seen as a reversal of socialism (Lin, 1992), the increase in output that year prompted the central government to allow the HRS experiment to expand into some of the poorest areas of China. However, central guidelines were largely ignored, and counties soon adopted the HRS in response to underlying economic forces (Lin, 1987). In 1982, the HRS was established as a national policy, officially reversing the policy of agricultural collectivisation that had been in place since 1953. Figure 1 shows the rollout of the HRS at the county level between 1978 and 1983, with the majority of counties (93.76% in our sample) having adopted the policy by 1981 (before the official announcement).[6]

[6] See Section 3.1.1 for further details of the sample of counties used in Figures 1 and 2.

Appendix B provides additional details on the commune system that was replaced by the HRS.

Figure 1: The rollout of the HRS



Notes: Figure 1 illustrates the rollout of the HRS. Black and gray lines outline provinces and counties, respectively. White areas are those where no data are available.

Following the bottom-up land reform under the HRS, households and individuals became the basic units responsible for agricultural production. Peasants obtained the right to use the land and could take full advantage of their rights as residual claimants, with increased procurement prices within the state commercial channel and newly opened free market.[7] Land reform significantly increased agricultural productivity and rural revenues (McMillan et al., 1989).

In the context of diminished government intervention, the cost of childbearing also returned to the individual household and away from communal childcare, potentially influencing fertility choices (Chen and Xie, 2025). Consequently, sex ratios may have changed as a results of possible

[7]For example, in 1979 the weighted average price within the state commercial system increased by 22.1% (Lin, 1992).

gender differences in the price elasticity of children. In addition, patriarchal cultural norms may generate additional utility for the birth of a son, which makes the demand for sons less elastic to prices than the demand for daughters. Notably, a family has several routes to satisfy a son preference. First, through increased fertility to ensure at least one son, and second, through sex selection technology (if available) for a given level of fertility. With the HRS there is a reduction in the economic feasibility of increased fertility as the cost of childbearing is no longer borne by the entire commune but by the individual, although this effect may be offset by an increase in household income.

2.3 China's One-Child Policy

Following agricultural collectivisation, China's population experienced rapid growth (rising from 588 million in 1953 to 830 million in 1970).[8] After the Great Famine (1959–1961), the total fertility rate in China remained above six throughout the 1960s, far above the replacement fertility rate of 2.1.[9] Consequently, a series of population policies were implemented. In 1973, the National Family Planning Conference advertised the campaign of "Later, Longer, Fewer".[10] In 1978, family planning was formally incorporated into the Chinese constitution, which initiated the gradual implementation of the One-Child Policy, restricting most couples to one legal child.[11,12] Figure 2 shows the rollout of the OCP across counties, with more than half of the counties in our sample adopting the OCP by 1979.

In practical terms, illegal children were excluded from free public education, and their parents faced financial penalties (Ebenstein, 2010). Specifically, unauthorised children in rural areas were denied grain rations during the era of agricultural collectivisation. Even in the period when private plots were briefly permitted, household landholding was not expanded to accommodate increased family sizes when the increase was due to an illegal birth. After the introduction of the HRS, population control measures were further escalated by establishing a "dual contract" system which

[8] Data are from National Bureau of Statistics of China. See `https://data.stats.gov.cn`.

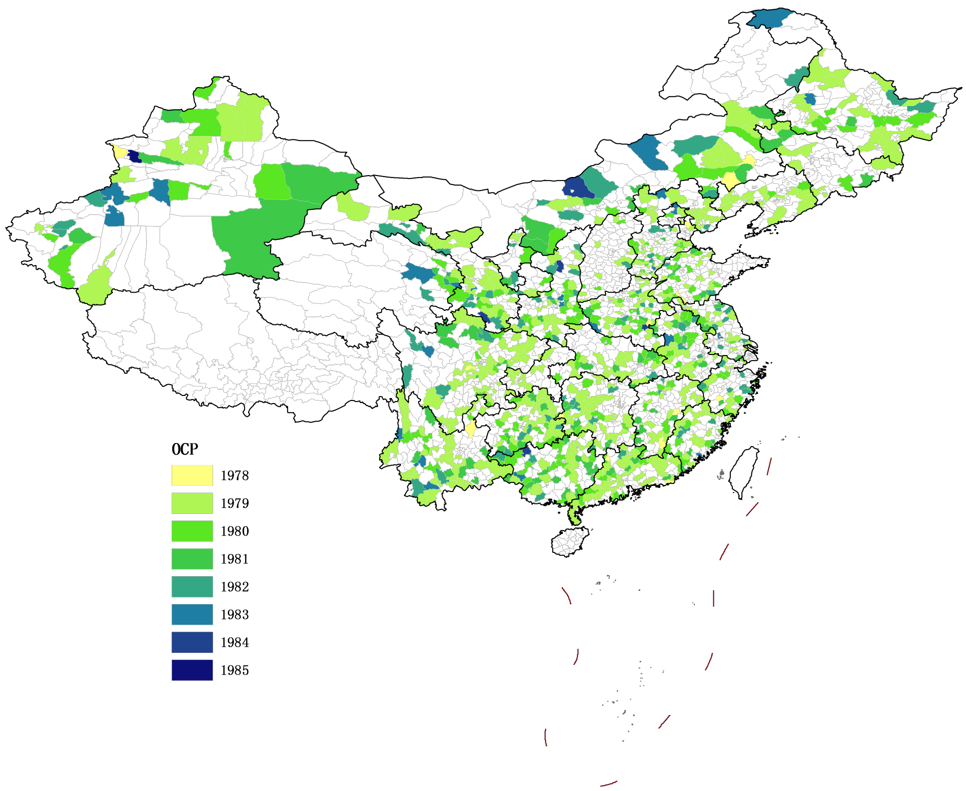
[9] Data are from the World Bank. See `https://data.worldbank.org/indicator/SP.DYN.TFRT.IN?end=1970&locations=CN&start=1960`.

[10] "Later" encouraged men to marry after the age of 25 and women after 23, with childbirth postponed until at least 24. "Longer" refers to a birth spacing of no less than three years, while "Fewer" encouraged couples to have no more than two children.

[11] A description of the evolution of China's family planning policies can be found at `https://www.gov.cn/zhengce/2015-02/09/content_2816919.htm`.

[12] Minority ethnic groups were not exempt from the OCP until 1984 (Scharping, 2013, p. 150).

Figure 2: The rollout of the OCP



Notes: Figure 2 illustrates the rollout of the OCP. Black and gray lines outline provinces and counties, respectively. White areas are those where no data are available.

integrated the family planning system with the production responsibility system (Banister, 1987). For example, Shaodong County tightened its coercive policies in 1982 such that farmers who had a second child were denied access to contract land for 14 years and were required to pay an excess birth fee (Editorial Committee of Shaodong County Gazetteer, 1993, p. 76).

Such stringent policies were likely to reduce fertility significantly, given the implied additional cost of having further children. However, the magnitude of the fertility decline and compliance with the OCP may be influenced by patriarchal norms, as these constitute an important factor in determining the price elasticity of children (of both genders). Further, the OCP plausibly strengthened the degree of substitutability between fertility and sex selection technology, as it significantly undermined the economic viability of relying on high fertility to achieve son preference by chance (Zeng, 2007).

In the late 1970s and early 1980s, population control officials introduced portable ultrasound machines to hundreds of cities for diagnostic use, which enabled sex identification at approximately 15 weeks of pregnancy and, as an unintended consequence, facilitated sex-selective abortions (Chen et al., 2013). Although such equipment was not deployed in the majority of rural areas, farmers could still travel to provincial capitals to access ultrasound technology (Almond et al., 2019).[13] Notably, abortions for births exceeding the quota could be obtained for free, with forced sterilisation and abortion serving as key measures to limit population growth under the OCP (Banister, 1987). The policies that co-existed with the OCP had the effect of reducing the effective cost of sex selection. Thus, the male sex ratio could plausibly increase after the OCP and more so in areas strongly influenced by patriarchal norms underpinning son preference.

3 Data and Descriptive Evidence

This section begins with a description of our data sources and the construction and validation of our index of patriarchal norms. Summary statistics for our various datasets are reported in Appendix Table C1. Subsequently, we turn to an exploration of basic descriptive patterns in the data, which we go on to investigate more formally in Section 5.

3.1 Data

3.1.1 One-Child Policy and Land Reform

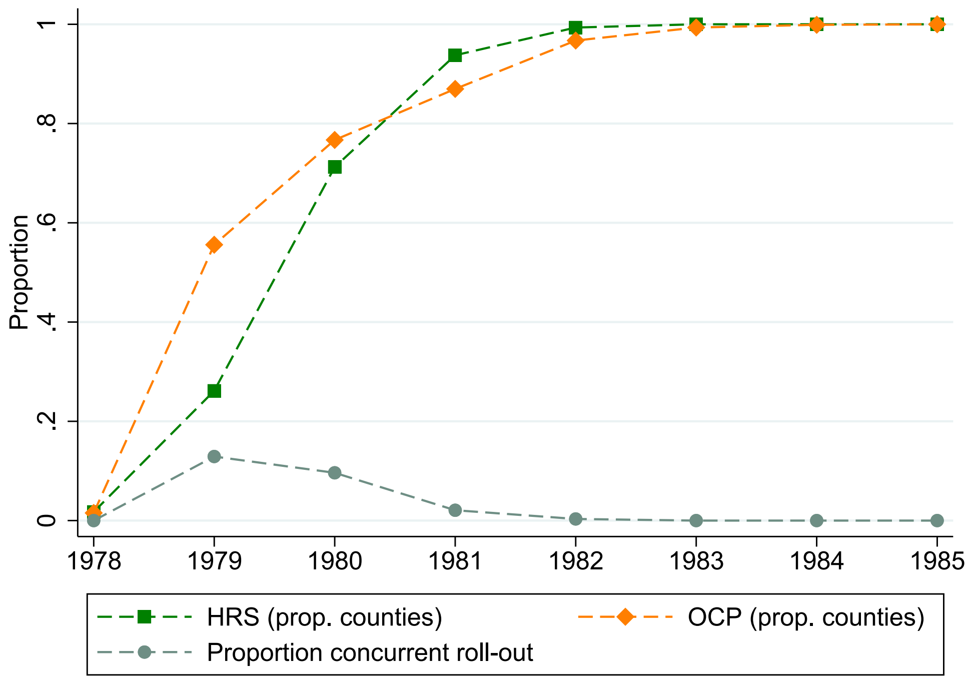
We draw on data from Almond et al. (2019) and compile county-level information on the OCP and HRS rollout from county gazetteers, covering 914 counties (around half of all counties in China) for which precise reform timing records are available.[14] Although ostensibly the policies happened at the same time, the staggered introduction of both policies means there are sufficient non-overlapping periods such that we are able to distinguish their separate effects, as shown in Figure 3. Here, green squares and orange diamonds plot the cumulative proportions of counties that adopted the OCP and HRS, respectively, over time. The grey circles indicate the proportion of counties that implemented both policies in the same year. As can be seen, rollout patterns are

[13] While not trivial, the cost of travelling to the provincial capital was much lower than the potential costs under the "dual contract" system (Almond et al., 2019).

[14] Almond et al. (2019) show that the 914 counties for which rollout data is available are comparable to other counties in terms of fertility and sex ratio patterns.

fairly distinct and in the majority of cases counties implemented the OCP before the HRS but not always. By 1983 the proportion of counties that had implemented both policies was effectively one.

Figure 3: The staggered rollout of two reforms: OCP and HRS



Notes: Figure 3 reports the proportion of counties in our sample in which the Household Responsibility System (HRS) and One-Child Policy (OCP) are rolled out, plotted using green squares and orange diamonds, respectively. Grey circles plot the proportion of counties in which the rollout of the two reforms was concurrent (occurring in the same year).

3.1.2 Fertility and Sex Ratios

Helpfully, the Almond et al. (2019) dataset also includes microdata from the 1 percent sample of the 1990 population census which are used to obtain information on the number and sex ratio of newborns between 1974-1986.[15] Migrants (0.63%) are excluded in order to approximate children's county of birth using their county of residence, as the census records only the location of enumeration, not an individual's birthplace.[16] The only difference between our sample and that of

[15] Focussing on births between 1974-1986 means we consider individuals aged 4-16 in 1990. As argued in Almond et al. (2019), this is to circumvent potential under-reporting of children under the age of 4.

[16] This approximation is reasonable as internal migration was strictly controlled (Almond et al., 2019). Migrants are classified as such in the 1990 data if they do not live in the same county as they did in 1985.

Almond et al. (2019) is that we focus on rural samples. Our focus on a rural sample is because this better captures the range of geographical locations affected by the HRS and is where agricultural production was reshaped by shifting incentives at the household level. The result is that we exclude only around 5 percent of second births. In robustness checks we show that the results are robust to retaining the urban sample.

3.1.3 Patriarchal Norms

As argued by Boserup (1970) and validated empirically on a global scale by Alesina et al. (2013), pre-industrial agricultural practices determined gendered divisions of labour and have shaped present-day gender norms. Following Alesina et al. (2013), we draw on this intuition to construct an index of patriarchal cultural norms based on the historic environmental suitability for use of the plough in agriculture.

We select three crops, wheat, barley and rye, as the representative crops that benefit the most from plough use to improve outputs (so-called "plough-positive" crops), in accordance with Alesina et al. (2013).[17] In the Chinese context, however, we need to consider the role played by wetland rice, which is the staple source of calories in southern China.[18] The advantage of using the plough in wetland rice cultivation relates to the role it plays in the double-cropping (and even triple-cropping) system, where the fertility of the soil can recover more quickly following shorter fallow periods if there is repeated ploughing to turn over the ground. However, the use of the plough for wetland rice cultivation deviates from the standard biological rationale regarding upper-body strength requirements for dry land ploughing, an attribute in which men generally exhibit a physiological advantage (Pryor, 1985). As a result, in some versions of our index, we include wetland rice as a plough-positive crop in our relative suitability measure to perform sensitivity analyses. The index excluding wetland rice nevertheless remains our baseline.[19]

Historically, the plough has been used in China since the Shang (ca. 1600 to 1046 BCE)

[17] Plough-positive crops are those that require not only extensive land preparation but also a considerable surface area to produce the food calories necessary to feed a family, or other crops that require the land to be prepared in a very short period of time (Pryor, 1985).

[18] These crops (wheat, barley, rye and wetland rice) accounted for approximately 73% of the total sown area of cereals in 1978, and 49% of the total sown area of all crops. Data are from *Statistical Compendium of China's Rural Economy, 1949–1986*.

[19] This is both to ensure consistency with the prior literature and to account for the fact that wetland rice is still regarded as less plough-positive than other plough-positive crops (Pryor, 1985).

and early Zhou (ca. 1046 to 256 BCE) Dynasties, initially in the forms of *Lei* and *Si*. Over time, the technology was improved to enable deeper tillage, for example, through the use of iron around 500 BCE, and the introduction of heavy triangular and quadrilateral frames during the Han Dynasty (206 BCE to 220 CE) (Bray, 1978). The centrality of the plough in traditional Chinese agriculture is symbolised by the imperial ritual of ceremonial ploughing, in which the emperor prays for a bountiful harvest on behalf of the nation (Armstrong, 1943). This ritual persisted until the fall of the Qing Dynasty in 1911 (Bishop, 1936). Such plough-based agricultural systems were heavily dependent on upper body strength in which men had a physiological advantage, and hence household production and consumption reinforced women's subordinate status (Macfarlane, 1978).

Data on crop suitability was retrieved from the Food and Agriculture Organization's Global Agro-Ecological Zones (GAEZ v4) database and are derived from climate, soil and terrain characteristics, under rain-fed and low-input conditions.[20] Using the underlying raster data (demarcated at 5 arc-minutes by 5 arc-minutes grid-cells), we were able to compute the average suitability for each crop at the county level for 901 counties in our sample.[21]

The index is constructed as the ratio of average plough-positive crop suitability to average cereal suitability, where the underlying rationale is that plough-positive crops would historically have been cultivated with the plough due to their higher comparative productivity relative to other cereals.[22] The suitability class ranges from very suitable (1) to not suitable (8), with smaller values indicating higher suitability. Thus, smaller values of the index are used as a proxy for higher levels of historical patriarchy. The geographic distribution of this patriarchy index is presented in Figure 4 and clearly shows a deeper patriarchal culture in the north of China.[23] This north-south gradient mirrors findings from previous research on regional variation in gender norms within China, such as

[20] See `https://gaez.fao.org/pages/theme-details-theme-4`.

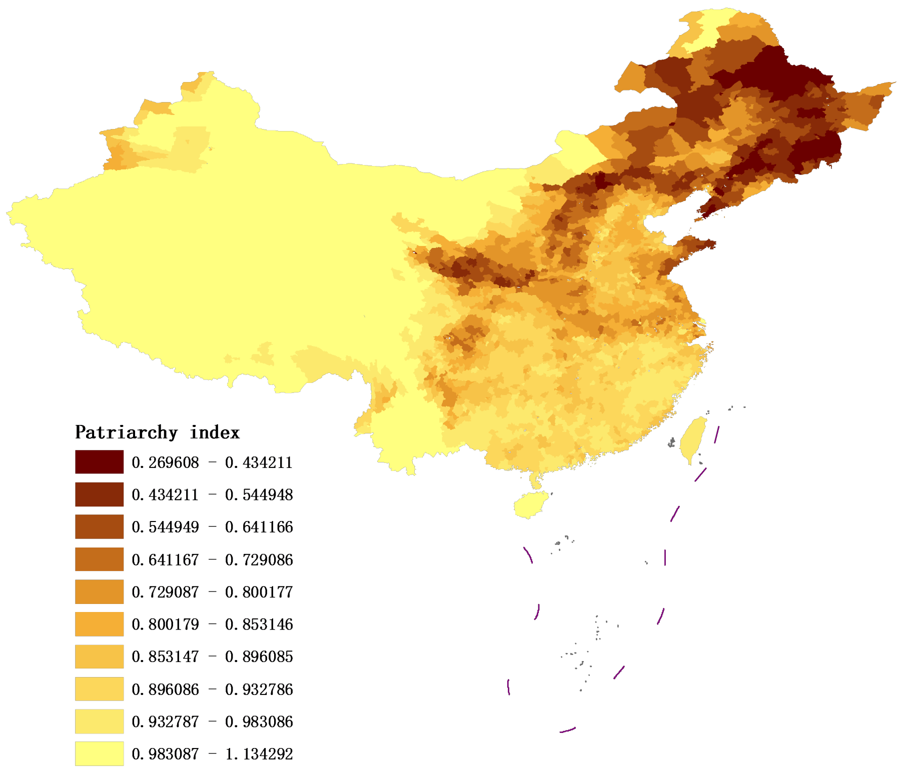
[21] Although there have been recent criticisms of the use of crop suitability indices to derive historical proxies because they are constructed using data from the twentieth century, making it possibly endogenous to technology and other factors (Rhode, 2024), in our econometric framework we rely on variation within spatial units over reform timing, which mitigates such endogeneity concerns as it is highly unlikely that unobserved confounding shocks are correlated with both crop suitability and the timing of the reform. Importantly, despite recent concerns about the precise parameters used to calculate crop suitability indices, these still demonstrate a meaningful degree of explanatory power for the actually cultivated area and yields, both in the US and China, with R-squared values ranging from approximately 0.2 to 0.4 (Almond et al., 2019; Rhode, 2024). Given that the GAEZ framework is based on rain-fed agriculture and low levels of human input (that is, excluding the influence of fertiliser and irrigation), this explanatory power is arguably substantial and reflects objective geographical and environmental conditions.

[22] Alesina et al. (2013) substantiate this point and show that higher plough-positive crop suitability is associated with a higher likelihood of historical plough adoption in agriculture.

[23] Appendix Figure C1 reports the geographic distribution of the index with the inclusion of wetland rice which results in areas of higher patriarchy in central China.

in the practice of foot-binding (Turner, 1997; Fan and Wu, 2023). For the majority of our analysis, we divide counties into high and low patriarchy by taking a median split of the patriarchy index across the whole of China.[24]

Figure 4: Spatial distribution of the patriarchy index



Notes: Figure 4 shows the spatial distribution of the patriarchy index. Colors indicate deciles; darker shades (smaller values of the index) represent stronger patriarchy. See text for details on the construction of the index.

In Appendix D, we cross-validate our index using measures proposed in the anthropological literature on historical patriarchy. Specifically, we show that our index correlates strongly with the gender domains of the Patriarchy Index (Szołtysek et al., 2017; Szołtysek et al., 2025), including proxies for male dominance, patrilocality and son preference.

In Appendix E, we provide a detailed description of our patriarchy index as a measure of cul-

[24]The median cut-off is calculated based on all counties in China that we have a patriarchy index for rather than the 901 counties in our sample. In a robustness check, we show that our findings are unchanged if we instead perform the split along the median of counties in our sample.

ture and provide a validation exercise which shows that the index proxies cultural, rather than economic, phenomena. A number of observations stand out. First, the index correlates strongly with present-day gender norms within China after controlling for economic factors. Highly patriarchal counties exhibit larger gender gaps in education, lower female labour force participation, and higher female risk aversion. Second, high-patriarchy counties do not see larger increases in agricultural productivity after the land reform. This suggests that the index is not contaminated by differential agricultural productivity in the contemporary context (perhaps unsurprisingly, given that the index considers relative historical suitability for plough-positive crops, rather than absolute soil quality). Together, these pieces of evidence suggest that our measure of patriarchy captures persistent cultural factors.

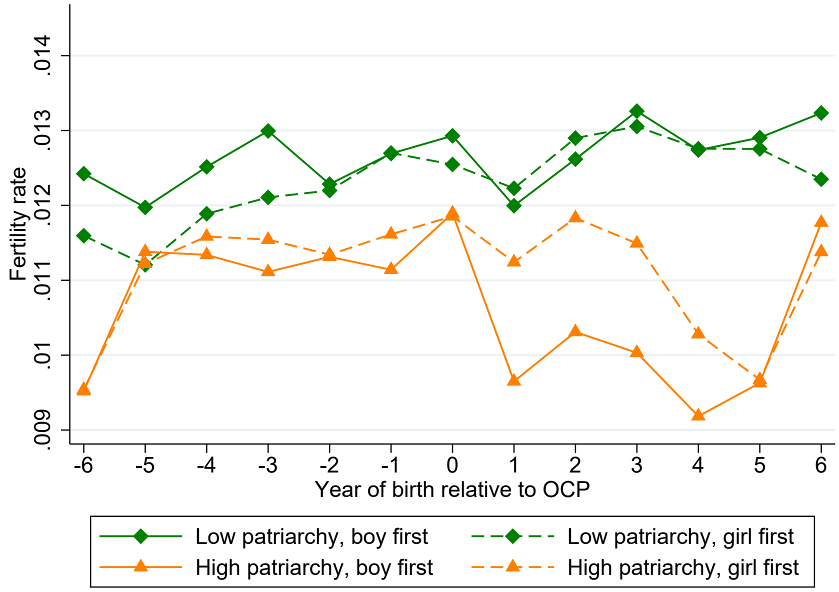
3.2 Descriptive Evidence

We now describe the basic patterns of fertility and the evolution of the sex ratio before and after the OCP and HRS. Figure 5(a) and Figure 5(b) plot the time series of average second-child fertility (the number of second births divided by the number of women of childbearing age in a given county) before and after the OCP and HRS reforms, respectively and separately for low-patriarchy counties (green diamonds) and high-patriarchy counties (orange triangles). For each set of counties, we further disaggregate fertility by whether the first child, which the second birth follows, was a boy (solid lines) or a girl (dashed lines).

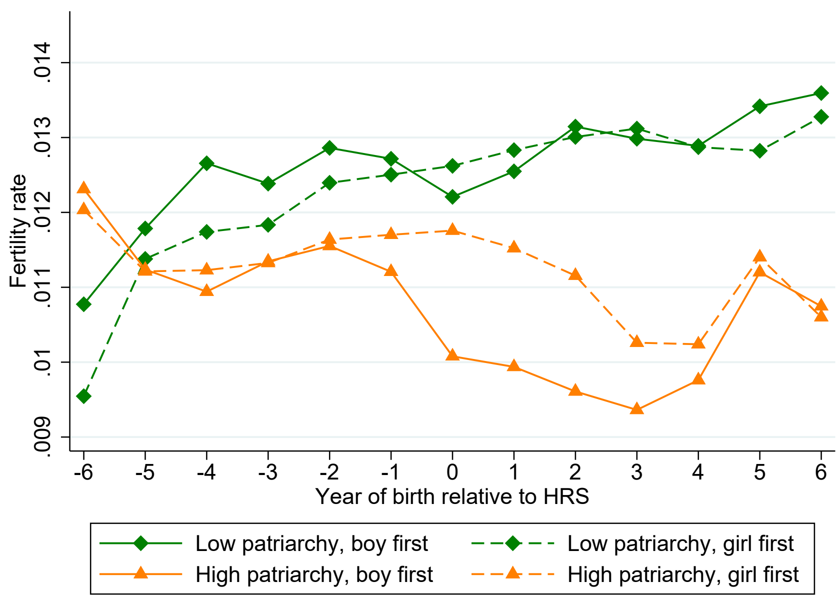
The fertility rate of second children shows a clear difference between low-patriarchy and high-patriarchy counties following the introduction of the OCP. After its implementation in low-patriarchy counties, second-child fertility evolves broadly along similar lines irrespective of whether the first child was male or female. In contrast, in high-patriarchy counties, second-child fertility drops following a firstborn boy and less so if the firstborn was a girl. Put differently, in high-patriarchy counties where son preference is stronger, the OCP does not appear effective in reducing second-child fertility if the first child was not a son. The changes in fertility rates before and after the HRS in Figure 5b are similar to those before and after the OCP, although the decline is more muted.

Turning to sex ratios, there is extensive evidence that sex selection is limited for first children even in the context of the One-Child Policy (Ebenstein, 2010; Li et al., 2011; Babiarz et al., 2019). As Figure 6 shows, this is confirmed in our data. We plot the average sex ratio of first children

Figure 5: Fertility rate of second children relative to timing of the OCP and HRS reforms



(a) Fertility rate of second children before and after the OCP



(b) Fertility rate of second children before and after the HRS

Notes: Average second-child fertility (second births in a county divided by the number of women of childbearing age) in each year before and after the OCP and HRS are rolled out. Averages are reported separately for low-patriarchy (green diamonds) and high-patriarchy (orange triangles) counties, and separately within each set of counties for second births that follow a firstborn boy (solid lines) or girl (dashed lines).

in the years before and after each reform, separately for low-patriarchy (grey circles) and high-patriarchy (grey squares) counties. There is no significant shift away from the biological ratio of 1.05, following the OCP or the HRS.

However, patterns change dramatically when we consider second-born children. The sex of the first child appears to exert a significant influence on subsequent parental choices (Zhang, 2017; Qian, 2018). The first observation is that the OCP and HRS do not appear to significantly affect the sex ratio of second children for families which already have a son, either in low- or high-patriarchy counties. Both solid lines (green diamonds and orange triangles) remain relatively stable. However, after the implementation of these policies, families with a first girl (dashed lines) severely distort the sex ratio of second children compared to those families with a first boy. The male skew in sex ratios is particularly pronounced in high-patriarchy counties where son preference is more salient.

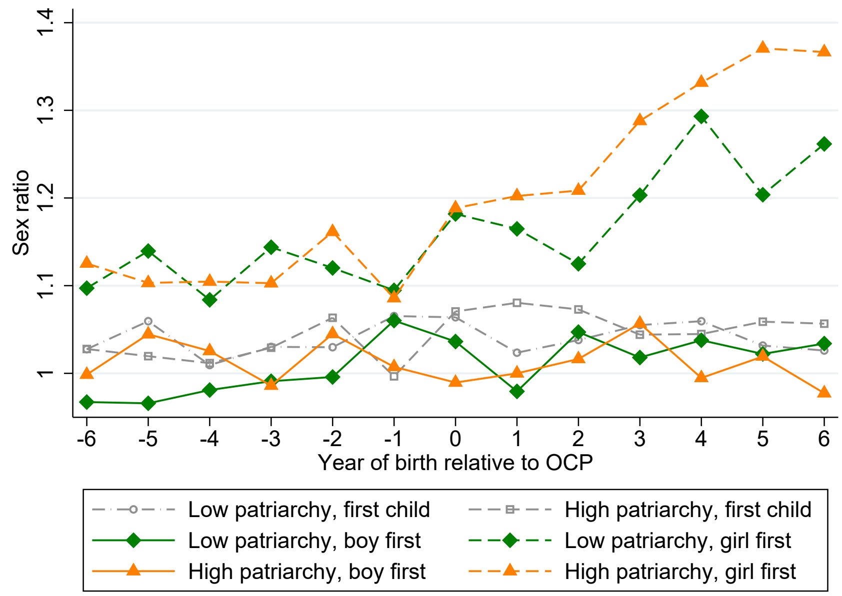
4 Econometric Framework

Our econometric framework uses a difference-in-differences estimator to trace the evolution of our outcomes of interest, second-child fertility and the male ratio, before and after the OCP and HRS reforms. We are interested, in particular, in the difference in the effect of the reforms on second births following a firstborn girl compared to when the firstborn was a boy and whether this differed between high- and low-patriarchy counties. The regression related to second-child fertility is specified at the county level as follows:

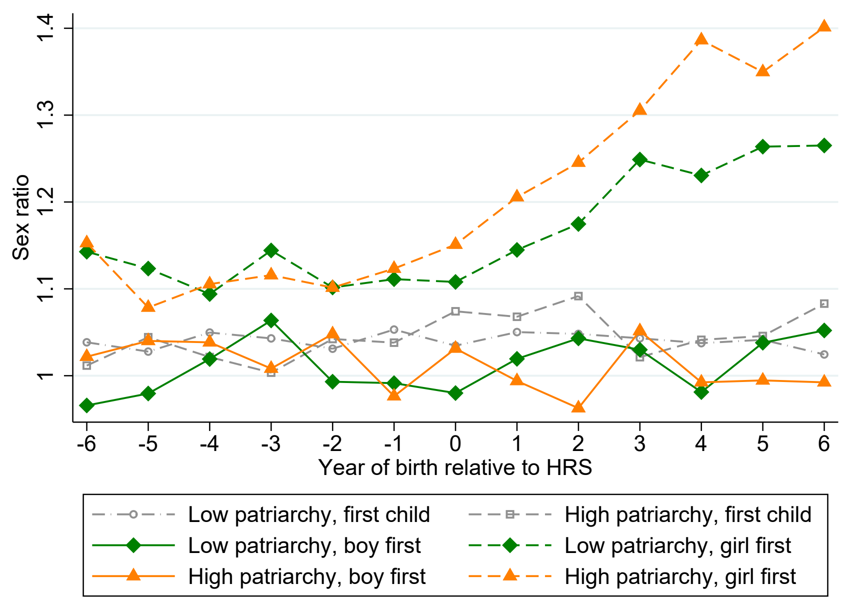
$$\ln(\text{births})_{gct} = \beta_1 \, \mathrm{G}_g + \beta_2 \, \mathrm{P}_c \times \mathrm{G}_g + \beta_3 \, \mathrm{R}_{ct} \times \mathrm{G}_g + \ \beta_4 \mathrm{R}_{ct} \times \mathrm{P}_c \times \mathrm{G}_g + \gamma_{ct} + \epsilon_{gct} \tag{1}$$

where $\ln(\text{births})_{gct}$ is the log of the number of second children following a firstborn girl or boy g in county c and year t. Note, in effect, this implies a data structure with two observations per county-year: one observation counting the number of second births following a firstborn girl in a given county-year, and one counting the number of second births following a firstborn boy. G_g is an indicator for observations counting the number of children following a firstborn girl. R_{ct} is a reform indicator, which switches on when a reform is implemented in a county (OCP, HRS or both, depending on the specification). Finally, P_c is an indicator for high-patriarchy counties, constructed according to our soil suitability approach described above. County-by-birth-year fixed effects γ_{ct}

Figure 6: Sex ratios of first and second children relative to the timing of the OCP and HRS reforms.



(a) Sex ratios before and after the OCP



(b) Sex ratios before and after the HRS

Notes: Average first- and second-child sex ratios (ratios of males to females among first and second children, respectively) in each year before and after the OCP and HRS are rolled out. Averages are reported separately for low-patriarchy (green diamonds) and high-patriarchy (orange triangles) counties, and separately within each set of counties for second births that follow a firstborn boy (solid lines) or girl (dashed lines). Averages are also shown for first children in low-patriarchy (grey circles) and high-patriarchy (grey squares) counties.

control for all other (possibly time-varying) county characteristics. Standard errors are clustered at the county level and regressions are weighted by the number of women of childbearing age in each county-year.[25,26]

Our interest lies in the estimates of β_3 and β_4. β_3 measures the difference in the response to the policies between families already with or without a son (in low-patriarchy counties). It captures the distinct fertility patterns induced by the reforms when interacted with a pre-existing son preference. β_4 captures the differential effect in high-patriarchy counties. Suppose that we expect a reform like the OCP to reduce second-child fertility, but that the induced reduction is smaller if the firstborn child is a girl (as a result of son preference). The implication would be $\beta_3 > 0$. Suppose further that son preference is particularly pronounced in high-patriarchy counties. This would lead us to predict $\beta_4 > 0$.[27]

The regression related to the sex ratio of second children is specified at the individual level as follows:

$$\text{Male}_{ict} = \beta_1\, \text{G}_{ict} + \beta_2\, \text{P}_c \times \text{G}_{ict} + \beta_3\, \text{R}_{ct} \times \text{G}_{ict} + \beta_4 \text{R}_{ct} \times \text{P}_c \times \text{G}_{ict} + \gamma_{ct} + \epsilon_{ict} \tag{2}$$

where Male_{ict} is an indicator of whether a newborn child i in county c and year t is a boy. The right-hand side variables are analogous to those in Equation (1). Standard errors remain clustered at the county level, but the weights used in Equation (1) are no longer applied, as individual newborns constitute the basic unit of analysis. β_3 captures the average change in the post-reform proportion of males, for families in low-patriarchy counties without a son. β_4 measures any additional distortion to the male sex ratios in high-patriarchy counties. As above, we expect $\beta_4 > 0$, showing a stronger son preference under patriarchal cultural norms.[28]

[25] This weighting scheme indirectly simulates fertility rates by giving each woman the same weight across counties, with the assumption that theoretically, more women in childbearing ages would have more births. This approach ensures that regression results are not overly influenced by extreme values. Unweighted regressions are discussed alongside other robustness checks further below.

[26] We observe some zero births following a first child of a specific gender in some county-year combinations. These observations are excluded to mitigate noise arising from extreme values due to sampling data. Nevertheless, we account for the potential influence of zero outcomes by estimating a Poisson pseudo-maximum likelihood model and the results are discussed as part of a series of robustness checks.

[27] Note that, in the presence of county-year fixed effects, we are netting out the average effects of the reform.

[28] Note that we can rule out another explanation for potentially increased male ratio following a first girl, sex composition preference (to have at least one child of each gender) in our regression settings (Angrist and Evans, 1998). Such an explanation would not be consistent with the fact that the data show a relative increase in fertility only following a firstborn girl, but not following a firstborn boy, after the implementation of the OCP (see Figure 5a).

In addition, we employ an event-study approach to examine the dynamics of the relationship between the reforms, patriarchal norms and fertility/male sex ratios. Separately for the set of high- and low-patriarchy counties, we estimate the following event study where we control for a series of dummy variables to trace out the year-by-year effects of the reforms:

$$\text{Outcome}_{gct/ict} = \sum_{\tau \neq -1} \beta_\tau \, \text{G}_{g/ict} \times \text{D}_\tau + \gamma_{ct} + \epsilon_{gct/ict}, \tag{3}$$

where D_τ equals one in the τth year relative to the introduction of each policy (OCP or HRS). We estimate β_τ for each event period ranging from $-6+$ to $6+$ years for both the OCP and HRS, taking $\tau = -1$ as the omitted reference year. Note that $-6+$ and $6+$ represent 6 or more years before and after these policies, with pooling at the endpoints to allow for more precise estimation given smaller sample sizes at the endpoints (Beck et al., 2010). A series of β_τ coefficients trace out the changes in behaviour relative to the omitted period. The other terms and settings are the same as those in Equations (1) and (2). We rely on an assumption of parallel trends to obtain a causal interpretation in Equations (1) and (2), and present evidence in support of this assumption further below.

5 Results

5.1 Patriarchy, Reforms and Fertility

The results from estimating Equation (1) are reported in Table 1. In Column (1), we first show the effect of each individual reform on fertility as a result of son preference and the role of patriarchy in son preference. Panels A and B respectively report the estimates for the OCP and HRS. As shown in row 3 of Panel A, following the OCP, second-child fertility is higher after a firstborn girl compared to after a firstborn boy. Row 4 shows this increase is differentially stronger in high-patriarchy counties. Similar patterns emerge when we consider land reform under the HRS instead in Panel B, which is unsurprising given the partial temporal overlap between the two reforms. Nevertheless, given the substantial subset of counties where the reforms do not coincide, we are able to tease apart the separate effects of the two reforms in Panel C, serving as the main results. These results suggest that the OCP is a more important driver of the fertility patterns we observe.

The results in Table 1 underscore the persistent impact of patriarchy on son preference and the potential for underlying cultural norms to shape behavioural responses to policy. Our estimates suggest that after the introduction of the OCP, in high-patriarchy counties, families with a firstborn girl have approximately 19.2% more second children compared to families that already have a son (row 4, Panel C). A back-of-the-envelope calculation suggests that these preferences led to an additional 4 million second children being born that would not otherwise have been born if the firstborn had been a boy.[29] Our findings suggest that these effects are much more muted (and statistically insignificant in Panel C) in low-patriarchy counties and highlights the existence of heterogeneities in cultural norms and preferences even within the same ethnicity in China.[30]

The results for subsamples from the south and north are presented in Columns (2) and (3) of Table 1. The pattern of son preference remains consistent with that of the full sample, although the absolute magnitude of the coefficient is larger in the south. The more muted effect in the north may reflect the fundamental role of its stricter state control, as evidenced by higher fine rates and a larger share of state-owned enterprise (SOE) employment, making son preference harder to achieve.[31] Finally, we include wetland rice suitability in the construction of our patriarchy index in Column (4). This is to address the concern that wetland rice historically benefitted from plough use and constitutes the main source of calories in southern China (as described in the background section). Results remain robust across subsamples and ways of constructing our patriarchy index.

5.2 Patriarchy, Reforms and the Sex Ratio

We now turn to our estimates from the individual-level regressions to explore the effect of the reforms, interacted with patriarchal norms, on the male sex ratio. The results are reported in Table 2, which follows an identical structure to Table 1. As shown in Column (1), following the reforms, the sex ratio of second children skews more male if the first child was female, and this effect

[29] There are 26,881 families with a first girl in high-patriarchy counties before the introduction of the OCP. Combined with the OCP's effect lasting up to 8 years in our sample (as shown in the later event study), we can calculate the maximum number of second births attributable to son preference as families $\times$ average effect size $\times$ years post-reform $\times$ 100 (to account for 1% sampling) $= 26,881 \times 0.192 \times 8 \times 100 \approx 4,128,921$.

[30] Individuals of Han ethnicity constitute approximately 90% of our regression sample. Results are robust when restricting the sample to those of Han ethnicity.

[31] In the most patriarchal northeastern provinces (as seen in Figure 4), the average fine rate of violating the OCP was 17.49% higher than the national average (excluding municipalities) within the sample period (Ebenstein, 2010). In addition, these provinces had an exceptionally high share of SOE employment, who were directly subject to stricter family-planning enforcement such as the threat of job dismissal (52% in 1978, compared with a national average of 23% in other provinces, excluding municipalities). Data are from *China Compendium of Statistics 1949-2008*.

Table 1: Patriarchy and the impact of OCP and HRS on the number of second births

	Dependent variable: ln(number of second births)			
	Full sample	South only	North only	South only (with wetland rice)
	(1)	(2)	(3)	(4)
Panel A: OCP only				
Girl first	-0.038**	-0.040**	-0.030	-0.032
	(0.017)	(0.018)	(0.038)	(0.024)
High patriarchy × girl first	0.057**	0.056*	0.050	0.017
	(0.023)	(0.031)	(0.042)	(0.030)
OCP × girl first	0.061**	0.061**	0.065	0.037
	(0.025)	(0.027)	(0.059)	(0.038)
OCP × high patriarchy × girl first	0.146***	0.212***	0.090	0.127**
	(0.044)	(0.072)	(0.070)	(0.053)
Panel B: HRS only				
Girl first	-0.032**	-0.033*	-0.025	-0.032
	(0.016)	(0.017)	(0.038)	(0.027)
High patriarchy × girl first	0.086***	0.094***	0.074*	0.045
	(0.022)	(0.031)	(0.043)	(0.032)
HRS × girl first	0.051**	0.049*	0.059	0.037
	(0.025)	(0.027)	(0.061)	(0.041)
HRS × high patriarchy × girl first	0.097**	0.147**	0.049	0.080
	(0.043)	(0.068)	(0.072)	(0.053)
Panel C: OCP and HRS				
Girl first	-0.038**	-0.040**	-0.032	-0.034
	(0.016)	(0.018)	(0.039)	(0.026)
High patriarchy × girl first	0.061***	0.060**	0.056	0.023
	(0.022)	(0.030)	(0.044)	(0.031)
OCP × girl first	0.062	0.066	0.048	0.020
	(0.043)	(0.051)	(0.075)	(0.070)
OCP × high patriarchy × girl first	0.192***	0.251**	0.154*	0.190*
	(0.072)	(0.117)	(0.093)	(0.098)
HRS × girl first	-0.001	-0.006	0.022	0.020
	(0.044)	(0.051)	(0.080)	(0.076)
HRS × high patriarchy × girl first	-0.057	-0.050	-0.078	-0.076
	(0.070)	(0.112)	(0.097)	(0.099)
Observations	22,065	12,778	9,287	12,778
Counties	901	515	386	515
Outcome mean	1.954	2.043	1.832	2.043
County × birth year FE	Y	Y	Y	Y

Notes: Estimates of Equation (1). Regressions weight by the number of women in childbearing ages within a county in each birth year. Standard errors in parentheses are clustered at the county level. Significance levels denoted by *** p<0.01, ** p<0.05, * p<0.1.

is particularly strong in high-patriarchy counties (Panels A and B). If we estimate the effects of both reforms simultaneously, results are driven by land reform under the HRS (Panel C). Our estimates suggest that the HRS increased the proportion of males among second children for families with a firstborn girl by 3.90 percentage points more in high-patriarchy than in low-patriarchy counties.

Again, a back-of-the-envelope calculation suggests that this resulted in approximately 0.89 million missing girls in high-patriarchy counties.[32]

When we turn to the subsamples presented in Columns (2) and (3), the coefficients become more muted in the north, although the pattern of son preference remains in the south.[33] In addition, Panel C shows that the results are robust in the southern subsamples when wetland rice is included in the patriarchy index, as shown in Column (4).

Combining the findings from Tables 1 and 2, the independent effects of the OCP and HRS highlight two distinct ways to achieve son preference. Where no second-child exception applied under the OCP, families with a firstborn girl could respond primarily by having an additional birth and accepting the uncertainty over its sex. In contrast, the HRS effectively increases the male ratio of second children following a firstborn girl, as higher household income after agricultural decollectivisation makes ultrasound technology more accessible (Almond et al., 2019).

To replicate this mechanism, we conduct further subsample analyses based on the availability of ultrasound machines in provincial capitals in the year preceding the birth of second children. These findings are reported in Table 3, and row 6 of Column (1) indicates that the rise in the male sex ratio occurs only in high-patriarchy counties in provinces where ultrasound technology was accessible. This provides further evidence for the mechanism underpinning the findings of Almond et al. (2019), showing that the HRS increased the male sex ratio due to sex-selective abortions, particularly in places where historical patriarchal norms give rise to son preference.

5.3 Positive Female Income Shocks Do Not Undo Son Preference

To explore how deep-rooted patriarchal cultural norms have become, we consider the positive female income shock which resulted from the large increase in tea prices after the HRS and test whether it eroded son preference. According to Qian (2008), in 1986 the price of tea increased so that it was around 2.5 times that of the 1979 price, leading to significantly higher relative prices compared to other crops. Women's comparative advantages in picking tea generated a positive female income

[32] There are 28,618 families with a first girl in high-patriarchy counties before the introduction of the HRS. If we imagine a counterfactual where the additional boys that were born as a result of the interaction between land reform and son preference in high-patriarchy counties were girls instead, we can calculate the number of missing girls as families $\times$ average effect size $\times$ years post-reform $\times$ 100 (to account for 1% sampling) $= 28,618 \times 0.039 \times 8 \times 100 \approx 892,881$.

[33] It is reasonable when combining with results in Table 1, given that the premise of influencing sex ratio is that a second child has already been born.

Table 2: Patriarchy and the impact of OCP and HRS on the sex ratio of second children

	Dependent variable: male second child = 1			
	Full sample	South only	North only	South only (with wetland rice)
	(1)	(2)	(3)	(4)
Panel A: OCP only				
Girl first	0.027***	0.026***	0.031**	0.019**
	(0.005)	(0.005)	(0.012)	(0.009)
High patriarchy × girl first	-0.001	0.004	-0.009	0.012
	(0.007)	(0.009)	(0.014)	(0.010)
OCP × girl first	0.015**	0.015**	0.015	0.017
	(0.007)	(0.007)	(0.016)	(0.012)
OCP × high patriarchy × girl first	0.021**	0.025**	0.018	0.007
	(0.009)	(0.012)	(0.018)	(0.014)
Panel B: HRS only				
Girl first	0.027***	0.027***	0.028**	0.023**
	(0.005)	(0.005)	(0.013)	(0.009)
High patriarchy × girl first	-0.005	-0.001	-0.008	0.005
	(0.007)	(0.009)	(0.014)	(0.010)
HRS × girl first	0.015**	0.014*	0.021	0.010
	(0.007)	(0.007)	(0.017)	(0.012)
HRS × high patriarchy × girl first	0.030***	0.037***	0.020	0.021
	(0.009)	(0.012)	(0.019)	(0.014)
Panel C: OCP and HRS				
Girl first	0.026***	0.026***	0.028**	0.020**
	(0.005)	(0.005)	(0.013)	(0.009)
High patriarchy × girl first	-0.004	0.000	-0.009	0.008
	(0.007)	(0.009)	(0.015)	(0.011)
OCP × girl first	0.007	0.011	-0.004	0.030
	(0.012)	(0.014)	(0.025)	(0.022)
OCP × high patriarchy × girl first	-0.011	-0.017	0.004	-0.035
	(0.016)	(0.020)	(0.029)	(0.025)
HRS × girl first	0.010	0.005	0.025	-0.016
	(0.012)	(0.014)	(0.027)	(0.022)
HRS × high patriarchy × girl first	0.039**	0.051**	0.016	0.050**
	(0.016)	(0.021)	(0.031)	(0.025)
Observations	225,772	142,533	83,239	142,533
Counties	901	515	386	515
Outcome mean	0.524	0.523	0.526	0.523
County × birth year FE	Y	Y	Y	Y

Notes: Estimates of Equation (2). Standard errors in parentheses are clustered at the county level. Significance levels denoted by *** p<0.01, ** p<0.05, * p<0.1.

shock, increasing women's relative economic value. Hence, we split our sample by the median of tea suitability and estimate our main specifications within each subsample.[34] In Table 4, we find that families with a firstborn girl still show higher second-child fertility and a higher male sex ratio

[34]The tea suitability measures the potential yield following Almond et al. (2019). We do not divide it by the average potential yield of other crops, as its relative price is rising in this context. Therefore, we assume that farmers will opt to cultivate tea once it is deemed suitable, rather than requiring it to be more suitable than other crops.

Table 3: Heterogeneity of effect on sex ratios by ultrasound access

	Dependent variable: male second child = 1	
	With access	Without access
	(1)	(2)
Girl first	0.027***	0.025***
	(0.007)	(0.007)
High patriarchy × girl first	-0.005	-0.002
	(0.009)	(0.011)
OCP × girl first	0.002	0.034
	(0.013)	(0.029)
OCP × high patriarchy × girl first	-0.004	-0.063
	(0.017)	(0.051)
HRS × girl first	0.013	0.001
	(0.014)	(0.027)
HRS × high patriarchy × girl first	0.036**	0.020
	(0.018)	(0.051)
Observations	178,376	47,396
Counties	901	385
Outcome mean	0.526	0.518
County × birth year FE	Y	Y

Notes: Estimates of Equation (2). Standard errors in parentheses are clustered at the county level. Significance levels denoted by *** p<0.01, ** p<0.05, * p<0.1.

in counties exposed to a considerable tea shock (see Columns (2) and (4)). If anything, we find stronger effects in counties highly suitable for tea. This demonstrates the persistence of culture and a limited role of short-term income shocks in undoing deep-seated norms. In other words, the OCP effect on second-child fertility and the HRS effect on the probability of a male second birth remain strong in high-tea-suitability counties.

5.4 Event Study Estimation

Finally, Figures 7 and 8 plot event-study estimates. In each figure, we report two event studies: one for low-patriarchy and one for high-patriarchy counties. In effect, these trace out the effects (before and after the event) of each reform, separately for each subsample of counties.

In both figures, we find consistently small and insignificant coefficients before and after the reforms in low-patriarchy counties. This is further evidence in favour of our interpretation above: the sex of first children does not affect second-child fertility or sex ratios in the absence of strong patriarchal norms. In high-patriarchy counties, we similarly find no evidence of diverging trends before the reforms.[35] Following each reform, however, we find that underlying patriarchal norms

[35]The null hypothesis of the Wald test (that the coefficients before the reforms are jointly equal to 0) are not

Table 4: Heterogeneity of effects by tea suitability

	Dependent variable:			
	ln(number of second births)		male second child = 1	
	Low tea suit. (1)	High tea suit. (2)	Low tea suit. (3)	High tea suit. (4)
Girl first	-0.033	-0.042**	0.030***	0.024***
	(0.036)	(0.018)	(0.010)	(0.005)
High patriarchy × girl first	0.058	0.042	-0.006	-0.001
	(0.039)	(0.032)	(0.012)	(0.011)
OCP × girl first	0.085	0.054	-0.014	0.017
	(0.063)	(0.056)	(0.022)	(0.014)
OCP × high patriarchy × girl first	0.155*	0.280**	0.014	-0.047*
	(0.090)	(0.112)	(0.025)	(0.027)
HRS × girl first	-0.007	0.004	0.028	0.002
	(0.067)	(0.055)	(0.025)	(0.013)
HRS × high patriarchy × girl first	-0.064	-0.066	0.013	0.078***
	(0.089)	(0.112)	(0.028)	(0.028)
Observations	11,308	10,553	113,868	109,434
Counties	472	420	472	420
Outcome mean	1.917	1.989	0.525	0.523
County × birth year FE	Y	Y	Y	Y

Notes: The regressions in columns (1) and (2) estimate Equation (1), and those in columns (3) and (4) estimate Equation (2). The regressions in columns (1) and (2) weight by the number of women in childbearing ages within a county in each birth year. Standard errors in parentheses are clustered at the county level. Significance levels denoted by *** p<0.01, ** p<0.05, * p<0.1.

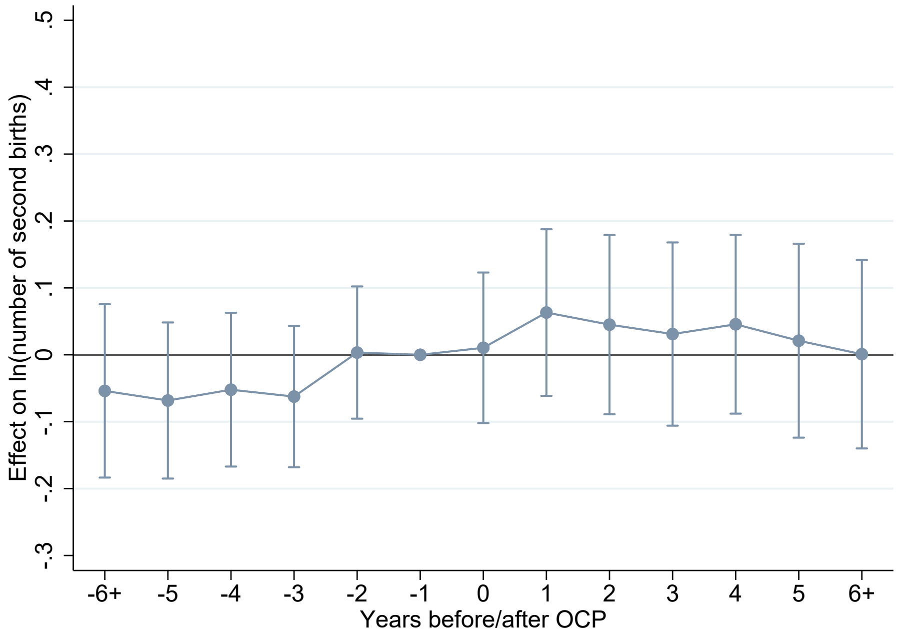
and implied son preference have significant impacts on fertility and the male-biased sex ratio. These effects are long-lasting, and we detect significant differences in the last post-reform period (6 or more years post-reform).

5.5 Robustness Checks

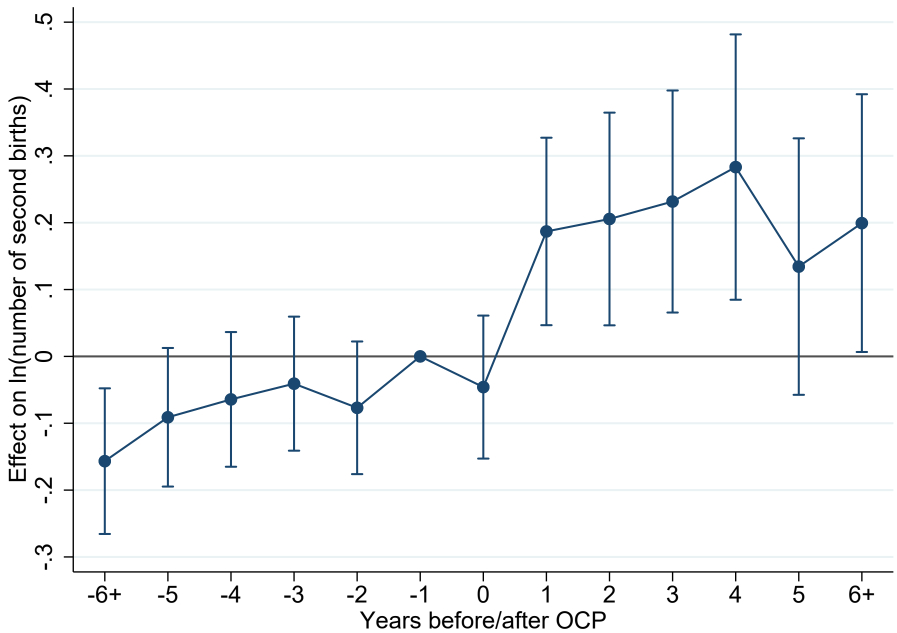
In Appendix F, we conduct a range of further checks to probe the robustness of our results for both the number of second births and the sex ratio. These include: 1) unweighted and Poisson pseudo-maximum likelihood regressions of second-child fertility, 2) considering higher-parity births (third children and above), 3) controlling for the "1.5"-Child Policy, 4) other variants of the patriarchy variable, 5) further probing the independent effects of the two reforms, controlling for their interactions or distinguishing their relative implementation time, 6) a discussion of spatial trends (via a placebo check by randomizing the distribution of our patriarchy index), 7) including the urban sample (excluded in the main regressions), 8) controlling for average slope, and 9) accounting for potentially dynamic and heterogeneous effects in staggered difference-in-differences designs (Sun

rejected (with p-values of 0.11 and 0.65 for the regressions on fertility and male ratio, respectively).

Figure 7: Event study: effect of OCP on number of second births



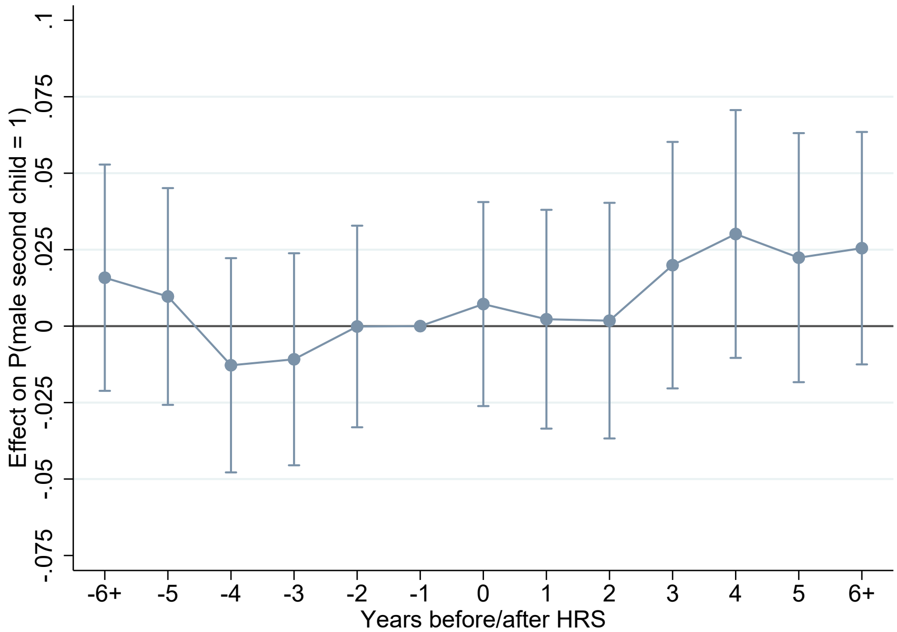
(a) Low-patriarchy counties



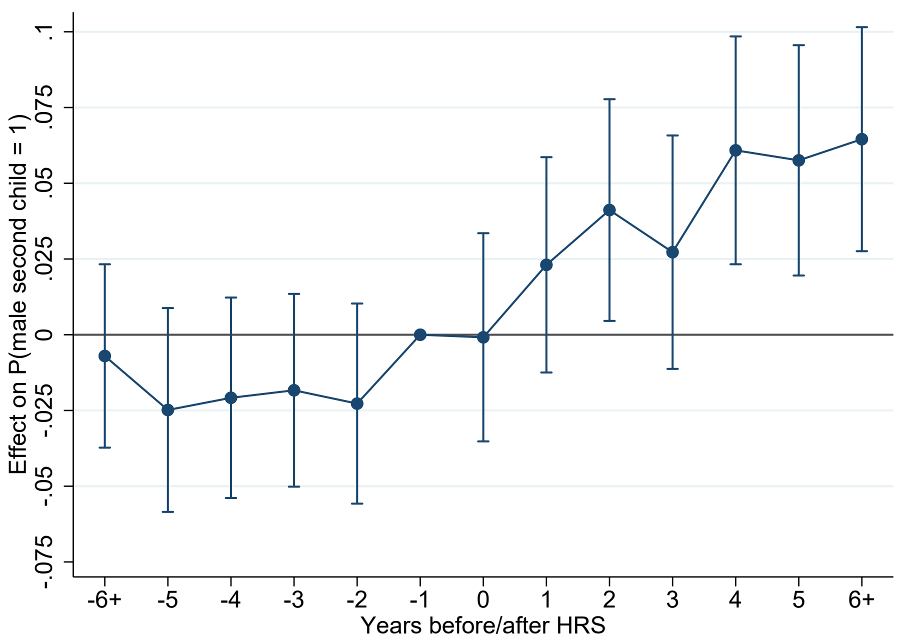
(b) High-patriarchy counties

Notes: Estimates of Equation (3), separately for low- and high-patriarchy counties. Standard errors are clustered at the county level and 95% confidence intervals are shown.

Figure 8: Event study: effect of HRS on the sex ratio of second children



(a) Low-patriarchy counties



(b) High-patriarchy counties

Notes: Estimates of Equation (3), separately for low- and high-patriarchy counties. Standard errors are clustered at the county level and 95% confidence intervals are shown.

and Abraham, 2021, Borusyak et al., 2024). The results are discussed in detail in Appendix F; the sign and magnitude of our main results are robust throughout.

6 Conclusion

By constructing a historical plough-use suitability index as a proxy for patriarchal cultural norms across China, we investigate the role of patriarchy in shaping son preference under the fertility constraints imposed by the One-Child Policy (OCP) and Household Responsibility System (HRS). Combining this index with data on reform rollout, fertility and sex ratios (Almond et al., 2019) and a difference-in-differences econometric design, we document differential effects of these reforms across high- and low-patriarchy counties.

We find that historical patriarchy shapes household responses to the OCP and HRS, influencing the distribution of contemporary son preference even within the same ethnicity. There is little evidence that son preference causes unintended policy effects in low-patriarchy counties: families with or without a son show similar patterns of fertility and sex selection after the introduction of the OCP and HRS. In contrast, in high-patriarchy counties, families with a firstborn girl tend to increase both second-child fertility and skew the second-child sex ratio in favour of males. Moreover, the OCP and HRS manifest household son preference in different ways. Families without a son (in high-patriarchy counties) tend to violate the OCP, but this does not necessarily lead to the birth of a son through natural means. Consequently, increased fertility but not an increased male sex ratio is observed after the OCP. The HRS, on the other hand, increases household income and thereby facilitates access to sex selection technology, making sex selection more feasible.

Thus, our study shows that contemporary choices and behaviours can be shaped by policy in a manner that is strongly contingent on ancestral culture. This pattern complicates the evaluation of the side-effects of policies. In the Chinese case, for instance, the phenomenon of missing girls constitutes a severe problem after the introduction of the OCP and HRS. The Three Child Policy, which allows families to have three legal children, was introduced in China in 2021. It is expected to decrease the number of missing girls, but in the presence of persistent patriarchal norms it may further exacerbate the gender gap in education and income via interactions between a "quantity-quality" trade-off and son preference (Becker and Lewis, 1973; Zhang and Ma, 2017). Such potential future side-effects signify the policy implications of our findings. That is, gender equality is unlikely to be achieved via strictly economic means. Direct efforts to shift norms may be required, too.

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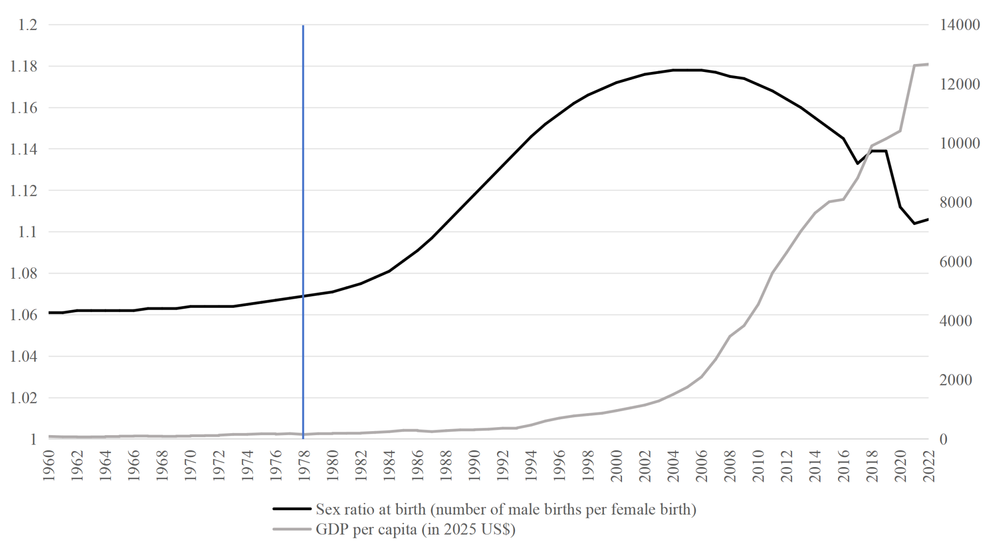
Appendix A: Figure Referenced in Introduction

Figure A1: Economic development and China's sex ratio



Notes: Data are from the World Bank. See `https://data.worldbank.org/indicator/NY.GDP.PCAP.CD?locations=CN` and `https://data.worldbank.org/indicator/SP.POP.BRTH.MF?locations=CN`.

Appendix B: Further Details on Household Responsibility System

To understand land reform under the HRS, one must consider first the commune system it replaced. The system of People's Communes was proposed by Chairman Mao in 1958 as a means to accelerate industrialisation and further advance the socialist cause. It sought to deepen the collectivisation of existing advanced agricultural cooperatives through greater economies of scale.[36] Communes replaced township governments as the basic unit of governance, thereby comprising not only an economic entity but also a political one. The result was a broadly consistent operational system, applied across China (Central Committee of the Communist Party of China, 1958b, 10 December).

The guiding principle, known as "One: Big; Two: Public" (*yi da er gong*), emphasised larger-scale production and a higher degree of public ownership (Central Committee of the Communist Party of China, 1958a, 10 September). By the end of 1958, a total of 26,576 People's Communes had been established, with the proportion of participating households reaching 99.1%. A commune consisted of 4,797 households on average (approximately 20,000 individuals) making it about 28 times larger than the average advanced agricultural cooperative (Liu, 2014, p. 69, 73–74). The small "private plots" that existed under the cooperative system (accounting for roughly 5% of arable land) were abolished and transferred to commune ownership.[37] As a result, the free markets that had emerged for outputs from private plots also disappeared (Henan Provincial Committee of the Communist Party of China, 1958, August 22; Perkins, 1988).

The wage system of work-points (*gongfen*) assessed both the quantity and quality of work and was designed to complement the basic food needs achieved by the grain ration within a commune. This gave an incentive for individual effort (Wortzel, 1981). However, high monitoring costs and a lack of incentives to monitor due to the absence of rights to residual claims by cadres made the system vulnerable to free-rider problems (Kung and Putterman, 1997).[38] Therefore, communal dining played a fundamental role in allocating agricultural output on a per capita basis. A natural consequence was that the cost of childbearing was collectively undertaken by the whole commune (Unger, 1985; Chen and Xie, 2025).

[36] By the end of 1956, 87.8% of households nationwide had joined the advanced cooperatives that formed the backbone of the People's Communes (Liu, 2014, p. 13).

[37] Although private plots were briefly allowed after the Great Famine (1959-1961), they soon returned to collective control (Chen and Xie, 2025).

[38] Cadres were also subject to the commune system and received grain ration only after mandatory state procurement and collective allocations were fulfilled, leaving them with little material incentive to monitor production.

In line with agricultural collectivisation, prices were embedded in a national procurement system (established in 1953), where both quota prices and above-quota prices existed. The mandatory quota was allocated to local production units in a unified system, which essentially imposed a heavy tax on agriculture due to artificially low prices, especially for more productive regions where a larger share of production was procured (Lin, 1992; Meng et al., 2015). In order to achieve overall production targets, the state executed meticulous control over crop cultivation, the areas under cultivation, input levels and planting techniques by crop irrespective of local comparative advantages, resulting in huge economic inefficiencies (Qian, 2008). Additional distortions were introduced by the excessive diversion of rural resources from agriculture to industry (Kung and Lin, 2003; Li and Yang, 2005).

Lin (1992) finds that agricultural productivity grew slowly during the collectivisation period, and there was even negative growth in some years. Despite the overall increase in crop output, the rural income per capita in the 1970s remained at the same level as the 1950s, with a poverty rate of 30.7% in rural areas in 1978 (Lardy, 1983; National Bureau of Statistics of China, 2002).

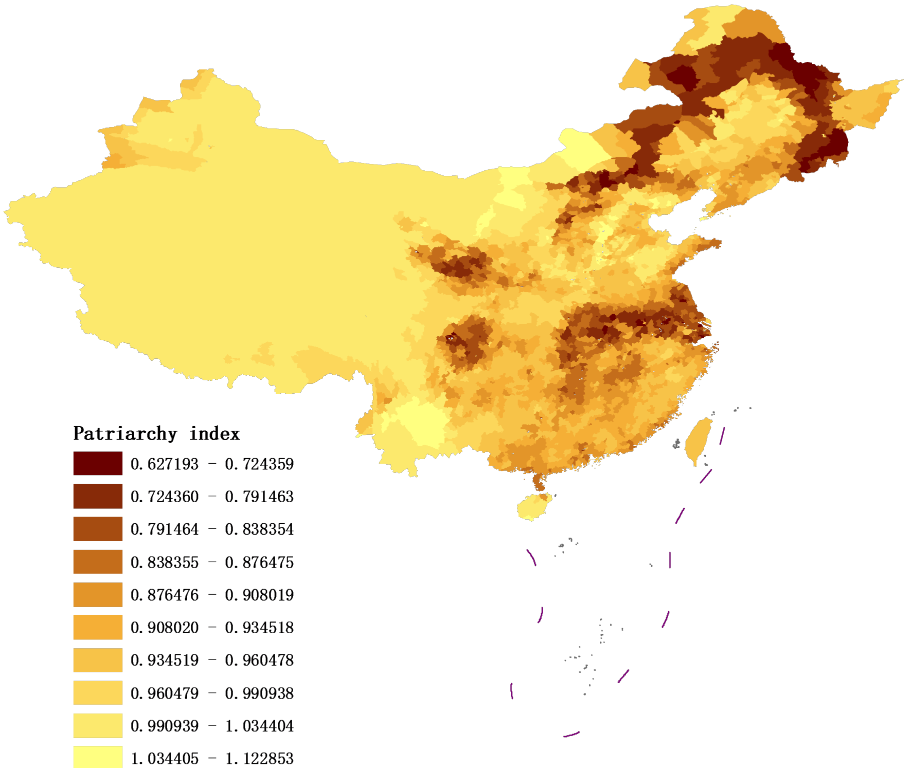
Appendix C: Additional Tables and Figures

Table C1: Summary statistics

	Obs.	Mean	St. Dev.	Min.	Max.
Panel A: observation is a county-year					
Second births	11432	19.75	17.19	1.00	171.00
Women of childbearing age	11432	850.86	662.60	15.00	5171.00
Second births per woman (fertility)	11432	0.02	0.01	0.00	0.19
Panel B: observation is a county					
Patriarchy index (rel. prod. plough-pos. crops)	901	0.85	0.13	0.28	1.07
Patriarchy index (incl. wet rice)	901	0.93	0.07	0.64	1.11
High patriarchy index	901	0.45	0.50	0.00	1.00
High patriarchy index (incl. wet rice)	901	0.51	0.50	0.00	1.00
In southern China	901	0.57	0.50	0.00	1.00
Average slope	901	7.11	8.61	0.16	55.64
Panel C: observation is an individual					
Male	225772	0.52	0.50	0.00	1.00
Patriarchy index (rel. prod. plough-pos. crops)	225772	0.85	0.11	0.28	1.07
Patriarchy index (incl. wet rice)	225772	0.91	0.07	0.64	1.11
High patriarchy index	225772	0.49	0.50	0.00	1.00
High patriarchy index (incl. wet rice)	225772	0.60	0.49	0.00	1.00
In southern China	225772	0.63	0.48	0.00	1.00
Average slope	225772	8.28	8.72	0.16	55.64

Notes: Panels A and B show county-year level second-child fertility outcomes and key geographic features involved in Equation (1), including various patriarchy indices, the North–South division, and land slope. Panel C presents analogous variables at the individual level used in Equation (2), which focuses on a male dummy for the second child as the outcome.

Figure C1: Spatial distribution of the patriarchy index (including wetland rice)



Notes: The map shows the spatial distribution of patriarchy when wetland rice is included in the index calculation. Colors indicate deciles; darker shades represent stronger patriarchy.

Appendix D: Cross-Validation Against the Patriarchy Index

The Patriarchy Index (Szołtysek et al., 2017) conceptualises "patriarchy" across four core domains: male dominance over females (*male domination*), intergenerational control by elders (*generational dominance*), patrilocal residence patterns (*patrilocality*), and a distinct son preference (*son preference*). Szołtysek et al. (2017) propose proxy indicators for each domain and Szołtysek et al. (2025) apply these to the Asian context. Using the 1990 population census, we construct these proxy measures as summarised in Table D1. Following Singh et al. (2022), each sub-measure is standardised to a 0–10 scale and domain scores are computed as the arithmetic mean of their respective sub-measures. The final Patriarchy Index is then calculated as the sum across all domains.

Subsequently, we regress the resulting Patriarchy Index on our measure of patriarchy at the county level (where a smaller value indicates a higher degree of suitability for "plough-positive" crops), as reported in Table D2. We focus on the domains of the Patriarchy Index with an explicit gender dimension. That is, we exclude components related to generational domination as this is a dimension of patriarchy less clearly related to our main mechanisms of interest. We present results for the full sample of all counties and separately for counties in South and North China. Across all samples, our index correlates strongly in the expected direction with the Patriarchy Index (recall that in our index, lower values are associated with higher relative suitability for plough-positive crops, and thus higher historical patriarchy). A 0.1-unit decrease in the plough-positive index (indicating greater relative plough suitability) is associated with approximately a 0.097-point increase in the Patriarchy Index.

Table D1: Four domains of the Patriarchy Index

Variable	**Description**	**Hypothesised Association**	**Definition / Formula**
Male domination			
Female heads	Proportion of female household heads	Male-only household leadership reflects patriarchy; thus, female headship is *negatively correlated* with patriarchy	$\frac{\text{Female household heads (20+)}}{\text{Household heads (20+)}}$
Young brides	Share of women married early	Male dominance is reinforced through early marriage of females; this metric is *positively correlated* with patriarchy	$\frac{\text{Ever-married brides aged 15–19}}{\text{Ever-married women aged 15–49}}$
Older wives	Share of wives older than their spouses	Patriarchy discourages men from marrying older women; this ratio is *negatively correlated* with patriarchy	$\frac{\text{Wives older than their husbands}}{\text{Couples with known ages}}$
Generational domination			
Younger household head	Share of senior males (65+) living in households where a younger male relative serves as the head	Senior males retain headship authority under patriarchy; thus, this metric is *negatively correlated* with patriarchy	$\frac{\text{Those under a younger male head}}{\text{Total elderly men}}$
Neo-local	Proportion of neo-local residence among young men	Sons continue to live with their parents even after marriage; this measure should be *negatively correlated* with patriarchy	$\frac{\text{Those who are heads}}{\text{Ever married men aged 20–29}}$
Patrilocality			
Married daughter	Proportion of elderly people living with married daughters	All daughters leave their father's household on marriage; this measure should be *negatively correlated* with patriarchy	$\frac{\text{Those living with parents}}{\text{Ever-married daughter (15-30)}}$
Son preference			
Boy as last child	Proportion of boys among the last children	Parents will try to have at least a son; this measure should be *positively correlated* with patriarchy	$\frac{\text{Boys among last children}}{\text{Total number of last births}}$
Sex ratio	Number of boys per 100 girls (aged 0–4)	Girls are not treated at par with boys; measure should be *positively correlated* with patriarchy	$100 * \frac{\text{Male aged 0-4}}{\text{Female aged 0-4}}$

Notes: Adapted from Szołtysek et al. (2017), Singh et al. (2022), and Szołtysek et al. (2025). "Those" refer to the same group as the denominator.

Table D2: Correlation between our index and gender domains of the Patriarchy Index

	Dependent variable: Patriarchy Index (gender domains)		
	Full sample (1)	South only (2)	North only (3)
Plough positive index	-0.965***	-1.929**	-0.777**
	(0.347)	(0.832)	(0.380)
Urbanisation rate	-1.608***	-1.823***	-1.437***
	(0.122)	(0.166)	(0.173)
Observations	2,386	1,192	1,194
Outcome mean	14.979	15.000	14.959
Province FE	Y	Y	Y

Notes: The dependent variable consists of the gender domains of the Patriarchy Index (Szołtysek et al., 2017, 2025): male domination, patrilocality and son preference. See Table D1 for details. Urbanisation rate is controlled for given the potential role of economic development in shaping culture and norms. Province fixed effects are controlled for in all columns. Standard errors in parentheses are clustered at the county level. Significance levels denoted by *** $p<0.01$, ** $p<0.05$, * $p<0.1$.

Appendix E: Our Index of Patriarchy as a Measure of Culture

E1. Correlation with Present-Day Gender Norms

To provide further evidence on the validity of our measure of patriarchy as a representation of cultural norms, we show a strong relationship between our index and indicators of female empowerment. Existing studies have emphasised the fundamental role of education and employment on female empowerment, whereby women have more power to achieve their preferred fertility rate, enhance health, and boost their political influence, thereby forming a virtuous cycle (Ross, 2008; Heath and Jayachandran, 2018).

First, we construct a measure of the gender gap in school enrolment for those of school age (between 6 and 18) at the county level using the 1990 population census following Qian (2024) which we then regress on our measure of patriarchy. We report the result in Column (1) of Table E1, which shows a significant positive correlation between patriarchal culture (indicated by smaller index value) and the gender gap of educational enrolment rates.

Second, drawing on the same census data we regress individual labour force participation (for those aged from 19 to 49, above the school age and below the maximum childbearing age). Results are presented in Column (2). As shown by the interaction term, patriarchy significantly exacerbates the gender gap in labour force participation (in counties where patriarchy is more of a cultural norm, participation rates are lower). Consider a plough-neutral county with a patriarchy index value of 1. In such a county, women are 3.6 percentage points ($= 0.298 - 0.262 \times 1$) less likely than men to participate in the labour force. In contrast, in the strongest patriarchy counties (where the value of the index is 0.27), this gap increases to 22.7 percentage points ($= 0.298 - 0.262 \times 0.27$).

Finally, we focus on whether patriarchy influences an individual's willingness to take risks (scale 0 to 10 indicates lower to higher willingness) using the 2008 wave of the Chinese Household Income Project Survey (CHIP).[39] Willingness to take risks is thought to be a key feature of entrepreneurship and the search for self-realisation (Caliendo et al., 2015; Rietveld and Patel, 2022). As shown by the interaction term, the results in Column (3) suggest that females in high-patriarchy counties

[39] This sampling but representative survey is conducted by Beijing Normal University to track the dynamics of income distribution in China, with the assistance of the National Bureau of Statistics. The 2008 wave was conducted in collaboration with the Australian National University and Institute of Labor Economics. See `https://dataverse.iza.org/dataset.xhtml?persistentId=doi:10.15185/izadp.7680.1`.

(with smaller values of the index) are less willing to take risks.

However, one could argue that the impact of patriarchy on these indicators of female empowerment is simply a reasonable response to consistent economic opportunities or institutions that are more beneficial to men, proxied by historical plough use (Alesina et al., 2013). This concern is mitigated by controlling for province fixed effects and county by birth year fixed effects in Appendix Table E1.

Table E1: Relationship between patriarchy index and contemporary gender norms

	Dependent variable:		
	Enrolment gap (1)	Labour force participation (2)	Risk preferences (3)
Patriarchy index	-0.035**		
	(0.016)		
Urbanisation rate	-0.086***		
	(0.005)		
Female		-0.298***	-1.999***
		(0.027)	(0.746)
Female × patriarchy index		0.262***	1.595*
		(0.031)	(0.872)
Agricultural hukou		0.033***	0.104
		(0.005)	(0.188)
Observations	2,387	5,069,753	11,609
Counties	2,387	2,389	77
Outcome mean	0.067	0.933	2.403
Province FE	Y		
County × birth year FE		Y	Y
Ethnicity FE		Y	Y

Notes: The enrolment gap measures the gender difference in school attendance rates at the county level. Labour force participation is a dummy variable indicating an individual's employment status. Risk preference is measured on a scale capturing an individual's willingness to take risks, with values from 0 to 10 indicating low to high risk tolerance. Urbanization rate is controlled for in Column (1) given the potential better educational infrastructure and overall higher education participation rate in urban areas. Agricultural Hukou and ethnicity are controlled for in Columns (2)-(3) because these characteristics capture important economic and cultural differences across individuals. Province and county by birth year fixed effects are controlled for in Column (1) and (2)-(3), respectively. Standard errors in parentheses are clustered at the county level. Significance levels denoted by *** p<0.01, ** p<0.05, * p<0.1.

The results presented in Appendix Table E1 strongly suggest that our index of patriarchy indeed captures a cultural dimension. It is therefore reasonable that patriarchal cultural norms stemming from historical plough use constitute an important root of son preference even under strict policies such as the OCP, and its enduring impact continues although male physical advantage in ploughing agriculture is a thing of the past.

E2. No Correlation with Effects of Land Reform on Agricultural Productivity

To confirm the cultural attributes associated with our patriarchy index, we now show that it does not appear to impact agricultural productivity when interacted with land reform. We regress grain output per capita from 1974 to 1984 on an indicator for the HRS and its interaction with the high-patriarchy indicator. Extensive evidence shows the important role played by land rights in incentives to increase agricultural output (Lin, 1988; Goldstein and Udry, 2008). It is therefore plausible that the majority of agricultural potential can be realised after the land reform granted residual claim rights to farmers. The boost to productivity driven by the new material incentives also holds in our data, forming a good foundation to test the relationship of our patriarchy index and agricultural productivity in the contemporary context. Table E2 presents the results estimating the impact of the HRS on grain output.

Table E2: Land reform, patriarchy and agricultural productivity

	Dependent variable: ln(grain output p. c.)	
	(1)	(2)
HRS	0.073***	0.078***
	(0.011)	(0.012)
HRS × high patriarchy		-0.012
		(0.022)
Observations	4,550	4,518
Counties	415	412
Outcome mean	5.784	5.781
County-specific trend	Y	Y

Notes: Dependent variable is the logarithm form of grain output per capita. County-specific linear time trends are included. Standard errors in parentheses are clustered at the county level. Significance levels denoted by *** p<0.01, ** p<0.05, * p<0.1.

As column (1) shows, the HRS improved grain output by approximately 7.3 percent. The result is similar after we control for the interaction of the HRS and our patriarchy index in Column (2). The results imply that there is no evidence that historical plough use positively influences agricultural output after the HRS, a period when efficient input utilisation was already in place (Sun and Chen, 2020). That is, historical plough use does not appear to constitute a key factor in improving agricultural productivity in a contemporary context.[40] Therefore, our baseline results

[40] This is perhaps unsurprising, given technical developments, especially in agricultural mechanisation, in the period of collectivisation (Hsu, 1979; Lin, 1987).

cannot be viewed as a consequence of the persistence of comparative male productivity in those counties heavily relying on plough use historically.

Appendix F: Robustness Checks

In this Appendix, we report a number of checks to probe the robustness of our results.

Unweighted and Poisson pseudo-maximum likelihood regressions. In our main county-level results (Table 1), we weight by the number of women of childbearing age in each county-year. In Table F1, we instead report unweighted regressions. Findings are very similar to those from the baseline. Similarly, Table F2 reports Poisson pseudo-maximum likelihood regressions that retain observations with zero values (with weights applied) rather than excluding them as in our baseline specification. The results remain consistent.

Higher parity births. Results are consistent when we turn to higher-parity births (at least third births). Appendix Table F3 shows, as in our main findings, that the OCP leads to larger increases in higher-parity fertility following female children in high-patriarchy counties. Similarly, the HRS leads to more male-skewed sex ratios of higher-parity children, particularly in high-patriarchy counties.

"1.5"-Child Policy. The "1.5"-Child Policy started in 1984 in rural China and allowed agricultural residents to have a legal second child if their first child was a girl (Feng and Hao, 1992). To show that this policy is not the main driver of our results on second children, we control for the policy at the provincial level and do not find any significant changes in our results (Table F4).

Continuous patriarchy index. In our main results, we use a binary indicator to divide counties into high- and low-patriarchy subsamples. In Tables F5 and F6, we use the underlying continuous index itself. As the index takes on smaller values for places that are more "plough-positive" (and therefore more patriarchal), these coefficients should be interpreted in an opposite fashion to those reported in our main findings. Nevertheless, across all specifications, results are highly consistent with the baseline results.

High patriarchy defined within estimating sample. In the main text, we define high and low patriarchy by taking a median split of the patriarchy index across all of China. In Tables F7 and F8, we show consistent results when we use the median of the patriarchy index in our estimating samples as the cut-off to split counties into high and low patriarchy.

Complementary effects of reforms. We explore whether the effects of the two policies (OCP and HRS) are complementary, and if so, whether our findings remain for the heterogeneous patterns of son preference across high- and low-patriarchy counties. Therefore, in Table F9 we further add the interaction of the two policies, as well as interactions with the "girl first" indicator and the high-patriarchy indicator. Our findings remain robust for the independent policy effects on the manifestation of son preference, both in terms of second-child fertility and male-skewed sex ratio. Note that while the coefficient on the interaction between HRS, high patriarchy and "girl first" is statistically insignificant in Column (2), it is virtually identical in magnitude compared to our main findings.

Effects of each reform before and after the other reform. To provide further evidence for the independent effects of the OCP and HRS, we divide the sample based on the relative implementation time of the two policies, with the OCP either before or after the HRS (and vice versa). The results are presented in Appendix Tables F10 and F11. Coefficients are very stable compared with the baseline results, though occasionally statistically insignificant.

Spatial trends. There is a concern, when studying enduring legacies of the past, that spatial trends and autocorrelation confound the estimated effects. Given our econometric design, it is unlikely to constitute a major problem in our setting. Our regressions control for county by birth year fixed effects, and standard errors are clustered at the county level. This in effect means that we are using variation within counties over time to identify our coefficients of interest. While spatial trends may well influence our patriarchy index, for them to invalidate our empirical design they would therefore have to have time-varying impacts that map precisely into reform timings. This seems unlikely. However, we further conduct a placebo test by randomizing the distribution of the patriarchy index 1000 times (see Figure F1). The t-values from our main findings appear unlikely to have arisen by pure chance.

Including urban sample. In our baseline regressions, we exclude urban samples. In Table F12, we show that our conclusions are unchanged when we retain the urban observations.

Controlling for average slope. One potential issue is that broad geographical factors may drive our results, although we have demonstrated the cultural nature of our patriarchy index in Appendix E. Notably, as shown in Almond et al. (2019), interactions between policies, gender, and grain suitability are statistically insignificant. If any geographic factors could become the primary driver, there would be little reason to observe strong results based on our weighted crop suitability (the construction of patriarchy index) while grain suitability itself shows no significance. To further rule out this competing explanation, we include interaction terms between policies, gender and slope in our baseline specification (see Table F13), and find that our baseline results remain robust.[41]

Potentially heterogeneous effects in staggered difference-in-differences designs. Potentially heterogeneous effect of roll-out policies across units and time would lead to biased estimation in staggered difference-in-differences designs (Sun and Abraham, 2021; Borusyak et al., 2024). Following the method of Sun and Abraham (2021), we individually estimate the average treatment effect in each time window relative to the implementation of policies to exclude contamination from "forbidden" comparisons. The adjusted results of the dynamic event-study approach are presented in Figures F2 and F3, and are highly comparable to the baseline results.

[41] The slope variable represents the percentage of land area (0-100) with a slope of less than 2 degrees, as cultivation costs tend to outweigh benefits on steeper terrain (Nunn and Puga, 2012). Data are from ASTER Global Digital Elevation Model V003. See `https://www.earthdata.nasa.gov/data/catalog/lpcloud-astgtm-003`.

Table F1: Unweighted county-level regressions

	Dependent variable: ln(number of second births)			
	Full sample	South only	North only	South only (with wetland rice)
	(1)	(2)	(3)	(4)
Panel A: OCP only				
Girl first	-0.049***	-0.053***	-0.031	-0.058**
	(0.017)	(0.019)	(0.042)	(0.028)
High patriarchy × girl first	0.051**	0.048	0.037	0.026
	(0.024)	(0.033)	(0.047)	(0.034)
OCP × girl first	0.042*	0.046*	0.023	0.038
	(0.024)	(0.025)	(0.062)	(0.038)
OCP × high patriarchy × girl first	0.133***	0.220***	0.109	0.095*
	(0.036)	(0.059)	(0.070)	(0.048)
Panel B: HRS only				
Girl first	-0.047***	-0.048**	-0.041	-0.059**
	(0.017)	(0.019)	(0.040)	(0.030)
High patriarchy × girl first	0.079***	0.082**	0.073	0.049
	(0.024)	(0.033)	(0.045)	(0.035)
HRS × girl first	0.038	0.037	0.045	0.041
	(0.024)	(0.026)	(0.061)	(0.041)
HRS × high patriarchy × girl first	0.081**	0.159***	0.040	0.049
	(0.036)	(0.058)	(0.068)	(0.050)
Panel C: OCP and HRS				
Girl first	-0.050***	-0.052***	-0.037	-0.060**
	(0.018)	(0.020)	(0.042)	(0.029)
High patriarchy × girl first	0.056**	0.051	0.048	0.032
	(0.024)	(0.033)	(0.047)	(0.035)
OCP × girl first	0.032	0.052	-0.029	0.013
	(0.045)	(0.050)	(0.103)	(0.073)
OCP × high patriarchy × girl first	0.198***	0.254**	0.223*	0.174*
	(0.066)	(0.107)	(0.115)	(0.094)
HRS × girl first	0.012	-0.007	0.068	0.031
	(0.046)	(0.051)	(0.101)	(0.078)
HRS × high patriarchy × girl first	-0.080	-0.043	-0.143	-0.095
	(0.066)	(0.107)	(0.113)	(0.097)
Observations	22,065	12,778	9,287	12,778
Counties	901	515	386	515
Outcome mean	1.954	2.043	1.832	2.043
County × birth year FE	Y	Y	Y	Y

Notes: Estimates of Equation (1). Standard errors in parentheses are clustered at the county level. Significance levels denoted by *** p<0.01, ** p<0.05, * p<0.1.

Table F2: Poisson pseudo-maximum likelihood county-level regressions

	Dependent variable: number of second births			
	Full sample	South only	North only	South only (with wetland rice)
	(1)	(2)	(3)	(4)
Panel A: OCP only				
Girl first	-0.040***	-0.044***	-0.018	-0.020
	(0.012)	(0.013)	(0.023)	(0.017)
OCP × girl first	0.049***	0.054***	0.023	0.037
	(0.018)	(0.020)	(0.035)	(0.033)
High patriarchy × girl first	0.056***	0.057***	0.037	0.003
	(0.015)	(0.020)	(0.026)	(0.021)
OCP × high patriarchy × girl first	0.068**	0.079**	0.077*	0.049
	(0.028)	(0.040)	(0.044)	(0.038)
Panel B: HRS only				
Girl first	-0.027**	-0.030**	-0.012	-0.012
	(0.011)	(0.013)	(0.024)	(0.020)
HRS × girl first	0.030	0.032	0.015	0.025
	(0.018)	(0.021)	(0.039)	(0.042)
High patriarchy × girl first	0.074***	0.080***	0.055**	0.022
	(0.015)	(0.019)	(0.027)	(0.024)
HRS × high patriarchy × girl first	0.036	0.036	0.048	0.017
	(0.028)	(0.039)	(0.047)	(0.046)
Panel C: OCP and HRS				
Girl first	-0.037***	-0.041***	-0.017	-0.019
	(0.012)	(0.013)	(0.024)	(0.019)
High patriarchy × girl first	0.060***	0.064***	0.040	0.008
	(0.015)	(0.019)	(0.027)	(0.022)
OCP × girl first	0.083***	0.094***	0.032	0.056
	(0.028)	(0.032)	(0.039)	(0.041)
OCP × high patriarchy × girl first	0.108**	0.130**	0.118**	0.108**
	(0.044)	(0.062)	(0.055)	(0.054)
HRS × girl first	-0.039	-0.047	-0.011	-0.022
	(0.029)	(0.033)	(0.050)	(0.063)
HRS × high patriarchy × girl first	-0.050	-0.065	-0.050	-0.071
	(0.044)	(0.061)	(0.064)	(0.071)
Observations	22,864	13,152	9,712	13,152
Counties	901	515	386	515
Outcome mean	9.875	10.837	8.571	10.837
County × birth year FE	Y	Y	Y	Y

Notes: Estimates of Equation (1). Standard errors in parentheses are clustered at the county level. Significance levels denoted by *** p<0.01, ** p<0.05, * p<0.1.

Table F3: Impact of reforms on higher-parity births and sex ratio

	Dependent variable:	
	ln(number of births)	male child = 1
	(1)	(2)
Girl first	-1.237***	0.048***
	(0.039)	(0.006)
High patriarchy × girl first	0.368***	0.001
	(0.064)	(0.009)
HRS × girl first	0.090	0.017
	(0.089)	(0.014)
HRS × high patriarchy × girl first	-0.054	0.071***
	(0.129)	(0.022)
OCP × girl first	0.401***	0.030**
	(0.096)	(0.014)
OCP × high patriarchy × girl first	0.241*	-0.012
	(0.132)	(0.023)
Observations	19,749	169,352
Counties	902	902
Outcome mean	1.631	0.529
County × birth year FE	Y	Y

Notes: Girl first indicates all previous births are girls. The regressions in Columns (1) and (2) estimate Equations (1) and (2), respectively. The regression in Column (1) weights by the number of women in childbearing ages within a county in each birth year. Standard errors in parentheses are clustered at the county level. Significance levels denoted by *** $p<0.01$, ** $p<0.05$, * $p<0.1$.

Table F4: Controlling for the “1.5”-child policy

	Dependent variable:	
	ln(number of second births)	male second child = 1
	(1)	(2)
Girl first	-0.034**	0.026***
	(0.016)	(0.005)
High patriarchy × girl first	0.056**	-0.004
	(0.022)	(0.007)
HRS × girl first	0.009	0.010
	(0.044)	(0.012)
HRS × high patriarchy × girl first	-0.067	0.036**
	(0.072)	(0.017)
OCP × girl first	0.069	0.007
	(0.043)	(0.012)
OCP × high patriarchy × girl first	0.184**	-0.013
	(0.072)	(0.016)
1.5-CP × girl first	-0.055**	-0.000
	(0.028)	(0.009)
1.5-CP × high patriarchy × girl first	0.065	0.020
	(0.058)	(0.015)
Observations	22,065	225,772
Counties	901	901
Outcome mean	1.954	0.524
County × birth year FE	Y	Y

Notes: The regressions in Columns (1) and (2) estimate Equations (1) and (2), respectively. The regression in column (1) weights by the number of women in childbearing ages within a county in each birth year. Standard errors in parentheses are clustered at the county level. Significance levels denoted by *** p<0.01, ** p<0.05, * p<0.1.

Table F5: Continuous patriarchy index (impact on number of second births)

	Dependent variable: ln(number of second births)			
	Full sample	South only	North only	South only (with wetland rice)
	(1)	(2)	(3)	(4)
Panel A: OCP only				
Girl first	0.111	0.332*	0.011	0.268
	(0.076)	(0.189)	(0.093)	(0.219)
Patriarchy index × girl first	-0.142	-0.398*	-0.001	-0.321
	(0.089)	(0.212)	(0.120)	(0.242)
OCP × girl first	0.664***	1.357***	0.578***	1.394***
	(0.134)	(0.361)	(0.163)	(0.455)
OCP × patriarchy index × girl first	-0.626***	-1.381***	-0.573***	-1.406***
	(0.154)	(0.390)	(0.205)	(0.491)
Panel B: HRS only				
Girl first	0.208***	0.524***	0.101	0.499**
	(0.080)	(0.194)	(0.098)	(0.221)
Patriarchy index × girl first	-0.231**	-0.591***	-0.088	-0.556**
	(0.094)	(0.218)	(0.126)	(0.246)
HRS × girl first	0.540***	1.059***	0.465***	0.983**
	(0.137)	(0.376)	(0.166)	(0.451)
HRS × patriarchy index × girl first	-0.521***	-1.082***	-0.472**	-0.986**
	(0.157)	(0.410)	(0.210)	(0.489)
Panel C: OCP and HRS				
Girl first	0.106	0.329*	0.007	0.295
	(0.077)	(0.185)	(0.095)	(0.212)
Patriarchy index × girl first	-0.133	-0.393*	0.006	-0.350
	(0.091)	(0.209)	(0.123)	(0.237)
OCP × girl first	0.603***	1.322*	0.546**	1.632**
	(0.217)	(0.707)	(0.234)	(0.797)
OCP × patriarchy index × girl first	-0.524**	-1.313*	-0.503*	-1.644*
	(0.251)	(0.779)	(0.294)	(0.867)
HRS × girl first	0.074	0.036	0.039	-0.303
	(0.223)	(0.740)	(0.240)	(0.809)
HRS × patriarchy index × girl first	-0.123	-0.074	-0.086	0.304
	(0.259)	(0.819)	(0.303)	(0.885)
Observations	22,065	12,778	9,287	12,778
Counties	901	515	386	515
Outcome mean	1.954	2.043	1.832	2.043
County × birth year FE	Y	Y	Y	Y

Notes: Estimates of Equation (1), with a continuous patriarchy index (lower values indicate stronger patriarchy) instead of an indicator for above-median patriarchy. Regressions weight by the number of women in childbearing ages within a county in each birth year. Standard errors in parentheses are clustered at the county level. Significance levels denoted by *** p<0.01, ** p<0.05, * p<0.1.

Table F6: Continuous patriarchy index (impact on sex ratio of second children)

	Dependent variable: male second child = 1			
	Full sample	South only	North only	South only (with wetland rice)
	(1)	(2)	(3)	(4)
Panel A: OCP only				
Girl first	0.014	0.062	0.001	0.083
	(0.026)	(0.053)	(0.033)	(0.064)
Patriarchy index × girl first	0.015	-0.039	0.029	-0.061
	(0.031)	(0.060)	(0.042)	(0.071)
OCP × girl first	0.091**	0.248***	0.041	0.163*
	(0.036)	(0.069)	(0.046)	(0.087)
OCP × patriarchy index × girl first	-0.078*	-0.254***	-0.015	-0.157
	(0.042)	(0.078)	(0.057)	(0.097)
Panel B: HRS only				
Girl first	0.004	0.028	0.002	0.038
	(0.025)	(0.056)	(0.031)	(0.067)
Patriarchy index × girl first	0.024	-0.002	0.025	-0.013
	(0.030)	(0.063)	(0.040)	(0.075)
HRS × girl first	0.124***	0.331***	0.050	0.258***
	(0.037)	(0.072)	(0.047)	(0.088)
HRS × patriarchy index × girl first	-0.111**	-0.343***	-0.018	-0.258***
	(0.043)	(0.081)	(0.059)	(0.098)
Panel C: OCP and HRS				
Girl first	0.007	0.034	0.001	0.049
	(0.026)	(0.055)	(0.033)	(0.068)
Patriarchy index × girl first	0.021	-0.008	0.026	-0.026
	(0.031)	(0.062)	(0.042)	(0.076)
OCP × girl first	-0.019	-0.057	0.006	-0.135
	(0.058)	(0.122)	(0.071)	(0.141)
OCP × patriarchy index × girl first	0.024	0.068	-0.009	0.155
	(0.069)	(0.139)	(0.089)	(0.158)
HRS × girl first	0.140**	0.378***	0.045	0.370***
	(0.060)	(0.129)	(0.073)	(0.142)
HRS × patriarchy index × girl first	-0.131*	-0.399***	-0.010	-0.387**
	(0.070)	(0.146)	(0.092)	(0.159)
Observations	225,772	142,533	83,239	142,533
Counties	901	515	386	515
Outcome mean	0.524	0.523	0.526	0.523
County × birth year FE	Y	Y	Y	Y

Notes: Estimates of Equation (2), with a continuous patriarchy index (lower values indicate stronger patriarchy) instead of an indicator for above-median patriarchy. Standard errors in parentheses are clustered at the county level. Significance levels denoted by *** p<0.01, ** p<0.05, * p<0.1.

Table F7: High patriarchy defined within estimating sample (number of second births)

	Dependent variable: ln(number of second births)			
	Full sample	South only	North only	South only (with wetland rice)
	(1)	(2)	(3)	(4)
Panel A: OCP only				
Girl first	-0.034**	-0.038**	-0.013	-0.028
	(0.016)	(0.017)	(0.041)	(0.023)
High patriarchy × girl first	0.045**	0.044	0.027	0.012
	(0.022)	(0.031)	(0.045)	(0.030)
OCP × girl first	0.038	0.041	0.019	0.027
	(0.025)	(0.028)	(0.064)	(0.036)
OCP × high patriarchy × girl first	0.172***	0.229***	0.139*	0.143***
	(0.042)	(0.065)	(0.073)	(0.052)
Panel B: HRS only				
Girl first	-0.029*	-0.032*	-0.014	-0.028
	(0.016)	(0.017)	(0.044)	(0.026)
High patriarchy × girl first	0.075***	0.081***	0.056	0.041
	(0.022)	(0.030)	(0.048)	(0.031)
HRS × girl first	0.030	0.032	0.020	0.028
	(0.027)	(0.029)	(0.070)	(0.039)
HRS × high patriarchy × girl first	0.123***	0.166***	0.093	0.094*
	(0.041)	(0.062)	(0.079)	(0.053)
Panel C: OCP and HRS				
Girl first	-0.034**	-0.038**	-0.014	-0.029
	(0.016)	(0.018)	(0.044)	(0.025)
High patriarchy × girl first	0.048**	0.048	0.032	0.017
	(0.022)	(0.030)	(0.048)	(0.030)
OCP × girl first	0.042	0.050	0.007	0.011
	(0.039)	(0.043)	(0.088)	(0.067)
OCP × high patriarchy × girl first	0.209***	0.264**	0.190*	0.206**
	(0.068)	(0.109)	(0.102)	(0.097)
HRS × girl first	-0.005	-0.010	0.015	0.019
	(0.041)	(0.045)	(0.100)	(0.072)
HRS × high patriarchy × girl first	-0.045	-0.044	-0.062	-0.075
	(0.067)	(0.104)	(0.112)	(0.097)
Observations	22,065	12,778	9,287	12,778
Counties	901	515	386	515
Outcome mean	1.954	2.043	1.832	2.043
County × birth year FE	Y	Y	Y	Y

Notes: Estimates of Equation (1), with the high patriarchy indicator defined within the estimating sample (rather than within all of China). Regressions weight by the number of women in childbearing ages within a county in each birth year. Standard errors in parentheses are clustered at the county level. Significance levels denoted by *** p<0.01, ** p<0.05, * p<0.1.

Table F8: High patriarchy defined within estimating sample (sex ratio of second children)

	Dependent variable: male second child = 1			
	Full sample	South only	North only	South only (with wetland rice)
	(1)	(2)	(3)	(4)
Panel A: OCP only				
Girl first	0.025***	0.025***	0.025*	0.017*
	(0.005)	(0.005)	(0.014)	(0.009)
High patriarchy × girl first	0.003	0.007	-0.002	0.015
	(0.007)	(0.009)	(0.016)	(0.010)
OCP × girl first	0.017**	0.016**	0.024	0.018
	(0.007)	(0.008)	(0.017)	(0.012)
OCP × high patriarchy × girl first	0.016	0.020*	0.007	0.005
	(0.009)	(0.012)	(0.019)	(0.014)
Panel B: HRS only				
Girl first	0.025***	0.026***	0.020	0.021**
	(0.005)	(0.005)	(0.015)	(0.009)
High patriarchy × girl first	-0.001	0.002	0.001	0.007
	(0.007)	(0.009)	(0.016)	(0.010)
HRS × girl first	0.017**	0.015*	0.033*	0.010
	(0.007)	(0.008)	(0.019)	(0.012)
HRS × high patriarchy × girl first	0.023**	0.031**	0.004	0.022
	(0.010)	(0.012)	(0.021)	(0.014)
Panel C: OCP and HRS				
Girl first	0.024***	0.025***	0.021	0.018**
	(0.005)	(0.005)	(0.015)	(0.009)
High patriarchy × girl first	0.000	0.003	-0.000	0.011
	(0.007)	(0.009)	(0.016)	(0.010)
OCP × girl first	0.007	0.011	-0.009	0.036*
	(0.013)	(0.014)	(0.029)	(0.021)
OCP × high patriarchy × girl first	-0.011	-0.017	0.009	-0.044*
	(0.016)	(0.020)	(0.032)	(0.024)
HRS × girl first	0.011	0.005	0.040	-0.021
	(0.013)	(0.014)	(0.031)	(0.021)
HRS × high patriarchy × girl first	0.032*	0.046**	-0.004	0.059**
	(0.017)	(0.021)	(0.034)	(0.024)
Observations	225,772	142,533	83,239	142,533
Counties	901	515	386	515
Outcome mean	0.524	0.523	0.526	0.523
County × birth year FE	Y	Y	Y	Y

Notes: Estimates of Equation (2), with the high patriarchy indicator defined within the estimating sample (rather than within all of China). Standard errors in parentheses are clustered at the county level. Significance levels denoted by *** p<0.01, ** p<0.05, * p<0.1.

Table F9: Interactive effects of OCP and HRS

	Dependent variable:	
	ln(number of second births)	male second child = 1
	(1)	(2)
Girl first	-0.047***	0.027***
	(0.017)	(0.005)
High patriarchy × girl first	0.060***	-0.004
	(0.023)	(0.007)
OCP × girl first	0.144***	0.003
	(0.051)	(0.016)
OCP × high patriarchy × girl first	0.183**	-0.011
	(0.082)	(0.021)
HRS × girl first	0.124*	0.003
	(0.069)	(0.017)
HRS × high patriarchy × girl first	-0.048	0.038
	(0.116)	(0.025)
OCP × HRS × girl first	-0.206**	0.010
	(0.086)	(0.024)
OCP × HRS × high patriarchy × girl first	-0.001	-0.000
	(0.137)	(0.033)
Observations	22,065	225,772
Counties	901	901
Outcome mean	1.954	0.524
County × birth year FE	Y	Y

Notes: The regressions in Columns (1) and (2) estimate Equations (1) and (2), respectively. The regression in column (1) weights by the number of women in childbearing ages within a county in each birth year. Standard errors in parentheses are clustered at the county level. Significance levels denoted by *** p<0.01, ** p<0.05, * p<0.1.

Table F10: Effect of OCP before and after HRS

	Dependent variable: ln(number of second births)	
	Before HRS	After HRS
	(1)	(2)
Girl first	-0.047***	0.077
	(0.017)	(0.067)
High patriarchy × girl first	0.060***	0.012
	(0.023)	(0.114)
OCP × girl first	0.144***	-0.061
	(0.051)	(0.070)
OCP × high patriarchy × girl first	0.183**	0.182
	(0.082)	(0.116)
Observations	11,921	10,144
Counties	900	900
Outcome mean	1.863	2.062
County × birth year FE	Y	Y

Notes: Estimates of Equation (1). Regressions weight by the number of women in childbearing ages within a county in each birth year. Standard errors in parentheses are clustered at the county level. Significance levels denoted by *** p<0.01, ** p<0.05, * p<0.1.

Table F11: Effect of HRS before and after OCP

	Dependent variable: male second child = 1	
	Before OCP	After OCP
	(1)	(2)
Girl first	0.027***	0.030*
	(0.005)	(0.016)
High patriarchy × girl first	-0.004	-0.014
	(0.007)	(0.020)
HRS × girl first	0.003	0.014
	(0.017)	(0.017)
HRS × high patriarchy × girl first	0.038	0.038*
	(0.026)	(0.022)
Observations	104,043	121,729
Counties	901	901
Outcome mean	0.516	0.531
County × birth year FE	Y	Y

Notes: Estimates of Equation (2). Standard errors in parentheses are clustered at the county level. Significance levels denoted by *** p<0.01, ** p<0.05, * p<0.1.

Table F12: Including urban sample

	Dependent variable:	
	ln(number of second births)	male second child = 1
	(1)	(2)
Girl first	-0.040***	0.028***
	(0.015)	(0.005)
High patriarchy × girl first	0.063***	-0.006
	(0.020)	(0.007)
OCP × girl first	0.068	0.003
	(0.041)	(0.012)
OCP × high patriarchy × girl first	0.185***	-0.012
	(0.070)	(0.016)
HRS × girl first	0.003	0.014
	(0.041)	(0.012)
HRS × high patriarchy × girl first	-0.042	0.039**
	(0.068)	(0.016)
Observations	22,275	239,035
Counties	904	904
Outcome mean	2.001	0.524
County × birth year FE	Y	Y

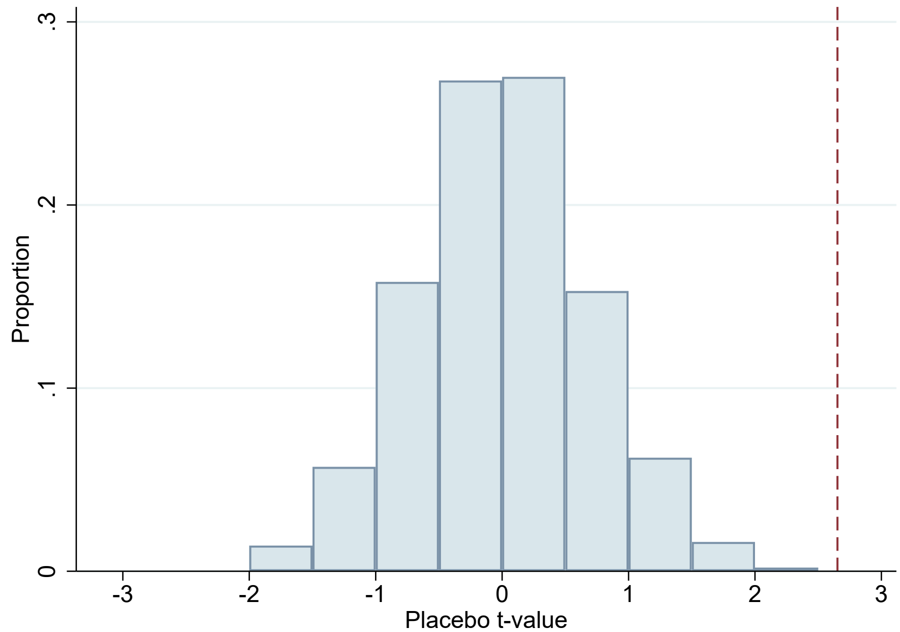
Notes: The regressions in Columns (1) and (2) estimate Equations (1) and (2), respectively. The regression in column (1) weights by the number of women in childbearing ages within a county in each birth year. Standard errors in parentheses are clustered at the county level. Significance levels denoted by *** p<0.01, ** p<0.05, * p<0.1.

Table F13: Controlling for average slope

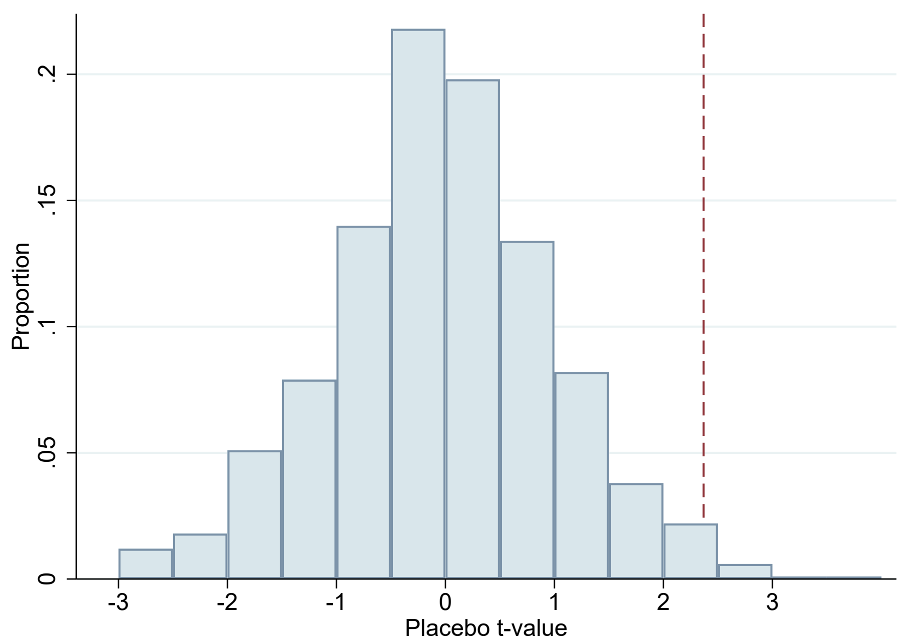
	Dependent variable:	
	ln(number of second births)	male second child = 1
	(1)	(2)
Girl first	-0.051***	0.031***
	(0.018)	(0.005)
High patriarchy × girl first	0.053**	-0.002
	(0.023)	(0.007)
OCP × girl first	0.031	0.017
	(0.058)	(0.014)
OCP × high patriarchy × girl first	0.186***	-0.008
	(0.065)	(0.016)
HRS × girl first	-0.006	-0.006
	(0.055)	(0.013)
HRS × high patriarchy × girl first	-0.071	0.033**
	(0.065)	(0.017)
Average slope × girl first	0.002	-0.001
	(0.001)	(0.000)
OCP × average slope × girl first	0.004	-0.001
	(0.005)	(0.001)
HRS × average slope × girl first	0.002	0.002**
	(0.005)	(0.001)
Observations	22,065	225,772
Counties	901	901
Outcome mean	1.954	0.524
County × birth year FE	Y	Y

Notes: The regressions in Columns (1) and (2) estimate Equations (1) and (2), respectively. The regression in column (1) weights by the number of women in childbearing ages within a county in each birth year. Standard errors in parentheses are clustered at the county level. Significance levels denoted by *** p<0.01, ** p<0.05, * p<0.1.

Figure F1: Permutation tests



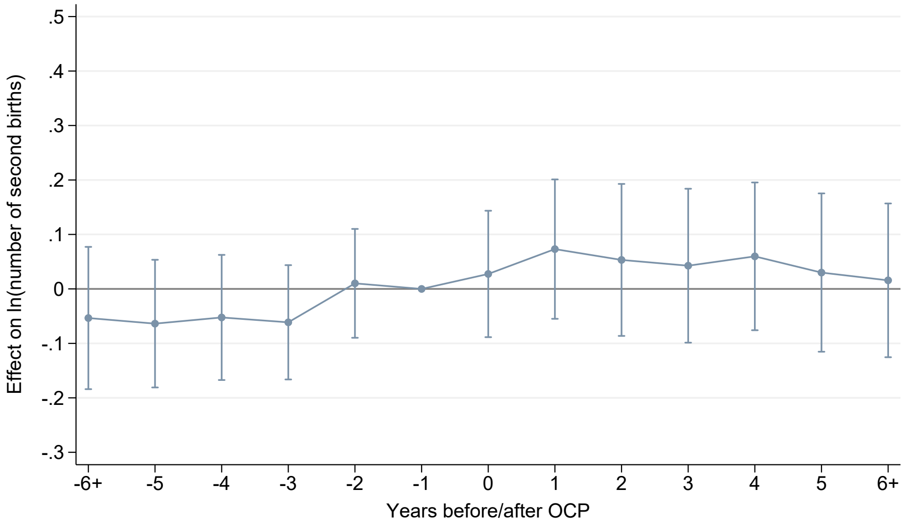
(a) Main coefficient (OCP × high patriarchy × girl first) for fertility results



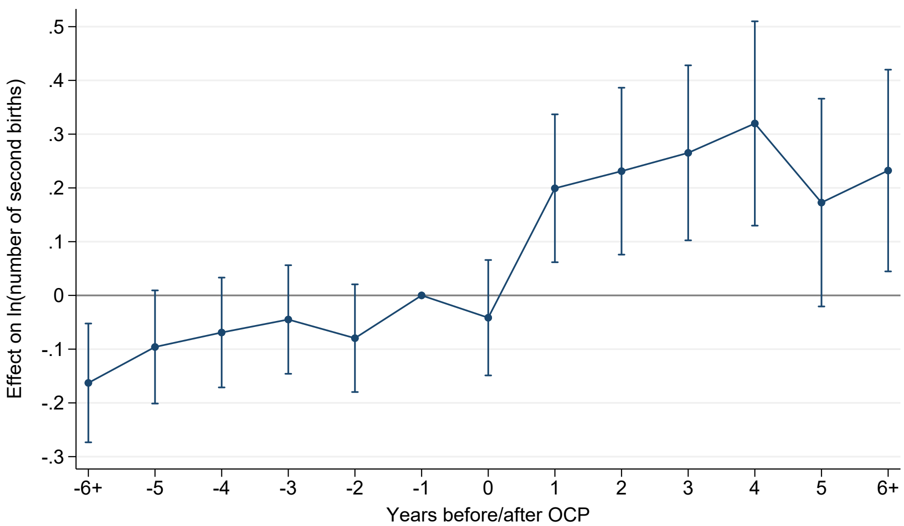
(b) Main coefficient (HRS × high patriarchy × girl first) for sex ratio results

Notes: The figures plot the distributions of t-values from the main coefficient of interest from Equations (1) and (2), where high- and low-patriarchy counties are assigned randomly. Each distribution is based on 1,000 independent draws. The t-value from the corresponding main result is indicated by the vertical dashed line.

Figure F2: Sun and Abraham (2021) estimator: effect of OCP on number of second births



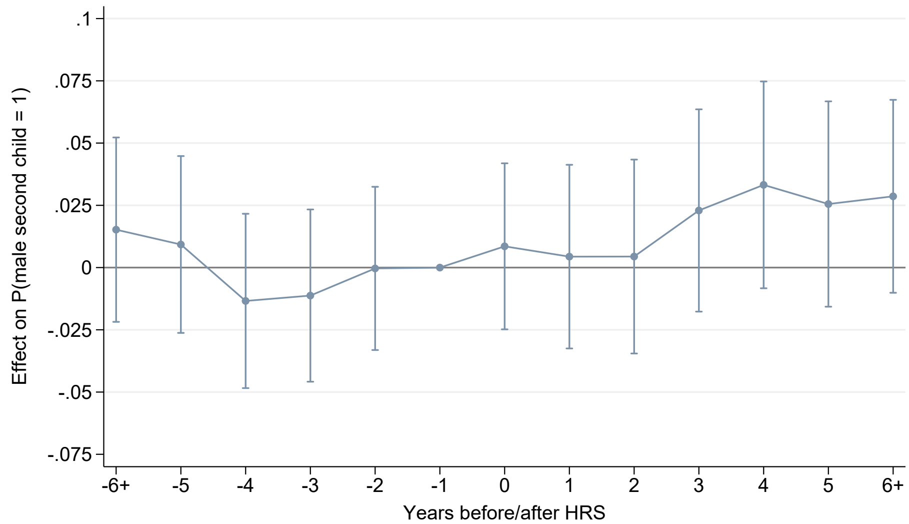
(a) Low-patriarchy counties



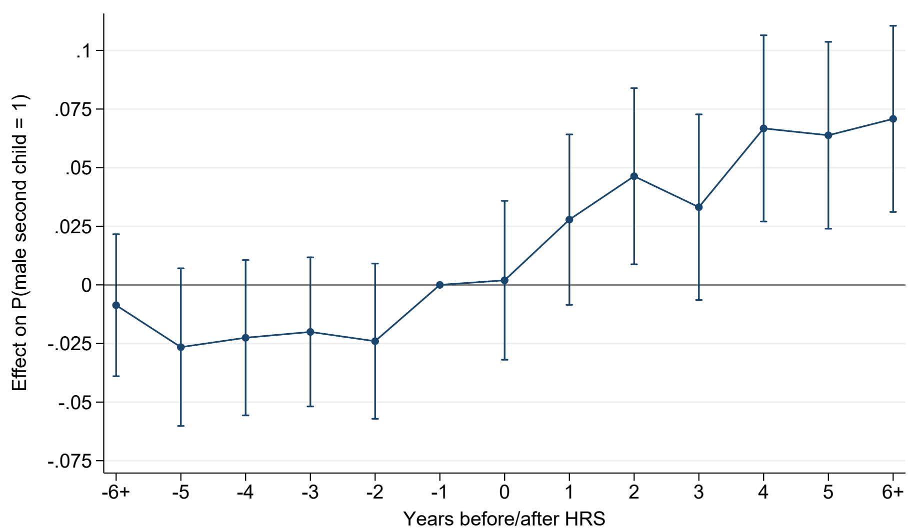
(b) High-patriarchy counties

Notes: Estimates of Equation (3) using the method of Sun and Abraham (2021), separately for low- and high-patriarchy counties. Standard errors are clustered at the county level and 95% confidence intervals are shown.

Figure F3: Sun and Abraham (2021) estimator: effect of HRS on the sex ratio of second children



(a) Low-patriarchy counties



(b) High-patriarchy counties

Notes: Estimates of Equation (3) using the method of Sun and Abraham (2021), separately for low- and high-patriarchy counties. Standard errors are clustered at the county level and 95% confidence intervals are shown.