

From Prohibition to Choice: The Impact of Abortion Legalization on Fertility and Child Investments in Nepal *

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November 2, 2025

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Abstract

We study whether expanding reproductive autonomy changes fertility and early-life investments in a son-preferring society, in the context of Nepal's 2002 abortion legalization. Using a triple-difference design comparing girls and boys across firstborn-sex families before and after the reform, we find that the abortion legalization substantially reduced son-biased fertility stopping: the sibling size gap between firstborn-girl and firstborn-boy families fell by nearly three-fifths. In terms of investments, daughters in firstborn-girl families gained about two months of breastfeeding, closing most of the pre-existing deficit, also consistent with reduced reliance on breastfeeding as a birth spacing tool. Effects on vaccination and under-five survival were limited, aligning with these inputs being less tied to fertility timing. Abortion access, therefore, relaxed fertility constraints and shifted investments on the birth spacing sensitive margin.

Keywords: Abortion, Fertility, Breastfeeding, Nepal

JEL codes: J13, J16, I18, I14

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1 Introduction

Fertility control is a first-order determinant of women’s well-being ([Goldin and Katz, 2002](#); [Bailey, 2006](#)). Restrictive abortion laws govern over 40% of the world’s population ([CRR, 2020](#)) and shape women’s lifetime outcomes and human-capital investments in children ([Londoño-Vélez and Saravia, 2025](#)). Abortion policy also shifts who is born and the conditions into which children arrive, affecting completed fertility and child outcomes ([Gruber et al., 1999](#)).

In South Asian societies with son preference, two facts organize the context. One, sex-selective abortion skews sex ratios at birth and is central to the “missing women” phenomenon ([Sen, 1990](#); [Anderson and Ray, 2010](#)). Access to prenatal sex-detection technology alters fertility behavior—weakens son-stopping, and changes postnatal resource allocations ([Anukriti et al., 2022](#)), consistent with the quantity–quality trade-off in fertility ([Becker, 1960](#); [Becker and Lewis, 1973](#); [Becker and Tomes, 1976](#)). Second, breastfeeding is both a core early-life investment, directly relevant to Sustainable Development Goals (SDGs) on ending malnutrition (SDG 2), child health and survival (SDG 3), and gender equality (SDG 5), and a common biological spacing mechanism where access to modern contraception is limited ([Jayachandran and Kuziemko, 2011](#)). How changes in legal access to abortion affect fertility and breastfeeding behavior in son-preferring societies remains an open question.

We answer this question in the context of Nepal’s 2002 abortion legalization by examining the joint impact on fertility response, specifically son-biased stopping behavior, and early-life investments. Before the reform, Nepal had one of the world’s most restrictive abortion laws. Abortion was a criminal offense in all circumstances, including rape and incest, and women could be imprisoned for seeking care. In 1997, about 20 percent of women in Nepali prisons had been convicted of abortion or infanticide ([CREHPA, 1996](#)). Unsafe, illegal procedures were a leading cause of maternal mortality ([Tamang, 1996](#); [Thapa et al., 2014](#)). After a decade of advocacy and legislative debate centered on high maternal morbidity and mortality, Nepal finally legalized abortion in 2002. First-trimester surgical services began in 2004, along with facility accreditation and provider training. Second-trimester abortion training began in 2007. Beginning in 2008, the government decentralized provision by authorizing and training nurses and auxiliary nurse-midwives ([Wariner et al., 2011](#)). In 2009, the landmark Supreme Court decision in *Lakshmi Dhikta v. Nepal* held the government accountable for building the necessary institutions and implementing policies to

make abortion services affordable and accessible, and medical abortion was introduced the same year. By the mid-2010s, access and utilization were widespread (Henderson et al., 2013; Adhikari, 2016).¹ To reflect this evolution in our analysis, we define Stage 1 as 2003–2007 and Stage 2 as 2008–2018. The rollout of abortion legalization provides a natural experiment to evaluate how expanded reproductive choice in a low-income, son-preference context affects fertility behavior and child well-being.

We use six nationally representative waves of the Nepal Demographic and Health Survey (1996, 2001, 2006, 2011, 2016, 2022), which provide complete retrospective birth histories for women aged 15–49. For fertility analysis, we construct a mother–year panel following women from marriage to the survey. Similarly, we use the birth histories to create child-level data to analyze early life investments. Our analysis combines three sources of variation. First, we compare families whose first child is a girl with those whose first child is a boy, which captures the incentive to continue childbearing to obtain a son. Quasi randomness of the gender of the firstborn child in the context of India is well-established in the literature (Alfano, 2017; Milazzo, 2018; Anukriti et al., 2022), and we also show that it holds in our data (see Figure A1) and provide details in Appendix B. Second, we compare cohorts born before and after legalization, which captures exposure to the reform. Third, we compare girls and boys among later-born children to measure gender gaps in investments and survival. We use these three sources of variation in a triple difference framework. The identifying assumption is that, absent the abortion law change, any secular trends in fertility or child investments would have affected first-daughter and first-son households similarly. Event-study evidence shows no differential pre-trends between firstborn girl vs. firstborn boy families. This approach accounts for general time trends in fertility and child health and isolates the reform’s impact on son-biased behaviors. The child-sex margin captures gender gaps in investments and survival, and differencing by gender purges unobservable trends that equally affect both boys and girls.

Our first main result is the effect of the policy on fertility, measured as the probability of a subsequent birth in the mother–year panel. Before 2002, Nepali couples who had a daughter as their first child were 4.6 percentage points (p.p.) more likely to keep having children (and to have a larger total number of births) than those whose first child was a son – a gap driven by the pursuit

¹Access to prenatal ultrasound also evolved over this period. We describe the timing and enforcement in Section 3 and report the effects of ultrasound separately in Table 11.

of a son. After abortion became legally available, this gap narrowed by roughly three-quarters. In other words, parents with firstborn girls became less likely to continue childbearing solely to attain a son, indicating a weakening of the son-biased fertility stopping norm. This suggests that expanding reproductive choice allowed families to better align fertility with their ideal family size rather than being dictated purely by son preference. We also find effects on birth spacing. Pre-legalization, families with a firstborn daughter had birth intervals that were 7.26 months shorter than families with a firstborn son. After legalization, spacing in firstborn-girl families lengthened by 1.93 months relative to the pre-legalization period.

Next, we estimate effects for postnatal investments. We find an increase in breastfeeding duration for girls relative to boys, contributing to a reduction in the gender gap in breastfeeding. Prior to the reform, Nepali girls were breastfed for shorter periods (for about 1.7–2.0 months less than sons) as parents hastened to try for a son. Post-reform, daughters in firstborn-girl families were breastfed about two months longer, nearly closing the baseline gap. Given the importance of breastfeeding for child survival and development, this is a meaningful improvement in girls' welfare. These results are consistent with the following two mechanisms: First, these patterns are consistent with reduced use of breastfeeding to accelerate the next conception once abortion became available ([Jayachandran and Kuziemko, 2011](#)). Second, abortion increases the likelihood that the girls who are born are more 'wanted' and receive increased parental investments ([Gruber et al., 1999](#); [Anukriti et al., 2022](#)).

Finally, we examine the effects on child vaccination coverage and under-five mortality. For vaccinations, we find muted effects with no significant changes in gender differences after the reform. Immunization coverage in Nepal was already relatively high for both genders ([Ashish et al., 2017](#)), and it appears to have continued to improve for all children, irrespective of sex. Likewise, under-five mortality rates fell over the 2000s due to general health progress, but we do not detect a significant shift in the girl–boy survival gap attributable to the abortion reform. The absence of a sizable impact on these latter outcomes suggests that the policy's primary short-run effects operated through fertility behavior and breastfeeding practices, rather than through broad changes in health service utilization or neglect. Taken together, our results indicate that Nepal's abortion liberalization attenuated the manifestations of son preference (fewer extra births and longer breastfeeding when a daughter was born), without harming, and possibly modestly

benefiting, girls' survival prospects.

We conduct a series of robustness checks to validate our results. First, event-study graphs and placebo tests show no evidence of differential pre-trends in the 1980–1990 window for our main outcomes, a period that precedes ultrasound diffusion. Second, we address concerns about concurrent changes in access to prenatal sex detection. We examine in isolation the effects of ultrasound on fertility. Using a proxy for the diffusion of prenatal sex detection, we replicate [Anukriti et al. \(2022\)](#) on the probability of an additional birth: for firstborn-girl families, the probability of a subsequent birth declines by about 1.1–2.0 percentage points in the early post-ultrasound period and by about 3.2–4.2 percentage points in the later period, closely matching prior magnitudes. In our main legalization estimates, we define 1990–2002 as the pre-period and treat ultrasound diffusion as part of the baseline environment, so the law's effect is identified net of ultrasound. Third, the findings are stable across alternative specifications. Results are robust to alternative fixed-effect structures, including child-gender-specific cohort fixed effects, cohort fixed effects that vary by firstborn-sex family type, and district-year fixed effects that absorb nationwide gender-differential trends and local shocks. They are unchanged when we redefine exposure using early and late post-reform phases, when we vary the clustering level across administrative units, and when we add controls for evolving preferences using mothers' reported ideal number of children and ideal sex composition and their interactions with cohort and firstborn sex.

A large body of literature studies how abortion policy changes affect fertility and the average conditions into which children are born ([Gruber et al., 1999](#); [Pop-Eleches, 2006](#); [Ananat et al., 2009](#)). In son-preferring societies, evidence centers on access to prenatal sex detection rather than on policy-driven legalization ([Almond et al., 2010](#); [Hu and Schlosser, 2015](#); [Anukriti et al., 2022](#)). In Taiwan, [Lin et al. \(2014\)](#) show that abortion legalization reduced fertility at third and higher parities and lowered neonatal (but not post-neonatal) female mortality at higher parities conditional on socioeconomic status; they did not examine postnatal health investments. In Mexico City, [Clarke and Mühlrad \(2016\)](#) find that decriminalization reduced births and maternal mortality. We contribute to the literature in two ways. First, we add evidence on abortion legalization showing that abortion access can reduce son-biased fertility continuation in a setting with son preference. Second, in line with the fertility-driven quantity-quality trade-off, we show effects on early-life investments driven by changes in breastfeeding duration.

The remainder of the paper proceeds as follows. Section 2 shows where the paper falls within the literature. Section 3 provides background on Nepal’s reform and demographic context. Section 4 describes the data and measurement. Section 5 presents the empirical strategy. Section 6 reports the results on fertility, breastfeeding, vaccination, and under-five mortality. Section 7 concludes with implications for policy and future research.

2 Literature Review and Contribution

This paper situates the study within two key strands of literature: (i) research on the consequences of expanding women’s reproductive autonomy on fertility and child outcomes, and (ii) the operation of the quantity–quality trade-off in contexts with entrenched son preference. Together, these literatures provide the theoretical and empirical foundation for understanding how abortion legalization in Nepal shaped fertility behavior and gender gaps in parental investments.

Evidence from developed countries points to large long-term effects of reproductive autonomy on children’s outcomes.² [Levine et al. \(1996\)](#) and [Gruber et al. \(1999\)](#) show that abortion legalization in the United States improved the circumstances of the “marginal child,” leading to higher average parental investments and better child outcomes.³ [Pop-Eleches \(2006\)](#) finds that the reversal of abortion access in Romania worsened children’s long-term educational and labor market trajectories, consistent with the idea that unwanted births dilute household resources. Collectively, this literature establishes that reproductive autonomy affects both the quantity of children and the quality of investments per child.

There is limited but growing empirical evidence on these long-term effects in low- and middle-income countries. [Clarke and Mühlrad \(2016\)](#) find that abortion legalization in Mexico reduced fertility and maternal mortality, while [Holcombe and Kidanemariam Gebru \(2022\)](#) document similar fertility declines in Ethiopia. In Taiwan, [Lin et al. \(2014\)](#) show that access to abortion in conjunction with prenatal sex detection altered sex ratios at birth and reduced relative female mortality at higher parities. Yet most of these studies focus on fertility and survival outcomes;

²For mothers in high-income settings, early access to oral contraceptives and legal abortion expanded women’s control over fertility, shifting education, work, and marriage timing and reducing unintended births; women could better align childbearing with life-cycle plans ([Goldin and Katz \(2002\)](#); [Bailey \(2006\)](#)).

³Prior studies on abortion legalization document substantial effects on fertility and child outcomes. [Levine et al. \(1996\)](#) find that U.S. birth rates declined by around 8% in early-repeal states, with the largest reductions among teens, older mothers, and unmarried women. [Gruber, Levine, and Staiger \(1999\)](#) show that the “marginal child” not born due to legalization would have been 40–60% more likely to live in poverty, receive welfare, live in a single-parent household, and die in infancy.

fewer examine how legalization affects postnatal parental investments such as breastfeeding, vaccination, or schooling.

Economic models of the family emphasize the trade-off between the number of children (quantity) and the investments made in each child (quality) ([Becker and Lewis, 1973](#); [Becker and Tomes, 1976](#)). In this framework, households face a budget constraint: as fertility rises, resources per child fall, while fertility decline creates scope for greater per-child investment. Empirical studies confirm this mechanism. In India, [Rosenzweig and Wolpin \(1980\)](#) show that exogenous fertility shocks lowered children's schooling, while [Angrist et al. \(2010\)](#) document similar effects in Israel. In China, [Li et al. \(2008\)](#) find that fertility reductions improved education. These findings underscore that fertility decline often coincides with greater human capital investments, consistent with the model's predictions.

The strength of the trade-off, however, depends on parental preferences, household resources, and cultural norms ([Rosenzweig and Zhang, 2009](#); [Black et al., 2005](#)). When families value sons over daughters, fertility reductions may disproportionately benefit boys, leaving girls disadvantaged even as overall resources per child increase. In South Asia, strong son preference has long shaped fertility and investment decisions ([Gupta \(1987\)](#); [Bhat and Xavier \(2003\)](#)). Sons are often seen as economic and social assets, while daughters are viewed as financial liabilities due to dowry and marriage practices ([Jayachandran and Pande \(2017\)](#)). This preference manifests in son-biased fertility stopping, where families continue childbearing until a desired number of sons are born. As a result, girls often grow up in larger sibships, facing diluted resources relative to boys.

A large body of empirical work documents the consequences of son preference for child outcomes. Girls in son-preferring households receive less schooling, fewer health inputs, and face higher mortality, despite the biological survival advantage of female infants ([Waldron \(1998\)](#); [Jayachandran and Kuziemko \(2011\)](#)). Access to prenatal sex-detection technologies has further complicated these dynamics. In India, [Anukriti et al. \(2022\)](#) find that access to abortion and sex selection narrowed gender gaps in breastfeeding and vaccination by increasing the "wantedness" of surviving girls. Closest to our study, they implement a triple-difference strategy, interacting pre/post access to prenatal sex detection with firstborn-sex and child sex, to show that ultrasound-enabled sex selection moderated son-biased stopping, narrowed gender gaps in breastfeeding and

immunization, and reduced excess female mortality into early childhood ⁴.

Similarly, [Hu and Schlosser \(2015\)](#) show that in areas with higher uptake of sex selection, girls' nutritional status and survival improved. In contrast, [Almond et al. \(2010\)](#) document that in some settings, prenatal sex determination increased female neonatal mortality without corresponding improvements in postnatal investments. These mixed results highlight that the consequences of sex selection are highly context-dependent, shaped by cultural norms, fertility preferences, and the availability of health services.

3 Fertility Trends in Nepal

In Nepal, son preference is deeply embedded in patriarchal traditions. Sons are regarded as economic and social assets, while daughters are often viewed as financial liabilities due to dowry obligations and the expectation of marriage outside the parental home ([Brunson, 2010](#)). This strong son preference shapes fertility behavior, leading to larger numbers of male children. According to the 2011 Census, among children under age 10, boys outnumbered girls by 2.2 percent, with disparities more pronounced in urban areas (5.6 percent) than rural areas (1.7 percent). Despite the biological survival advantage of female newborns ([Waldron, 1998](#)), Nepal's mortality patterns diverge: male neonatal mortality exceeds female rates (37 versus 33 per thousand live births), but female post-neonatal mortality (ages 28 days to one year) surpasses that of males (19 versus 17 per thousand) ([Ministry of Health, 2012](#)).

The spread of prenatal sex-determination technology, particularly ultrasound, has further enabled couples to act on son preference through sex-selective abortion, contributing to elevated sex ratios at birth (SRB) and broader population imbalances ([Guilmoto, 2009](#)). Ultrasound was first introduced in Nepal in 1988 at Bir Hospital in Kathmandu, when the government of India donated an ultrasound unit ([Subedi and Sharma, 2013](#); [Mukhiya and Mishra, 2025](#)). Since then, services have expanded rapidly and are now widely accessible, with scans costing as little as USD 6 even in rural and remote areas ([NHSP, 2013](#)). Because ultrasound diffusion may itself affect fertility behavior and child investments, we explicitly test its impact in Table 11. In our main analysis,

⁴[Anukriti et al. \(2022\)](#) show that the probability of a subsequent birth for firstborn-girl families fell by about 2.0–2.3 percentage points in the early post-ultrasound period and by about 3.7–4.3 percentage points in the later post period, relative to firstborn-boy families. Using our ultrasound proxy, we find a decline of 1.1–2.0 p.p. in the early post period and 3.2–4.2 p.p. in the later post period (Table 11), magnitudes that are strikingly similar. This replication reinforces that ultrasound availability reduced son-biased fertility continuation in both India and Nepal.

however, we treat 1990 as the baseline year so that the estimated treatment effect captures the impact of abortion legalization, net of any influence from ultrasound availability.

At the same time, fertility in Nepal has declined sharply—from 4.6 births per woman in 1996 to 2.6 in 2011 ([Ministry of Health, 2012](#)). The legalization of abortion in 2002 (policy details below) and subsequent expansion of safe abortion services have further altered reproductive decision-making. Together, the availability of sex-detection technology, entrenched son preference, expanded abortion access, and rapid fertility decline create conditions conducive to sex-selective abortion, skewed sex ratios, and long-term demographic challenges, including shortages of marriageable women, and potentially heightened risks of gender-based violence, abduction, and trafficking ([Hesketh and Xing, 2006](#); [Bien et al., 2013](#)).

National Safe Abortion Policy and Strategy

Prior to 2002, abortion was considered a criminal act (homicide) in the *Muluki Ain* (the National Legal Code). Due to the restrictive law, most of the abortions were conducted illegally by unqualified personnel. These service providers used many barbaric procedures, like inserting cow dung, unknown medicines, or herbal mixtures into the uterus ([Tamang, 1996](#)). Deaths from abortion-related complications accounted for more than half of the maternal deaths that occurred in major hospitals ([Thapa et al., 2014](#)). In 1997, 20% of women in Nepali jails had been convicted on charges of abortion or infanticide. Mostly poor and illiterate, they were kept in miserable conditions, unable to afford legal assistance or even to understand what had happened to them. Only women were imprisoned; their male partners and the abortion providers were not held accountable ([CREHPA, 1996](#)).

In 2002, in response to mounting evidence of maternal deaths and injuries from unsafe abortions and to expand women's reproductive autonomy, the Government of Nepal amended the *Muluki Ain* 1959, which had previously prohibited abortion under all circumstances and classified it as an offense against life. The amendment paved the way for the National Safe Abortion Policy and Strategy 2002, which guarantees access to safe abortion services under specific conditions: up to 12 weeks of gestation with the pregnant woman's consent; up to 18 weeks in cases of rape or incest; and at any stage of pregnancy if it endangers the woman's life, physical, or mental health, or in the case of a severe fetal anomaly, with the recommendation of a medical practitioner and

the woman's consent ([Ministry of Health, 2002](#)). Sex-selective abortion is “prohibited”⁵, and adult consent is required for girls less than 16 years old.

In 2004, Nepal's first certified abortion clinic opened, marking the beginning of a steady expansion of services. Strong government leadership ensured the implementation of safe abortion services in Nepal. The Ministry of Health established a multisector Abortion Task Force that brought together public agencies and national and international nongovernmental organizations to draft policy and design strategies for provider training and service delivery. The Nepal Society of Obstetricians and Gynecologists provided technical support in developing standardized clinical protocols and training guidelines ([Wu et al., 2017](#)). In 2004, the Abortion Task Force was dissolved and replaced by the Technical Committee for the Implementation of Comprehensive Abortion Care, which oversaw accreditation, training, and quality assurance, and ensured that policies were evidence-based and updated as new data became available ([Samandari et al., 2012](#)). Initially, training in manual vacuum aspiration (MVA) was limited to physicians; however, beginning in 2008, staff nurses and auxiliary nurse midwives were permitted to perform MVA for pregnancies up to eight weeks. Second-trimester abortion training and certification for physicians commenced in 2007, and medication abortion was incorporated into the safe abortion program in 2009. Also in 2009, the landmark Supreme Court decision in *Lakshmi Dhikta v. Nepal* held the government accountable for building the necessary institutions and implementing policies to make abortion services affordable and accessible ([Samandari et al., 2012](#)). Data shows that the policy was implemented successfully nationwide. Between 2011 and 2016 reported over 400,000 abortions were performed at legal, safe abortion sites ([Ministry of Health, 2016](#)). Maternal mortality in Nepal decreased from 548 deaths per 100,000 live births in 2000 to 258 deaths per 100,000 live births in 2015. Research also indicates a substantial long-term reduction in maternal health risks, such as infection and injury, following the reform, with declines in sepsis observed even during the early implementation phase ([Henderson et al., 2013](#)).

⁵Despite sex-selective abortion being banned, evidence suggests that it is still prevalent ([Frost et al., 2013](#); [Lamichhane et al., 2011](#)) because of difficulty ascertaining whether families are seeking abortion for sex-selection purposes, and fear that women will resort to unsafe abortion if they are under pressure to bear sons but unable to access safe abortion services.

4 Data and Measurement

4.1 Data

This study draws on six waves of the Nepal Demographic and Health Survey (NDHS), conducted in 1996, 2001, 2006, 2011, 2016, and 2022. The NDHS is nationally representative and provides complete retrospective birth histories for women aged 15–49, including children’s year of birth, birth order, age at death, and maternal characteristics at birth. The 1996 survey interviewed 8,429 ever-married women; the 2001 survey interviewed 8,726 women; the 2006 survey interviewed 10,793 women; the 2011 survey interviewed 12,674 women; the 2016 survey interviewed 12,862 women; and the 2022 survey interviewed 14,845 women.

For the fertility analysis, we pool all surveyed women to create a woman-level dataset. The analytic sample excludes women younger than age 13 and those who had not given birth by the time of the survey. We also restrict the sample to women who began childbearing in 1980 or later,⁶ yielding a final sample of 18,350 mothers. For analyses of fertility timing, we construct a mother–year panel in which women enter in their year of marriage and remain until the year of the survey.

For the mortality and postnatal investment analyses, we pool all births of surveyed women to construct a child-level dataset. To measure under-five mortality, we exclude children younger than five at the time of survey (to allow full exposure to mortality risk) and those older than 15 (to mitigate recall bias).⁷ These restrictions yield 36,298 children for the mortality sample. Breast-feeding analyses focus on the last two surviving births of each mother and are limited to children at least two years old to account for censoring.⁸ This leaves 5,858 children in the breastfeeding sample.

4.2 Measurement

The NDHS includes women’s complete birth histories and their children’s year of birth and other outcomes. Our first fertility analysis relies on the mother’s birth profile. We construct a variable *Birth* which observes each woman from the year of her marriage until the survey year and assign a

⁶We keep all women whose first child was born on or after 1980.

⁷In Appendix D, we also show results for mortality at each age interval between birth and age five. For each outcome, children younger than the cutoff age are excluded to ensure complete exposure. Results are robust to alternative age cutoffs.

⁸We further restrict the sample to children born within 20 years of the interview date.

value of one if the mother gave birth in a certain year and zero if not. Our second fertility analysis uses the number of children that the mother had during the time of the survey from the fertility module.

We build child-health investment outcomes from mothers' DHS birth histories and child health modules. For breastfeeding, we use the mother's report of "How many months did you breastfeed your x^{th} child?" and take breastfeeding duration in months as the outcome. To limit censoring and recall error, we restrict to the last two surviving births and to children who are at least 24 months old at interview. For vaccinations, we use DHS records (from vaccination cards or maternal recall) on whether the child received each of the eight standard EPI doses⁹, and construct a simple count from 0 to 8. Finally, under-five mortality is measured with a child-level indicator equal to 1 if the child died before 60 months of age and 0 otherwise. To ensure full exposure to risk and reduce recall concerns, we restrict the mortality sample to children who are at least 60 months and at most 15 years old at the time of the survey.

The summary statistics for fertility, child health investments, and mortality are provided in Table 1. The average mother in the sample is 29 years old, with ages ranging from 15 to 49. A large majority (87 percent) identify as Hindu. Only 43 percent of mothers and 45 percent of fathers have at least a primary education. The mean household wealth index falls around the middle of the distribution (3.1 on a scale of 1 to 5). On average, mothers report having nearly five children, with family sizes ranging from 1 to 8. Children are breastfed for about 31 months on average, though the duration varies widely across families. The under-five mortality rate is 8.9 percent, and children receive an average of 6.7 vaccinations out of a maximum of 8.

5 Empirical Strategy

Our identification strategy relies on two key facts: (a) the sex of the firstborn child is effectively random, which we demonstrate empirically rather than assume in panel (a) of Figure A1, and (b) sex-selective abortion at second and higher-order births occurs primarily in families whose firstborn is a daughter (panels (b) and (c) of Figure A1). The central identifying assumption, tested directly in our analysis in Figure A2, is that, absent access to ultrasound technology, trends in outcomes would have evolved similarly for families with firstborn sons and firstborn daughters.

⁹The eight vaccinations include BCG, DPT 1, DPT 2, DPT 3, Polio 1, Polio 2, Polio 3, and Measles.

Table 1: Summary Statistics

| | N | Mean | Std. Dev | Min | Max |
|---------------------------------------|-------|-------|----------|-----|-----|
| Mother's age | 18350 | 28.82 | 7.54755 | 15 | 49 |
| Mother's Education (At least Primary) | 18350 | 0.43 | .4957616 | 0 | 1 |
| Father's Education (At least Primary) | 18350 | 0.45 | .4970647 | 0 | 1 |
| Wealth Quintile | 18350 | 3.10 | 1.415311 | 1 | 5 |
| 1 [Hindu] | 18350 | 0.87 | .3358624 | 0 | 1 |
| Total no. of children | 18350 | 2.02 | 1.064824 | 1 | 8 |
| Months of breastfeeding | 5858 | 31.07 | 6.959548 | 0 | 35 |
| Under-five mortality (in %) | 36298 | 8.90 | 28.47667 | 0 | 100 |
| No. of vaccinations | 13404 | 6.69 | 2.335607 | 0 | 8 |

Notes: Data source: Nepal Demographic and Health Survey (NDHS).

We estimate a triple-difference specification that exploits the 2002 legalization of abortion. Our design compares changes in fertility and child investments across families by the sex of the first-born child, treating 1990–2002 as the pre-period so that the estimated treatment effect captures the causal impact of abortion legalization while holding constant the contemporaneous diffusion of ultrasound technology. The triple interaction combines birth cohort variation in exposure to the reform with the sex of the firstborn child, a well-established proxy for parents' propensity to engage in sex selection, and the sex of the child. This structure allows us to test whether legalization narrowed the disadvantages faced by daughters in firstborn-girl families relative to other children.

Abortion was legalized nationwide in 2002, but service availability expanded gradually.¹⁰ To capture temporal variation in access, we classify the data into three periods: pre-legalization (1990–2002), early legalization (2003–2007), and late legalization (2008–2018), when access and utilization became widespread. Differentiating between the two post-legalization periods is important, as the late period saw the introduction of second-trimester abortion training, the authorization of nurses to provide services, and the integration of medical abortion into the national program.

We examine the impact of the legalization of abortion in gender gaps in fertility in two ways. First, we test if the legalization changed the probability of birth in a given year for mothers with firstborn girls versus firstborn boys. To do this, we construct a retrospective mother-year panel, where each woman is observed from the year of her marriage until the survey year. We estimate

¹⁰Table A1 provides further details.

$$Birth_{it} = \alpha + \beta_1 FirstbornGirl_i \times Post_t^1 + \beta_2 FirstbornGirl_i \times Post_t^2 + \gamma FirstbornGirl_i + \omega_t + X_i' \tau + \phi_a + \psi_b + \sigma_r + \delta_d + \nu_d FirstbornGirl_i + \theta_{dt} + \epsilon_{it} \quad (1)$$

for mother i from district d , aged a in year t , who has had $b - 1$ children by year t and whose last birth occurred r years ago. The outcome variable $Birth_{it}$ equals one if the mother gave birth in year t . The variable $FirstbornGirl_i$ equals one if mother j 's first child is a girl, and zero otherwise. $Post_t^1$ indicates that year t falls within the early years after the abortion legalization (2003-2007), while $Post_t^2$ corresponds to the later years after the abortion ban was lifted (2008-2018). The vector X_i comprises indicators for household wealth quintiles, the mother's and her husband's education level, region of residence, residence in a rural area, and the mother's year of birth. The fixed effects include year (ω_t), district (δ_d), birth parity (ψ_b), years since last birth (σ_r), district-specific firstborn-girl fixed effects ($\nu_d FirstbornGirl_i$), and district-specific year fixed effects (θ_{dt}). The standard errors are clustered by district, and we have seventy-five districts in our sample.

The second analysis examines whether the legalization of abortion had an effect on the total number of children a woman had at the time of the survey. We estimate

$$N_{jt} = \alpha + \beta_1 FirstbornGirl_j \times Post_t^1 + \beta_2 FirstbornGirl_j \times Post_t^2 + \gamma FirstbornGirl_j + \sigma Post_t^1 + \psi Post_t^2 + X_j' \tau + \delta_d + \nu_d FirstbornGirl_j + \theta_d Post_t^1 + \omega_d Post_t^2 + \epsilon_{jt}. \quad (2)$$

for mother j from district d who has had N_{jt} children as of the year of the survey t . To ensure consistent exposure, we restrict the sample to mothers who had all their births within a single period—either pre-legalization, early legalization (2003–2007), or late legalization (2008-2018). Thus, $Post_t^1$ and $Post_t^2$ indicate that the woman began and completed childbearing during the early or late legalization periods, respectively. As before, $FirstbornGirl_j$ is an indicator for whether the woman's first child was a girl. The vector X_j includes controls for household wealth quintiles, the education levels of the woman and her husband, rural residence, and the woman's birth year. The coefficients β_1 and β_2 test our hypothesis that there was less son bias in fertility decisions after the legalization.

We examine three child investment outcomes: months of breastfeeding, number of vaccinations, and under-five mortality. For all three, we estimate the following equation

$$\begin{aligned}
Y_i = & \alpha + \beta_1 FirstbornGirl_j \times Female_i \times Post_t^1 + \beta_2 FirstbornGirl_j \times Female_i \times Post_t^2 \\
& + \gamma FirstbornGirl_j \times Female_i + \omega_t FirstbornGirl_j + \sigma_t Female_i + \mathbf{X}_{ijt}' \tau + \delta_r Female_i \\
& + \nu_r FirstbornGirl_j + \psi_b Female_i + \xi_b FirstbornGirl_j + \rho_{bt} + \eta_{bd} + \phi_{dt} + \epsilon_i \quad (3)
\end{aligned}$$

for child i of birth order b born to mother j in year t and district d . The dependent variable Y_i is an indicator for either the months of breastfeeding, number of vaccinations, or mortality for child i . $Post_t^1$, $Post_t^2$, and $FirstbornGirl_j$ are defined as earlier. β_1 and β_2 , the coefficients of the two triple interaction terms are our coefficients of interest. This equation is estimated for second- and higher-order births, making pre-ultrasound births and second- and higher-order births to mothers whose firstborn is a boy our control group. The standard errors are clustered by district.

We do not observe large differences in the socioeconomic characteristics of firstborn-boy and firstborn-girl families (Table 2), so selection on the sex of the firstborn is of limited concern. However, we still control for socioeconomic conditions like wealth, education, and location, and consider sensitivity of the estimates to conditioning upon a mother's stated desired fertility and desired sex composition of children.

6 Results

This section presents the effects of abortion legalization on fertility and child investments in Nepal. We first examine fertility outcomes, focusing on son-biased fertility stopping behavior, and then turn to parental investments, including breastfeeding, vaccination, and child survival.

6.1 Effects on Fertility

In Tables 3 and 4, we present the impacts of the introduction of abortion legalization on son-biased fertility stopping. The results provide clear evidence that abortion legalization significantly reduced son-biased fertility stopping. Prior to legalization, families with a firstborn daughter were

Table 2: Test of Balance in Samples by the Sex of the Firstborn

| | 1990-2002 | | 2003-2007 | | 2008-2018 | | All years |
|------------------------------|-----------|------|-----------|------|-----------|------|-----------|
| | (1) | (2) | (3) | (4) | (5) | (6) | (7) |
| | FB | FG | FB | FG | FB | FG | FB-FG |
| Rural | 0.67 | 0.67 | 0.58 | 0.59 | 0.48 | 0.48 | -0.0083 |
| <i>Mother's education</i> | | | | | | | |
| No education | 0.63 | 0.64 | 0.41 | 0.41 | 0.22 | 0.22 | -0.00013 |
| Incomplete secondary | 0.31 | 0.30 | 0.47 | 0.47 | 0.56 | 0.57 | 0.00055 |
| Secondary or higher | 0.06 | 0.06 | 0.12 | 0.12 | 0.22 | 0.21 | -0.00039 |
| <i>Father's education</i> | | | | | | | |
| No education | 0.23 | 0.24 | 0.15 | 0.15 | 0.09 | 0.09 | 0.0010 |
| Incomplete secondary | 0.30 | 0.30 | 0.32 | 0.31 | 0.31 | 0.32 | 0.0016 |
| Secondary or higher | 0.47 | 0.46 | 0.54 | 0.54 | 0.59 | 0.59 | -0.0026 |
| <i>Mother's birth cohort</i> | | | | | | | |
| 1960-1975 | 0.35 | 0.37 | 0.01 | 0.01 | 0.00 | 0.00 | -0.0061 |
| 1975-1990 | 0.50 | 0.48 | 0.95 | 0.95 | 0.37 | 0.39 | 0.0030 |
| 1990-2005 | 0.00 | 0.00 | 0.02 | 0.02 | 0.63 | 0.61 | 0.0059 |
| <i>Mother's age at birth</i> | | | | | | | |
| 12-15 | 0.04 | 0.03 | 0.03 | 0.03 | 0.02 | 0.02 | 0.0022 |
| 16-18 | 0.33 | 0.33 | 0.28 | 0.30 | 0.26 | 0.25 | -0.0087 |
| 19-24 | 0.56 | 0.56 | 0.59 | 0.59 | 0.59 | 0.60 | -0.0028 |
| 25-30 | 0.07 | 0.07 | 0.09 | 0.08 | 0.11 | 0.11 | 0.0025 |
| 31-49 | 0.01 | 0.01 | 0.01 | 0.01 | 0.02 | 0.02 | -0.0011 |
| <i>Household wealth</i> | | | | | | | |
| Second quintile | 0.19 | 0.20 | 0.20 | 0.19 | 0.20 | 0.20 | -0.0034 |
| Third quintile | 0.20 | 0.19 | 0.19 | 0.19 | 0.21 | 0.20 | 0.0061 |
| Fourth quintile | 0.20 | 0.18 | 0.18 | 0.19 | 0.20 | 0.20 | 0.0077* |
| Richest quintile | 0.20 | 0.20 | 0.18 | 0.18 | 0.14 | 0.15 | -0.0026 |

Notes: This table compares the socioeconomic characteristics of firstborn-boy (FB) and firstborn-girl (FG) families during the pre-legalization period and the two post-legalization periods in the DHS sample. The sample is restricted to first births, as only these are quasi-random. Column (7) shows the difference in sample means for the entire sample. $p < 0.1$ * $p < 0.05$ ** $p < 0.01$ ***

substantially more likely to continue childbearing than those with a firstborn son. The coefficient of *FirstbornGirl_i* is positive and significant, confirming that the women whose first child was a girl were 4 percentage points (p-value < 0.01) more likely to give birth in a given year. This behavior reflects the well-documented “fertility-stopping rule” in South Asia, where households pursue additional births until they achieve their desired number of sons (Gupta (1987); Bhat and Zavier (2003)). Consistent with this pattern, our estimates show that mothers with a firstborn daughter were more likely to give birth in any given year and had, on average, more children over their reproductive span compared to firstborn-boy mothers.

After the 2002 reform, these gaps narrowed dramatically. The difference in sibship size between firstborn-girl and firstborn-boy families declined by nearly three-quarters (p-value < 0.01), suggesting that families no longer needed to rely as heavily on continued childbearing to achieve desired composition. The effects are already noticeable in the early legalization period and become more pronounced in the late legalization period.

Our empirical strategy incrementally addresses several concerns. Column (1) includes only the baseline triple-difference terms. Column (2) adds child gender-specific cohort fixed effects to absorb nationwide shifts in gender gaps, such as declining son preference or improvements in maternal health that might differentially benefit boys. Column (3) adds firstborn-sex specific cohort fixed effects, ensuring that trends particular to firstborn-girl families (e.g., changes in stopping behavior) are not driving our results. In subsequent specifications, we incorporate district-year fixed effects, birth order controls, and years-since-last-birth controls, which together account for local shocks, biological differences across parity, and fertility spacing dynamics. Across all specifications, the coefficients of interest remain stable, increasing confidence that the results reflect the effect of abortion legalization rather than omitted trends.

The findings of Table 4 confirm the results of Table 3, and we see that the coefficient of *FirstbornGirl_j* is positive and significant. Women with a firstborn girl child had 0.257 (p-value < 0.01) more births than women with a firstborn son. Column (3) of the table shows that the pre-legalization gap in the number of births declined by 0.099 to 0.160, or by 3% to 7%, compared to the baseline mean of 2.28 births.

Because access to abortion can reduce unwanted births, we test whether legalization brought actual fertility closer to desired fertility. We construct excess fertility as actual minus desired fertil-

Table 3: Probability of Birth

| | (1) | (2) | (3) | (4) | (5) |
|--|------------------------|------------------------|------------------------|------------------------|------------------------|
| FirstbornGirl | 0.0440*** (0.0021) | 0.0463*** (0.0022) | | | |
| FirstbornGirl X Post 1 | -0.0234*** (0.0027) | -0.0275*** (0.0029) | -0.0274*** (0.0029) | -0.0272*** (0.0029) | -0.0271*** (0.0030) |
| FirstbornGirl X Post 2 | -0.0292*** (0.0027) | -0.0353*** (0.0028) | -0.0358*** (0.0029) | -0.0353*** (0.0029) | -0.0351*** (0.0030) |
| FirstbornGirl X Ideal fraction of sons | | | | | 0.0587*** (0.0080) |
| FirstbornGirl X Ideal number of children | | | | | -0.0036* (0.0019) |
| N | 566,868 | 566,867 | 566,867 | 560,993 | 560,993 |
| <i>Demographic Controls</i> | × | × | × | × | × |
| <i>Year FEs</i> | × | × | × | × | × |
| <i>District FEs</i> | × | × | × | × | × |
| <i>Parity FEs</i> | × | × | × | × | × |
| <i>District × year FEs</i> | × | × | × | × | × |
| <i>Years since last birth FEs</i> | | × | × | × | × |
| <i>FirstbornGirl × district FEs</i> | | | × | × | × |

Notes: Coefficients from specification 1 estimated using OLS regression on the mother-year sample from the year of their marriage to the year of interview. The dependent variable is an indicator for whether a mother gave birth in a given year. Post 1 indicates the early legalization period from 2003-2007 and Post 2 indicates the late legalization period from 2008-2018. The sample includes all mothers who have ever given birth. SEs in parentheses are clustered by district. Data: NDHS. $p < 0.1$ * $p < 0.05$ ** $p < 0.01$ ***

ity. Column (4) of Table 4 shows large reductions in excess fertility after legalization. At baseline, the first row indicates that firstborn-girl families had 0.124 more births than desired relative to firstborn-boy families. The post-legalization coefficients for excess fertility closely match those for actual fertility, indicating that the decline in births among firstborn-girl families reflects fewer unwanted births rather than a change in desired fertility.

Table 4 also shows that actual fertility increases with both the desired number of children and the desired son-to-daughter ratio. By contrast, excess fertility (actual minus desired) declines as the desired son share rises. This pattern implies that desired fertility rises more steeply with son preference than realized fertility. The interpretation is straightforward: parents who value sons plan for more births to reach their target composition through son-biased stopping, but conception, timing, and sex realization are not fully under their control, so actual fertility increases less than proportionately relative to stated desire, and excess fertility falls with the desired son share.

Table 4: Fertility

| | Number of Births | | | Excess Fertility | |
|--|------------------------|------------------------|------------------------|------------------------|------------------------|
| | (1) | (2) | (3) | (4) | (5) |
| FirstbornGirl | 0.2019*** (0.0188) | 0.1750*** (0.0192) | 0.2571*** (0.0780) | 0.1419*** (0.0216) | 0.1235*** (0.0214) |
| FirstbornGirl X Post 1 | -0.1811*** (0.0270) | -0.1518*** (0.0319) | -0.1600*** (0.0349) | -0.2017*** (0.0359) | -0.1053*** (0.0380) |
| FirstbornGirl X Post 2 | -0.0842*** (0.0241) | -0.0907*** (0.0264) | -0.0997*** (0.0279) | -0.1058*** (0.0299) | -0.0556* (0.0296) |
| Ideal number of children | | 0.1782*** (0.0189) | 0.1985*** (0.0313) | | |
| Ideal fraction of sons | | 0.0654*** (0.0246) | 0.0603* (0.0315) | | -0.5622*** (0.0216) |
| FirstbornGirl X Ideal fraction of sons | | | 0.0013 (0.0409) | | |
| FirstbornGirl X Ideal number of children | | | -0.0342 (0.0404) | | |
| N | 18,347 | 13,121 | 13,121 | 18,269 | 13,121 |
| Baseline mean | 2.28 | 2.28 | 2.28 | -.0252 | -.0252 |

Notes: Coefficients from specification 2 estimated using OLS regression. The dependent variable in columns (1) to (3) is the number of births at the time of interview, and the dependent variable in columns (4) to (5) is the excess fertility, which equals the number of births minus the self-reported ideal number of children. Sample includes all mothers who had both their first birth and last birth within the Pre, Post1, and Post2 periods. Post 1 indicates the early legalization period from 2003-2007 and Post 2 indicates the late legalization period from 2008-2018. SEs in parentheses are clustered by district. Data: NDHS. $p < 0.1$ * $p < 0.05$ ** $p < 0.01$ ***

Birth Spacing

If it is the availability of sex-selective abortion rather than something else that is driving our findings, then we should see an increase in birth spacing in the post-legalization period among families with firstborn girls relative to other families. We check for this in Table 5. The estimates show that before legalization, firstborn-girl mothers had shorter birth intervals between their first and second children, about seven months sooner on average, relative to firstborn-boy mothers. After legalization, this gap narrowed significantly, suggesting that families relied less on manipulating birth spacing once abortion became available as a fertility control tool. This adjustment is consistent with the dynamic margin highlighted in the framework in Appendix A, where investments such as breastfeeding reduce the probability of conception, and legalization reduces the incentive to manipulate such investments to achieve desired composition.

Table 5: Birth spacing: Months between first and second births

| | (1) No. of months |
|-----------------------|-----------------------|
| Firstborn girl | -7.265*** (0.2880) |
| Firstborn girl * Post | 1.933** (0.7827) |
| N | 11,200 |

Notes: OLS regression. The sample includes mothers who had at least 2 births at the time of the interview. The dependent variable is the number of months between births. Post refers to the post-legalization period as a whole (2003-2018). Standard errors in parentheses are clustered by district. Data: NDHS. $p < 0.1$ * $p < 0.05$ ** $p < 0.01$ ***

6.2 Effects on Child Investments

We present the estimates of the impact of abortion legalization on parental health investments in Tables 6 to 8, presented in specification (3).

6.2.1 Breastfeeding

Turning to parental investments, we find strong evidence that abortion legalization narrowed gender gaps in breastfeeding duration. Table 6 reports the estimated impact of legalization on breastfeeding behavior. Additional control variables are sequentially incorporated across columns from left to right. The coefficient on *FirstbornGirl* \times *Female* for breastfeeding in the table confirms

that before the reform, daughters in firstborn-girl families were breastfed for 1.7–2.0 months less (p-value < 0.01) than their brothers, reflecting a behavioral pattern documented elsewhere in South Asia (Jayachandran and Kuziemko (2011)). This is consistent with the notion that breastfeeding doubles as a form of birth spacing: families eager to conceive again may shorten breastfeeding for daughters to accelerate the next pregnancy, particularly if they have not yet had a son.

The triple interaction coefficients, $FirstbornGirl_j \times Female_i \times Post_t^1$ and $FirstbornGirl_j \times Female_i \times Post_t^2$ indicate that this disadvantage declined significantly after legalization. In the late post-reform period, daughters in firstborn-girl families breastfed about two months longer (p < 0.01), nearly closing the baseline gap. These findings suggest that with abortion available as an alternative means of fertility control, parents no longer need to manipulate breastfeeding to pursue son-biased fertility goals.

Table A3 decomposes the results by wealth, rural residence, and maternal education. The gap in breastfeeding duration between firstborn-girl and firstborn-boy families is concentrated among poorer, rural, and less-educated households. This pattern is consistent with breastfeeding functioning as a low-cost spacing strategy where access to modern contraception is limited and reliance on traditional methods is higher.

To mitigate any censoring bias, we restrict the sample to the last two surviving births and to children older than two. We also report decomposed estimates by firstborn-sex families to confirm that the observed narrowing is not driven by compositional differences. Table 7 presents these decomposed estimates. Among firstborn-girl families, daughters were breastfed 2.21 months less before legalization (p < 0.001) but gained 2.20 months afterward (p < 0.001), effectively eliminating the gap. In firstborn-boy families, the baseline difference between sons and daughters was small (−0.19 months, p = 0.67) and did not change significantly after legalization (−1.12 to +0.29 months, p = 0.26–0.78). The main effects in Table 6 are therefore driven by improvements in firstborn-girl families.

6.2.2 Vaccination and Mortality

For other dimensions of child health, the effects are less pronounced. Table 8 shows that before legalization, vaccination rates were only slightly higher for boys in firstborn-girl families, and

Table 6: Breastfeeding as a Function of Abortion Legalization and the Sex of the Firstborn

| | (1) | (2) | (3) | (4) | (5) |
|---|------------------------|------------------------|-----------------------|------------------------|-----------------------|
| Firstborn girl * Female | -1.9927*** (0.6728) | -1.9507*** (0.6616) | -1.7811** (0.6758) | -1.8940*** (0.6559) | -1.7360** (0.6710) |
| Firstborn girl * Female * Post1 | 1.7713* (0.9853) | 1.8381* (0.9636) | 1.8185* (0.9578) | 1.7643* (0.9562) | 1.7534* (0.9505) |
| Firstborn girl * Female * Post2 | 1.9835*** (0.6736) | 1.9423*** (0.6680) | 1.7800** (0.6741) | 1.8093*** (0.6702) | 1.6616** (0.6785) |
| N | 5,856 | 5,851 | 5,809 | 5,851 | 5,809 |
| Baseline mean | 19.3 | 19.3 | 19.3 | 19.3 | 19.3 |
| X_{ijt} | × | × | × | × | × |
| <i>District FEs</i> | × | × | × | × | × |
| <i>Year FEs</i> | × | × | × | × | × |
| <i>Age FEs</i> | × | × | × | × | × |
| <i>FirstbornGirl</i> × <i>birthyearFEs</i> | | | × | | × |
| <i>Female</i> × <i>birthorderFEs</i> | | | × | | × |
| <i>Birthorder</i> × <i>birthyearFEs</i> | | | × | | × |
| <i>Birthorder</i> × <i>regionFEs</i> | | | × | | × |
| <i>FirstbornGirl</i> × <i>regionFEs</i> | | | × | | × |
| <i>FirstbornGirl</i> × <i>birthorderFEs</i> | | | × | | × |
| <i>Additional X_{ijt}</i> | | | | × | × |

Notes: This table reports breastfeeding effects (in months) for children of second- and higher-order birth order. Results are based on the last two surviving births of a mother, and we restrict the sample to children above age two. Coefficients are from the specification 3 estimated using OLS regression. Post 1 indicates the early legalization period from 2003-2007, and Post 2 indicates the late legalization period from 2008-2018. SEs in parentheses are clustered by district. Data: NDHS. $p < 0.1$ * $p < 0.05$ ** $p < 0.01$ ***

Table 7: Breastfeeding by Firstborn Sex

| | Firstborn girl family | | | Firstborn boy family | | |
|----------------|------------------------|------------------------|------------------------|----------------------|---------------------|---------------------|
| | (1) | (2) | (3) | (4) | (5) | (6) |
| Female | -2.2052*** (0.5157) | -2.2006*** (0.5100) | -2.1790*** (0.5093) | -0.1877 (0.4505) | -0.2179 (0.4492) | -0.2525 (0.4478) |
| Female * Post1 | 0.7787 (0.7358) | 0.7839 (0.7189) | 0.7784 (0.7223) | -1.0048 (0.6669) | -1.1174 (0.6832) | -1.0769 (0.6836) |
| Female * Post2 | 2.1996*** (0.5185) | 2.1989*** (0.5166) | 2.1225*** (0.5225) | 0.2444 (0.4419) | 0.2869 (0.4366) | 0.3485 (0.4430) |
| N | 3,079 | 3,077 | 3,077 | 2,772 | 2,767 | 2,767 |
| Baseline mean | 18.8 | 18.8 | 18.8 | 19.3 | 19.3 | 19.3 |

Notes: This table reports breastfeeding effects (in months) for children of second- and higher-order birth order. Results are based on the last two surviving births of a mother, and we restrict the sample to children above age two. Each column is a separate OLS regression. Post 1 indicates the early legalization period from 2003-2007 and Post 2 indicates the late legalization period from 2008-2018. SEs in parentheses are clustered by district. Data: NDHS. $p < 0.1$ * $p < 0.05$ ** $p < 0.01$ ***

this modest difference did not narrow significantly after the reform. The small initial gap helps explain the absence of large post-reform changes: unlike breastfeeding, where girls faced a clear disadvantage, vaccination practices were already relatively equitable across family types.

This stability likely reflects Nepal's long-standing national immunization efforts. The National Immunization Program (NIP), launched in 1979 as part of the World Health Organization's Expanded Program on Immunization (EPI), sought to reduce childhood morbidity and mortality from vaccine-preventable diseases (Ministry of Health, 1979). Because all EPI vaccines are provided free of charge, coverage among children under 12 months increased sharply, with the largest gains among poorer and less-educated households. These improvements substantially narrowed equity gaps in vaccination over time (Ashish et al., 2017). This is further evident from the heterogeneity analysis in Table A5. We see that most of the gap in vaccination rates between firstborn girl versus firstborn boy families are concentrated among families in the top 40% of the wealth quintiles.

Table 8: Vaccination as a Function of Abortion Legalization and the Sex of the Firstborn

| | (1) | (2) | (3) | (4) | (5) |
|--|----------------------|----------------------|----------------------|----------------------|----------------------|
| Firstborn girl * Female | -0.1936* (0.1083) | -0.1948* (0.1078) | -0.2134* (0.1083) | -0.1882* (0.1054) | -0.2052* (0.1063) |
| Firstborn girl * Female * Post1 | 0.1725 (0.2159) | 0.1702 (0.2135) | 0.1447 (0.1962) | 0.1955 (0.2084) | 0.1681 (0.1925) |
| Firstborn girl * Female * Post2 | 0.2061 (0.1832) | 0.2057 (0.1800) | 0.1891 (0.1767) | 0.2175 (0.1808) | 0.1990 (0.1764) |
| N | 13,404 | 13,403 | 13,372 | 13,403 | 13,372 |
| Baseline mean | 5.84 | 5.84 | 5.84 | 5.84 | 5.84 |
| X_{ijt} | × | × | × | × | × |
| <i>District FEs</i> | × | × | × | × | × |
| <i>Year FEs</i> | × | × | × | × | × |
| <i>Age FEs</i> | × | × | × | × | × |
| <i>FirstbornGirl</i> × <i>birthyear</i> FEs | | | × | | × |
| <i>Female</i> × <i>birthorder</i> FEs | | | × | | × |
| <i>Birthorder</i> × <i>birthyear</i> FEs | | | × | | × |
| <i>Birthorder</i> × <i>region</i> FEs | | | × | | × |
| <i>FirstbornGirl</i> × <i>region</i> FEs | | | × | | × |
| <i>FirstbornGirl</i> × <i>birthorder</i> FEs | | | × | | × |
| <i>Additional</i> X_{ijt} | | | | × | × |

Notes: This table reports vaccination effects for children of second- and higher-order birth order. Coefficients are from the specification 3 estimated using OLS regression. Post 1 indicates the early legalization period from 2003-2007 and Post 2 indicates the late legalization period from 2008-2018. SEs in parentheses are clustered by district. Data: NDHS. $p < 0.1$ * $p < 0.05$ ** $p < 0.01$ ***

Turning to survival outcomes, Table 9 shows that before legalization, girls with a firstborn

sister were significantly more likely to die before age five than those with a firstborn brother. This excess female mortality gap narrowed in the early post-legalization period, although the improvement did not persist in the long run. These short-lived effects mirror findings from Taiwan ([Lin et al., 2014](#)) and parts of India ([Hu and Schlosser, 2015](#)), where access to abortion or prenatal sex detection improved relative female survival temporarily but not consistently across all contexts. In Nepal, legalization may have mitigated the most severe forms of postnatal discrimination immediately after the reform, yet entrenched cultural norms appear to have constrained longer-term progress.

Child mortality is shaped by a range of structural factors beyond fertility behavior, including health system expansion, sanitation improvements, and broader declines in under-five mortality. To isolate the effect of abortion legalization, we include child gender-specific cohort fixed effects, which absorb national trends in survival that differ by gender, such as improvements in prenatal or neonatal care that disproportionately benefit boys. District-year fixed effects further control for localized health initiatives and infrastructure expansion. These adjustments ensure that the observed short-run reduction in excess female mortality among firstborn-girl families reflects the impact of abortion legalization rather than concurrent nationwide improvements in survival.

Taken together, these results suggest that while legalization eased fertility pressures and improved breastfeeding outcomes, it did not generate broad-based equalization across all child health inputs. This pattern aligns with the model's prediction that investment responses are strongest for inputs directly linked to fertility behavior, whereas those less tied to fertility decisions exhibit weaker or no effects.

6.3 Robustness Checks

The baseline results establish that abortion legalization reduced son-biased fertility stopping and narrowed certain gender gaps in parental investments, particularly breastfeeding. In this section, we present a series of robustness checks to address potential concerns about exposure, measurement, and alternative mechanisms.

Table 9: Excess Female Under-Five Mortality

| | (1) | (2) | (3) | (4) | (5) |
|---|-----------------------|-----------------------|----------------------|-----------------------|----------------------|
| Firstborn girl * Female | 0.0190** (0.0090) | 0.0181* (0.0092) | 0.0183** (0.0088) | 0.0177* (0.0091) | 0.0179** (0.0087) |
| Firstborn girl * Female * Post1 | -0.0424** (0.0188) | -0.0412** (0.0189) | -0.0378* (0.0190) | -0.0405** (0.0189) | -0.0370* (0.0190) |
| Firstborn girl * Female * Post2 | 0.0028 (0.0157) | 0.0043 (0.0158) | 0.0014 (0.0159) | 0.0053 (0.0157) | 0.0024 (0.0157) |
| N | 37,089 | 37,088 | 37,060 | 37,087 | 37,059 |
| Baseline mean | .102 | .102 | .102 | .102 | .102 |
| X_{ijt} | × | × | × | × | × |
| <i>District FEs</i> | × | × | × | × | × |
| <i>Year FEs</i> | × | × | × | × | × |
| <i>Age FEs</i> | × | × | × | × | × |
| <i>FirstbornGirl</i> × <i>birthyearFEs</i> | | | × | | × |
| <i>Female</i> × <i>birthorderFEs</i> | | | × | | × |
| <i>Birthorder</i> × <i>birthyearFEs</i> | | | × | | × |
| <i>Birthorder</i> × <i>regionFEs</i> | | | × | | × |
| <i>FirstbornGirl</i> × <i>regionFEs</i> | | | × | | × |
| <i>FirstbornGirl</i> × <i>birthorderFEs</i> | | | × | | × |
| <i>Additional</i> X_{ijt} | | | | × | × |

Notes: Sample of second- and higher-order births. Each column is a separate OLS regression. The outcome is an indicator of death before age five. We drop children who are less than five years old to allow each child in the sample full exposure to the risk of under-five mortality. Post 1 indicates the early legalization period from 2003-2007 and Post 2 indicates the late legalization period from 2008-2018. SEs in parentheses are clustered by district. Data: NDHS. $p < 0.1$ * $p < 0.05$ ** $p < 0.01$ ***

Consistent exposure to legalization

A first concern is that women whose reproductive spans straddled different policy regimes may introduce bias if part of their fertility occurred before legalization and part after. To address this, we restrict the sample to mothers whose entire fertility occurred within a single period, either pre-legalization, early post-legalization (2003–2007), or late post-legalization (2008–2018). As shown in Table 10, the results remain qualitatively similar: firstborn-girl mothers had higher fertility than firstborn-boy mothers in the pre-legalization period, but this gap narrowed substantially after legalization. This strengthens confidence that the main results are not driven by inconsistent exposure across cohorts.

Table 10: Probability of Birth (Fertility Sample)

| | (1) | (2) | (3) | (4) | (5) |
|--|------------------------|------------------------|------------------------|------------------------|------------------------|
| FirstbornGirl | 0.0302*** (0.0022) | 0.0334*** (0.0023) | | | |
| FirstbornGirl X Post 1 | -0.0282*** (0.0051) | -0.0404*** (0.0053) | -0.0389*** (0.0054) | -0.0389*** (0.0054) | -0.0343*** (0.0054) |
| FirstbornGirl X Post 2 | -0.0064 (0.0043) | -0.0135*** (0.0046) | -0.0152*** (0.0045) | -0.0152*** (0.0045) | -0.0113** (0.0045) |
| FirstbornGirl X Ideal fraction of sons | | | | | 0.0849*** (0.0140) |
| FirstbornGirl X Ideal number of children | | | | | 0.0115*** (0.0027) |
| N | 214,070 | 214,070 | 214,070 | 214,070 | 211,970 |
| X_i | × | × | × | × | × |
| Year FEs | × | × | × | × | × |
| District FEs | × | × | × | × | × |
| Age FEs | × | × | × | × | × |
| Parity FEs | × | × | × | × | × |
| District × year FEs | × | × | × | × | × |
| Years since last birth FEs | × | × | × | × | × |
| FirstbornGirl × district FEs | | | × | × | × |

Notes: Coefficients from specification 1 estimated using OLS regression on the mother-year sample from the year of their marriage to the year of interview. The sample is restricted to mothers whose entire fertility occurred within a single period, either pre-legalization, early post-legalization (2003–2007), or late post-legalization (2008–2018). The dependent variable is an indicator for whether a mother gave birth in a given year. The sample includes all mothers who have ever given birth. SEs in parentheses are clustered by district. Data: NDHS. $p < 0.1$ * $p < 0.05$ ** $p < 0.01$ ***

Effects of ultrasound on fertility

Second, we address concerns about concurrent changes in access to prenatal sex detection in Table 11. Using a proxy for the diffusion of prenatal sex detection, we replicate Anukriti et al. (2022)

on the probability of an additional birth: for firstborn-girl families, the probability of a subsequent birth declines by about 1.1–2.0 percentage points in the early post-ultrasound period and by about 3.2–4.2 percentage points in the later period, closely matching prior magnitudes. In our main legalization estimates, we define 1990–2002 as the pre-period and treat ultrasound diffusion as part of the baseline environment, so the law’s effect is identified net of ultrasound.

Table 11: Probability of Birth (Effects of Ultrasound)

| | (1) | (2) | (3) | (4) | (5) |
|--|------------------------|------------------------|------------------------|------------------------|------------------------|
| FirstbornGirl | 0.0498*** (0.0050) | 0.0589*** (0.0055) | | | |
| FirstbornGirl X Post 1 | -0.0109** (0.0049) | -0.0204*** (0.0055) | -0.0203*** (0.0056) | -0.0197*** (0.0055) | -0.0196*** (0.0054) |
| FirstbornGirl X Post 2 | -0.0324*** (0.0053) | -0.0428*** (0.0059) | -0.0427*** (0.0060) | -0.0422*** (0.0058) | -0.0419*** (0.0057) |
| FirstbornGirl X Ideal fraction of sons | | | | | 0.0567*** (0.0120) |
| FirstbornGirl X Ideal number of children | | | | | -0.0015 (0.0022) |
| N | 323,020 | 323,018 | 323,018 | 319,136 | 319,136 |

Notes: Coefficients from specification 1 estimated using OLS regression on the mother-year sample from the year of their marriage to the year of interview. The dependent variable is an indicator for whether a mother gave birth in a given year. Post 1 indicates the early ultrasound diffusion period from 1990-1994 and Post 2 indicates the late ultrasound diffusion period from 1995-2002. The sample includes all mothers who have ever given birth. SEs in parentheses are clustered by district. Data: NDHS. $p < 0.1$ * $p < 0.05$ ** $p < 0.01$ ***

Alternative definitions of sex composition

Finally, we examine whether the fertility effects are robust to alternative definitions of household gender composition. We use the alternative definitions of gender composition from [Alfano \(2017\)](#) in Table 12, which presents specifications using the gender of the most recent child, the ratio of girls to boys in the family, and a dummy for whether daughters outnumber sons. Across all three measures, families with more daughters exhibited higher fertility in the pre-legalization period, consistent with son-biased stopping behavior. After legalization, these gaps diminished significantly. This robustness exercise shows that the results do not depend solely on defining incentives by the sex of the firstborn but rather reflect a broader pattern of son preference in fertility behavior.

Taken together, these robustness checks reinforce the central conclusion: abortion legalization reduced son-biased fertility behavior, both by lowering overall fertility gaps and by altering the timing and spacing of births. The consistency of the findings across different definitions of exposure

and composition further strengthens the interpretation that the reform shifted family formation dynamics in a way consistent with the quantity–quality trade-off model, where constraints on fertility are relaxed and the marginal cost of reducing family size declines.

Table 12: Alternative definitions of children’s gender composition

| Dependent variable: Birth indicator | (1) | (2) | (3) | (4) | (5) | (6) |
|-------------------------------------|--------------------------|---------------------|------------------------|----------------------|----------------------|----------------------|
| | Measurement of G_i | | Ratio of girls to boys | | More girls than boys | |
| | Youngest child is female | | | | | |
| G_i | 0.025*** (0.002) | 0.028*** (0.002) | 0.050*** (0.001) | 0.070*** (0.002) | 0.050*** (0.002) | 0.0559*** (0.003) |
| $G_i * Post$ | | -0.007* (0.003) | | -0.037*** (0.002) | | -0.012*** (0.004) |
| District specific trends | yes | yes | yes | yes | yes | yes |
| Observations | 190,481 | 190,481 | 131,588 | 131,588 | 206,881 | 206,881 |

Notes: Youngest child is female is a dummy variable taking the value 1 if the youngest child born before year t is a girl; Ratio of girls to boys is the total number of girls born by year t divided by the total number of boys; More girls than boys is a dummy variable taking the value 1 if in year t the total number of girls exceeds the total number of boys. SEs in parentheses are clustered by district. Data: NDHS.
 $p < 0.1^* p < 0.05^{**} p < 0.01^{***}$

7 Conclusion

This paper has examined the effects of abortion legalization in Nepal on fertility and gender gaps in child investments, situating the analysis within the broader framework of the quantity–quality trade-off. Consistent with [Becker and Lewis \(1973\)](#), the results show that expanding women’s reproductive autonomy allowed families to reduce fertility, particularly among those with a first-born daughter, who previously faced stronger incentives to continue childbearing. The narrowing of sibship size gaps between firstborn-girl and firstborn-boy families by as much as three-quarters provides direct evidence that abortion reform mitigated son-biased fertility stopping. By lowering the effective “price” of limiting fertility while maintaining desired composition, legalization shifted households closer to their fertility ideals without the same costs for daughters.

On the quality dimension, abortion legalization narrowed gender gaps in breastfeeding, a domain of investment strongly linked to both fertility timing and survival ([Jayachandran and Kuziemko, 2011](#)). Daughters in firstborn-girl families, who were previously disadvantaged, experienced substantial gains in breastfeeding duration after legalization. However, effects on other investments like vaccinations were muted, and mortality improvements were modest and short-lived. Taken together, these patterns suggest that abortion legalization’s main effects operated through the fertility channel, with more selective improvements in child investments. These findings extend the Becker–Lewis model to a setting with strong son preference, highlighting how re-

productive autonomy reshapes both fertility behavior and within-household investment patterns, consistent with the mechanisms described by ([Jayachandran and Kuziemko, 2011](#)).

From a policy perspective, the results indicate that legal reforms expanding reproductive rights can promote gender equity indirectly by reshaping fertility behavior, but their effects on child investments depend on broader social norms. Reproductive autonomy is a necessary but not sufficient condition for closing gender gaps. Complementary policies that strengthen maternal and child health systems, expand immunization coverage, and challenge discriminatory norms are essential for ensuring that gains in autonomy translate into lasting improvements in girls' well-being.

These results also carry implications for global health goals. Breastfeeding—identified by the World Health Organization and UNICEF as central to achieving the Sustainable Development Goals—responds not only to health policy but also to underlying reproductive constraints. By easing those constraints, abortion legalization fostered a more equitable environment for early-life investments, particularly for daughters. Yet achieving the SDG targets on child survival and nutrition ultimately requires aligning legal reforms with social change, so that gains in autonomy lead to sustained improvements in health and gender equality.

Finally, the findings speak to ongoing demographic transitions in South Asia and other settings where declining fertility coincides with persistent son preference. As access to reproductive technologies expands, the interaction between autonomy and cultural norms will remain central to shaping gender equity. Future research should explore how these reforms influence later outcomes such as education, labor market participation, and women's bargaining power, as well as their intergenerational consequences. Nepal's experience demonstrates both the transformative potential of legal access to abortion and the limits imposed by enduring social norms.

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A Appendix A: Conceptual Framework

We present a general model of fertility and parental investments. The model is agnostic to context and can be specialized to different settings by adding environment-specific assumptions. We then discuss how the Nepal setting maps into this structure.

General Model: Fertility–Investment Trade-off

Preferences and technology. Parents choose the number of children $n \in \mathbb{R}_+$ and per-child investment $e \in \mathbb{R}_+$ to maximize

$$U = U(C, n, q), \quad q = f(e),$$

where C is consumption, q is average child “quality,” $U_C, U_n, U_q > 0$, and U is strictly concave. The quality production function satisfies $f'(e) > 0$, $f''(e) < 0$.

Budget and time costs. Each child involves a resource cost that depends on e . Let the total cost of rearing one child be $k(e)$ with $k'(e) > 0$, $k''(e) \geq 0$. Income is Y . The static resource constraint is

$$C + nk(e) \leq Y.$$

A time cost formulation is equivalent if time is priced at the wage; the key object is the *shadow price* of a child, $k(e)$, which rises in e .

Optimality conditions. Let λ be the multiplier on the budget. The first-order conditions (FOCs) are

$$\frac{\partial U / \partial n}{\lambda} = k(e), \tag{A1}$$

$$\frac{\partial U / \partial q}{\lambda} f'(e) = nk'(e), \tag{A2}$$

together with the budget constraint. Equation (A1) equates the marginal utility of an additional child to its shadow price. Equation (A2) equates the marginal benefit of raising per-child investment to the marginal resource cost across n children. The pair (n, e) solves a standard quantity–quality trade-off.

Comparative statics (general). Let \tilde{k} denote a shock to the shadow price of fertility, holding e fixed. If \tilde{k} rises (children become more costly), the optimal n falls and, through (A2), e tends to rise since fewer children spread the marginal cost $nk'(e)$ over a smaller base. If an investment subsidy lowers $k'(e)$, then for given n the optimal e rises; the effect on n is in general ambiguous and depends on cross-substitution in U .

Dynamic Extension

The static framework can be extended to a dynamic setting in which fertility unfolds across periods $t = 1, \dots, T$. In each period, parents decide whether to attempt another birth and how much to invest in existing children.

Let π_t denote the probability of conception in period t if parents try for a child. Some investments, such as time spent breastfeeding, reduce the time or biological readiness for conception. This can be represented by $\pi_t = \pi(e_t)$ with $\pi'(e_t) < 0$, meaning that higher investment reduces the probability of conceiving in the next period.

The household's problem is to maximize the present value of utility

$$\sum_{t=1}^T \beta^t U(C_t, n_t, q_t),$$

subject to the budget constraint

$$C_t + n_t k(e_t) \leq Y_t,$$

and the law of motion for n_t governed by π_t .

The dynamic structure reproduces the static trade-off each period but also adds an intertemporal margin: parents weigh the benefits of investing more in current children against the fertility cost of delaying or reducing the chance of having additional children. This makes the model especially suited to settings where some investments double as fertility-spacing tools.

Specialization to Contexts

The general model becomes empirically informative once context-specific primitives are introduced. Two common specializations in the literature are compositional preferences and fertility-control technologies.

Compositional preferences. Suppose parents value the probability of having at least one son. Let $s \in [0, 1]$ denote this probability and augment preferences as $U(C, n, q, s)$ with $U_s > 0$. In a static representation, s is increasing in n if sex is i.i.d. at birth, which raises the marginal utility of additional births when no son is present. In the dynamic extension, U_s creates a state variable indicating whether a son has arrived.

Fertility-control technology. Let $\phi \geq 0$ be the effective cost parameter for abortion or contraception. A reduction in ϕ lowers the shadow cost of achieving desired family size and composition. In the dynamic extension, a lower ϕ raises the option value of a fertility attempt and weakens the complementarity between low e_t and trying for a birth when composition targets are unmet.

Implications. With $U_s > 0$ and high ϕ , families may select higher n when a son is absent and may tilt e toward lower levels for a time-intensive subset of investments. When ϕ falls, parents can

achieve composition targets with fewer births, which reduces excess fertility and relaxes the need to adjust e for timing reasons. The sign and magnitude of the investment response are ambiguous in general and depend on which investments compete most with fertility effort and on preference curvature.

Mapping to Nepal

To specialize the model to Nepal, take the following context-driven assumptions that are consistent with descriptive facts:

1. Strong son preference. Model via $U(C, n, q, s)$ with $U_s > 0$ and a state that records whether a son has arrived.
2. Legal reform. Abortion legalization is a reduction in ϕ in 2002. Ultrasound diffusion affects the precision with which composition can be targeted, which further lowers the effective cost of achieving s .
3. Investment margins. Some investments are time or spacing sensitive (for example, breast-feeding duration) and therefore enter $\pi(e_t)$ with $\pi'(e_t) < 0$. Other investments (for example, vaccination take-up) are less tied to timing and mainly shift $k'(e)$ without affecting $\pi(e_t)$.

Under these assumptions the framework yields two broad empirical implications. First, a decline in ϕ weakens the link between composition and fertility behavior, which reduces excess fertility among families with a firstborn daughter. Second, investment patterns may change because fertility constraints are relaxed, but the direction and magnitude are theoretically ambiguous ex ante and depend on which investments compete most with fertility effort. The empirical analysis examines these margins directly.

B Appendix B: Policy Details and Identification Tests

Timeline

Table A1: Timeline of Abortion Services Rollout

| Year | Reform |
|--------------|--|
| 1988 | Ultrasound services introduced. |
| 2002 | Abortion legalized under specific conditions. |
| 2004 | First-trimester surgical abortion services launched. |
| 2007 | Second-trimester abortion services introduced. |
| 2009 | Medical abortion introduced, rural area expansion. |
| 2009 | Supreme Court ruled abortion as a human right. |
| 2018 | Free abortion service launched in all govt. hospitals. |
| 2020-Present | Ongoing efforts to expand services and reduce stigma. |

The rollout of abortion and related reproductive health services in Nepal occurred gradually over several decades. Ultrasound services were first introduced in 1988 ([Subedi and Sharma, 2013](#); [Mukhiya and Mishra, 2025](#)), enabling prenatal sex determination well before abortion became legal. In 2002, abortion was legalized under specific conditions, marking a critical policy shift. Service provision expanded in stages: first-trimester surgical abortion became available in 2004, followed by the introduction of second-trimester procedures in 2007. Medical abortion was rolled out in 2009, alongside efforts to extend access to rural areas. That same year, the Supreme Court of Nepal ruled abortion to be a constitutional right, strengthening the legal foundation of the policy. Subsequent reforms further broadened access, including the 2018 launch of free abortion services in all government hospitals. Since 2020, ongoing initiatives have aimed to expand coverage and reduce social stigma, consolidating the reform into a nationwide reproductive health program.

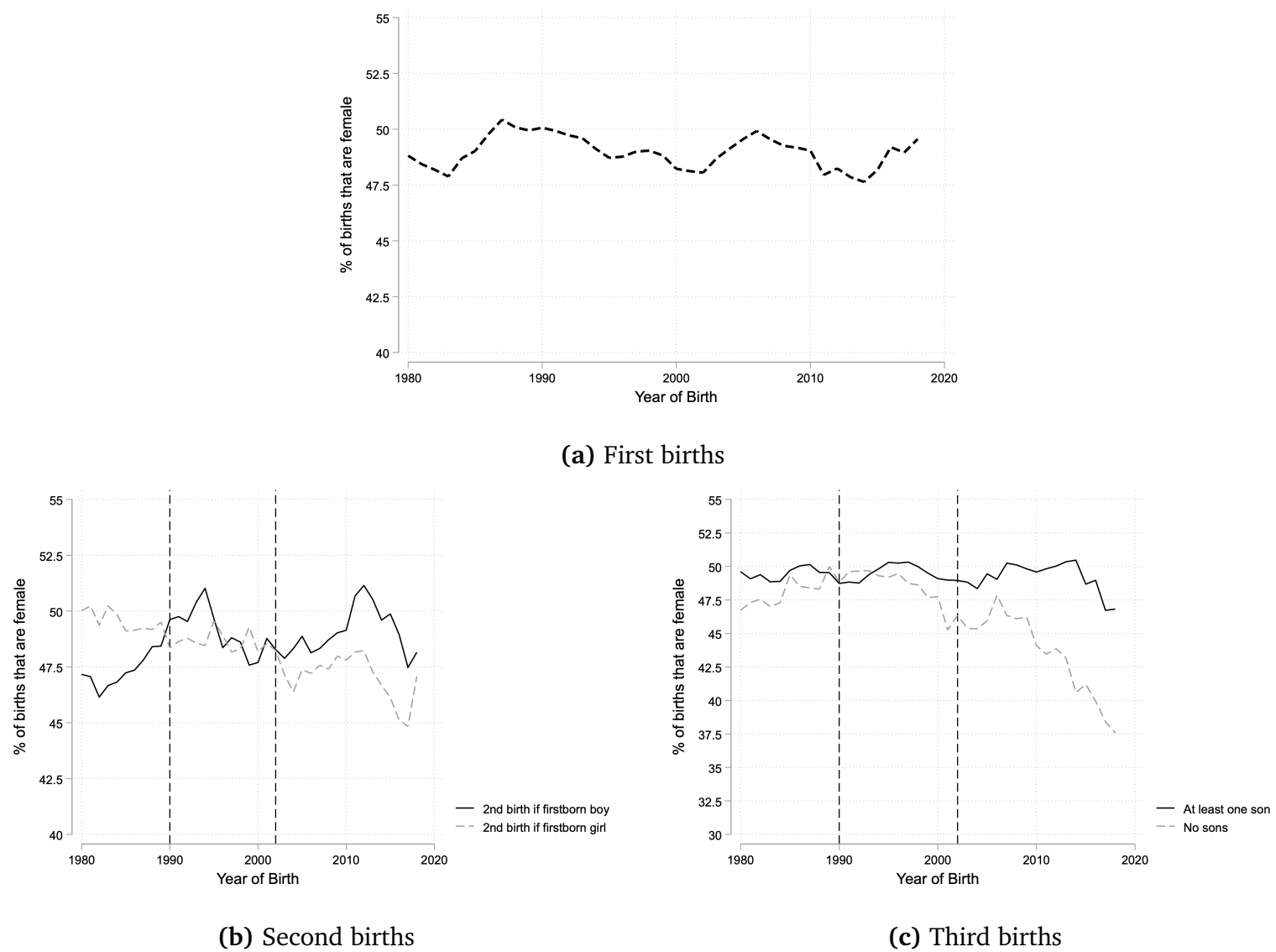
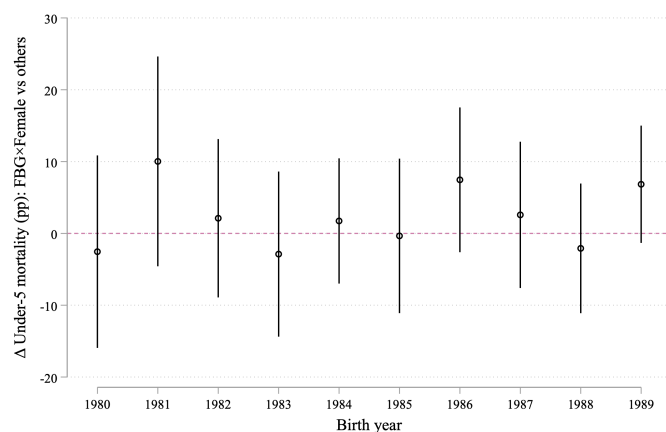


Figure A1: Trends in Proportion of Females at Birth by Birth Order and Sex Composition of Older Siblings

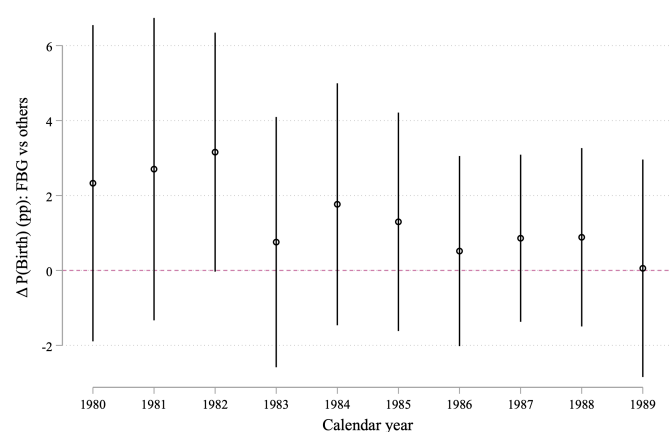
Notes: Panel (a) shows the evolution of the percentage of females among first births over time. In panels (b) and (c) the trend in the percentage of births that are female is respectively plotted for second and third births separately for families that have at least one son and families with no sons at the time of the respective birth. In all cases, the y axis shows the five-year moving average of percentage of births that are female. The figures show that, despite ultrasound availability and abortion legalization, the sex ratio of first births has remained normal. It also shows that the sex ratio at birth in families without sons starts diverging from the sex ratio in families with sons after the introduction of ultrasound and there is a clear divergence after the legalization of abortion.

Other Identification Tests

We test that the sex of the firstborn child is effectively random in panel (a) of Figure A1 and that the sex-selective abortion at second and higher-order births occurs primarily in families whose firstborn is a daughter in Figure A1. The figures show that, despite ultrasound availability and abortion legalization, the sex ratio of first births has remained normal. Therefore, it supports our assumption that the sex of the firstborn child is quasi-random. Panels (b) and (c) show that the sex ratio at birth in families without sons starts diverging from the sex ratio in families with sons after the introduction of ultrasound and there is a clear divergence after the legalization of abortion. This allows us to assume that the firstborn-sex margin captures households' incentives to continue childbearing to obtain a son.



(a) Under-five Mortality



(b) Probability of Birth

Figure A2: Test of Differential Pre-Trends in Outcomes by the Sex of the Firstborn

Parallel Trends Tests

To assess the validity of the identifying assumption, we test for pre-trends using data from 1980 to 1990, a period that predates the introduction of ultrasound technology. For fertility, we regress the probability of birth on interactions between firstborn sex and year dummies. The joint test yields an F-statistic of 1.02 with a p-value of 0.43, providing no evidence of differential pre-trends across firstborn-girl and firstborn-boy families. For under-five mortality, the analogous test produces an F-statistic of 1.20 with a p-value of 0.30, again consistent with parallel trends. For breastfeeding duration, data are only available beginning with the 1996 survey. Using these years, the joint test yields an F-statistic of 2.10 with a p-value of 0.064, suggesting marginal evidence of divergence across family types even before abortion legalization. Taken together, these results support the parallel trends assumption for fertility and mortality, while indicating greater caution in interpreting breastfeeding estimates.

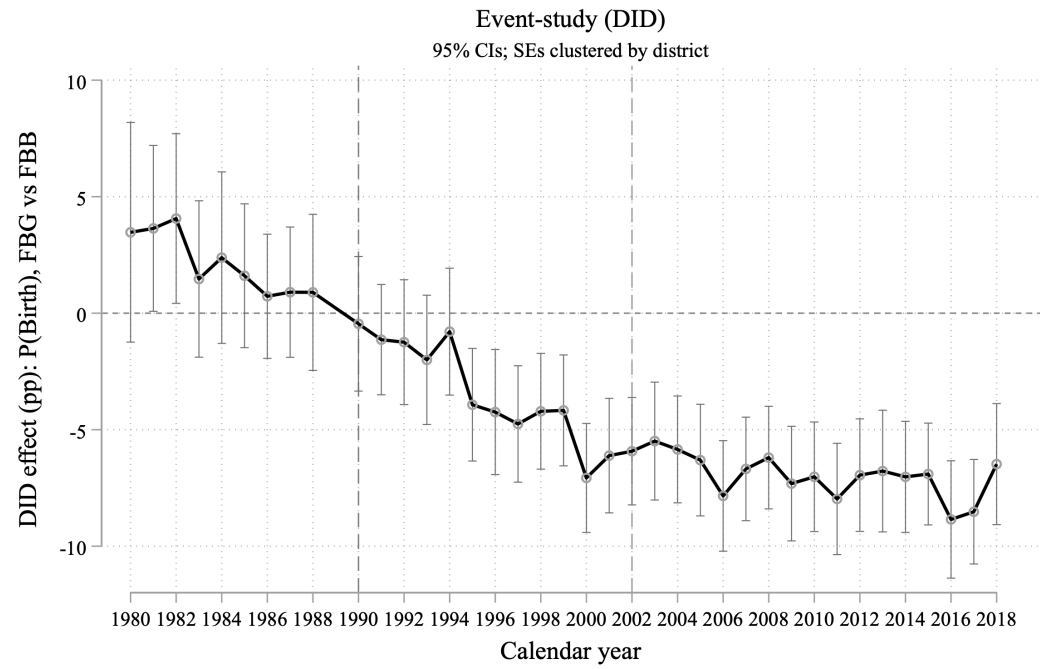


Figure A3: Probability of Birth

C Appendix C: Heterogeneity Tests

Table A2: Probability of Birth: Heterogeneity by Socioeconomic Status

| | Mother's education | | Wealth | | Rurality | |
|-------------------------|------------------------|------------------------|------------------------|------------------------|------------------------|------------------------|
| | (1) Illiterate | (2) Literate | (3) Bottom 40% | (4) Top 40% | (5) Rural | (6) Urban |
| Firstborn girl * Female | 0.0323*** (0.0023) | 0.0613*** (0.0032) | 0.0505*** (0.0028) | 0.0408*** (0.0036) | 0.0365*** (0.0026) | 0.0590*** (0.0032) |
| Firstborn girl * Post1 | -0.0145*** (0.0029) | -0.0267*** (0.0040) | -0.0316*** (0.0037) | -0.0188*** (0.0047) | -0.0181*** (0.0036) | -0.0314*** (0.0036) |
| Firstborn girl * Post2 | -0.0111*** (0.0031) | -0.0421*** (0.0035) | -0.0371*** (0.0032) | -0.0241*** (0.0046) | -0.0224*** (0.0041) | -0.0404*** (0.0034) |
| N | 347,792 | 219,061 | 212,659 | 243,834 | 340,478 | 226,377 |

Notes: This table reports the probability of birth for children of second- and higher-order birth order for mothers with different socioeconomic status. Coefficients are from the specification 3. Each column within a panel is a separate OLS regression. Post 1 indicates the early legalization period from 2003-2007, and Post 2 indicates the late legalization period from 2008-2018. SEs in parentheses are clustered by district. Data: NDHS.
 $p < 0.1$ * $p < 0.05$ ** $p < 0.01$ ***

Table A3: Breastfeeding: Heterogeneity by Socioeconomic Status

| | Mother's education | | Wealth | | Rurality | |
|---------------------------------|------------------------|---------------------|-----------------------|---------------------|-----------------------|---------------------|
| | (1) Illiterate | (2) Literate | (3) Bottom 40% | (4) Top 40% | (5) Rural | (6) Urban |
| Firstborn girl * Female | -2.0889*** (0.7709) | -0.2412 (1.5559) | -2.4216** (1.1857) | -1.2244 (0.8494) | -1.6887** (0.7279) | 0.4566 (1.4490) |
| Firstborn girl * Female * Post1 | 2.6492** (1.1633) | -0.8513 (1.7588) | 1.7131 (1.7832) | 1.5379 (1.1356) | 1.5255 (1.0265) | -0.6409 (2.3371) |
| Firstborn girl * Female * Post2 | 2.0511** (0.8383) | 0.4293 (1.5634) | 2.4880** (1.1988) | 1.4420* (0.8618) | 1.6455** (0.7570) | -0.8010 (1.5082) |
| N | 3,848 | 1,905 | 1,642 | 3,052 | 4,436 | 1,313 |
| Baseline mean | 18.6 | 19.1 | 18.8 | 18.9 | 18.6 | 20.2 |

Notes: The dependent variable is breastfeeding (in months) for children of second- and higher-order birth order. Results are based on the last two surviving births of a mother, and we restrict the sample to children above age two. Coefficients are from the specification 3. Each column within a panel is a separate OLS regression. Post 1 indicates the early legalization period from 2003-2007, and Post 2 indicates the late legalization period from 2008-2018. SEs in parentheses are clustered by district. Baseline mean is the mean breastfeeding duration during the pre-legalization period by mothers with the specific socioeconomic status. Data: NDHS. $p < 0.1$ * $p < 0.05$ ** $p < 0.01$ ***

Table A4: Under-5 Mortality: Heterogeneity by Socioeconomic Status

| | Mother's education | | Wealth | | Rurality | |
|---------------------------------|---------------------|-----------------------|---------------------|---------------------|-----------------------|---------------------|
| | (1) Illiterate | (2) Literate | (3) Bottom 40% | (4) Top 40% | (5) Rural | (6) Urban |
| Firstborn girl * Female | 0.0158 (0.0101) | 0.0271 (0.0173) | 0.0076 (0.0140) | 0.0256* (0.0136) | 0.0224** (0.0101) | -0.0075 (0.0205) |
| Firstborn girl * Female * Post1 | -0.0301 (0.0225) | -0.0648** (0.0264) | -0.0334 (0.0335) | -0.0381 (0.0251) | -0.0569** (0.0244) | 0.0072 (0.0252) |
| Firstborn girl * Female * Post2 | -0.0085 (0.0219) | 0.0093 (0.0205) | 0.0390* (0.0221) | -0.0062 (0.0220) | -0.0108 (0.0201) | 0.0361 (0.0234) |
| N | 26,479 | 10,536 | 10,322 | 19,819 | 25,221 | 11,794 |
| Baseline mean | .115 | .0581 | .0695 | .123 | .108 | .0802 |

Notes: The dependent variable is an indicator of death before age five. Sample of second- and higher-order births. Each column is a separate OLS regression. We drop children who are less than five years old to allow each child in the sample full exposure to the risk of under-five mortality. Post 1 indicates the early legalization period from 2003-2007 and Post 2 indicates the late legalization period from 2008-2018. SEs in parentheses are clustered by district. . Baseline mean refers to under-five mortality for children born in the pre-ultrasound period to mothers with the specific socioeconomic status. Data: NDHS. $p < 0.1$ * $p < 0.05$ ** $p < 0.01$ ***

Table A5: No. of Vaccinations: Heterogeneity by Socioeconomic Status

| | Mother's education | | Wealth | | Rurality | |
|---------------------------------|-----------------------|---------------------|---------------------|------------------------|----------------------|---------------------|
| | (1) Illiterate | (2) Literate | (3) Bottom 40% | (4) Top 40% | (5) Rural | (6) Urban |
| Firstborn girl * Female | -0.2632** (0.1274) | 0.2006 (0.1722) | 0.1588 (0.1688) | -0.4652*** (0.1575) | -0.2075* (0.1105) | 0.0846 (0.2238) |
| Firstborn girl * Female * Post1 | 0.2328 (0.2414) | -0.1910 (0.2552) | 0.0150 (0.3007) | 0.3398 (0.2796) | 0.1870 (0.2101) | -0.0523 (0.4584) |
| Firstborn girl * Female * Post2 | 0.2107 (0.2600) | -0.0950 (0.2184) | -0.1285 (0.2623) | 0.3969 (0.2437) | 0.0957 (0.2042) | 0.2294 (0.3104) |
| N | 8,812 | 4,513 | 3,845 | 7,067 | 10,682 | 2,642 |
| Baseline mean | 5.67 | 6.86 | 6.54 | 5.53 | 5.88 | 6.86 |

Notes: This table reports vaccination effects for children of second- and higher-order birth order. Coefficients are from the specification 3. Each column within a panel is a separate OLS regression. Post 1 indicates the early legalization period from 2003-2007, and Post 2 indicates the late legalization period from 2008-2018. SEs in parentheses are clustered by district. Baseline mean is the mean number of vaccinations received during the pre-legalization period by the children of mothers with the specific socioeconomic status. Data: NDHS. $p < 0.1$ * $p < 0.05$ ** $p < 0.01$ ***

D Appendix D: Other Mortality Analyses

Table A6: Excess Female Under-One Mortality

| | (1) | (2) | (3) | (4) | (5) |
|---------------------------------|-----------------------|-----------------------|----------------------|-----------------------|----------------------|
| Firstborn girl * Female | 0.0109 (0.0071) | 0.0103 (0.0072) | 0.0097 (0.0071) | 0.0099 (0.0071) | 0.0094 (0.0070) |
| Firstborn girl * Female * Post1 | -0.0255** (0.0117) | -0.0246** (0.0117) | -0.0219* (0.0115) | -0.0245** (0.0116) | -0.0217* (0.0115) |
| Firstborn girl * Female * Post2 | 0.0030 (0.0101) | 0.0035 (0.0102) | 0.0029 (0.0099) | 0.0041 (0.0102) | 0.0034 (0.0099) |
| N | 50,354 | 50,353 | 50,329 | 50,352 | 50,328 |
| Baseline mean | .0772 | .0772 | .0772 | .0772 | .0772 |

Notes: The dependent variable is an indicator of death before age one. Sample of second- and higher-order births. Each column is a separate OLS regression. We drop children who are less than one year old to allow each child in the sample full exposure to the risk of under-five mortality. Post 1 indicates the early legalization period from 2003-2007, and Post 2 indicates the late legalization period from 2008-2018. SEs in parentheses are clustered by district. Data: NDHS. $p < 0.1$ * $p < 0.05$ ** $p < 0.01$ ***

Table A7: Excess Female Under-Two Mortality

| | (1) | (2) | (3) | (4) | (5) |
|---------------------------------|-----------------------|-----------------------|----------------------|-----------------------|----------------------|
| Firstborn girl * Female | 0.0123 (0.0075) | 0.0118 (0.0076) | 0.0115 (0.0073) | 0.0114 (0.0076) | 0.0111 (0.0073) |
| Firstborn girl * Female * Post1 | -0.0288** (0.0132) | -0.0282** (0.0132) | -0.0248* (0.0129) | -0.0279** (0.0131) | -0.0244* (0.0128) |
| Firstborn girl * Female * Post2 | 0.0020 (0.0117) | 0.0026 (0.0119) | 0.0010 (0.0114) | 0.0037 (0.0118) | 0.0021 (0.0114) |
| N | 47,418 | 47,417 | 47,396 | 47,416 | 47,395 |
| Baseline mean | .0904 | .0904 | .0904 | .0904 | .0904 |

Notes: The dependent variable is an indicator of death before age two. Sample of second- and higher-order births. Each column is a separate OLS regression. We drop children who are less than two years old to allow each child in the sample full exposure to the risk of under-five mortality. Post 1 indicates the early legalization period from 2003-2007, and Post 2 indicates the late legalization period from 2008-2018. SEs in parentheses are clustered by district. Data: NDHS. $p < 0.1$ * $p < 0.05$ ** $p < 0.01$ ***

Table A8: Excess Female Under-Three Mortality

| | (1) | (2) | (3) | (4) | (5) |
|---------------------------------|-----------------------|-----------------------|-----------------------|-----------------------|-----------------------|
| Firstborn girl * Female | 0.0129* (0.0077) | 0.0124 (0.0078) | 0.0124 (0.0076) | 0.0120 (0.0078) | 0.0120 (0.0075) |
| Firstborn girl * Female * Post1 | -0.0329** (0.0148) | -0.0321** (0.0148) | -0.0294** (0.0145) | -0.0316** (0.0147) | -0.0289** (0.0144) |
| Firstborn girl * Female * Post2 | 0.0039 (0.0127) | 0.0048 (0.0128) | 0.0027 (0.0122) | 0.0056 (0.0127) | 0.0035 (0.0120) |
| N | 44,263 | 44,262 | 44,241 | 44,261 | 44,240 |
| Baseline mean | .0955 | .0955 | .0955 | .0955 | .0955 |

Notes: The dependent variable is an indicator of death before age three. Sample of second- and higher-order births. Each column is a separate OLS regression. We drop children who are less than three years old to allow each child in the sample full exposure to the risk of under-five mortality. Post 1 indicates the early legalization period from 2003-2007, and Post 2 indicates the late legalization period from 2008-2018. SEs in parentheses are clustered by district. Data: NDHS. $p < 0.1$ * $p < 0.05$ ** $p < 0.01$ ***

Table A9: Excess Female Under-Four Mortality

| | (1) | (2) | (3) | (4) | (5) |
|---------------------------------|-----------------------|-----------------------|----------------------|-----------------------|----------------------|
| Firstborn girl * Female | 0.0172** (0.0082) | 0.0165* (0.0083) | 0.0165** (0.0081) | 0.0162* (0.0082) | 0.0161** (0.0080) |
| Firstborn girl * Female * Post1 | -0.0375** (0.0168) | -0.0364** (0.0169) | -0.0337* (0.0170) | -0.0360** (0.0169) | -0.0332* (0.0169) |
| Firstborn girl * Female * Post2 | 0.0041 (0.0142) | 0.0054 (0.0143) | 0.0024 (0.0138) | 0.0062 (0.0142) | 0.0032 (0.0137) |
| N | 40,701 | 40,698 | 40,677 | 40,697 | 40,676 |
| Baseline mean | .0991 | .0991 | .0991 | .0991 | .0991 |

Notes: The dependent variable is an indicator of death before age four. Sample of second- and higher-order births. Each column is a separate OLS regression. We drop children who are less than four years old to allow each child in the sample full exposure to the risk of under-five mortality. Post 1 indicates the early legalization period from 2003-2007, and Post 2 indicates the late legalization period from 2008-2018. SEs in parentheses are clustered by district. Data: NDHS. $p < 0.1$ * $p < 0.05$ ** $p < 0.01$ ***