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

Jijeebisha Bhattarai Ajinkya Keskar

October 2, 2025

Abstract

We examine the impact of the 2002 abortion legalization reform in Nepal on fertility decisions and child investments. Nepal's abortion laws pre-2002 were one of the world's most stringent, criminalizing pregnancy terminations and prosecuting and imprisoning women and their family members. We exploit the quasi-random variation in the sex of the firstborn child to estimate how the reform shaped family size, breastfeeding, vaccination, and child survival. Legalization substantially reduced son-biased fertility stopping, narrowing sibship-size disparities between firstborn-girl and firstborn-boy families by nearly three-quarters. It also closed much of the pre-existing gender gap in breastfeeding duration, consistent with reduced reliance on breastfeeding as a fertility-spacing tool. By contrast, effects on vaccinations and survival were limited, reflecting that these inputs were less closely tied to fertility behavior. These findings show that abortion access relaxed fertility constraints and altered investments linked to fertility timing, while broader forms of gender inequality persisted. The results extend the literature on the quantity–quality trade-off by showing how reproductive autonomy interacts with cultural preferences to shape fertility and child well-being.

Keywords: Abortion, Fertility, Breastfeeding, Nepal

JEL codes: J13, J16, I18, I14

*Authors: Bhattarai: Binghamton University (SUNY). Email: jbhatta1@binghamton.edu. Keskar: Binghamton University (SUNY). E-mail: akeskar@binghamton.edu.

1 Introduction

Fertility choices and investments in children are closely linked. Parents allocate limited resources between the number of children they raise and the amount devoted to each child's health, education, and well-being. Classic models of the family formalize this "quantity–quality trade-off," predicting that fertility decline frees resources for greater per-child investment (Becker and Lewis, 1973; Becker and Tomes, 1976). Evidence from several settings supports this view: exogenous reductions in family size raise schooling, improve health, and enhance other measures of child welfare (Rosenzweig and Wolpin, 1980; Angrist et al., 2010; Li et al., 2008).

In settings with strong son preference, however, fertility behavior follows a different logic. Parents often continue childbearing until they reach their desired number of sons, a practice known as son-biased fertility stopping (Gupta, 1987; Bhat and Zavier, 2003). This strategy disadvantages daughters in two ways: they are more likely to grow up in larger sibships, diluting household resources, and they often receive fewer direct investments in survival and health (Jayachandran and Pande, 2017). Where contraceptive options are limited, the combination of declining desired family size and persistent son preference amplifies these gendered inequities.

Policies governing reproductive autonomy shape the interaction between fertility decisions, family composition, and investments in children. Where abortion and contraception are restricted, parents have fewer tools to realize their preferred family composition, often resorting to discriminatory postnatal practices. Where access is expanded, parents may achieve their fertility and composition goals with fewer births, potentially reducing the disadvantages faced by daughters.[1] Against this backdrop, Nepal provides a compelling setting to study these dynamics. The country had one of the world's most restrictive abortion laws: abortion was fully prohibited until 2002, and women faced imprisonment for seeking terminations. Unsafe procedures were a leading cause of maternal mortality (Tamang, 1996; Thapa et al., 2014). In 2002, the National Safe Abortion Policy legalized abortion under specified conditions, followed by a rapid rollout of services through public–private partnerships (Samandari et al., 2012; Henderson et al., 2013). Legalization represented a dramatic expansion of reproductive autonomy in a country with deep-rooted son preference.

The reform coincided with two other sweeping changes: a sharp fertility decline (from 4.6 births per woman in 1996 to 2.6 in 2011; MoHP 2012) and the diffusion of ultrasound technology,

[1]Evidence from high-income countries shows that access to contraception and abortion expanded women's autonomy, influencing fertility timing, marriage, and labor-market trajectories Goldin and Katz (2002); Bailey (2006).

which enabled relatively cheap prenatal sex detection (Guilmoto, 2009; Frost et al., 2013). These overlapping shifts meant that parents could now reconcile a smaller desired family size with son preference by turning to abortion, including sex-selective abortion. Nepal thus provides a natural experiment for examining how abortion legalization interacts with cultural preferences and the quantity–quality trade-off in a son-preferring society.

Our analysis reveals three main findings. First, abortion legalization sharply reduced son-biased fertility stopping: families with a firstborn daughter had much larger sibship sizes before the reform, but this gap narrowed by nearly three-quarters afterward. Second, legalization narrowed gender disparities in breastfeeding duration, as daughters in firstborn-girl families gained significantly. This pattern is consistent with parents relying less on breastfeeding as a fertility-spacing tool. Third, legalization did little to change other health investments. Gender gaps in vaccination were already small before the reform, and excess female mortality declined only in the short run without lasting improvements.

These findings advance three strands of literature. First, they provide the first causal evidence on abortion legalization's consequences for both fertility and child investments in Nepal, using quasi-random variation in firstborn sex as a proxy for parental incentives to engage in sex selection. Second, they extend the literature on the quantity–quality trade-off by showing that its operation is mediated by cultural preferences: legalization reduced fertility disparities and narrowed breastfeeding gaps, but left vaccination and survival gaps largely intact. Third, they contribute to research on reproductive rights in low- and middle-income countries by offering new evidence from a setting with both rapid fertility decline and persistent son preference.

Finally, this paper situates Nepal's reform within the broader literature on abortion legalization, which documents diverse impacts on fertility, sex selection, and child outcomes across countries. In Taiwan, abortion legalization alongside sex detection raised sex ratios at birth but reduced female mortality at higher parities (Lin et al., 2014). In India, access to sex selection improved some postnatal investments in girls by reducing the number of "unwanted" daughters (Anukriti et al., 2016; Hu and Schlosser, 2015), though in other contexts it exacerbated female disadvantage (Almond et al., 2010). Evidence from Mexico (Clarke and Mühlrad, 2016) and Ethiopia (Holcombe and Kidanemariam Gebru, 2022) suggests that abortion legalization reduced fertility and improved maternal and early child outcomes, though causal evidence on direct postnatal investments remains limited. By analyzing Nepal, we add new evidence from a low-income, son-preferring setting where these dynamics are particularly stark.

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; 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

[2] For 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)).

[3] Prior 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.

(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 Zavier (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. (2016) 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].

[4] Anukriti et al. (2016) 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 A2), magnitudes that are strikingly similar. This replication reinforces that ultrasound availability reduced son-biased fertility continuation in both India and Nepal.

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) (MoHP (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 (NHSPP, 2013). Because ultrasound diffusion may itself affect fertility behavior and child investments, we explicitly test its impact in Table A2. In our main analysis, 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 (MoHP (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 (MoHP (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 ex-

[5]Despite 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.

pansion of services. In partnership with non-governmental organizations, the government launched a nationwide program to train abortion providers and ensure the safety and accessibility of care (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. Research 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)).

4 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

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

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

$$C + n\,k(e) \leq Y.$$

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

$$\frac{\partial U/\partial n}{\lambda} = k(e), \tag{1}$$

$$\frac{\partial U/\partial q}{\lambda} f'(e) = n\,k'(e), \tag{2}$$

together with the budget constraint. Equation (1) equates the marginal utility of an additional child to its shadow price. Equation (2) 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 (2), e tends to rise since fewer children spread the marginal cost $n\,k'(e)$ 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 $t = 1, \ldots, T$. 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.

5 Data and Measurement

5.1 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. Breastfeeding 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.

5.2 Measurement

Abortion was legalized nationwide in 2002, but service availability expanded gradually.[9] 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.

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

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

[7] In the appendix, 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.

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

[9] Table A1 provides further details.

variable *Birth* which observes each woman from the year of her marriage until the survey year and assign a value of one if the mother gave birth in a certain year and zero if not. The variable *FirstbornGirl* is used directly from the survey data and takes the value of one if the mother's first child is a girl. We interact this variable with indicators for post-legalization to create our main outcome of interest, $FirstbornGirl \times Post$. Our second fertility analysis uses the variable *N*, which uses the number of children that the mother had during the time of the survey.

For the quality analyses, we use the mother's reported answers for the months of breastfeeding and the number of vaccinations that a child has received during the time of the survey. Breastfeeding analyses focus on the last two surviving births of each mother and are limited to children at least two years old. Finally, the under-five mortality is defined as the percentage of births that die before the age of five. We interact the variable *FirstbornGirl* from before with indicators for post-legalization and a dummy variable for whether the child is a female to create our main outcome of interest, $FirstbornGirl \times Post \times Female$.

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.

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
$\mathbb{1}$ [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).

6 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 Figure A2, and (b) sex-selective abortion at second and higher-order births occurs primarily in families whose firstborn is a daughter (Figure A4). The central identifying assumption, tested directly in our analysis in Figure A1, is that, absent access to ultrasound technology, trends in outcomes would have evolved similarly for families with firstborn sons and firstborn daughters.

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 firstborn 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.

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

$$\begin{aligned} Birth_{it} = {} & \alpha + \beta_1 FirstbornGirl_i \times Post^1_t + \beta_2 FirstbornGirl_i \times Post^2_t + \\ & \gamma FirstbornGirl_i + \omega_t + \boldsymbol{X}'_i \tau + \phi_a + \psi_b + \sigma_r + \delta_d + \nu_d FirstbornGirl_i + \theta_{dt} + \epsilon_{it} \quad (3) \end{aligned}$$

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. $Post^1_t$ indicates that year t falls within the early years after the abortion legalization (2003-2007), while $Post^2_t$ corresponds to the later years after the abortion ban was lifted (2008-2018). The vector $\boldsymbol{X}_i$ 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

$$
\begin{aligned}
N_{jt} = \alpha + \beta_1 FirstbornGirl_j \times Post_t^1 + \beta_2 FirstbornGirl_j \times Post_t^2 + \gamma FirstbornGirl_j + \sigma Post_t^1 \\
+ \psi Post_t^2 + \boldsymbol{X}_j' \tau + \delta_d + \nu_d FirstbornGirl_j + \theta_d Post_t^1 + \omega_d Post_t^2 + \epsilon jt. \quad (4)
\end{aligned}
$$

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 $\boldsymbol{X}_j$ 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 quality outcomes for children: months of breastfeeding, number of vaccinations, and under-five mortality. For all three, we estimate the following equation

$$
\begin{aligned}
Y_i = \alpha + \beta_1 FirstbornGirl_j \times Female_i \times Post_t^1 + \beta_2 FirstbornGirl_j \times Female_i \times Post_t^2 \\
+ \gamma FirstbornGirl_j \times Female_i + \omega_t FirstbornGirl_j + \sigma_t Female_i + \boldsymbol{X}_{ijt}' \tau + \delta_r Female_i \\
+ \nu_r FirstbornGirl_j + \psi_b Female_i + \xi_b FirstbornGirl_j + \rho_{bt} + \eta_{bd} + \phi_{dt} + \epsilon_i \quad (5)
\end{aligned}
$$

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.

7 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.

7.1 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 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

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
Contraception Use	0.46	0.52	0.47	0.52	0.51	0.57	-0.0565***
Mother's education							
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
Father's education							
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
Mother's birth cohort							
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
Mother's age at birth							
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
Household wealth							
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. $p < 0.1^{*} p < 0.05^{**} p < 0.01^{***}$

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.

Table 3: Probability of Birth

	(1)	(2)	(3)	(4)	(5)
FirstbornGirl	0.0440***	0.0463***			
	(0.0021)	(0.0022)			
FirstbornGirl X Post 1	-0.0234***	-0.0275***	-0.0274***	-0.0272***	-0.0271***
	(0.0027)	(0.0029)	(0.0029)	(0.0029)	(0.0030)
FirstbornGirl X Post 2	-0.0292***	-0.0353***	-0.0358***	-0.0353***	-0.0351***
	(0.0027)	(0.0028)	(0.0029)	(0.0029)	(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
X_i	×	×	×	×	×
$Year\ FEs$	×	×	×	×	×
$District\ FEs$	×	×	×	×	×
$Age\ FEs$	×	×	×	×	×
$Parity\ FEs$	×	×	×	×	×
$District \times year\ FEs$	×	×	×	×	×
$Years\ since\ last\ birth\ FEs$	×	×	×	×	×
$FirstbornGirl \times district\ FEs$			×	×	×

Notes: Coefficients from specification 3 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. $p < 0.1^{*} p < 0.05^{**} p < 0.01^{***}$

The findings of Table 4 confirm the results of Table 3, and we see that the coefficient of $FirstbornGirl_j$ is positive and significant. This indicates that the 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-ultrasound gap in the number of births declined by 0.099 to 0.160, or by 40% to 60%.

These findings resonate with prior work on the fertility consequences of abortion reforms in other contexts mentioned in Section 2. Our results extend this literature by showing that in a setting with strong son preference, legalization specifically attenuated son-biased stopping behavior, aligning with the predictions of the quantity–quality trade-off model (Becker and Lewis (1973); Becker and Tomes (1976)). By enabling families to achieve smaller, desired families, legalization reduced the costs borne disproportionately by daughters in larger sibships.

Table 4: Fertility

	Number of Births			Excess Fertility	
	(1)	(2)	(3)	(4)	(5)
FirstbornGirl	0.2019***	0.1750***	0.2571***	0.1235***	0.0596
	(0.0188)	(0.0192)	(0.0780)	(0.0214)	(0.0577)
FirstbornGirl X Post 1	-0.1811***	-0.1518***	-0.1600***	-0.1053***	-0.0975**
	(0.0270)	(0.0319)	(0.0349)	(0.0380)	(0.0402)
FirstbornGirl X Post 2	-0.0842***	-0.0907***	-0.0997***	-0.0556*	-0.0466
	(0.0241)	(0.0264)	(0.0279)	(0.0296)	(0.0320)
Ideal number of children		0.1782***	0.1985***		
		(0.0189)	(0.0313)		
Ideal fraction of sons		0.0654***	0.0603*	-0.5622***	-0.5854***
		(0.0246)	(0.0315)	(0.0216)	(0.0275)
FirstbornGirl X Ideal fraction of sons			0.0013		0.0490
			(0.0409)		(0.0386)
FirstbornGirl X Ideal number of children			-0.0342		
			(0.0404)		
N	18,347	13,121	13,121	13,121	13,121
Baseline mean	2.28	2.28	2.28	-.0252	-.0252

Notes: Coefficients from specification 4 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

A second check considers whether legalization influenced not only the number of children but also the timing of births. Table 5 shows that before legalization, firstborn-girl mothers had shorter birth intervals between their first and second children, about three months sooner on average, relative to firstborn-boy mothers. This is consistent with literature suggesting that early weaning of daughters was used as a strategy to accelerate the next pregnancy in the pursuit of a son (Jayachandran and Kuziemko (2011)). 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, 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	-2.995***
	(0.2055)
Firstborn girl * Post	1.243*
	(0.5999)
N	27,891

Notes: OLS regression. The sample includes mothers who had at least 2 births at the time of the interview. The dependent variable is the average 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. $p < 0.1^{*} p < 0.05^{**} p < 0.01^{***}$

7.2 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 (5).

7.2.1 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 $FirstbornGirl_j \times Female_i \times Post_t^2$ show that this difference was significantly reduced in the post-legalization periods. In the late post-reform period, daughters in firstborn-girl families experienced significant increase of 2 months (p-value < 0.01) in breastfeeding duration, nearly offsetting the entire baseline deficit. These results suggest that with abortion available as an alternative fertility-control mechanism, families no longer needed to manipulate breastfeeding practices to pursue son-biased fertility goals.

Breastfeeding is both a health investment and a fertility-spacing mechanism, which raises concerns about confounding. Families eager to conceive again may shorten breastfeeding, particularly for daughters, leading to gender gaps that could reflect fertility preferences rather than direct discrimination. Censoring is another concern, as younger children may not yet have completed breastfeeding at the time of the survey. We address these issues by restricting the breastfeeding sample to the last two surviving births and to children older than age two, minimizing censoring bias. In addition, we show decomposed estimates by firstborn-sex families, ensuring that the observed narrowing of gaps is not driven by baseline differences across households. These design choices allow us to interpret post-reform improvements in daughters' breastfeeding duration as reflecting reduced reliance on breastfeeding for fertility spacing, rather than mechanical age effects or compositional differences.

To document the underlying variation and highlight behavioral differences between families with firstborn boys and girls, Table 7 decomposes the triple-difference estimates. It reports a double-difference version of specification 5 separately for firstborn-girl and firstborn-boy families. Among firstborn-girl families, the baseline female disadvantage was –2.21 months ($p < 0.001$), but daughters gained 2.20 months post-legalization ($p < 0.001$), effectively eliminating the gap. In contrast, in firstborn-boy families, the baseline difference between sons and daughters was –0.19 months ($p = 0.67$), and the coefficients on the post-reform interactions were small and statistically insignificant (–1.12 to +0.29 months, $p = 0.26$–0.78). So much of the effect that we see in Table

6 is coming from an increase in breastfeeding in firstborn girl families.

This finding complements evidence from India, where access to sex selection has been shown to improve postnatal investments in surviving daughters by reducing their "unwantedness" (Anukriti et al. (2016); Hu and Schlosser (2015)). At the same time, it contrasts with studies showing that access to sex determination technology sometimes worsened female survival in contexts where postnatal discrimination persisted (Almond et al. (2010)). The Nepal case thus highlights how legalization can ease constraints on fertility while also shifting intrahousehold allocation toward more equitable treatment, at least in domains closely tied to fertility timing.

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***	-1.9507***	-1.7811**	-1.8940***	-1.7360**
	(0.6728)	(0.6616)	(0.6758)	(0.6559)	(0.6710)
Firstborn girl * Female * Post1	1.7713*	1.8381*	1.8185*	1.7643*	1.7534*
	(0.9853)	(0.9636)	(0.9578)	(0.9562)	(0.9505)
Firstborn girl * Female * Post2	1.9835***	1.9423***	1.7800**	1.8093***	1.6616**
	(0.6736)	(0.6680)	(0.6741)	(0.6702)	(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}	×	×	×	×	×
$District\,FEs$	×	×	×	×	×
$Year\,FEs$	×	×	×	×	×
$Age\,FEs$	×	×	×	×	×
$FirstbornGirl \times birthyearFEs$			×		×
$Female \times birthorderFEs$			×		×
$Birthorder \times birthyearFEs$			×		×
$Birthorder \times regionFEs$			×		×
$FirstbornGirl \times regionFEs$			×		×
$FirstbornGirl \times birthorderFEs$			×		×
$Additional\,X_{ijt}$				×	×

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 5 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. $p < 0.1^{*} p < 0.05^{**} p < 0.01^{***}$

7.2.2 Vaccination and Mortality

For other child health investments, the evidence is more muted. Table 8 shows that vaccination rates were only slightly higher for boys in firstborn-girl families before legalization, and this modest gap does not narrow significantly after the reform. The small baseline difference helps explain the absence of large post-reform changes: unlike breastfeeding, where girls faced a clear disadvantage,

Table 7: Breastfeeding by Firstborn Sex

	Firstborn girl family			Firstborn boy family		
	(1)	(2)	(3)	(4)	(5)	(6)
Female	-2.2052***	-2.2006***	-2.1790***	-0.1877	-0.2179	-0.2525
	(0.5157)	(0.5100)	(0.5093)	(0.4505)	(0.4492)	(0.4478)
Female * Post1	0.7787	0.7839	0.7784	-1.0048	-1.1174	-1.0769
	(0.7358)	(0.7189)	(0.7223)	(0.6669)	(0.6832)	(0.6836)
Female * Post2	2.1996***	2.1989***	2.1225***	0.2444	0.2869	0.3485
	(0.5185)	(0.5166)	(0.5225)	(0.4419)	(0.4366)	(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. $p < 0.1^{*} p < 0.05^{**} p < 0.01^{***}$

vaccination practices were already more equal across family types. These null results suggest that while legalization reduced fertility pressures and improved breastfeeding practices, it did not translate into broad-based equalization across all health inputs. This pattern is in line with the framework's prediction that investment responses are theoretically ambiguous and strongest for inputs directly tied to fertility behavior, whereas less fertility-sensitive investments show muted effects.

In line with the preceding results, Table 9 indicates that, during the pre-legalization period, the probability of dying before age five was significantly higher for girls with a firstborn sister compared to those with a firstborn brother. This gap narrowed significantly in the early-legalization period, though this effect does not persist in the long run. This echoes 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 in some periods but not consistently across all settings. In Nepal, legalization may have reduced the most extreme cases of postnatal discrimination in the immediate aftermath, but the cultural norms appear to have limited sustained gains.

Child mortality is shaped by multiple factors beyond fertility behavior, including health system expansion, improvements in sanitation, and overall declines in child mortality. A concern is that observed changes in excess female mortality could simply reflect these broader secular trends. To address this, we incorporate child gender-specific cohort fixed effects, which absorb nationwide shifts in mortality that differ by gender, such as improvements in prenatal or neonatal care that

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*	-0.1948*	-0.2134*	-0.1882*	-0.2052*
	(0.1083)	(0.1078)	(0.1083)	(0.1054)	(0.1063)
Firstborn girl * Female * Post1	0.1725	0.1702	0.1447	0.1955	0.1681
	(0.2159)	(0.2135)	(0.1962)	(0.2084)	(0.1925)
Firstborn girl * Female * Post2	0.2061	0.2057	0.1891	0.2175	0.1990
	(0.1832)	(0.1800)	(0.1767)	(0.1808)	(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}	×	×	×	×	×
$District\,FEs$	×	×	×	×	×
$Year\,FEs$	×	×	×	×	×
$Age\,FEs$	×	×	×	×	×
$FirstbornGirl \times birthyearFEs$			×		×
$Female \times birthorderFEs$			×		×
$Birthorder \times birthyearFEs$			×		×
$Birthorder \times regionFEs$			×		×
$FirstbornGirl \times regionFEs$			×		×
$FirstbornGirl \times birthorderFEs$			×		×
$Additional\,X_{ijt}$				×	×

Notes: This table reports vaccination effects for children of second- and higher-order birth order. Coefficients are from the specification 5 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. $p < 0.1^{*} p < 0.05^{**} p < 0.01^{***}$

disproportionately benefit boys. District-year fixed effects further account for localized health interventions and infrastructure improvements. By controlling for these trends, our design ensures that the observed short-run reduction in excess female mortality among firstborn-girl families is attributable to abortion legalization, rather than concurrent nationwide improvements in survival.

Table 9: Excess Female Under-Five Mortality

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

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. $p < 0.1^{*} p < 0.05^{**} p < 0.01^{***}$

7.3 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.

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***	0.0334***			
	(0.0022)	(0.0023)			
FirstbornGirl X Post 1	-0.0282***	-0.0404***	-0.0389***	-0.0389***	-0.0343***
	(0.0051)	(0.0053)	(0.0054)	(0.0054)	(0.0054)
FirstbornGirl X Post 2	-0.0064	-0.0135***	-0.0152***	-0.0152***	-0.0113**
	(0.0043)	(0.0046)	(0.0045)	(0.0045)	(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	×	×	×	×	×
$Year\ FEs$	×	×	×	×	×
$District\ FEs$	×	×	×	×	×
$Age\ FEs$	×	×	×	×	×
$Parity\ FEs$	×	×	×	×	×
$District \times year\ FEs$	×	×	×	×	×
$Years\ since\ last\ birth\ FEs$	×	×	×	×	×
$FirstbornGirl \times district\ FEs$			×	×	×

Notes: Coefficients from specification 3 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. $p < 0.1^{*} p < 0.05^{**} p < 0.01^{***}$

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 compostion from Alfano (2017) in Table 11, 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 11: Alternative definitions of children's gender composition

	(1)	(2)	(3)	(4)	(5)	(6)
Dependent variable: Birth indicator						
	Measurement of G_i					
	Youngest child is female		Ratio of girls to boys		More girls than boys	
G_i	0.025***	0.028***	0.050***	0.070***	0.050***	0.0559***
	(0.002)	(0.002)	(0.001)	(0.002)	(0.002)	(0.003)
$G_i * Post$		-0.007*		-0.037***		-0.012***
		(0.003)		(0.002)		(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. $p < 0.1^{*} p < 0.05^{**} p < 0.01^{***}$

8 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 demonstrate that expanding women's reproductive autonomy allowed families to reduce fertility, particularly among those with a firstborn 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. Consistent with the general fertility–investment framework, 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, we find that abortion legalization narrowed gender gaps in breastfeeding, a domain of investment strongly linked to both fertility timing and survival (Jayachandran and Kuziemko (2011)). In particular, daughters in firstborn-girl families, who were previously disadvantaged, experienced large gains in breastfeeding duration post-legalization. However, effects on other investments like vaccinations and institutional deliveries are muted, and mortality improvements are only modest and present in the short-run. Taken together, these patterns suggest that abortion legalization's main effect operated through the fertility channel, with more selective improvements in child investments.

From a policy perspective, the results suggest that legal reforms expanding reproductive rights can improve gender equity indirectly by reshaping fertility behavior, but their impact on child investments is contingent on broader social norms. Reproductive autonomy is a necessary but not sufficient condition for closing gender gaps. Complementary policies, including strengthening maternal and child health systems, expanding immunization coverage, and challenging discriminatory norms are needed to ensure that improvements in autonomy translate into lasting gains for girls' well-being.

Finally, the findings speak to ongoing demographic transitions in South Asia and beyond. As fertility declines and access to reproductive technologies expands, the interaction of autonomy and son preference will remain central to shaping child outcomes. Future research should explore the longer-term consequences for education, labor market participation, and gender equity across cohorts. Nepal's experience illustrates both the transformative potential of abortion legalization and the persistence of cultural constraints that temper its effects.

References

ALFANO, M. (2017): "Daughters, dowries, deliveries: The effect of marital payments on fertility choices in India," *Journal of Development Economics*, 125, 89–104. [25]

ALMOND, D., L. EDLUND, H. LI, AND J. ZHANG (2010): "Long-term effects of early-life development: Evidence from the 1959 to 1961 China famine," in *The economic consequences of demographic change in East Asia*, University of Chicago Press, 321–345. [2], [5], [21]

ANGRIST, J., V. LAVY, AND A. SCHLOSSER (2010): "Multiple experiments for the causal link between the quantity and quality of children," *Journal of Labor Economics*, 28, 773–824. [1], [4]

ANUKRITI, S., S. R. BHALOTRA, AND E. H. F. TAM (2016): "On the quantity and quality of girls: New evidence on abortion, fertility, and parental investments," . [2], [4], [21]

BAILEY, M. J. (2006): "More power to the pill: The impact of contraceptive freedom on women's life cycle labor supply," *The quarterly journal of economics*, 121, 289–320. [1], [3]

BECKER, G. S. AND H. G. LEWIS (1973): "On the interaction between the quantity and quality of children," *Journal of political Economy*, 81, S279–S288. [1], [4], [18], [26]

BECKER, G. S. AND N. TOMES (1976): "Child endowments and the quantity and quality of children," *Journal of political Economy*, 84, S143–S162. [1], [4], [18]

BHAT, P. M. AND A. F. ZAVIER (2003): "Fertility decline and gender bias in Northern India," *Demography*, 40, 637–657. [1], [4], [15]

BIEN, C. H., Y. CAI, M. E. EMCH, W. PARISH, AND J. D. TUCKER (2013): "High adult sex ratios and risky sexual behaviors: A systematic review," *PloS one*, 8. [6]

BLACK, S. E., P. J. DEVEREUX, AND K. G. SALVANES (2005): "The more the merrier? The effect of family size and birth order on children's education," *The Quarterly Journal of Economics*, 120, 669–700. [4]

BRUNSON, J. (2010): "Son preference in the context of fertility decline: limits to new constructions of gender and kinship in Nepal," *Studies in family planning*, 41, 89–98. [5]

CLARKE, D. AND H. MÜHLRAD (2016): "The impact of abortion Legalization on fertility and maternal mortality: new evidence from Mexico," Tech. rep., CINCH Series. [2], [3]

CREHPA (1996): *Opinion Poll Survey on Abortion Rights for Women*. [6]

FROST, M. D., M. PURI, AND P. R. A. HINDE (2013): "Falling sex ratios and emerging evidence of sex-selective abortion in Nepal: evidence from nationally representative survey data," *BMJ open*, 3, e002612. [2], [6] Goldin, C. and L. F. Katz (2002): "The power of the pill: Oral contraceptives and women's career and marriage decisions," *Journal of political Economy*, 110, 730–770. [1], [3]

Gruber, J., P. Levine, and D. Staiger (1999): "Abortion legalization and child living circumstances: who is the "marginal child"?" *The Quarterly Journal of Economics*, 114, 263–291. [3]

Guilmoto, C. Z. (2009): "The Sex Ratio Transition in Asia," *Population and Development Review*, 35, 519–549. [2], [5]

Gupta, M. D. (1987): "Selective discrimination against female children in rural Punjab, India," *Population and development review*, 77–100. [1], [4], [15]

Henderson, J. T., M. Puri, M. Blum, C. C. Harper, A. Rana, G. Gurung, N. Pradhan, K. Regmi, K. Malla, S. Sharma, et al. (2013): "Effects of abortion legalization in Nepal, 2001–2010," *PloS one*, 8, e64775. [1], [7]

Hesketh, T. and Z. W. Xing (2006): "Abnormal sex ratios in human populations: causes and consequences," *Proceedings of the National Academy of Sciences*, 103, 13271–13275. [6]

Holcombe, S. J. and S. Kidanemariam Gebru (2022): "Agenda setting and socially contentious policies: Ethiopia's 2005 reform of its law on abortion," *Reproductive Health*, 19, 218. [2], [3]

Hu, L. and A. Schlosser (2015): "Prenatal sex selection and girls' well-being: Evidence from India," *The Economic Journal*, 125, 1227–1261. [2], [5], [21], [22]

Jayachandran, S. and I. Kuziemko (2011): "Why do mothers breastfeed girls less than boys? Evidence and implications for child health in India," *The Quarterly Journal of Economics*, 126, 1485–1538. [4], [19], [20], [27]

Jayachandran, S. and R. Pande (2017): "Why are Indian children so short? The role of birth order and son preference," *American Economic Review*, 107, 2600–2629. [1], [4]

Lamichhane, P., T. Harken, M. Puri, P. D. Darney, M. Blum, C. C. Harper, and J. T. Henderson (2011): "Sex-selective abortion in Nepal: a qualitative study of health workers' perspectives," *Women's Health Issues*, 21, S37–S41. [6]

Levine, P. B., A. B. Trainor, and D. J. Zimmerman (1996): "The effect of Medicaid abortion funding restrictions on abortions, pregnancies and births," *Journal of Health Economics*, 15, 555–578. [3]

Li, H., J. Zhang, and Y. Zhu (2008): "The quantity-quality trade-off of children in a developing country: Identification using Chinese twins," *Demography*, 45, 223–243. [1], [4]

Lin, M.-J., J.-T. Liu, and N. Qian (2014): "More missing women, fewer dying girls: The impact of sex-selective abortion on sex at birth and relative female mortality in Taiwan," *Journal of the European Economic Association*, 12, 899–926. [2], [3], [22] MoHP (2002): *National Safe Abortion Policy*. [6]

MoHP, N. E. (2012): *Nepal Demographic and Health Survey 2011*. [5]

Mukhiya, G. and N. Mishra (2025): "The history and current scenario of interventional radiology in Nepal," *Journal of Medical Imaging and Interventional Radiology*, 12, 8. [5], [31]

NHSPP, N. H. S. S. P. (2013): *Antenatal Rural Ultrasound Programme: A Pilot Intervention in Dhading District, Nepal: Advocacy Brief*. [5]

Pop-Eleches, C. (2006): "The impact of an abortion ban on socioeconomic outcomes of children: evidence from Romania," *Journal of Political Economy*, 114, 744–773. [3]

Rosenzweig, M. R. and K. I. Wolpin (1980): "Testing the quantity-quality fertility model: The use of twins as a natural experiment," *Econometrica: journal of the Econometric Society*, 227–240. [1], [4]

Rosenzweig, M. R. and J. Zhang (2009): "Do population control policies induce more human capital investment? Twins, birth weight and China's "one-child" policy," *The Review of Economic Studies*, 76, 1149–1174. [4]

Samandari, G., M. Wolf, I. Basnett, A. Hyman, and K. Andersen (2012): "Implementation of legal abortion in Nepal: a model for rapid scale-up of high-quality care," *Reproductive Health*, 9, 7. [1], [7]

Subedi, K. S. and P. Sharma (2013): "Development of radiology in Nepal: gearing up for mountainous challenges," *Journal of the American College of Radiology*, 10, 291–295. [5], [31]

Tamang, A. (1996): "Induced Abortions and Subsequent Reproductive Behaviour Among Women in Urban Areas of Nepal 26 (3 & 4)," *Social Change*, 26, 271–285. [1], [6]

Thapa, S., S. K. Sharma, and N. Khatiwada (2014): "Women's knowledge of abortion law and availability of services in Nepal," *Journal of biosocial science*, 46, 266–277. [1], [6]

Waldron, I. (1998): "Sex differences in infant and early childhood mortality: major causes of death and possible biological causes Too young to die genes or gender. 1998 New York United Nations, Department of Economic and Social Affairs," *Population Division*, 64–83. [4], [5]

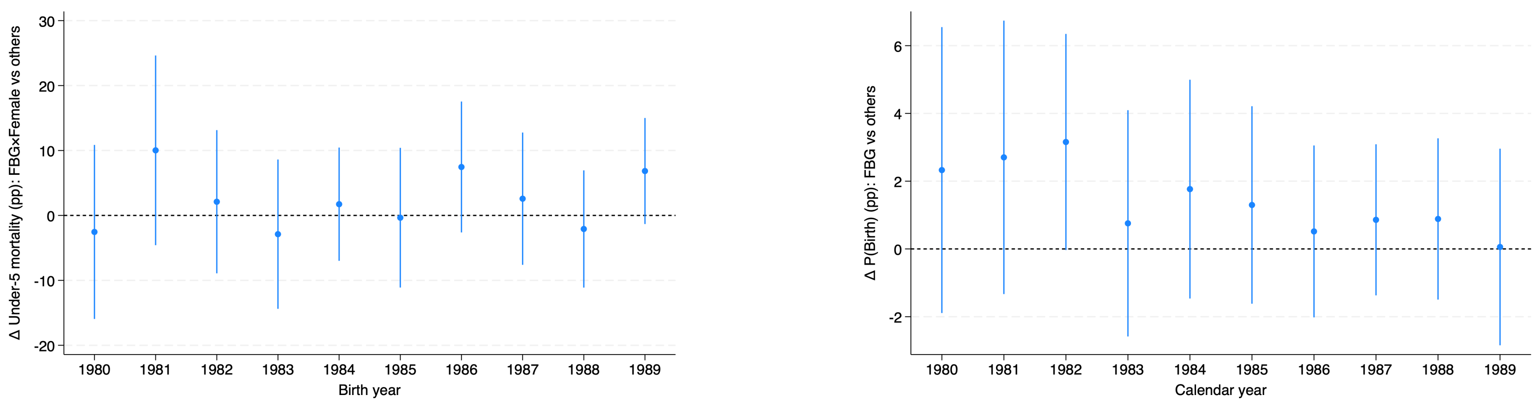
Appendix

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.



(a) Under-five Mortality **(b)** Probability of Birth

Figure A1: 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.

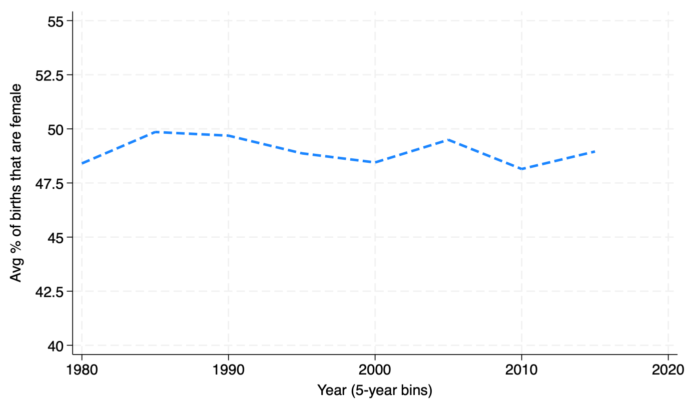
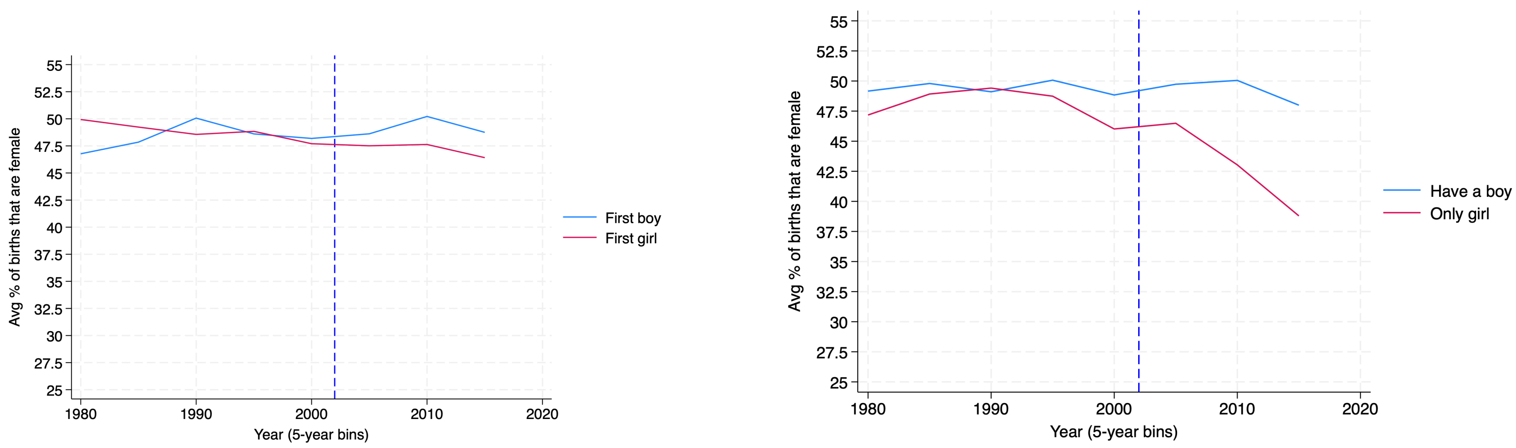


Figure A2: First births



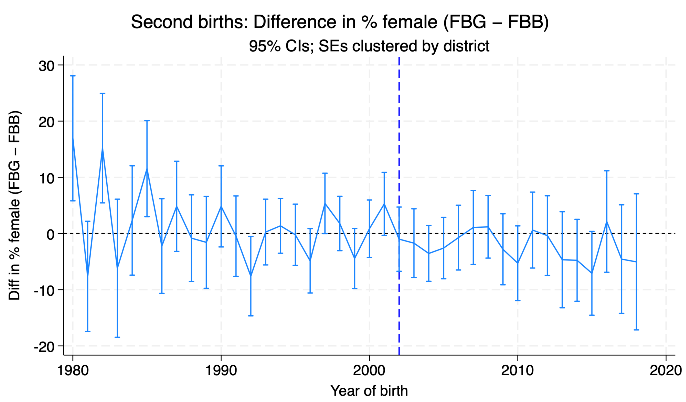
(a) Second births

(b) Third births

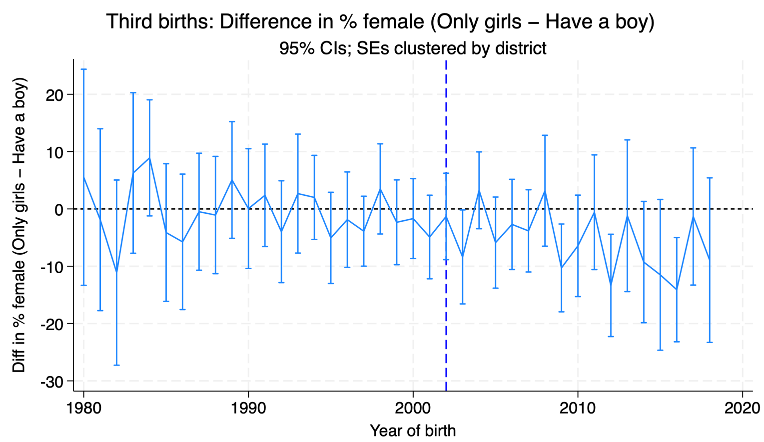
Figure A3: Trends in Proportion of Females at Birth by Birth Order and Sex Composition of Older Siblings

Other Identification Tests

We test that the sex of the firstborn child is effectively random in Figure A2 and that the sex-selective abortion at second and higher-order births occurs primarily in families whose firstborn is a daughter in Figure A4.



(a) Second births



(b) Third births

Figure A4: Trends in Proportion of Females at Birth by Birth Order and Sex Composition of Older Siblings

Table A2: Probability of Birth (Effects of Ultrasound)

	(1)	(2)	(3)	(4)	(5)
FirstbornGirl	0.0498***	0.0589***			
	(0.0050)	(0.0055)			
FirstbornGirl X Post 1	-0.0109**	-0.0204***	-0.0203***	-0.0197***	-0.0196***
	(0.0049)	(0.0055)	(0.0056)	(0.0055)	(0.0054)
FirstbornGirl X Post 2	-0.0324***	-0.0428***	-0.0427***	-0.0422***	-0.0419***
	(0.0053)	(0.0059)	(0.0060)	(0.0058)	(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 3 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 1998-1994 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. $p < 0.1^{*} p < 0.05^{**} p < 0.01^{***}$

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***	-0.2412	-2.4216**	-1.2244	-1.6887**	0.4566
	(0.7709)	(1.5559)	(1.1857)	(0.8494)	(0.7279)	(1.4490)
Firstborn girl * Female * Post1	2.6492**	-0.8513	1.7131	1.5379	1.5255	-0.6409
	(1.1633)	(1.7588)	(1.7832)	(1.1356)	(1.0265)	(2.3371)
Firstborn girl * Female * Post2	2.0511**	0.4293	2.4880**	1.4420*	1.6455**	-0.8010
	(0.8383)	(1.5634)	(1.1988)	(0.8618)	(0.7570)	(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

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	0.0271	0.0076	0.0256*	0.0224**	-0.0075
	(0.0101)	(0.0173)	(0.0140)	(0.0136)	(0.0101)	(0.0205)
Firstborn girl * Female * Post1	-0.0301	-0.0648**	-0.0334	-0.0381	-0.0569**	0.0072
	(0.0225)	(0.0264)	(0.0335)	(0.0251)	(0.0244)	(0.0252)
Firstborn girl * Female * Post2	-0.0085	0.0093	0.0390*	-0.0062	-0.0108	0.0361
	(0.0219)	(0.0205)	(0.0221)	(0.0220)	(0.0201)	(0.0234)
N	26,479	10,536	10,322	19,819	25,221	11,794
Baseline mean	.115	.0581	.0695	.123	.108	.0802

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**	0.2006	0.1588	-0.4652***	-0.2075*	0.0846
	(0.1274)	(0.1722)	(0.1688)	(0.1575)	(0.1105)	(0.2238)
Firstborn girl * Female * Post1	0.2328	-0.1910	0.0150	0.3398	0.1870	-0.0523
	(0.2414)	(0.2552)	(0.3007)	(0.2796)	(0.2101)	(0.4584)
Firstborn girl * Female * Post2	0.2107	-0.0950	-0.1285	0.3969	0.0957	0.2294
	(0.2600)	(0.2184)	(0.2623)	(0.2437)	(0.2042)	(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

Table A6: Excess Female Under-One Mortality

	(1)	(2)	(3)	(4)	(5)
Firstborn girl * Female	0.0109	0.0103	0.0097	0.0099	0.0094
	(0.0071)	(0.0072)	(0.0071)	(0.0071)	(0.0070)
Firstborn girl * Female * Post1	-0.0255**	-0.0246**	-0.0219*	-0.0245**	-0.0217*
	(0.0117)	(0.0117)	(0.0115)	(0.0116)	(0.0115)
Firstborn girl * Female * Post2	0.0030	0.0035	0.0029	0.0041	0.0034
	(0.0101)	(0.0102)	(0.0099)	(0.0102)	(0.0099)
Firstborn girl * Post1	0.0162*	0.0159*	0.0000	0.0159*	0.0000
	(0.0088)	(0.0086)	(.)	(0.0085)	(.)
Firstborn girl * Post2	0.0028	0.0023	0.0000	0.0020	0.0000
	(0.0061)	(0.0062)	(.)	(0.0062)	(.)
Female * Post1	0.0166**	0.0162**	0.0000	0.0158**	0.0000
	(0.0073)	(0.0073)	(.)	(0.0072)	(.)
Female * Post2	-0.0013	-0.0016	0.0000	-0.0021	0.0000
	(0.0063)	(0.0064)	(.)	(0.0063)	(.)
N	50,354	50,353	50,329	50,352	50,328
Baseline mean	.0772	.0772	.0772	.0772	.0772

Table A7: Excess Female Under-Two Mortality

	(1)	(2)	(3)	(4)	(5)
Firstborn girl * Female	0.0123	0.0118	0.0115	0.0114	0.0111
	(0.0075)	(0.0076)	(0.0073)	(0.0076)	(0.0073)
Firstborn girl * Female * Post1	-0.0288**	-0.0282**	-0.0248*	-0.0279**	-0.0244*
	(0.0132)	(0.0132)	(0.0129)	(0.0131)	(0.0128)
Firstborn girl * Female * Post2	0.0020	0.0026	0.0010	0.0037	0.0021
	(0.0117)	(0.0119)	(0.0114)	(0.0118)	(0.0114)
Firstborn girl * Post1	0.0180*	0.0175*	0.0000	0.0173*	0.0000
	(0.0095)	(0.0094)	(.)	(0.0093)	(.)
Firstborn girl * Post2	0.0037	0.0029	0.0000	0.0023	0.0000
	(0.0070)	(0.0070)	(.)	(0.0071)	(.)
Female * Post1	0.0154*	0.0152*	0.0000	0.0146*	0.0000
	(0.0081)	(0.0081)	(.)	(0.0080)	(.)
Female * Post2	-0.0037	-0.0040	0.0000	-0.0049	0.0000
	(0.0077)	(0.0078)	(.)	(0.0078)	(.)
N	47,418	47,417	47,396	47,416	47,395
Baseline mean	.0904	.0904	.0904	.0904	.0904

Table A8: Excess Female Under-Three Mortality

	(1)	(2)	(3)	(4)	(5)
Firstborn girl * Female	0.0129*	0.0124	0.0124	0.0120	0.0120
	(0.0077)	(0.0078)	(0.0076)	(0.0078)	(0.0075)
Firstborn girl * Female * Post1	-0.0329**	-0.0321**	-0.0294**	-0.0316**	-0.0289**
	(0.0148)	(0.0148)	(0.0145)	(0.0147)	(0.0144)
Firstborn girl * Female * Post2	0.0039	0.0048	0.0027	0.0056	0.0035
	(0.0127)	(0.0128)	(0.0122)	(0.0127)	(0.0120)
Firstborn girl * Post1	0.0193*	0.0183*	0.0000	0.0181*	0.0000
	(0.0104)	(0.0103)	(.)	(0.0102)	(.)
Firstborn girl * Post2	0.0022	0.0012	0.0000	0.0008	0.0000
	(0.0078)	(0.0079)	(.)	(0.0080)	(.)
Female * Post1	0.0145	0.0141	0.0000	0.0133	0.0000
	(0.0091)	(0.0091)	(.)	(0.0090)	(.)
Female * Post2	-0.0062	-0.0069	0.0000	-0.0076	0.0000
	(0.0082)	(0.0083)	(.)	(0.0083)	(.)
N	44,263	44,262	44,241	44,261	44,240
Baseline mean	.0955	.0955	.0955	.0955	.0955

Table A9: Excess Female Under-Four Mortality

	(1)	(2)	(3)	(4)	(5)
Firstborn girl * Female	0.0172**	0.0165*	0.0165**	0.0162*	0.0161**
	(0.0082)	(0.0083)	(0.0081)	(0.0082)	(0.0080)
Firstborn girl * Female * Post1	-0.0375**	-0.0364**	-0.0337*	-0.0360**	-0.0332*
	(0.0168)	(0.0169)	(0.0170)	(0.0169)	(0.0169)
Firstborn girl * Female * Post2	0.0041	0.0054	0.0024	0.0062	0.0032
	(0.0142)	(0.0143)	(0.0138)	(0.0142)	(0.0137)
Firstborn girl * Post1	0.0230*	0.0221*	0.0000	0.0219*	0.0000
	(0.0117)	(0.0116)	(.)	(0.0114)	(.)
Firstborn girl * Post2	0.0040	0.0031	0.0000	0.0027	0.0000
	(0.0090)	(0.0091)	(.)	(0.0091)	(.)
Female * Post1	0.0149	0.0144	0.0000	0.0136	0.0000
	(0.0100)	(0.0100)	(.)	(0.0099)	(.)
Female * Post2	-0.0079	-0.0086	0.0000	-0.0094	0.0000
	(0.0093)	(0.0095)	(.)	(0.0094)	(.)
N	40,701	40,698	40,677	40,697	40,676
Baseline mean	.0991	.0991	.0991	.0991	.0991

Table A10: Contraceptive Use

	Frequency	Percent
Not Using	81287	50.01
Pill	5163	3.18
IUD	1501	0.92
Injections	15752	9.69
Diaphragm /Foam/Jelly	21	0.01
Condom	4512	2.78
Female Sterilization	27641	17.01
Male Sterilization	12469	7.67
Periodic Abstinence	1957	1.20
Withdrawal	7650	4.71
Norplant	4308	2.65
Other	266	0.16
Total	162527	100.00

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