






Maternal education and child survival: causal evidence from Kenya

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Abstract

This study leverages the educational reform of 1985 as a source of exogenous variation in female education, providing insights into the effect of maternal schooling on the probability of child mortality at age one or younger and age five or younger. Utilizing data from the five waves of the KDHS conducted in 1993, 1998, 2003, 2008, and 2014, we employ a two-stage least-squares (2SLS) approach. Our findings indicate that women exposed to the 1985 policy change, on average, have approximately 1.87 more years of schooling compared to their counterparts. Moreover, each additional year of maternal schooling leads to a reduction in the likelihood of a child's death at age one or younger and at age five or younger by 0.6 and 0.9 percentage points, respectively. These results remain robust across a spectrum of robustness checks. Furthermore, we explore various potential mechanisms elucidating the influence of maternal education on child mortality. These mechanisms include examining fertility behavior, the likelihood of maternal engagement in the labor force, maternal health-seeking behaviors for children, and maternal involvement in household decision-making.

Keywords DHS · Maternal education · Children mortality rates · Kenya · Infant · Under-five

JEL classification I15 · I25 · J13

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1 Introduction

Globally, child mortality has drastically decreased since the Millennium Development Goals (MDG) were announced in 2002. An estimated 5.30 million children met their demise before reaching the age of five in 2019, marking a notable decline from the approximately 9.92 million deaths recorded for children under five years of age in 2000 (Perin et al., 2022). This reduction can be attributed to concerted efforts across diverse platforms by numerous organizations and governments during the MDG period from 2000 to 2015, extending into the subsequent era characterized by the Sustainable Development Goals (SDG). However, the advancement achieved displays disparities both among and within countries, as highlighted by Kimani-Murage et al. (2014) and Mejía-Guevara et al. (2019). Children born in Sub-Saharan Africa confront the highest risk of early mortality globally, exemplified by an under-five mortality rate (children deceased before their fifth birthday) of 74 deaths per 1,000 live births documented in 2021. This rate is 15 times higher than the risk observed for children in Europe and Northern America and 19 times higher than the corresponding rate in Australia and New Zealand, as substantiated by UNICEF (2022). Similarly, fatalities occurring before the age of one are predominantly concentrated in Sub-Saharan Africa, particularly in East Africa, exhibiting the highest infant mortality rate (92.2 deaths per 1000 live births) when juxtaposed with figures from developed countries (8 deaths per 1000 live births) (Tesema et al., 2022).

Kenya emerges as a significant focal area of concern in the global context of child mortality. According to Abuya et al. (2011), child mortality remains a significant health challenge in Kenya. The country falls short of achieving its MDG targets for child mortality, displaying inconsistent progress in trends spanning the period from 1990 to 2015 (Daniel et al., 2021). Despite recent improvements, the under-five mortality rate in Kenya stands at 37 deaths per 1000 live births in 2021, as documented by the World Bank (2023). In the endeavor to reduce childhood mortality, Kenya exhibits a comparative lag in relation to its East African counterparts, experiencing a 52% decline in under-five mortality, as opposed to 70 and 71% for Tanzania and Uganda, respectively. This discrepancy signifies that while Tanzania and Uganda successfully meet and surpass their MDG targets, Kenya falls short of achieving its prescribed goal (Machio, 2018). Presently, the under-five mortality rate is projected to approach the 2030 SDG target of 25 deaths per 1,000 live births, based on the prevailing annual rate of reduction (Keats et al., 2018). Effectively addressing child mortality necessitates intensified efforts to attain the SDG3 targets, striving for a rate no greater than 25 under-five deaths per 1000 live births by 2030.

To advance ongoing efforts aimed at reducing childhood mortality in Kenya, it is imperative to identify factors robustly that exhibit a robust association with child mortality. Machio (2018) identifies factors contributing to the slow decline rates in child mortality in Kenya compared to its East African neighbors, attributing it to the insufficient utilization of maternal healthcare services and proposing solutions, including increasing maternal education and decreasing average distances to health facilities. Omariba et al. (2007) find that children of mothers with no education face an increased risk of death of 16% in their infant time and approximately 33% in their childhood. Using the Kenya Demographic Health Survey (KDHS) in 2014, infants whose mothers had no education were 1.4 times more likely to die in the first 28 days

compared to mothers with higher education levels (Imbo et al., 2021). The aforementioned findings underscore the pivotal role of maternal education in influencing child mortality rates. Nevertheless, existing literature mainly concentrates on establishing correlations, thereby impeding the elucidation of a causal link between maternal education and their children's health. Accordingly, our study aims to address and bridge this research gap.

Aligned with prior studies that utilize educational policy alterations to address the endogeneity issues in education (Andriano & Monden, 2019; Breierova & Duflo, 2004; Grépin & Bharadwaj, 2015; Nguyen-Phung, 2023), this study capitalizes on the educational policy shift in 1985 as a source of exogenous variation in female schooling. To examine the impact of maternal education on child health outcomes, we use two health indicators, including a child deceased at age one or younger (infant mortality) and a child deceased at age five or younger (under-five mortality). Drawing data from the five waves of the KDHS in 1993, 1998, 2003, 2008, and 2014, we apply the two-stage least-squares (2SLS). Our findings show that, on average, women exposed to the policy change in 1985 have approximately 1.87 years of schooling in comparison to their counterparts. One additional year of schooling leads to a reduction in the likelihood of a child being dead at age one or younger and at age five or younger by 0.6 and 0.9 percentage points, respectively. Our results are robust through a range of robustness checks. We also examine the possible channels explaining the impacts of maternal education on child mortality, including fertility behavior, the mother's likelihood to engage in the labor force, the mother's health-seeking behaviors for her children, and the mother's involvement in decision-making in the household.

This paper makes substantial contributions to the existing literature on the inter-generational effects of women's education on the well-being of their offspring. First, the study provides empirical insights into the causal relationship between mothers' education and the mortality rates of their children within the Kenyan context. Second, in addressing potential endogeneity concerns and establishing a causal link, we employ the educational reform of 1985 as an instrumental variable (IV). The constructed IV is novel as it incorporates factors such as late enrollment and the grade repetition rate prevalent in Kenya. Third, the research explores the pathways through which maternal education in Kenya influences the development of children. Finally, given Kenya's commitment to achieving the objectives outlined in SDG3, a nuanced comprehension of the significance of maternal education in fostering children's well-being is crucial for policymakers.

The remainder of this article is as follows. The next section discusses a literature review on maternal education and their children's health outcomes, highlighting various indicators and a brief of the 1985 policy. Next are the methods and the identification strategy. The results section presents a synthesis of the main findings. Lastly, the discussion section concludes and recommends policy implications.

2 Related literature

2.1 Conceptual framework and hypotheses

Maternal influence on the longevity and vitality of offspring is a complex and significant subject thoroughly investigated in a multitude of studies (Currie, 2009;

Grossman, 2006; Strauss & Thomas, 1995). Previous research has illuminated the enduring ramifications of early-life health on later-life outcomes, underscoring the pivotal role that mothers, as the archetypal caregivers, play in sculpting these outcomes (Dewey & Begum, 2011; Martorell, 1999). The health and nutrition of children are profoundly affected by maternal actions within an environment largely shaped by maternal choices and socio-economic conditions, which decisively impacts their survival and growth (Currie, 2009; Hart et al., 2010; Papas et al., 2009; Robinson et al., 2007).

The level of maternal education stands as a formidable determinant of child health and mortality (Caldwell, 1993; Hobcraft et al., 1985; Li et al., 2020; Minh et al., 2016; Wolde et al., 2015). It has been documented that mothers who have attained higher levels of education are more likely to have offspring with superior health outcomes due to more informed decisions regarding health and nutrition (Aslam & Kingdon, 2012; De Neve & Subramanian, 2018; Ghuman et al., 2005; Glewwe, 1999; Güneş, 2015; Grytten et al., 2014; Keats, 2018; Lundborg et al., 2014; Robinson et al., 2007). This correlation is accentuated in terms of the mortality rate of children. Gakidou et al. (2010) and Le & Nguyen (2020) investigate on a global scale and in the developing countries context, respectively, and demonstrate that each additional year of maternal schooling is associated with a significant decrease in child mortality risks. Although there is only a limited number of studies on a country-wide scale, this positive impact of maternal education on reducing children's mortality has been observed in several developing regions involving Taiwan (Chou et al., 2010), Vietnam (Nga et al., 2012; Nguyen-Phung, 2023) and Indonesia (Breierova & Dufflo, 2004; Wang, 2021). Moreover, research on African nations is in lack of robust evidence. As Andriano & Monden (2019), Makate (2016), Makate & Makate (2016), and Grépin & Bharadwaj (2015) prove significant relationships in Malawi, Uganda, and Zimbabwe, while Ali & Elsayed (2018) and McCrary & Royer (2011) suggest a null result.

Grounded in the aforementioned studies, we hypothesize:

Hypothesis 1: The enhancement of maternal education is inversely related to child mortality rates in Kenya.

The current empirical research on the relationship between maternal education and children's mortality has not fully explored the potential pathways. Therefore, this study posits four channels through which maternal education influences child mortality.

The first pathway relates to maternal fertility behavior. Increased educational attainment is associated with a postponement of initial sexual activity (Black et al., 2008; Masuda & Yamauchi, 2020), deferred marriage (Ali & Gurmu, 2018; Field & Ambrus, 2008; Parsons & McCleary-Sills, 2014), and an extension of the age at first childbirth (Ali & Gurmu, 2018; Grépin & Bharadwaj, 2015). Literature affirms that children born to very young mothers or as firstborns encounter a higher risk of mortality (Hobcraft, 1993; Hobcraft et al., 1985). Moreover, there is a consistent demonstration of a reduction in fertility rates and mean birth interval correlated with the heightened educational levels of women (Ali & Gurmu, 2018; Becker et al., 2013; Grépin & Bharadwaj, 2015; Keats, 2018; Le & Nguyen, 2020; Makate & Makate, 2016; Osili & Long, 2008). As Becker & Lewis (1973) explore the interplay between the "quantity" and "quality" of children, elucidating a positive correlation

between a woman's education and the "quality" of her offspring, educational advancements indeed result in healthier childbirth practices and child-rearing environments (Currie & Moretti, 2003).

The second pathway is through labor force engagement. Maternal academic achievements translate into improved labor market opportunities, equipping women with essential skills for earning an income (Alam et al., 2011; Fasih, 2008; Goldberg & Smith, 2013; Herman, 2012; Krishnakumar & Nogales, 2020), therefore explain part of the improvement on children's health outcome (Le & Nguyen, 2020). Women with higher educational credentials command increased earnings, which enables them to invest in child health care services. Since women are central to household consumption decisions, particularly regarding diet and nutrition, their active participation in the workforce empowers them to invest in the nutritional and developmental needs of their children (Moehling, 1995; Smith et al., 2003). Mothers with greater economic resources can procure higher-quality nutrition and provide superior care for their children, contributing to enhanced child survival rates (Amugsi et al., 2016; Smith et al., 2003). From a broader perspective, children who benefit from improved nutrition and care are likely to develop better cognitive skills and achieve more favorable labor outcomes in adulthood (Currie, 2009; Kar et al., 2008).

The third pathway is health-seeking behavior. Educational endeavors serve as a vehicle for dispelling ignorance and augmenting mothers' capabilities in attaining a broad spectrum of information, leading to higher maternal health literacy (Duflo, 2012; Samarakoon & Parinduri, 2015; Thomas et al., 1991). This broader comprehension encompasses vital aspects of disease causation, prevention, identification, modern treatment, and a deeper understanding of nutritional needs that directly influence child mortality (Frost et al., 2005). Furthermore, education grants women access to and enhances their interpretation of health communications and advisories from varied sources, thus bolstering maternal health literacy (Ahmadi & Karamitanha, 2023; Gibbs et al., 2016; Phommachanh et al., 2021). As a result, educated women are better positioned to safeguard their children from infections through improved sanitary practices, diminish the risk of diseases through adequate nutrition, and expedite recovery from illnesses with more effective healthcare interventions (Ahmadi & Karamitanha, 2023; Andriano & Monden, 2019; Caldwell, 1979; Gibbs et al., 2016). Smoking is one of those detrimental practices that cause a higher mortality rate and is more likely to be avoided during pregnancy for women with higher educational levels (Currie & Moretti, 2003; Lindeboom et al., 2009). Moreover, women are more likely to pay more visit to prenatal care visits and deliver in health institutes so as to promise a facilitated environment for childbirth (Cleland & Van Ginneken, 1988; Grépin & Bharadwaj, 2015; Le & Nguyen, 2020; Makate & Makate, 2016; Nga et al., 2012). After the birth of infants, women with higher education levels intend to prepare babies with better immune systems through vaccination in order to fight against malnutrition and disease that lead to 10 percent of under-5 deaths (Keats, 2018; Makate & Makate, 2016).

The fourth pathway is women's involvement in decision-making. Education bestows confidence, independence, awareness, and self-worth, which are vital for initiating changes in women's decision-making abilities to manifest their intrinsic beliefs and surmount internal barriers (Chanana, 1996; Orenstein, 2013; Rivera-Garrido, 2022; Samarakoon & Parinduri, 2015). Women's empowerment, especially

as primary caretakers, exerts a substantial impact on child survival, allowing educated women to impose greater control over health choices for their children (Bhagowalia et al., 2012; Caldwell, 1979; Nguyen et al., 2013; Smith et al., 2003). Research conducted in Ghana indicates that women with significant autonomy in household decision-making are more likely to secure a diverse diet for their families, thereby contributing to their children's health (Amugsi et al., 2016). This dynamic is particularly evident among women with advanced education. Women's ability to make informed choices and procure resources as an empowerment dimension has a considerable correlation with the availability of a nutritious diet at home, which is a key factor in reducing child mortality (Aslam & Kingdon, 2012; Larsson & Stanfors, 2014; Nguyen-Phung & Nthenya, 2023). Consequently, educational attainments that confer bargaining power empower females to make informed household decisions, ultimately contributing to the enhancement of childbirth quality.

Accordingly, the formulated second hypothesis is:

Hypothesis 2. Maternal education's impact on child mortality in Kenya is facilitated through shifts in fertility behavior, labor market engagement, increased informational access, and women's involvement in household decision-making.

2.2 The 1985 education reform

Kenya's educational architecture, prior to the reform of 1985, included a sequential framework of seven years of primary education, followed by secondary and higher education, culminating in graduation. In response to critiques from the private sector about the education system's academic focus being disconnected from the labor market and development needs, the Kenyan government reformed its education policy in 1985, adding an extra year to primary schooling (Eshiwani, 1990; Wanjohi, 2011).¹ Maintaining the cumulative 16-year educational span, this policy shift modified the structure from 7 + 4 + 2 + 3 to 8 + 4 + 4 as depicted in Fig. 1, resulting in a significant rise in primary school enrollment and retention, particularly enhancing educational attainment for girls (Somerset, 2007). Consequently, seventh-grade students in 1985 progressed to an additional primary grade instead of advancing to secondary school. All students who were enrolled in a lower Grade than Grade 7 would attain an extra year of primary education compared with the older cohort who had already completed primary education at the time of reform. The reform also integrated practical skills such as arts and crafts into the curriculum, aiming to

¹ The 1985 educational reform introduced substantial alterations in curriculum and instruction, particularly at the primary level, where the number of subjects expanded from seven to thirteen, with nine subjects evaluated in the Kenya Certificate of Primary Education (Somerset, 2009). The redesigned curriculum emphasized prevocational subjects such as art and crafts, agriculture, and home science, aiming to provide students with essential knowledge and skills pertinent to the labor market (Kim, 2020; Sifuna, 1992). This realignment was consistent with the reform's goal of implementing the 8-4-4 system, which emphasizes a practical curriculum geared towards equipping graduates with employable skills, addressing the demand for skill-oriented education. A key change in instruction was the language used, particularly in primary education, where the reform mandated the use of the local vernacular for the first three years, transitioning to English thereafter. Moreover, under the 8-4-4 system, English and Swahili became mandatory and examinable subjects in primary schools (Piper et al., 2016), marking a departure from the previous 7-4-2-3 system where English was the primary language of instruction from the beginning (Anderson, 1965; Kim, 2020).

Age	Grade system* (before reform in 1985)	Grade system* (after reform in 1985)
4	Pre-school	Pre-school
5		
6	Primary Grade 1	Primary Grade 1
7	Primary Grade 2	Primary Grade 2
8	Primary Grade 3	Primary Grade 3
9	Primary Grade 4	Primary Grade 4
10	Primary Grade 5	Primary Grade 5
11	Primary Grade 6	Primary Grade 6
12	Primary Grade 7	Primary Grade 7
13	Secondary Grade 1	Primary Grade 8
14	Secondary Grade 2	Secondary Grade 1
15	Secondary Grade 3	Secondary Grade 2
16	Secondary Grade 4	Secondary Grade 3
17	Upper Secondary Grade 5	Secondary Grade 4
18	Upper Secondary Grade 6	University Grade 1
19	University Grade 1	University Grade 2
20	University Grade 2	University Grade 3
21	University Grade 3	University Grade 4
22	Graduated	Graduated
23		

* Grade in the normal case without late enrolment and grade repetition.

Fig. 1 Kenya’s education structure in pre-reform and post-reform

inculcate vocational skills at the primary level (Somerset, 2009). This period also saw prevalent late enrollments and grade repetitions across Africa, with Kenya experiencing similar trends, leading to older students or extended durations for completing primary education.

The reform had an extensive impact on those commencing primary education in 1979 or thereafter or individuals under 14 years of age in 1985 (i.e., those born after 1971). It partially affected women born between 1965 and 1971, with no discernible impact on those born between 1950 and 1964. The forthcoming sections provide an in-depth exploration of the reform’s influence on child mortality, with a particular focus on the role of improved maternal education.

3 Research data

3.1 Data sources

We utilize data from five waves of the KDHS conducted in 1993, 1998, 2003, 2008, and 2014. The surveys employ a two-step implementation process. Initially, enumeration areas are derived from the population census, followed by the selection of

households within these chosen areas. To ensure national representativeness, we adjust the survey weights. The survey encompasses comprehensive information, including details about women's backgrounds and their children's mortality.

Furthermore, our study investigates the likelihood of a student advancing to secondary school before the introduction of the educational policy, utilizing data from two distinct sources. First, to accommodate repetition rates and Grade One enrollment age, we incorporate information from Somerset (2007). Second, for an evaluation of the transition from primary to secondary school levels, we extract progression rates from Ohba (2009).

3.2 Variable definitions

To examine the influence of maternal education on child mortality rates, we utilize the number of years of schooling completed during adulthood as the primary indicator of female education. Our study incorporates additional education-related measures, and we conduct robustness checks to ensure the reliability of our findings. The subset of data used in this analysis is drawn from individual records of women born between 1950 and 1980, ensuring that their educational choices have been finalized.

Child mortality rate metrics are derived from birth records consisting of information about the child's gender, birth details, and age at death. We employ two binary measures of child mortality rates, specifically death at age one or younger and death at age five or younger. First, we create a dummy indicator representing death at age one or younger, assigning 1 if the child died within the first year of life and 0 otherwise. Second, we establish a binary indicator for death at age five or younger, where 1 represents the child's death before or at five years old and 0 otherwise.

Table 1 provides a comprehensive overview of descriptive statistics. In Panel A, we present the mean and standard deviation of the outcomes. On average, women complete approximately 6.8 years of schooling. The percentages of children who have deceased at age one or younger and at age five or younger are 6.5 and 7.2%, respectively. Moving to Panel B, we present the mean and standard deviation of control variables. Nearly half (49%) of the children are female, and 19.1% reside in urban areas. The ethnic distribution reveals that approximately 71.6% of groups belong to Kalenjin, Kamba, Kikuyu, Luhya, and Luo. Geographically, 72.7% of children are located in the Eastern, Nyanza, Rift Valley, and Western regions.

4 Research methodology

4.1 Probability of belonging to an educational cohort

It is well-documented that grade repetition and late entry in Kenya are extensively prevalent (Somerset, 2009), influencing diverse birth cohorts through educational policy. Therefore, the cohort years can be categorized into three different groups: the fully treated group, comprising cohorts exposed to the new educational system; the partially treated group, encompassing cohorts potentially affected by policy change

Table 1 Descriptive statistics

	Mean	SD	Min	Max
Panel A: Outcome variables				
Years of schooling	6.796	3.879	0	23
Dead at age one or younger	0.065	0.247	0	1
Dead at age five or younger	0.072	0.259	0	1
Panel B: Control variables				
Child is female	0.493	0.500	0	1
Urban status	0.191	0.393	0	1
Age of mother	31.151	6.755	15	49
<i>Ethnicity:</i>				
Kalenjin	0.141	0.348	0	1
Kamba	0.122	0.327	0	1
Kikuyu	0.164	0.371	0	1
Kisii	0.060	0.237	0	1
Luhya	0.164	0.370	0	1
Luo	0.125	0.331	0	1
Meru/Embu	0.060	0.237	0	1
Mijikenda/Swahili	0.061	0.239	0	1
Somali	0.027	0.162	0	1
Other	0.077	0.266	0	1
<i>Regional area</i>				
Nairobi	0.061	0.239	0	1
Central	0.102	0.303	0	1
Coast	0.091	0.287	0	1
Eastern	0.166	0.372	0	1
Nyanza	0.163	0.369	0	1
Rift Valley	0.261	0.439	0	1
Western	0.137	0.344	0	1
Northeastern	0.020	0.138	0	1

SD standard deviation

owing to grade repetition and delayed entry; and the control group, representing cohorts whose ages exceeded 20 years, thereby casting doubt on their susceptibility to the reform's impact. In response to the challenge associated with the partially treated group, certain academic investigations choose to completely omit this cohort from their analysis (Agiéro & Bharadwaj, 2014; Makate & Makate, 2016). Nonetheless, this method results in a considerable reduction in the quantity of observations and has the potential to generate an upward bias in the estimated outcomes. Other researchers, on the other hand, use a binary allocation of 0 or 1, failing to take into consideration the noteworthy obstacles posed by delayed enrollment and grade competition (Ferré, 2009). Rather than designating a binary for our cohorts or eliminating the partially treated group, we utilize the data from Somerset (2007) and Ohba (2009) to apply percentages of repetition and advancement to lower secondary

education and derive a probability distribution where each birth cohort is assigned to education cohorts via a probabilistic treatment assignment.²

In a more detailed breakdown, our sample is categorized into three different groups based on birth years, each of which is ascribed to a corresponding probability. The fully treated group, comprising female respondents born between 1972 and 1980, is designated a probability of 1. Women born between 1965 and 1971 fall into the partially treated group, and a probability between 0 and 1 is assigned to each individual within this category.³ As for the control group, a probability of 0 is allocated to women respondents born between 1950 and 1964.

$$pr_{-iv_b} = \begin{cases} 1 & \text{if } 1971 < b \\ Pr(\text{partially treated}) & \text{if } 1964 < b < 1972 \\ 0 & \text{if } b < 1965 \end{cases} \quad (1)$$

In this context, b symbolizes the birth cohort, Pr denotes the probability, and pr_{-iv_b} signifies the probability assigned to a female respondent being affiliated with the designated birth cohort b .

4.2 Identification strategy

We present the following econometric model to examine the influence of maternal education on their child's mortality rates:

$$mortality_{ib} = \alpha_0 + \alpha_1 maternal_educ_{ib} + \alpha_2 X_{ib} + \varepsilon_{ib} \quad (2)$$

where $mortality_{ib}$, $maternal_educ_{ib}$, and X_{ib} denote the mortality rates for a child i from birth cohort b , maternal education, and a set of control variables, respectively. In this study, we control for the gender of the child, the mother's ethnicity, the age of the mother, the status of the rural residents, and regional areas. α_0 , α_1 , and α_2 are associated parameters and ε_{ib} is the error term.

To alleviate the potential endogeneity issues inherent in the aforementioned ordinary least square (OLS) regression, we employ the instrumental variable (IV) regression. To be more specific, we utilize the previously derived probability of exposure to the policy (pr_{-iv_b}) as an IV. The rationale for employing pr_{-iv_b} as the IV is explicated as follows: Initially, confined to the designated range of ages, it is evident that the influence of educational reform on a child's mortality rates is anticipated to be mediated exclusively through its effect on their mother's education. It is crucial to maintain this exclusion restriction, considering that the implementation of the policy does not explicitly target child mortality rates. While there is a potential for the violation of this assumption due to the direct influence of other environmental or policy shifts, women within the

² Note that as opposed to the binary method, wherein each cohort is definitively assigned values of 0 or 1 to denote certain exposure to the new educational reform, we alleviate this constraint by endowing each birth cohort, particularly the partially treated group, with a probability of experiencing the reform. This methodological paradigm, originally proposed by Angrist & Lavy (1999), has undergone subsequent refinement and application in recent scholarly research (Chicoine, 2012; Lindskog & Durevall, 2021; Nguyen-Phung & Nthenya, 2023).

³ Interested readers can refer to Appendix A for detailed computations.

specified age range remain unaffected by these alterations.⁴ Second, our IV exhibits a robust association with maternal education. Specifically, the relevance of the IV is scrutinized in Table 3, signifying the absence of weakness in our chosen IV.

We use the two-stage least-squares (2SLS) approach to investigate the influence of maternal education on their child's mortality rates. In the first stage, we regress maternal education on our IV and other covariates:

$$\mathit{maternal_educ}_{ib} = \alpha_{f0} + \alpha_{f1}pr_iv_b + \alpha_{f2}X_{ib} + \varepsilon_{fib} \quad (3a)$$

where f symbolizes the first stage, α_{f0} , α_{f1} , and α_{f2} are parameters, and ε_{fib} is an error term. Next, we regress the child's mortality rates on the estimated value of maternal education and other covariates:

$$\mathit{mortality}_{ib} = \alpha_{s0} + \alpha_{s1}\widehat{\mathit{maternal_educ}}_{ib} + \alpha_{s2}X_{ib} + \varepsilon_{sib} \quad (3b)$$

where s signifies the second stage, α_{s0} , α_{s1} , and α_{s2} are parameters, and ε_{sib} is an error term.

5 Empirical results

5.1 Two-stage least-squared estimations

5.1.1 The impact of educational policy on female education

Table 2 presents the results of the first-stage estimation. Specifically, women exposed to the educational reform have, on average, approximately 1.87 more years of formal education compared to their counterparts.

The impact of educational reforms on female schooling is evident, in contrast to the statistically insignificant effects observed in the context of exposure to fictional policies. To conduct this comparison, we examine the educational achievements of individuals exposed to the fictional policy versus those unexposed, categorized by women's years of birth. Specifically, the fictional policy implementations occurred in 1951, 1957, 1977, and 1982, aligning with the corresponding birth years of women (1948–1953), (1954–1959), (1974–1979), and (1979–1984), respectively. The estimated coefficients, presented in Table 3, indicate no significant effects of fictional reforms on female schooling. Consequently, these findings affirm the conclusion that the educational policy has an insignificant impact on the educational attainment of unexposed cohorts.

5.1.2 The impact of maternal schooling on their child mortality rates

The main findings derived from the 2SLS estimation are presented in Table 4.⁵ Leveraging the predicted values in the first stage regression, we assess the

⁴ It is pertinent to acknowledge that assessing the validity of this assumption directly is hindered by the inherent complexities. Therefore, its validity is scrutinized through a range of robustness checks, and the detailed findings will be presented in subsequent sections.

⁵ The results of the baseline OLS estimation are reported in Appendix B for further details.

Table 2 The impact of educational reform on female schooling

	Years of schooling
Exposed to the reform	1.872*** (0.141)
R^2	0.297
<i>Observations</i>	22,489

The sample includes female respondents born from 1950 to 1980. All estimations include control variables of child's gender, mother's ethnicity, age of the mother, status for rural resident, and regional areas. Standard errors are in parentheses and clustered at the birth cohort

*** Significant at 1 per cent

Table 3 Fictional policies

	Years of schooling
Panel A: Policy implementation in 1951 and birth years of women (1948–1953)	0.118 (0.578)
R^2	0.152
<i>Observations</i>	481
Panel B: Policy implementation in 1957 and birth years of women (1954–1959)	0.114 (0.258)
R^2	0.189
<i>Observations</i>	1346
Panel C: Policy implementation in 1977 and birth years of women (1974–1979)	0.017 (0.104)
R^2	0.344
<i>Observations</i>	8336
Panel D: Policy implementation in 1982 and birth year of women (1979–1984)	−0.109 (0.124)
R^2	0.344
<i>Observations</i>	8952

The sample includes female respondents born from 1950 to 1980. All estimations include control variables of child's gender, mother's ethnicity, age of the mother, status for rural resident, and regional areas. Standard errors are in parentheses and clustered at the birth cohort

causal impacts of maternal education on two mortality indicators. To elaborate, each additional year of female education results in a decrease of 0.6 percentage points in the probability of a child's death at age one or younger and a reduction of 0.9 percentage points in the likelihood of a child's death at age five or younger.⁶

⁶ We add the survey fixed effects in the main estimation and report the results in Panel A in Appendix Table 15. We include a squared term of age trend in this model to account for continuous changes in the outcomes over birth cohorts, capturing both linear and quadratic trends (See Panel B in Appendix Table 15). Additionally, we incorporate a control for school attendance from the Kenya Census 1999 (See Panel C in Appendix Table 15). The estimated results from these exercises corroborate our key findings.

Table 4 The impact of female education on their child mortality rates

	Dead at age one or younger	Dead at age five or younger
Years of schooling	-0.006** (0.003)	-0.009*** (0.003)
Centered R^2	0.018	0.020
Observations	22,489	22,489
Weak id.	175.341	175.341

The sample includes female respondents born from 1950 to 1980. All estimations include control variables of child's gender, mother's ethnicity, age of the mother, status for rural resident, and regional areas. Standard errors are in parentheses and clustered at the birth cohort

** Significant at 5 per cent; *** significant at 1 per cent. Weak id. refers to Kleibergen-Paap rk Wald F statistics

We perform an array of robustness checks on the main findings presented in Table 4. First, we revisit the main results using two alternative measures of female education, ensuring the findings remain unaffected by variations in variable definitions or the functional form of the estimation model. The first measure represents the probability of completing at least 8 years of schooling, while the second measure involves the natural logarithm of female schooling. The outcomes of this re-estimation are detailed in Table 5. In Panel A, mothers with an education attainment of at least 8 years diminish the likelihood of their child being deceased at age one or younger and at age five and younger by 4.2 and 5.9 percentage points, respectively. Moving to Panel B, a one-percent increase in maternal schooling results in a reduced probability of her child being deceased at age one or younger and at age five and younger by 4.2 and 5.3 percentage points, respectively. These results consistently reinforce our main findings in Table 4.

Second, in response to widespread challenges associated with late enrollment and repetition rates, prior research typically eliminates the partially treated group and undertakes the analysis using exclusively the fully treated and untreated samples (Grépin & Bharadwaj, 2015; Masuda & Yamauchi, 2020). Adhering to this methodological approach, we replicate our primary findings by excluding the partially treated group, and the estimated results are detailed in Table 6. Notably, a persistent trend is evident, demonstrating a consistently negative and statistically significant influence of maternal education on child mortality rates.

Third, as previously stated in Section 4.1, we initially employed a probability scale ranging from 0 to 1 for the partially treated group to mitigate the prevalent challenges associated with late enrollment and repetition rates in Kenya. However, in our revised methodology, we no longer consider this issue. Instead, we categorize women exposed to the educational system changes with a value of 1, while those unexposed receive a value of 0. The findings are detailed in Table 7. We consistently observe negative and statistically significant impacts of maternal education on two measures of child mortality, specifically, the probability of their child's death before reaching the age of one or younger, and before reaching the age of five.

Moreover, when computing the original IV, we utilize the pre-reform information percentages of repetition and advancement to lower secondary education from

Table 5 Alternative measures of female education

	Dead at age one or younger	Dead at age five or younger
a) Completion of at least 8 years of schooling	-0.042** (0.021)	-0.059*** (0.022)
<i>Centered R²</i>	0.018	0.022
<i>Observations</i>	22,489	22,489
<i>Weak id.</i>	434.768	434.768
b) Logarithm of female education	-0.042*** (0.013)	-0.053*** (0.014)
<i>Centered R²</i>	0.009	0.009
<i>Observations</i>	22,489	22,489
<i>Weak id.</i>	128.263	128.263

The sample includes female respondents born from 1950 to 1980. All estimations include control variables of child's gender, mother's ethnicity, age of the mother, status for rural resident, and regional areas. Standard errors are in parentheses and clustered at the birth cohort

** Significant at 5 per cent; *** significant at 1 per cent. Weak id. refers to Kleibergen-Paap rk Wald F statistics

Table 6 Exclusion of the partially treated group

	Dead at age one or younger	Dead at age five or younger
Years of schooling	-0.010*** (0.003)	-0.012*** (0.003)
<i>Centered R²</i>	0.012	0.012
<i>Observations</i>	15,452	15,452
<i>Weak id.</i>	199.127	199.127

The sample includes female respondents born from 1950 to 1980. All estimations include control variables of child's gender, mother's ethnicity, age of the mother, status for rural resident, and regional areas. Standard errors are in parentheses and clustered at the birth cohort

*** Significant at 1 per cent. Weak id. refers to Kleibergen-Paap rk Wald F statistics

Table 7 Alternative measure of IV: exclusion of late enrollment and grade repetition

	Dead at age one or younger	Dead at age five or younger
Years of schooling	-0.008** (0.004)	-0.012*** (0.004)
<i>Centered R²</i>	-0.005	-0.013
<i>Observations</i>	22,489	22,489
<i>Weak id.</i>	85.956	85.956

The sample includes female respondents born from 1950 to 1980. All estimations include control variables of child's gender, mother's ethnicity, age of the mother, status for rural resident, and regional areas. Standard errors are in parentheses and clustered at the birth cohort.

** Significant at 5 per cent; *** significant at 1 per cent. Weak id. refers to Kleibergen-Paap rk Wald F statistics

Somerset (2007) and Ohba (2009). We now employ the fraction of individuals who completed at least 8 years of schooling from the Kenya Census 1999 — which is post-reform information — as our alternative IV. As a robustness check, we employ the same information from the Kenya Census 2009. The results are outlined in Table 8 and are similar to those in the main findings.

Table 8 Alternative measure of IV: using post-reform information from Kenya Census

	Dead at age one or younger	Dead at age five or younger
Panel A: Utilizing cohort-level data on the fraction of individuals who completed at least 8 years of schooling (Kenya Census 1999)		
Years of schooling	-0.007** (0.003)	-0.008*** (0.003)
Centered R^2	0.017	0.021
Observations	22,489	22,489
Weak id.	287.350	287.350
Panel B: Utilizing cohort-level data on the fraction of individuals who completed at least 8 years of schooling (Kenya Census 2009)		
Years of schooling	-0.005** (0.002)	-0.007*** (0.003)
Centered R^2	0.019	0.023
Observations	22,489	22,489
Weak id.	363.015	363.015

The sample includes female respondents born from 1950 to 1980. All estimations include control variables of child's gender, mother's ethnicity, age of the mother, status for rural resident, and regional areas. Standard errors are in parentheses and clustered at the birth cohort

** Significant at 5 per cent; *** significant at 1 per cent. Weak id. refers to Kleibergen-Paap rk Wald F statistics

In addition, to address cohort heterogeneity, we incorporate data on gender attitudes from World Values Survey Wave 6 into our main analyses.⁷ We utilize responses to the question, "Is a university education more essential for boys than for girls?" as a measure of gender attitude, where a value of 1 signifies disagreement with the statement and 0 otherwise. The outcomes are detailed in Table 9 and align with main findings.

Furthermore, our analysis integrates the control function approach and utilizes the bi-probit and iv-probit techniques to obtain our findings. In the control function approach, following the methodology of Rivers & Vuong (1988), we conduct an OLS regression in the first stage estimation, where female education is regressed on our instrument and explanatory variables. Subsequently, we calculate residuals and apply a probit model to regress child mortality indicators. This model incorporates maternal education, the residuals, and explanatory variables. In the alternative approach, we employ the bi-probit and iv-probit methods to handle the binary measures of child mortality rates. The outcomes of these analyses are detailed in Table 10. Our results consistently highlight a negative impact of maternal education on the likelihood of a child being deceased at age one or younger and at age five or younger.

⁷ We also include gender attitude in the model using the new instrumental variable, where 1 is assigned for the treated, and 0 otherwise. Additionally, we incorporate this control for another instrumental variable that utilizes post-reform information from the Kenya Census 1999 and 2009. The results remain robust (See Appendix Table 16).

Table 9 Inclusion of gender attitude

	Dead at age one or younger	Dead at age five or younger
Years of schooling	-0.009** (0.004)	-0.013*** (0.004)
Centered R^2	-0.008	-0.016
Observations	22,489	22,489
Weak id.	126.482	126.482

The sample includes female respondents born from 1950 to 1980. All estimations include control variables of child's gender, mother's ethnicity, age of the mother, status for rural resident, regional areas and gender attitude. Standard errors are in parentheses and clustered at the birth cohort

** Significant at 5 per cent; *** significant at 1 per cent. Weak id. refers to Kleibergen-Paap rk Wald F statistics

Table 10 Different estimations

	Dead at age one or younger	Dead at age five or younger
Panel A: Control function approach using Rivers & Vuong (1988)'s method		
Years of schooling	-0.008*** (0.002)	-0.010*** (0.002)
Observations	22,489	22,489
Panel B: Bi-probit estimation		
Years of schooling	-0.004*** (0.001)	-0.005*** (0.001)
Observations	22,489	22,489
Panel C: Iv-probit estimation		
Years of schooling	-0.070*** (0.022)	-0.082*** (0.020)
Observations	22,489	22,489

The sample includes female respondents born from 1950 to 1980. All estimates are average marginal effects. All estimations include control variables of child's gender, mother's ethnicity, age of the mother, status for rural resident, and regional areas. Standard errors are in parentheses and clustered at the birth cohort

*** Significant at 1 per cent

Finally, the crucial condition of meeting the exclusion restriction, a prerequisite for instrumental variables, is susceptible to violations in the presence of other factors influencing children's mortality rates. To address the endogeneity challenge, we apply Lewbel's (2012) 2SLS approach, a widely employed methodology in empirical studies for assessing the robustness of outcomes using external instruments (Chatterji et al., 2014; Ma, 2019; Tranchant et al., 2020; Wang & Cheng, 2022). This method relies on heteroskedastic covariance restrictions to create internal instrumental variables, thereby mitigating the need to strictly adhere to the exclusion restriction.

The results derived from Lewbel's (2012) approach are outlined in Table 11. We present two sets of Lewbel 2SLS results. The first exclusively relies on internal

Table 11 Lewbel 2SLS results

	Internal IV	Both internal and external IV
Dead at age one or younger	-0.002*** (0.001)	-0.002*** (0.001)
<i>Control variables</i>	Yes	Yes
<i>Number of observations</i>	22,489	22,489
<i>Weak id.</i>	37.429	49.070
Dead at age five or younger	-0.003*** (0.001)	-0.003*** (0.001)
<i>Control variables</i>	Yes	Yes
<i>Number of observations</i>	22,489	22,489
<i>Weak id.</i>	37.429	49.070

The sample includes female respondents born from 1950 to 1980. All estimations include control variables of child’s gender, mother’s ethnicity, age of the mother, status for rural resident, and regional areas. Standard errors are in parentheses and clustered at the birth cohort

*** Significant at 1 per cent. Weak id. refers to Kleibergen-Paap rk Wald F statistics

instruments, while the second integrates the probability of exposure to the policy change with internally generated instruments. These results further reinforce our main findings that maternal education has adverse effects on the likelihood of children being deceased at age one or younger and at age five or younger in Kenya.

5.2 Plausibly exogenous IV

Even though our IV, signifying the probability of exposure to the policy (pr_iv_b), displays a noteworthy association with our main variable ($maternal_educ_{ib}$), the ability to rigorously test the exclusion restriction is hindered by the intrinsic unobservability of the error term. To alleviate this challenge, we adopt the methodological framework known as the union of all confidence intervals (UCI), as originally introduced by Conley et al. (2012). This strategy is applied to assess our key results, considering a relaxation of the exclusion restriction.⁸ In the traditional IV regression, it is posited that the influence of the IV on the dependent variable is exclusively mediated through the main variable, as indicated by $\delta = 0$ in Eq. 4b provided below:

$$maternal_educ_{ib} = \gamma_0 + \gamma_1 pr_iv_b + \gamma_2 X_{ib} + \epsilon_{fib} \tag{4a}$$

$$mortality_{ib} = \lambda_0 + \beta maternal_educ_{ib} + \delta pr_iv_b + \lambda_1 X_{ib} + \theta_{ib} \tag{4b}$$

Nonetheless, if the IV (pr_iv_b) is considered plausibly exogenous ($\delta \approx 0$), it suggests a potential, albeit marginal, direct association between the IV and $mortality_{ib}$. Specifically, we allow the direct impact of pr_iv_b on $mortality_{ib}$ (δ) to vary within a predefined interval, thereby constructing a credible range of potential values (β) for the influence of $maternal_educ_{ib}$ on $mortality_{ib}$.

⁸ This approach has been widely utilized in prior scholarly studies (Azhgaliyeva et al., 2023; Clarke & Matta, 2018; Dittmar & Meisenzahl, 2020; Guo, 2020; Karadja & Prawitz, 2019; Le & Azhgaliyeva, 2023; Mariella, 2023; Nguyen-Phung & Le, 2023; Yang et al., 2023).

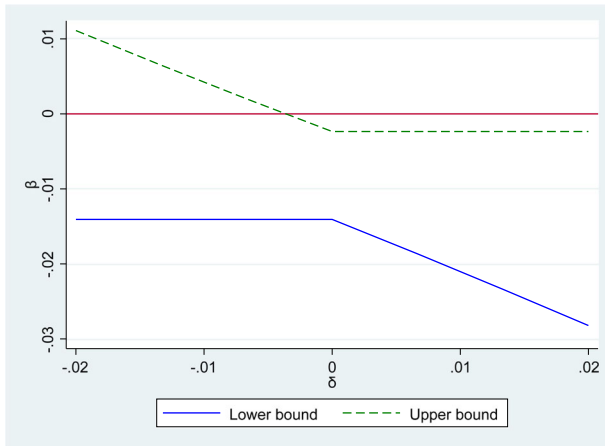


Fig. 2 Plausibly exogenous IV: The impacts of maternal education on child mortality rate (dead at age one or younger) using the UCI approach of Conley et al. (2012)

Figure 2 displays the estimated effects of maternal education on child mortality rates (dead at age one or younger) when the IV, the probability of exposure to the policy, is plausibly exogenous. The solid blue and dashed green lines indicate the lower and upper bounds of the 95% confidence interval for maternal education, respectively. In general, the 2SLS estimates regarding the effects of maternal education on the likelihood of a child's death at age one or younger fall within the 95% confidence intervals computed for different values of δ .

More precisely, if δ assumes a positive value, the 95% confidence interval for maternal education coefficient (β) does not include 0. This implies that maternal education has a negative and statistically significant impact on child mortality rates when a greater probability of exposure to the policy (IV) is correlated with elevated child mortality rates. It is worth noting that the 95% confidence interval for β exclusively includes 0 when δ is less than -0.004 . It is noteworthy that the reduced-form estimate indicates an approximate impact of -0.014 on a child's death at age one or younger as a result of exposure to the policy.⁹ Therefore, the selected threshold of -0.004 , which encapsulates roughly 30% of the total direct influence of IV on the likelihood of a child's death at age one or younger, is considered suitable.¹⁰ To conclude, our results consistently affirm the existence of a negative and statistically significant influence of maternal education on the probabilities of a child's death at age one or younger, even when accommodating a substantial deviation from the assumption of perfect exogeneity.¹¹

⁹ Interested readers can refer to Table 17 of Appendix C for further details.

¹⁰ Note that Guo (2020) allows for slightly smaller deviations, amounting to 25% of the reduced form estimated coefficient, from the state of perfect exogeneity. Similarly, Nguyen-Phung & Le (2023) allow for comparable departures from perfect exogeneity.

¹¹ Similar results are derived for a child's death at age five or younger. Figure 3 of Appendix D displays the result.

5.3 Potential pathways

Table 12 investigates the underlying paths contributing to the nexus between female education and their children's mortality rates. The first pathway involves fertility. As indicated in Panel A of Table 12, a one-year increase in female schooling extends the age at first marriage by 0.5 years and age at first intercourse by 0.4 years, similar to

Table 12 Potential mechanisms

Panel A: Fertility					
	Age at first marriage (1)	Age at first intercourse (2)	Age at first birth (3)	Number of children ever born (4)	
Years of schooling	0.504*** (0.049)	0.407*** (0.037)	0.358*** (0.046)	-0.471*** (0.028)	
Centered R^2	0.250	0.242	0.254	0.509	
Observations	14,908	15,879	15,933	15,933	
Weak id.	171.788	175.834	175.175	175.175	
Panel B: Labor market engagement					
	Currently working (1)	Worked in the last 12 months (2)			
Years of schooling	0.058*** (0.010)	0.051*** (0.007)			
Centered R^2	0.006	0.048			
Observations	10,039	13,862			
Weak id.	102.959	146.706			
Panel C: Health-seeking behaviors for children					
	Smoking (1)	Delivery in hospital (2)	Has child been vaccinated (3)		
Years of schooling	-0.028*** (0.006)	0.046*** (0.006)	0.032*** (0.006)		
Centered R^2	-0.139	0.241	0.032		
Observations	13,099	22,199	8,646		
Weak id.	87.817	172.761	138.288		
Panel D: Measures of women's involvement in decision making					
	Her health decision (1)	Large purchase decision (2)	Daily purchase decision (3)	Family visits decision (4)	Food purchase decision (5)
Years of schooling	0.125*** (0.015)	0.183*** (0.021)	0.144*** (0.021)	0.091*** (0.015)	0.050*** (0.009)
Centered R^2	-0.469	-1.086	-0.743	-0.192	-0.214
Observations	6572	6571	4929	6571	6570
Weak id.	90.880	90.657	60.931	90.640	90.937

The sample includes female respondents born from 1950 to 1980. All estimations include control variables of child's gender, mother's ethnicity, squared age of the mother, status for rural resident, and regional areas. Standard errors are in parentheses and clustered at the birth cohort

*** Significant at 1 per cent. Weak id. refers to Kleibergen-Paap rk Wald F statistics

Grépin & Bharadwaj's (2015) results for Zimbabwe, which are approximately 0.52 and 0.63 years, respectively. The timing of the first birth is increased by 0.3 years, consistent with Le & Nguyen's (2020) results of 0.11 years in 68 developing countries across five continents. Additionally, an extra year of schooling reduces the number of children ever born by approximately 0.5, which is in line with Samarakoon & Parinduri's (2015) result of around 0.3 live births. The second pathway focuses on labor market engagement. The results in Panel B reveal that an additional year of education increases the likelihood of having a job and working in the last 12 months by 5.8 and 5.2 percentage points, respectively, consistent with the findings of Le & Nguyen (2020) and Aslam & Kingdon (2012). The third pathway involves the mother's health-seeking behaviors for her children. Results in Panel C demonstrate that a one-year increase in female schooling decreases the mother's likelihood of smoking by 2.8 percentage points and improves her chances of birth delivery in a hospital and vaccinating her child by 4.6 and 3.2 percentage points, respectively, consistent with Makate & Makate's (2016) finding of a 52.5 percentage point improvement in prenatal care visits and Grépin & Bharadwaj's (2015) result of a 5.3 percentage point increase in cesarean section deliveries. The last pathway is associated with women's involvement in decision-making. We measure her decision-making in the household via five binary indicators, with 1 indicating that she is involved in the decision-making process and 0 otherwise. Panel D results show that an extra year of female schooling enhances her decision-making in health, large purchases, daily purchases, family visits, and food purchases by approximately 12.5, 18.3, 14.4, 9.1, and 5 percentage points, respectively, consistent with previous studies (Amugsi et al., 2016; Bhagowalia et al., 2012; Keats, 2018; Nguyen et al., 2013; Smith et al., 2003).

6 Conclusions and policy recommendations

This paper investigates the impacts of maternal education on her children's mortality rates in Kenya. Utilizing five waves of nationally representative data from KDHS, we explore how a mother's education influences her offspring's health and development. This impact is not only supported by the MDG and SDG frameworks but also substantiated by extant literature (Grytten et al., 2014; Lundborg et al., 2014; Robinson et al., 2007; Nguyen-Phung, 2023). Education, particularly the educational obtainments of mothers, emerges as a critical determinant in enhancing the well-being of children. This impact aligns with the broader global agenda, notably SDG3, which aspires to ensure healthy lives and promote well-being for all at every stage of life. Recognizing the pivotal role that maternal education plays in achieving these ambitious goals, our research seeks to contribute crucial perspectives to the ongoing discourse on improving child health outcomes in Kenya.

We utilize the educational reform of 1985 as a source of exogenous variation in female education, providing insights into the impact of maternal schooling on the probability of child mortality at age one or younger and age five or younger. Drawing data from the five waves of the KDHS conducted in 1993, 1998, 2003, 2008, and 2014, we apply a two-stage least-squares (2SLS) approach. Our findings

illustrate that women exposed to the 1985 policy change, on average, have approximately 1.87 more years of schooling than their counterparts. Moreover, an additional year of maternal schooling results in a reduction in the likelihood of a child's death at age one or younger and age five or younger by 0.6 and 0.9 percentage points, respectively. These results remain robust across a series of robustness checks. Additionally, we explore various potential mechanisms elucidating the impact of maternal education on child mortality. These mechanisms include examining fertility behavior, the likelihood of maternal engagement in the labor force, maternal health-seeking behaviors for children and maternal involvement in household decision-making.

Our findings are in line with previous studies which signify that education is a key factor through which women can reduce the mortality rate of their children (Andriano & Monden, 2019; Breierova & Duflo, 2004; Gakidou et al., 2010; Grépin & Bharadwaj, 2015; Le & Nguyen, 2020; Makate, 2016; Makate & Makate, 2016; Nguyen-Phung, 2023). We further confirm the mechanisms through which maternal education influences various aspects of her life, including fertility behavior (Le & Nguyen, 2020; Samarakoon & Parinduri, 2015), labor participation (Le & Nguyen, 2020; Aslam & Kingdon, 2012), health-seeking behaviors for children (Andriano & Monden, 2019; Grépin & Bharadwaj, 2015; Makate & Makate, 2016; Makate & Makate, 2016), and involvement in decision-making (Amugsi et al., 2016; Bhagowalia et al., 2012; Keats, 2018; Nguyen et al., 2013; Smith et al., 2003), leading to an improvement in their children's survival rate.

Our findings have crucial policy implications. First, considering the inter-generational effects of maternal education on their children's health and development, educational system reforms or programs that give special attention to enhancing access to school or ensuring extended years of schooling for females are essential for promoting the health of the next generation. Second, awareness campaigns and accessible healthcare services should encourage maternal health-seeking behaviors, including smoking cessation, hospital deliveries, and child vaccinations, contributing to positive health practices. Finally, initiatives should be implemented to enhance maternal involvement in household decision-making, including awareness programs and policies that promote gender equality within households, ensuring that women have a voice in decisions affecting their children's well-being.

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Compliance with ethical standards

Conflict of interest The authors declare no competing interests.

7 Appendix A

In this section, we provide detailed computations to ascertain the probability associated with the partially treated group. While the legal age for a student enroll in primary school is six years old, late entry often leads to variations in enrollment age, thereby introducing a substantial source of measurement error in the assignment variable. Utilizing data sourced from Somerset (2007) delineating the distribution of entry ages for Grade one, a conditional probability of entering the primary school level is computed specifically for the age range of 6 to 10.

Kenya exhibits a noteworthy and prevalent occurrence of repetition in primary school level, permitting students to prolong their study duration beyond the seven-year schedule stipulated by the old system (Somerset, 2009). To mitigate this issue, data on repetition rates from Somerset (2007) and transition rates to secondary school level from Ohba (2009) are gathered. Afterwards, we designate a probability to each birth cohort, demonstrating the possibility of repeating a grade.¹² Considering the two conditional probabilities, the computation of the probability pertaining to the partially treated cohorts is presently undertaken in the following manner:

$$Pr(\text{partially treated}) = 1 - Pr(\text{completion of primary school prior to 1985})$$

where the probability of completion of primary school is calculated as follows:

$$\begin{aligned} Pr(\text{completion of primary school prior to 1985}) &= \sum_{s=(y+12)}^{1983} Pr(b \cap s) \\ &= \sum_{en=6}^{(1977-y)} \sum_{r=0}^{(1977-y-en)} Pr(\text{primary entry} = en|b) \times Pr(\text{primary completion}|repetition = r) \end{aligned}$$

Here, s , y , en , and r denote the schooling cohort associated with an individual's completion of primary education, an individual's birth year, the entry age to Grade 1, and the number of the repeated year(s), respectively.

8 Appendix B

Tables 13, 14

The baseline OLS findings, detailed in Appendix Table 13, highlight a statistically significant correlation between maternal education and the likelihood of a child experiencing death at age one or younger, as well as death at age five or younger. To be specific, higher maternal schooling is linked to a decrease of approximately 0.03 percentage points in both the chances of a child being deceased at age one or younger and deceased at age five or younger.

Additionally, our analysis addresses concerns pertaining to potential omitted variable bias through the application of Oster's (2019) bounding approach, a method extensively employed in the literature (Bevis et al., 2023; Deshpande and

¹² It is noteworthy that we do not account for the dropout rate due to a scarcity of data available within this designated timeframe.

Table 13 OLS estimate

	Dead at age one or younger	Dead at age five or younger
Years of schooling	-0.003*** (0.001)	-0.003*** (0.001)
R^2	0.020	0.025
Observations	22,489	22,489

The sample includes female respondents born from 1950 to 1980. All estimations include control variables of child’s gender, mother’s ethnicity, age of the mother, status for rural resident, and regional areas. Standard errors are in parentheses and clustered at the birth cohort

*** Significant at 1 per cent

Table 14 Oster (2019) bound estimates

	$[\tilde{\beta}, \beta^*_{(R_{\max}=\min\{1.3\tilde{R}, 1\}, \delta=1)}]$ Identified sets (1)	Excludes 0? (2)	$\delta_{(R_{\max}=\min\{1.3\tilde{R}, 1\})}$ (3)	$\delta > 1$ (4)
Dead at age one or younger	[-0.0023, -0.0027]	Yes	10.022	Yes
Dead at age five or younger	[-0.0029, -0.0033]	Yes	9.087	Yes

The sample includes female respondents born from 1950 to 1980. All estimations include control variables of child’s gender, mother’s ethnicity, age of the mother, status for rural resident, and regional areas

Ramachandran, 2022; Fujii et al., 2018). This method evaluates the robustness of results against omitted variable bias, assuming that the connection between observed covariates and the treatment variable provides insights into the relationship between the treatment variable and unobservable factors.

Oster’s (2019) method assesses the influence of female education on children’s mortality rates across a range of values denoted as $\tilde{\beta}$ to β^* . Two critical pieces of information, δ (indicating the relative importance of observable versus unobservable variables in introducing bias) and R_{\max} (the R^2 from a model encompassing both observable and unobservable variables), play a pivotal role. Oster’s estimation assumes $\delta = 1$ and imposes constraints on $R_{\max} = \min\{1.3\tilde{R}, 1\}$, with \tilde{R} derived from empirical evidence. Our analysis calculates R_{\max} following Oster’s guidelines, determining $R_{\max} = \min\{1.3\tilde{R}, 1\}$.

Appendix Table 14 presents the empirical findings derived from Oster’s methodology. In Column 1, where we assess both measures of children’s mortality rates, the identified sets do not include zero. This illustrates the strength and reliability of our estimates even when unobserved variables are not considered. Moving to Column 3, the δ coefficient surpasses 1, which aligns with Oster’s (2019) recommended threshold. This indicates that omitted variables would need a more substantial impact than observed variables to negate the significant influence of maternal education on child mortality rates.

9 Appendix C

Tables 15–17

Table 15 The impact of female education on their child mortality rates with more controls

	Dead at age one or younger	Dead at age five or younger
Panel A: inclusion of survey fixed effects		
Years of schooling	-0.006** (0.003)	-0.009*** (0.003)
Centered R ²	0.018	0.021
Observations	22,489	22,489
Weak id.	198.345	198.345
Panel B: inclusion of quadratic trend of age trend		
Years of schooling	-0.007** (0.003)	-0.009*** (0.003)
Centered R ²	0.018	0.020
Observations	22,489	22,489
Weak id.	204.027	204.027
Panel C: inclusion of school attendance from Kenya Census 1989		
Years of schooling	-0.009** (0.004)	-0.012*** (0.004)
Centered R ²	-0.007	-0.015
Observations	22,489	22,489
Weak id.	88.351	88.351

The sample includes female respondents born from 1950 to 1980. All estimations include control variables of child’s gender, mother’s ethnicity, age of the mother, status for rural resident, and regional areas. Standard errors are in parentheses and clustered at the birth cohort

** Significant at 5 per cent; *** significant at 1 per cent. Weak id. refers to Kleibergen-Paap rk Wald F statistics

Table 16 Inclusion of gender attitude

	Dead at age one or younger	Dead at age five or younger
Panel A: Alternative measure of IV: do not account for late enrollment and grade repetition		
Years of schooling	-0.009* (0.004)	-0.013*** (0.005)
Centered R ²	-0.007	-0.017
Observations	22,489	22,489
Weak id.	74.776	74.776
Panel B: Alternative measure of IV: using post-reform information - utilizing cohort-level data on the fraction of individuals who completed at least 8 years of schooling (Kenya Census 1999)		
Years of schooling	-0.009** (0.003)	-0.011*** (0.004)
Centered R ²	-0.007	-0.012
Observations	22,489	22,489
Weak id.	206.884	206.884
Panel C: Alternative measure of IV: using post-reform information - utilizing cohort-level data on the fraction of individuals who completed at least 8 years of schooling (Kenya Census 2009)		
Years of schooling	-0.006** (0.003)	-0.008*** (0.003)
Centered R ²	0.017	0.021
Observations	22,489	22,489
Weak id.	291.712	291.712

The sample includes female respondents born from 1950 to 1980. All estimations include control variables of child’s gender, mother’s ethnicity, age of the mother, status for rural resident, regional areas and gender attitude. Standard errors are in parentheses and clustered at the birth cohort

** Significant at 5 per cent; *** significant at 1 per cent. Weak id. refers to Kleibergen-Paap rk Wald F statistics

Table 17 Reduced form estimation

	Dead at age one or younger	Dead at age five or younger
Exposure to the reform	-0.014** (0.006)	-0.016*** (0.006)
R^2	0.019	0.024
Observations	22,489	22,489

The sample includes female respondents born from 1950 to 1980. All estimations include control variables of child’s gender, mother’s ethnicity, age of the mother, status for rural resident, and regional areas. Standard errors are in parentheses and clustered at the birth cohort

* Significant at 10 per cent; ** significant at 5 per cent; *** significant at 1 per cents

10 Appendix D

Figure 3

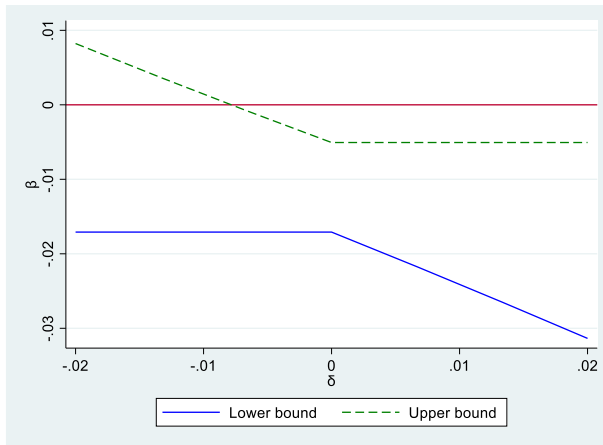


Fig. 3 Plausibly exogenous IV: The impacts of maternal education on their child’s mortality rate (dead at age five or younger) using the UCI approach of Conley et al. (2012)

This section presents the estimated effects of maternal education on child mortality rates (dead at age five or younger) when the IV, the probability of exposure to the policy, is plausibly exogenous. The results confirm the existence of a negative and statistically significant influence of maternal education on the likelihood of a child’s death at age five or younger, even when accommodating a substantial deviation from the assumption of perfect exogeneity.

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