DISCUSSION PAPER SERIES

DP15128

THE STOPPING RULE AND GENDER SELECTIVE MORTALITY: WORLD EVIDENCE.

Jean-Marie Baland, Guilhem Cassan and Francois Woitrin

DEVELOPMENT ECONOMICS

ISSN 0265-8003

THE STOPPING RULE AND GENDER SELECTIVE MORTALITY: WORLD EVIDENCE.

Jean-Marie Baland, Guilhem Cassan and Francois Woitrin

Discussion Paper DP15128
Published 03 August 2020
Submitted 30 June 2020

Centre for Economic Policy Research
33 Great Sutton Street, London EC1V 0DX, UK
Tel: +44 (0)20 7183 8801
www.cepr.org

THE STOPPING RULE AND GENDER SELECTIVE MORTALITY: WORLD EVIDENCE.

Abstract

The stopping rule in population studies refers to a behaviour by which parents continue child bearing till they have their desired number of boys. We first show that, under this rule, girls are, on average, exposed to a larger number of younger siblings than boys. This increased exposure to sibling competition may result in a higher mortality for girls, even in the absence of any other forms of discrimination. We then propose a new method to detect the prevalence of the stopping rule in a given society. This method allows us to identify countries in which the stopping rule prevails, some of which have been largely ignored in the literature. We also identify countries in which the stopping rule targets a desired number of girls rather than boys. We estimate the extent to which the stopping rule leads to a higher mortality among children through sibling competition. We show that this specific mechanism explains a non trivial share of mortality among young girls (for instance, 10% of the under 5 female mortality in India, and up to 35% in Armenia) and that this share is increasing over time.

JEL Classification: J13, O53

Keywords: son preference, target rule, stopping rule, missing girls

Jean-Marie Baland - jean-marie.baland@fundp.ac.be
University of Namur and CEPR

Guilhem Cassan - guilhem.cassan@unamur.be
university of namur

Francois Woitrin - francois.woitrin@unamur.be
university of namur

Acknowledgements
We thank Vincenzo Verardi and Paola Villar for useful suggestions. We thank the audiences of ULB, UNamur and the Indian Statistical Institute. Research on this project was financially supported by the Excellence of Science (EOS) Research project of FNRS O020918F.

The Stopping Rule and Gender selective mortality: World Evidence.*

Jean-Marie Baland†, Guilhem Cassan‡ and Francois Woitrin§

June 30, 2020

Abstract. The *stopping rule* in population studies refers to a behaviour by which parents continue child bearing till they have their desired number of boys. We first show that, under this rule, girls are, on average, exposed to a larger number of younger siblings than boys. This increased exposure to sibling competition may result in a higher mortality for girls, even in the absence of any other forms of discrimination. We then propose a new method to detect the prevalence of the stopping rule in a given society.

This method allows us to identify countries in which the stopping rule prevails, some of which have been largely ignored in the literature. We also identify countries in which the stopping rule targets a desired number of girls rather than boys. We estimate the extent to which the stopping rule leads to a higher mortality among children through sibling competition. We show that this specific mechanism explains a non trivial share of mortality among young girls (for instance, 10% of the under 5 female mortality in India, and up to 35% in Armenia) and that this share is increasing over time.

Keywords: Son preference, target rule, stopping rule, missing girls.

JEL classification: J13, O53

*We thank Vincenzo Verardi and Paola Villar for useful suggestions. We thank the audiences of ULB, UNamur and the Indian Statistical Institute. Research on this project was financially supported by the Excellence of Science (EOS) Research project of FNRS O020918F.

†CRED, DEFIPP, University of Namur.

‡CRED, DEFIPP, CEPREMAP, University of Namur.

§CRED, DEFIPP, University of Namur.

1 Introduction

Because of numerous forms of discrimination, women are at a disadvantage in terms of mortality relative to men. Building on Sen (1990)'s concept of missing women, Anderson & Ray (2010) for example show that in 2000 close to 5 million women were missing. These deaths could have been prevented if women were treated in the same manner as men. Among the reasons often put forward to explain this shocking pattern is a strong preference for sons, which often leads to active discrimination against girls. Many cultural factors may account for such discrimination including patrilocality (Ebenstein, 2014), old age support (Ebenstein & Leung, 2010; Lambert & Rossi, 2016), or the burden of the dowry (Arnold et al., 1998), among others (see Williamson (1976), Das Gupta et al. (2003) or Jayachandran (2015) for a detailed review of the various causes of preferences for sons).

In this paper, we explore a mechanism that explains a higher mortality among girls in the absence of active forms of discrimination. Building on the intuition of Arnold et al. (1998) and Ray (1998), we show that, where the preference for sons manifests itself under the form of a stopping rule - the continuation of childbearing until a given number of sons is obtained -, girls on average may suffer from higher mortality rates despite being treated in the same way as their brothers. In the words of Jensen (2003), which studied how the stopping rule may affect girls' education, girls may have "unequal outcomes" despite "equal treatment". The intuition is the following: because parents continue to have children until they reach their desired number of boys, girls end up having more younger siblings. If sibling competition leads to higher mortality, then, on aggregate, girls have a higher mortality rate than boys, despite being treated equally within the family.

We develop a simple formalization of the stopping rule which demonstrates that, in stopping rule countries, girls will have *more* younger siblings than boys, but the *same* number of older siblings. In other words, girls are *born* with the same number of siblings[1] as boys but may end up *dying* with more. Our formalization, while taking the perspective of the child rather than that of a family with a completed fertility, also replicates some well-known consequences of the stopping rule: total fertility may be higher than desired (Sheps, 1963), the total number of siblings is higher for girls than for boys (Yamaguchi, 1989; Basu & de Jong, 2010); the birth order of girls within families is, on average, lower for girls than for boys (Basu & de Jong, 2010) and the materialization of the stopping rule depends negatively on total desired fertility and positively on the total number of sons desired (Sheps, 1963). As a matter of fact, all these are direct consequences of girls having more younger siblings than boys under the stopping rule.

This formalisation also provides a simple method to identify countries in which the stopping rule prevails: it is based on detecting countries in which girls, at any birth rank, have more younger siblings than boys of the same

[1] And same gender composition.

rank.[2] Compared to other methods such as the sex ratio of the last born (Jayachandran, 2015), there is no need to refer to a natural sex ratio at birth, which has been shown to vary across time and space.(Chahnazarian, 1988; Waldron, 1998; Hesketh & Xing, 2006). Our method also allows to investigate families which have not completed their fertility and thereby reflect recent, instead of past, behaviors (Haughton & Haughton, 1998). As we will show, besides countries in South Asia and Northern Africa, many Central Asian and European countries do implement a stopping rule. Our approach being gender neutral, we also provide some evidence of a stopping rule in favour of girls in some, essentially African, countries.

Taking advantage of all the available DHS surveys, we first identify the countries in which the stopping rule prevails, and document the more intense sibling competition faced by young girls. We then estimate the impact of this competition on under 5 mortality. We measure this effect at the world level and show that, under the stopping rule, this mechanism explains a non trivial share of mortality among young girls in many countries even in the absence of any form of active discrimination against girls.[3] For instance, it accounts for close to 10% of under 5 female mortality in India, 16% in Nepal, 35% in Armenia and more than 20% in Azerbaidjan.

The practice of sex selective abortion has developed rapidly, particularly in Asia, as a method to obtain the desired gender composition in the family (Park & Ho, 1995; Arnold et al., 2002; Abrevaya, 2009; Jayachandran, 2017; Dimri et al., 2019). With respect to the stopping rule, parents can now directly interrupt pregnancies and control the gender of their children. While both sex selective abortion and the stopping rule have the same underlying causes, their demographic consequences are entirely different. Indeed, if parents fully control the gender of their children, they do not need to practice the stopping rule, and girls born under sex selective abortion are the result of a choice (Qian et al., 2014). Hence, these girls are not exposed to a stronger sibling competition. As a result, one may think that the widespread adoption of sex selective abortion should eradicate the stopping rule and its sibling competition consequences. This view is in fact not correct when sex selective abortion does not lead to a perfect control of the gender of the child, but simply affects the relative probability for each gender to be born. In that case, young girls will still be exposed to a more intense younger sibling competition, even though the difference with young boys will be mitigated. As a matter of fact, we find that the share of deaths explained by our mechanism is increasing over time.

The structure of the paper is as follows: we first present our formalization of the stopping rule. Following the method developed in that section, we identify the countries in which the stopping rule prevails. For this set of countries, we then measure at the country level the excess female mortality that is due to the increased sibling competition generated by the stopping rule. Finally, we investigate the evolution of the phenomenon over time.

[2] As discussed later in the paper, this method can only identify countries in which the stopping rule has an impact on the observed fertility behaviour.

[3] We of course acknowledge that girls are exposed to severe discrimination (Barcellos et al., 2014). What we propose here is to explore a specific channel that may explain part of the higher mortality of girls in the absence of active discrimination conditional on being born. The stopping rule in itself is a form of discrimination (fertility behavior differs by gender), but our approach measures the extent to which the mortality of girls can be explained despite being treated in the same manner as their brothers, once they are born.

2 The demographic consequences of the stopping rule

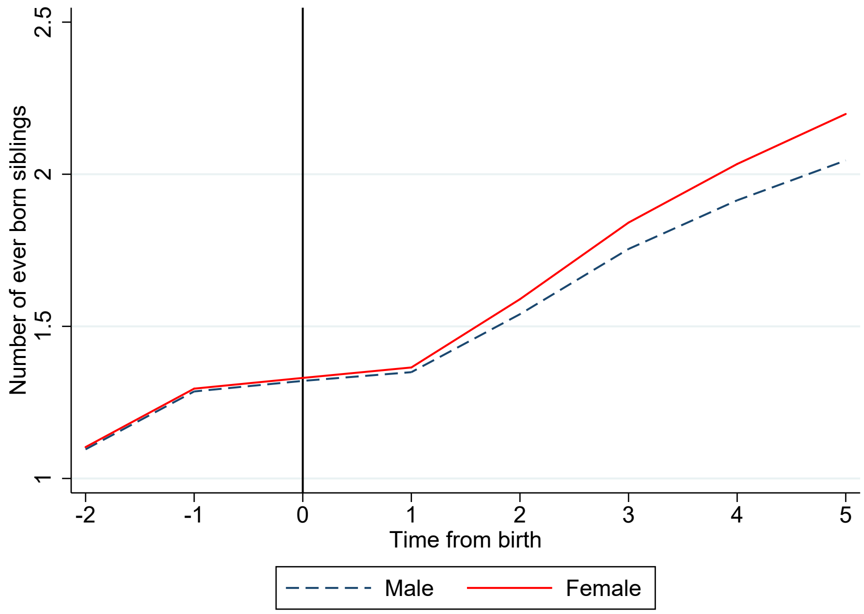
2.1 Gender at birth is a lottery

The stopping rule refers to a behavior by which parents cease childbearing once they reach the number of children of a specific gender they desire (sons, in general). The theoretical literature in demography has extensively studied the consequences of this behavior (e.g Sheps (1963); Yamaguchi (1989); Clark (2000); Basu & de Jong (2010)) by focusing on outcomes at the family level, such as total fertility or sex ratios among children. Our approach differs by taking the perspective of an individual child. While essentially replicating the results from the literature, this perspective drastically simplifies the modelling effort and delivers more precise empirical predictions.

The main intuition of our model goes as follows: suppose that the only reason why parents have children is to reach a desired number of boys. Each child is considered as a draw in a lottery in which having a boy is a "success", while having a girl is a "failure". When a boy is "drawn", parents are one unit closer to their objective. When a girl is "drawn", parents have made no progress, and additional children (draws) will be required in order to compensate for this failed attempt. This is true at each birth rank. Therefore, *at any birth rank, compared to a boy, a girl is a failed lottery draw, which does not contribute to reaching the desired number of boys. As a result, a girl of a particular rank will have exactly the same number of younger siblings as a boy of the same rank plus the expected number of additional draws required to have the boy that she is not.* In this respect, the death of a boy has the same consequences as the birth of a girl: it is a birth that does not contributes towards the desired number of boys, and leads to additional births. Under the stopping rule, the birth of a girl is equivalent to the death of a boy.

Figure 1 illustrates the model's main prediction with Indian data, India being a country in which the stopping rule is considered as pervasive. For all children who are at least 5 years old at the time of the survey, we have computed, at each age between -2 and 5, the average number of ever-born siblings for boys and girls separately. Before being born (at age -2 to 0), Indian boys and girls have the same number of (ever-born) older siblings. It is only after their births that the number of siblings for a girl becomes higher than for a boy, the more so the older they get. Note that if a child is not yet born (negative ages), she can only have elder siblings but when she is born (positive ages), her siblings will be either older or younger siblings. Therefore, the fact that the divergence in the number of ever born siblings emerges only after the child is born indicates that the divergence is driven *only* by younger siblings.

Figure 1: Number of ever-born siblings by age and gender in India



Data source: DHS India 2015, all children aged 5+ at the time of the survey.
Reading: at age 5, the average Indian boy has 2.05 ever-born siblings and the average Indian girl has 2.2 ever-born siblings.

2.2 The stopping rule with an unlimited number of children

We now formally investigate the impact of the stopping rule on the family structure. To this end, let us first assume that couples want to have a given number b^* of boys, and can have an unlimited number of children. The probability of having a boy at each birth is given and equal to p. At any birth rank, parents have p chances to have a boy and $(1-p)$ chances to have a girl. As a result, in a 'large' population and at any birth rank, for each male birth, there is exactly $\frac{1-p}{p}$ female birth. In other words, the (male to female) sex ratio at any rank in this population is constant and equal to $\frac{p}{1-p}$. As a result, over all ranks, the 'stopping rule' has no effect on the sex ratio at birth (Sheps, 1963). (Of course, by its very definition, the stopping rule determines the gender and therefore the sex ratio of the last born.)

By definition, a boy and a girl of a given birth rank k have the same number $k-1$ of elder siblings. The distribution of the sex composition of these elder siblings is also identical, as the sex composition of all siblings prior to rank k is independent of the sex of the child of rank k. For instance, if we assume that parents want at least 3 boys and focus on a child at rank 3, there are four possible combinations of the elder siblings: $(girl-girl, girl-boy, boy-girl, boy-boy)$. These events occur with probability $\big((1-p)^2, (1-p)p, p(1-p), p^2\big)$, which is the distribution faced by a child at rank 3, independently of whether she is a girl or a boy.

As a result, the only difference boys and girls of the same rank can face under the stopping rule comes from their younger siblings. The critical difference between the two is the fact that having a boy implies that parents are one unit closer to their desired number of boys. A girl has more younger siblings as one more boy among them is needed to compensate. In expected terms, this implies that parents need to have $1/p$ more children to make up for this difference in gender at rank k. Since this is true at any rank, a girls is expected to have $1/p$ more younger siblings than a boy. In particular, if $p = 1/2$, a girl will, on average, have 2 more younger siblings than a boy.

Note that the reasoning above does not exclude the practice of sex selective abortion, as long as it is not perfect. Sex selective abortion can be interpreted as an increase in p: for a higher p, girls still have more younger siblings than boys, but the difference between the two will be reduced. As p tends towards 1, i.e. parents perfectly choose the sex of their child and the number of additional younger siblings of girls tends towards 1. Therefore, the practice of sex selective abortion is not a theoretical problem for our approach, but will lead to increased difficulty in detecting the stopping rule empirically.

2.3 The stopping rule with a limited number of children

We have assumed that parents could have an unlimited number of children. We now assume, more realistically, that the number of children in a family cannot exceed a maximum, $\overline{N}$, which is exogenously given. This additional constraint implies that some families will not reach their desired number of boys, and this occurs more frequently at lower values of $\overline{N}$.

Consider a child of rank k and of gender $i = b, g$ who has e older brothers, with $e \leq k - 1$ and $e + 1 \leq b^*$. The last inequality indicates that the family has not yet reached her desired number of boys, b^*, before having a child of rank k . We denote by $E\left(Y_i\left(k, e\right)\right)$ his or her expected number of younger siblings. Generalizing the reasoning developed in the previous section, we have:

Proposition 1: At any rank k,with $k < \bar{N}$ and for any number of elder brothers e, with $e + 1 \leq b^*$, the expected number of younger siblings is strictly larger for a girl than for a boy:

$$E\left(Y_g\left(k, e\right)\right) > E\left(Y_b\left(k, e\right)\right),\ \forall k < \bar{N}, e \leq b^* - 1.$$

Proof: See Appendix A

The existence of a constraint on family size does not change our main result: at a given rank (smaller than $\bar{N}$), girls always have more younger siblings than boys. Note that, as the proposition holds for each rank, we also have, by summing over all ranks, that a girl on average has a larger expected number of younger siblings. It is easy to show that this difference is increasing in the maximum family size. More precisely, for a given number of desired boys, b^*, the difference in the expected number of younger siblings, at any rank $k < \bar{N}$, is monotonically increasing in $\bar{N}$. Conversely, for a given $\bar{N}$, it is also monotonically decreasing in the number of desired boys, b^*. Relatedly, the (male to female) sex ratio of the last born monotonically increases with the maximum number of children, $\bar{N}$.

2.4 The stopping rule with a desired family size

Some demographers follow a slightly different approach than the one discussed above (see Sheps (1963), for example). While they still assume that parents desire a given number of boys, b^*, parents also have a preference over their total number of children, n^*, which corresponds to their ideal family size. If, with n^* children, they don't have b^* boys, they continue to have children till they reach their desired number of boys. In other words, these parents have lexicographic preferences in n^* and b^*, with $0 < b^* \leq n^*$. To analyze this alternative model, we first assume away a constraint on the maximum number of children so that parents, if needed, have as many children as they need to reach the desired number of boys.

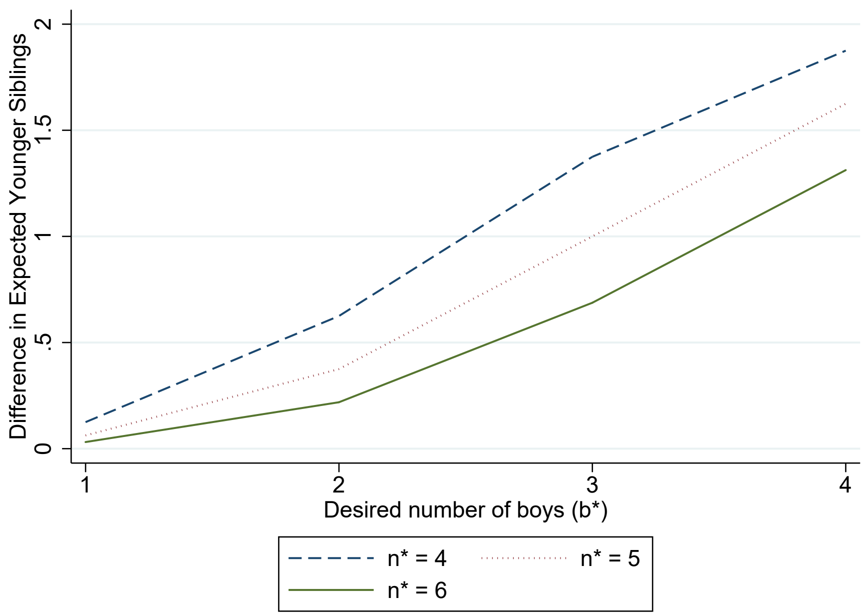
Consider first a family that succeeds in having at least b^* boys with n^*children. In such families, at any rank k, girls and boys have exactly the same number of younger siblings, which is equal to $(n^* - k)$. The proportion of such families in a large population is equal to the probability of having at least b^* 'successes' (boys) in n^* trials (children), which we denote as above $\sum_{j=b^*}^{n^*} B(j, n^*)$. All other families need more than n^*children to reach their desired number of boys. In such families, at any rank k, a girl will have $1/p$ more younger siblings than a boy, $1/p$ corresponding to the expected number of children necessary to have one extra boy. The proportion of such families is given by $\sum_{j=0}^{b^*-1} B(j, n^*)$. We therefore have:

Proposition 2: In families with lexicographic preferences over (b^*, n^*), at any rank, girls have in expected terms $\left(\frac{1}{p}\sum_{j=0}^{b^*-1} B(j, n^*)\right)$ more younger siblings than boys of the same rank.

A direct consequence of this proposition is that a girl on average (i.e., over all ranks) will also have $\left(\frac{1}{p}\sum_{j=0}^{b^*-1} B(j, n^*)\right)$ more younger siblings than a boy. A closer examination of this expression is illustrated in Figure 2: the difference in the expected number of younger siblings is larger for a smaller desired family size and for a larger desired number of boys. Note for example how, for a given desired number of boys an increase in the ideal family size leads to a decrease in the difference in younger siblings (a change in curve). Not also how, for a given ideal family size, an increase in the number of desired boys increases the difference in younger siblings (a change along the curve). As a result, it is likely that societies undergoing a demographic transition display a stronger differential in younger siblings than societies characterized by larger family sizes (Jayachandran, 2017), provided the desired number of boys does not vary too much.

Finally, imposing a constraint on family sizes in this setting does not change our main results. Assume again that family size cannot exceed a given level $\bar{N}$. Clearly, this constraint is only binding for families that needed more than n^* children to have their desired number of boys, b^*. Among this subset however, Proposition 1 above applies. More precisely, at any rank $k > n^*$,with $n^* < k < \bar{N}$ and for any number of elder brothers e, with $e \leq b^* - 1$, the expected number of younger siblings is strictly larger for a girl than for a boy.

Figure 2: Difference in expected number of younger siblings between girls and boys with lexicographic preferences in b^*



3 Identifying countries implementing the stopping rule

A large literature investigates the consequences of the stopping rule, particularly in terms of fertility, but focuses on particular cases, such as India or South Korea, in which its practice is widespread (Sheps, 1963; Arnold, 1985; Das Gupta, 1987; Yamaguchi, 1989; Arnold et al., 1998; Clark, 2000; Basu & de Jong, 2010; Jayachandran, 2017; Jayachandran & Pande, 2017). To broaden the focus to a wider and more general level, we need a systematic way to identify countries implementing the stopping rule, without relying on partial evidence. The empirical measure we present in this section serves this purpose. As implied by our theoretical approach, it is based, at the child level, on the number of younger siblings by gender. In the following, we briefly describe our data and discuss traditional approaches of the stopping rule before introducing our measure.

3.1 The data

For our main analysis, we use each country's most recent available Demographic and Health Survey (DHS). This represents 82 countries surveyed between 1985 and 2018, with observations on 1,255,711 mothers and 4,037,609 births. Note that because the survey year can widely differ across countries, results can not easily be compared across country, since the period they cover is widely different. These data can be interpreted as the most recent information available for each country. In the last section of this paper, when analyzing the evolution over time, we need to have comparable time periods for all countries studied. We select countries for which the DHS surveys offer information on births in the 1980s, 1990s and 2000s. This represents 68 countries, 2,533,137 mothers and 7,975,046 births covering 3 decades. Note that while the number of countries covered has decreased, the number of mothers and births covered has increased. This is because in this second sample, we use several DHS survey per country.

The DHS are particularly valuable to us as they are comparable across countries, and record the fertility history of ever married women aged 13 to 49. We can therefore reconstruct for each child at any age the number of siblings, older or younger, she had. In addition, we also know whether and when the child died, which allows us to precisely look into the link between mortality, sibling competition and the stopping rule. More precisely, we make use of this data in two separate ways. First, to identify countries implementing the stopping rule, we record information for each child at ages -2 to +5 and compute how many ever born or younger siblings she has at each age. Under this approach, we observe each child 8 times and can therefore implement a child fixed effect in our regressions. To investigate the consequences of the stopping rule in terms of mortality, we use the database in a more standard fashion and look at child survival at age 5 by keeping only one observation per child. Appendix B lists the countries and surveys we used, as well as their number of observations.[4]

3.2 Measuring the stopping rule

3.2.1 Previous measures

In the literature, the most popular measure of the stopping rule is based on a literal interpretation of the rule: the last born in the family tends to be a boy. As a result, countries in which the stopping rule is prevalent should display a disproportionate number of sons among the last born. While intuitive, we think that this approach, unlike our measure, suffers from some important shortcomings. It first requires a benchmark on what the natural sex ratio at birth would be in the absence of a stopping rule. There is no universal natural sex ratio at birth: depending on the country and the period, it can vary betwen 103 and 108 boys per 100 girls (see for instance Hesketh & Xing (2006); Chahnazarian (1988); Waldron (1998)). Moreover, this approach naturally focusses on families which have completed their fertility (i.e., mothers for which we know the gender

[4] As children need to be born more than 5 years before the survey to be able to survive up to age 5, our analysis is restricted to children born 5 years or more before the survey.

of the last born). As a result, it necessarily describes the behavior of older cohorts of mothers. Finally, this measure neglects the fact that the stopping rule also affects children at lower birth orders through their number of younger siblings. (In Appendix D, we show how the detection of countries implementing the stopping rule according to our method differs from the detection using the sex ratio of the last born).

Another closely related method used in demography is the "parity progression ratio" (Ben-Porath & Welch, 1976; Williamson, 1976; Arnold, 1997; Arnold et al., 1998; Norling, 2015). It evaluates, at a given birth rank, the relative probability to continue childbearing given the gender of the child at that rank. In other words, it measures the probability that a child of a given rank and gender is the last born. This measure is therefore very close to the "sex ratio of the last born" method but does not rely on a natural sex ratio at birth. However, it suffers from a number of limitations. First, it is a rank-specific measure and there is no obvious way to aggregate it over ranks. As a result, it does not provide a direct measure of the prevalence of the stopping rule at a more aggregate level (e.g. by country, region or ethnic group). Second, children of all ranks below the 'desired number of boys', b^*, will necessarily have younger siblings irrespective of their gender. Thus, if parents want, for instance, at least 2 boys, the first born of the family will necessarily have a younger sibling. It is only for ranks larger than b^* that the parity progression ratio can detect a stopping rule behavior. This is problematic for cross countries studies, as b^* may vary across countries and over time. This characteristic of the parity progression ratio follows from its focus on the gender of the last born, without taking into consideration the fact that the stopping rule also affects the younger sibling composition of all children, irrespective of their ranks.

3.2.2 The number of younger siblings as a measure of the stopping rule

Our theory offers a more precise and straightforward measure than the sex ratio of the last born or the parity progression ratio to identify countries in which the stopping rule is prevalent. This measure simply compares for boys and girls their number of younger siblings. In Section 2, we showed that the stopping rule affects the number of younger siblings of children of all ranks. Therefore, in a country in which the stopping rule does not apply, the difference in the number of younger siblings between boys and girls of any rank will be exactly zero. This benchmark does not depend on a natural sex ratio at birth, nor on whether a particular child is a last born or not. By contrast, where the stopping rule prevails, this difference is necessarily strictly greater than zero, if the stopping rule is in favour of boys, or smaller than zero if it is in favour of girls. In addition, the difference in the number of younger siblings emerges as soon as families have reached a number of births exceeding their desired number of 'boys', b^* (or girls), which occurs several births before they have completed their fertility. Therefore, our measure can capture behavioral changes sooner than a measure relying on the sex ratio of the last born. Finally, as discussed in Section 2, it can be aggregated across children and families in a straightforward manner.

To demonstrate our theoretical finding, we use a simple empirical test. We restrict our sample to all children, whether alive or not, born at least 5 years before the survey[5] We then observe her ever born siblings at each age between two years before and five year after her birth (in the spirit of Figure 1), and compute the number of ever-born siblings at these ages. As a result, each child in our sample is observed 8 times, with an age-varying number of ever born siblings. We then run the following regression separately for each country in our sample:

$$nb\_siblings_{it} = \sum_{t=-2}^{5} (\alpha_t * age\_t_{it} + \beta_t * female_i * age\_t_{it}) + \delta_i + \epsilon_{it} \tag{1}$$

where $nb\_siblings_{it}$ is the number of ever-born siblings of child i at age t, $female_i$ a dummy indicating that the child is a female, and δ_i a child fixed effect. Finally, we cluster standard errors at the child level and weight each observation by the DHS sample weight divided by the number of children in the family (this is so that each mother has the same weight in the regression: given that the stopping rule is a parental decision, there is no reason to give more weight to families with more children).

This specification implements, in a regression framework with child fixed effect, the approach pictured in Figure 1. Our theoretical discussion implies that the only difference between boys and girls in stopping rule countries comes from younger siblings. By focusing on ever born siblings, we can demonstrate this characteristics by showing that the number of ever born siblings does not differ between boys and girls before they are born, and start diverging after their birth (i.e when they start having younger siblings).

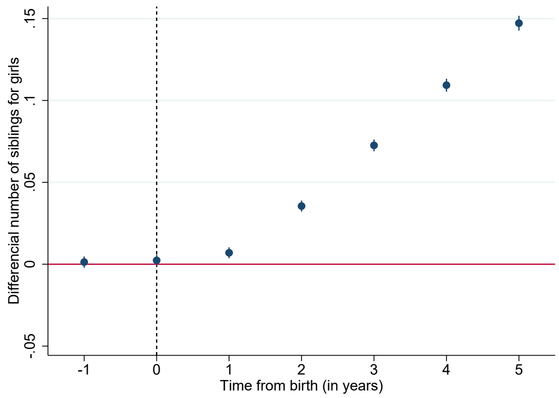
In a country in which the stopping rule prevails, girls have the same number of older siblings but more younger siblings than boys: β_1 to β_5 are expected to be positive, while coefficients β_{-1} to β_0 are expected to be zero in all countries. In Figure 3, we present the results of this regression for India and Paraguay. In India (panel (a)), girls and boys have the same number of older siblings before they are born. However, even a year after their birth, girls already tend to have more younger siblings. At the age of 5, they have on average 0.15 more siblings than boys. We do not observe such a pattern in Paraguay (panel (b)): at all ages, the number of siblings remains the same for boys and girls. According to our test, the stopping rule does not prevail in Paraguay.

Since we believe that it is now clear that the difference in the number of siblings due to the stopping rule is only driven by younger siblings, for the remainder of the paper, we will only focus on younger siblings. To detect the implementation of the stopping rule, we therefore run the following regression:

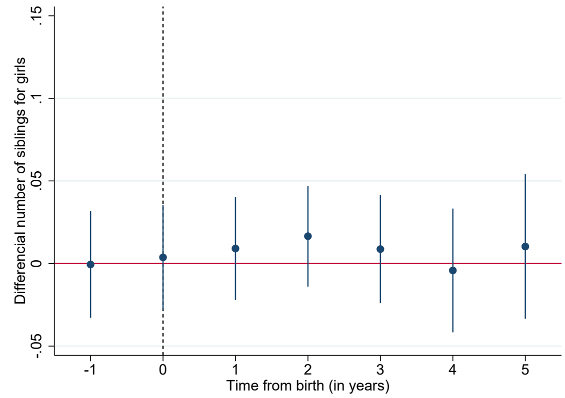
$$nb\_younger\_siblings_{it} = \sum_{t=-2}^{5} (\alpha_t * age\_t_{it} + \beta_t * female_i * age\_t_{it}) + \delta_i + \epsilon_{it} \tag{2}$$

[5] Of course, other thresholds can be used. We chose here to focus on under five mortality.

Figure 3: Differential number of ever-born siblings by age and gender in India and Paraguay



(a) Data source: DHS India 2015, all children aged 5+ at the time of the survey.
Reading: at age 5, the average Indian girl had 0.15 more siblings that the average Indian boy.



(b) Data source: DHS Paraguay 1990, all children aged 5+ at the time of the survey.
Reading: at age 5, the average Paraguayan girl had the same number of siblings as the average Paraguayan boy.

With the same notation as above and $nb\_younger\_siblings_{it}$ the number of ever born younger siblings of child i at age t.

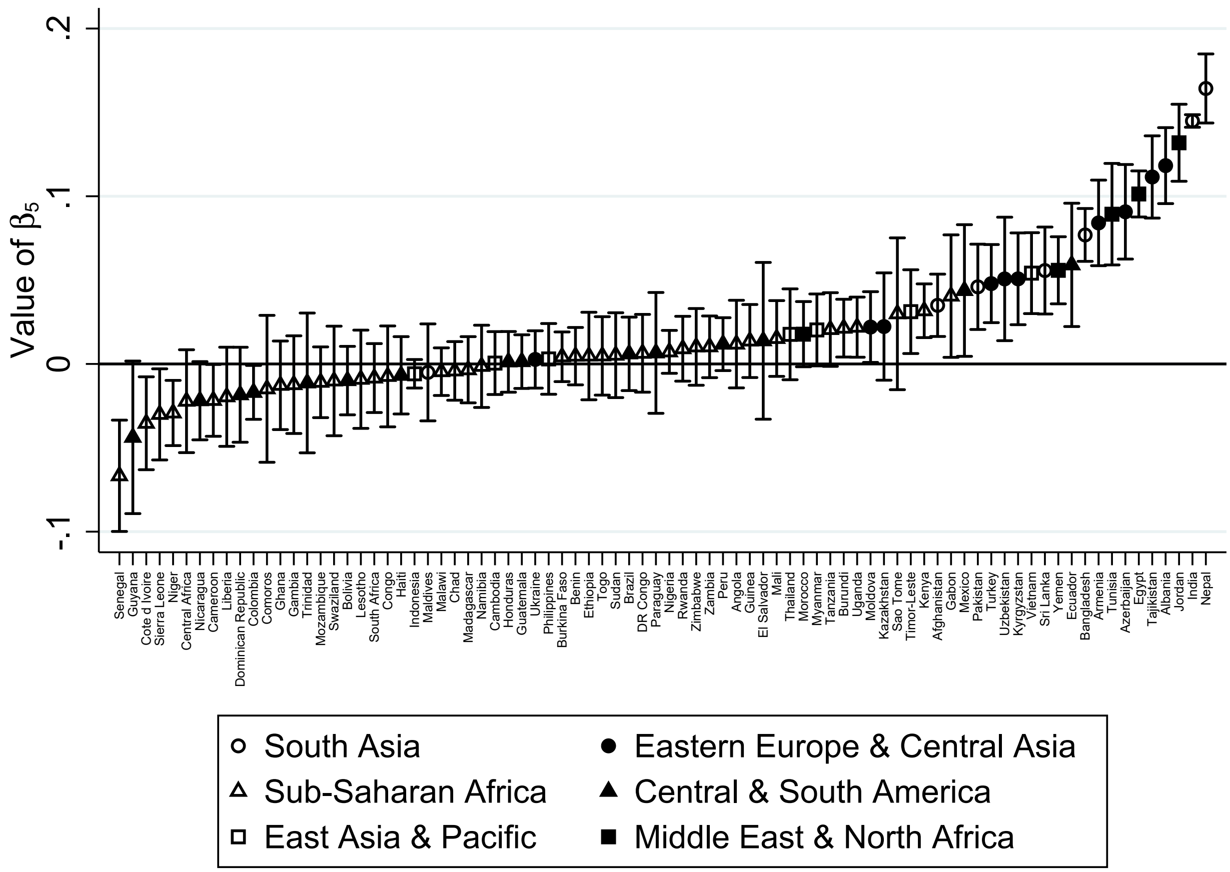
In Figure 4, we report the β_5 coefficients obtained for all the countries present in our sample, by increasing order.[6] As reported on the right hand side of the Figure, there is a substantial cluster of countries with a very high difference in the number of younger siblings, indicating the prevalence of the stopping rule in these countries. The latter does not only include the 'usual suspects', such as Nepal, India, Pakistan or Bangladesh but also countries of Eastern Europe and Central Asia, such as Albania or Azerbaidjan, and Northern Africa, such as Egypt or Tunisia. This has been essentially ignored in the literature (Ebenstein (2014) is a notable exception). Second, there is a smaller group of countries, essentially from Sub Saharan Africa, in which these coefficients are negative, indicating the presence of a stopping rule favoring girls and not boys. This possibility is hardly mentioned in the literature (Williamson, 1976), but our gender neutral approach allows to identify such cases. For expositional simplicity, we will in the following continue to write about girls, keeping in mind that, for certain countries, the stopping rule favors girls, and not boys.

We now focus on countries for which the β_5 coefficient is statistically different from zero (at the 5% level) and are therefore identified as stopping rule countries.[7] The stopping rule countries favouring boys are: Afghanistan, Albania, Armenia, Azerbaijan, Bangladesh, Burundi, Ecuador, Egypt, Gabon, India, Jordan, Kenya, Kyrgyzstan, Mexico, Moldova, Nepal, Pakistan, Sri Lanka, Tajikistan, Timor-Leste, Tunisia, Turkey, Uganda, Uzbek-

[6] Results using the number of ever born siblings rather that the number of ever born younger siblings can be found in Appendix C.

[7] Clearly, by definition of statistical significance, we wrongly classify 5% of countries as practicing the stopping rule. In addition, since the sample size of DHS surveys vary widely across countries, we detect more often countries in which the DHS sample is large. We do not know of a better classification method that would fit this data constraint.

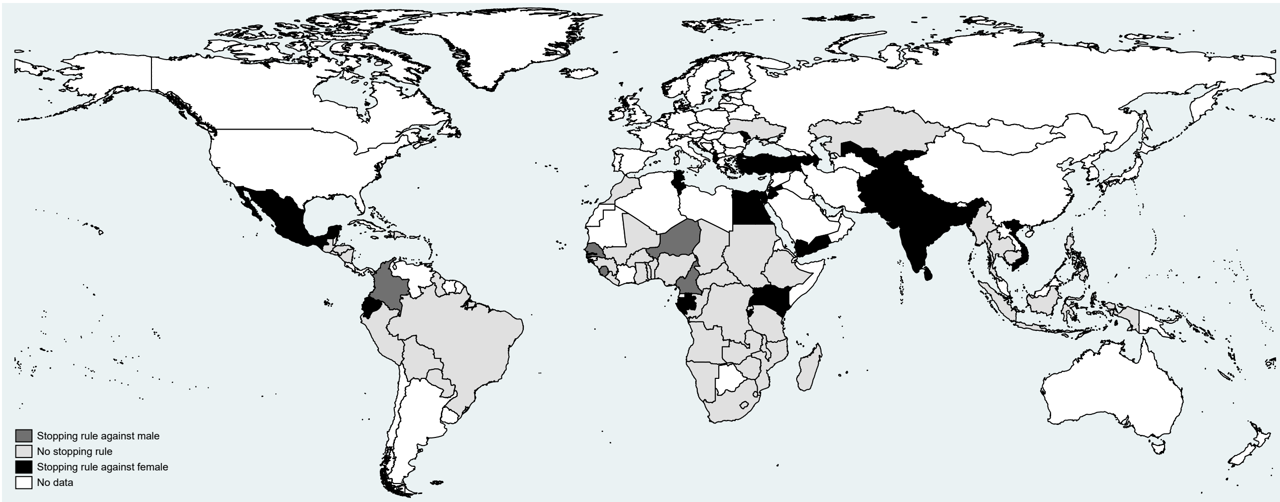
Figure 4: Differential number of younger siblings of girls at age 5, by country



Data source: DHS data, all children aged 5+ at the time of the survey.
Reading: at age 5, the average Nepalese girl has 0.164 more younger siblings than the average Nepalese boy.

istan, Vietnam and Yemen, while those favoring girls are: Cameroon, Colombia, Cote d'Ivoire, Niger, Senegal and Sierra Leone. We present in Figure 5 a map locating these countries.

Figure 5: Countries bounded by the stopping rule



3.2.3 Interpreting our measure: "absence of evidence is not evidence of absence'

The test proposed above provides a sufficient condition for the prevalence of the stopping rule, and this is the way we interpreted it so far. However, without further information on $\bar{N}$ or b^*, it does not provide a necessary condition of its existence as the stopping rule can remain undetected by the test proposed.[8] A difference between the number of younger siblings for boys and girls is only informative about the fact that the stopping rule is *prevalent*. Consider a society in which all families want exactly two sons but can have an infinite number of children. In that society, girls will systematically have 2 more younger siblings. Now, imagine that, in the same society, families can only have two children. In that case, all families have exactly two children. In spite of the existence of a strong stopping rule, the stopping rule is never binding. Girls and boys have exactly the same number of younger siblings (0.5 on average). Relatedly, as illustrated in Figure 2, for a given preference for sons, societies in which the desired fertility is increasing (n^* falls) exhibit a smaller difference in younger siblings by gender, which may also remain undetected by our test.

Moreover, our measure does not capture the *intensity* of a preference for sons, as the latter depends on the preferred or maximal family size as well as the number of boys desired. As a result, the difference in the number of younger siblings cannot be compared across countries to assess the relative intensity of the stopping rule, but simply as a sufficient condition for its prevalence. In the following, when referring to a stopping rule country, we therefore refer to a country in which the prevalence of the stopping rule can be detected through our measure. As we will focus on the implications of the stopping rule for sibling competition, these are the countries relevant for our investigation.

[8] The same remark holds for the traditional measures of the stopping rule.

4 Female mortality and the stopping rule

In this section, we investigate the implications on mortality of the more intense sibling competition faced by young girls in stopping rule countries. A large literature already documented various forms of discrimination against girls as an important source of differential mortality by gender (Qian, 2008; Barcellos et al., 2014). These mechanisms, based on active forms of discrimination, are not the focus of this paper. We instead aim at precisely quantifying the additional mortality caused by the increased sibling competition in countries in which the stopping rule prevails. To this end, we proceed in two steps. We first estimate at the country level the impact of an additional younger sibling on child mortality. The previous section measured for each country the number of younger siblings faced by girls versus boys. Combining these two measures, we can then compute, country by country, the number of child deaths caused by the sibling competition effect of the stopping rule.

4.1 Sibling competition and mortality

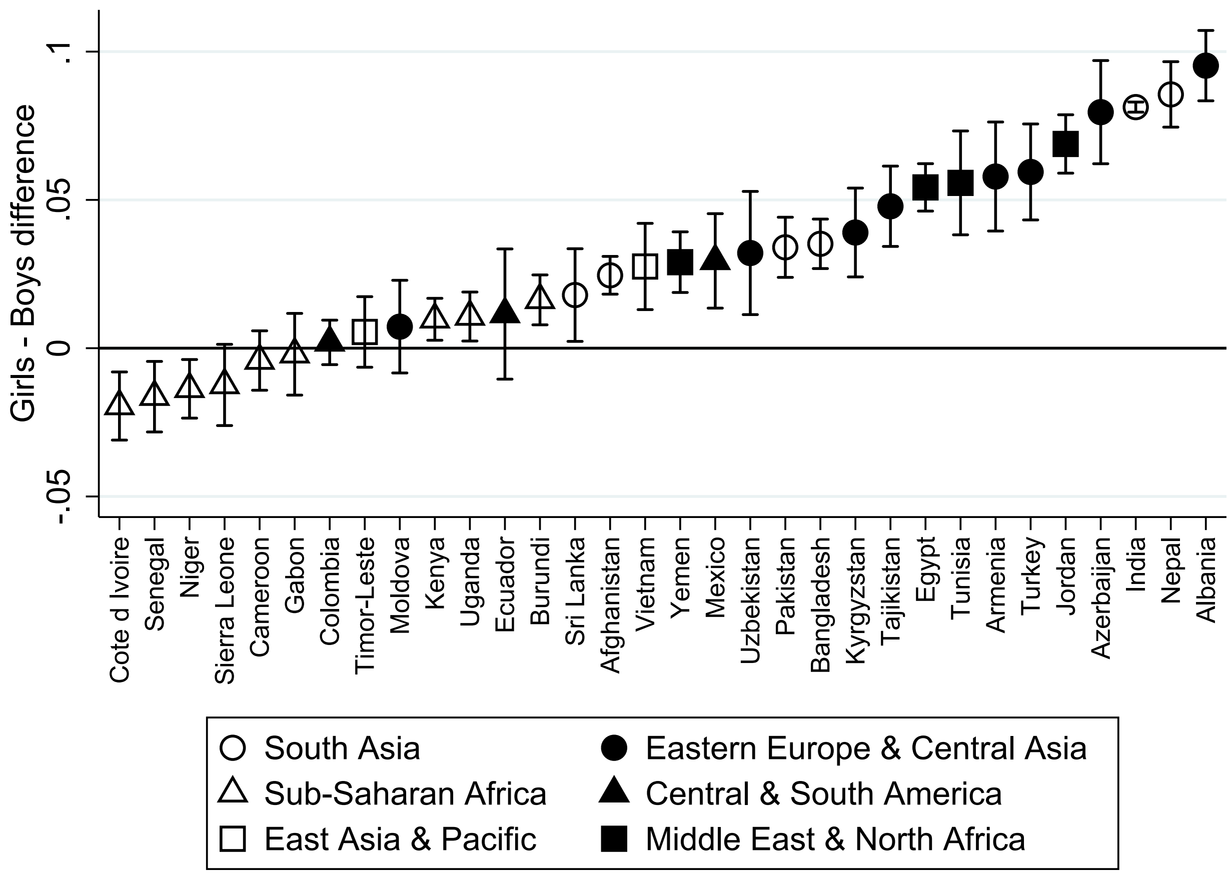
The estimation we would ideally like to implement is the following:

$$Mortality_i = \alpha + \delta Sibling\_Competition_i + \epsilon_i \tag{3}$$

Where $Mortality_i$ is a dummy indicating if child i was dead at age 5 and $Sibling\_Competition_i$ the average number of younger siblings-years child i had been exposed to from age 0 to 5. Note that the latter is slightly different from the number of younger siblings at age 5 that we have been using to detect the stopping rule. As we are interested in effective sibling competition, $Sibling\_Competition_i$ is the average number of younger siblings per year a child has been exposed to. This measure takes into account the number of years a child is exposed to competition. Thus, a younger sibling born when the child of interest is one year old will compete with that child for 4 years, while a sibling born at age 4 will only compete for a year. This measure also takes into account child death: a sibling born at age 1 and dying at age 4 will only compete for 3 years. In other words, $Sibling\_Competition_i$ is the average number of alive younger siblings per year that a child is exposed to between 0 and 5.[9] Figure 6 presents the difference in $Sibling\_Competition_i$ between gender for each stopping rule country. Note that some of these differences are not significant. This means that living in a country which implements the stopping does not always translate in a significant difference in younger siblings competition.

[9]The results presented below are essentially unchanged when using the number of younger siblings at age 5 instead our measure of siblings competition (these results are reported in Appendix G).

Figure 6: Gender difference in sibling competition, 0-5



4.1.1 The twin instrument

Having a child is an endogenous choice and there is an obvious omitted variable issue: individual and parental characteristics may determine both mortality and the number of younger siblings. In addition, the gender of the child may determine birth spacing to the next child, which itself can cause mortality (Jayachandran & Kuziemko, 2011; Rossi & Rouanet, 2015). Reverse causality is also an issue: parents may have younger siblings to "replace" an older sibling who died. We therefore choose an instrumentation strategy and follow the literature in using twin birth as an instrument for total fertility (Angrist & Evans, 1998; Black et al., 2005; Angrist & Schlosser, 2010). However, we adapt that instrument to our setting: because we need to instrument the number of younger siblings of a child, our instrument is a dummy indicating if child i has had twins among her younger siblings born before she was 5.

We run the following IV regression separately for each country:

$$Mortality_{im} = \alpha + \delta Sibling\_Competition_{im} + M_m + C_{im} + \epsilon_{im} \tag{4}$$

where $Sibling\_Competition_{im}$ is instrumented by the presence of twins among any birth following child i between age 1 and 5. C_{im} are various child level controls as several mechanisms relate different fertility behavior to mortality and we need to control for these. In particular, birth spacing has a direct impact on mortality (Palloni & Millman, 1986; Retherford et al., 1989; Jayachandran & Kuziemko, 2011). We therefore control for a set of fixed effects measuring birth spacing to the previous child. We can not however control for birth spacing to the following child, as this would be endogenous, but our instrument deals with that issue. Also, Jayachandran & Pande (2017) show that the first born (in particular boys) may be treated differently than the other children. We therefore control for whether the child is the first born boy or a first born daughter: we control for birth rank FE and for gender birth rank FE. We also control for whether the child is part of a twin birth, her year of birth (by fixed effect), the age of the mother at her birth (by fixed effect), and whether she is the last born. M_m are household level controls, which include the environment (rural or urban), the year of birth of the mother and her education level. A complete description of the variables used Equation 4 is given in Appendix F. Finally, we cluster standard errors at the mother level and weight each observation by the DHS sample weight.

Figure 7 presents the estimated δ coefficients for each country in our sample[10]

These coefficients measure the additional probability for a child to die before age 5 caused by an increase in her younger sibling competition. For instance, in India, our estimate indicates that one additional younger sibling increases the probability of death before age 5 by 7.25 percentage points, on average. Most of the estimates are not significant at the standard levels, indicating that in many countries, sibling competition is not a cause of early death.

4.2 Computing the number of deaths caused by the stopping rule

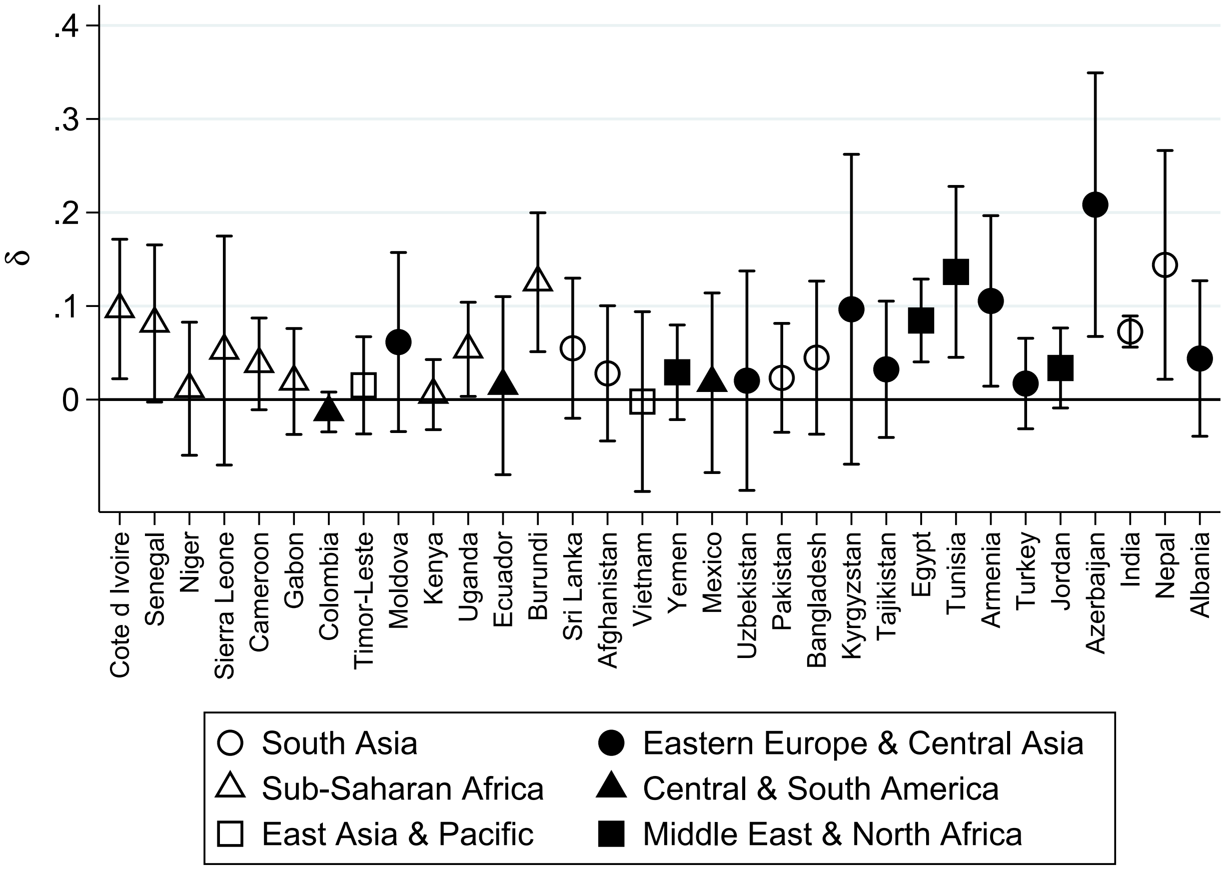
We are now in a position to compute the number of deaths caused by the stopping rule at the country level in a straightforward manner. We simply multiply the δ coefficients estimated from regression 4 with the average difference in sibling competition by gender (Figure 6). This product directly provides the additional probability for girls to die due to the stopping rule through the sibling competition channel. We then multiply this additional probability with the total number of girls[11] to estimate the total number of deaths caused by this mechanism:

$$\text{Dead girls} = \Big[\delta * \text{Sibling Competition Differential}\Big] * \text{Number of girls} \tag{5}$$

Table 1 presents the numbers of estimated deaths among girls and boys for the ten countries for which our estimates were statistically significant. We estimated the sibling competition mechanism to have caused the death of more than three million children, most of them in India (given its population size).

[10] The detailed results for each country of the first and second stages are reported in Appendix G.
[11] Taken from the World Bank databank. Health Nutrition and Population Statistics: Population estimates and projections.

Figure 7: Country level estimates of mortality caused by sibling competition

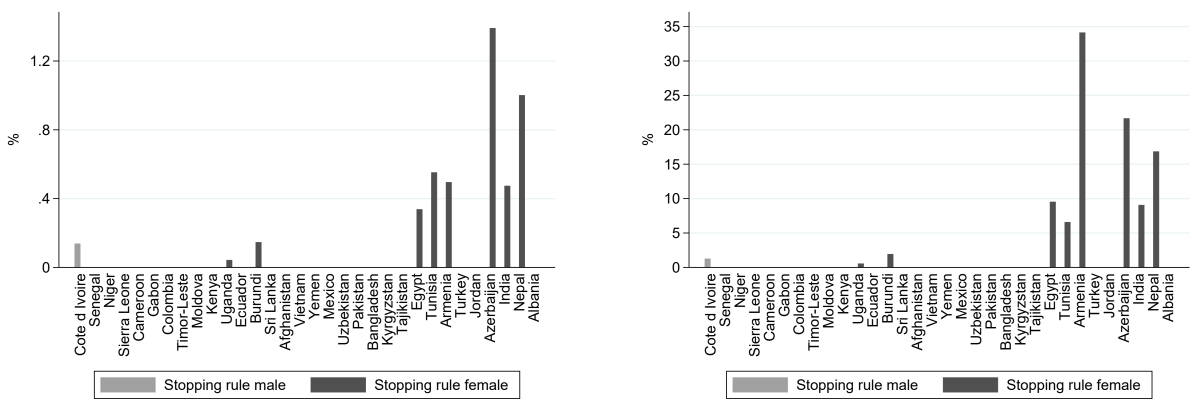


In panel (a) of Figure 8, we present the mortality rates by country. In Nepal for instance, 1% of girls die because of the stopping rule while this number gets as high as 1.3% in Azerbaijan. Panel (b) depicts the share of deaths that this mechanism explain. In Armenia, our mechanisms represents around 35% of the deaths of girls below 5. These numbers are also quite high in Azerbaijan and Nepal at 21 and 16% while they are close to 10% in India and Egypt. The contribution to mortality of the stopping rule through sibling competition is substantial.

Table 1: Number of deaths caused by the Stopping Rule

	Number of dead girls	Number of dead boys	Survey years
India	2,958,743		2015
Egypt	147,313		2014
Nepal	133,275		2017
Azerbaijan	55,773		2015
Tunisia	21,372		1988
Cote d'Ivoire		15,042	2012
Uganda	8,319		2016
Burundi	7,641		2016
Armenia	6,847		2016

Figure 8: Deaths and Stopping Rule



(a) Probability of dying before 5 because of the stopping rule.

(b) Share of deaths caused by the stopping rule.
.

5 The evolution over time

We now investigate the evolution over time of the impact of the stopping rule on mortality. This evolution is a priori ambiguous since, on the one hand, as countries get richer, sibling competition should have less dramatic consequences and, on the other hand, the desired family size has probably decreased over time which exacerbates the sibling competition effect by increasing the differential number of younger siblings by gender (see Figure 2). In the following, we select all the countries for which information on children born in all three decades between 1980 and 2010 is available, and replicate the estimation process described in the previous Section.[12].

A country is considered as a stopping rule country if the β_5 coefficient obtained from replicating Equation 1 on the sub-sample of children born between 1980 and 2010 is significant at the 5% level. [13] We then investigate, for those countries, the evolution of the mortality consequences of the stopping rule over time. When aggregating

[12]This is to allow us to meaningfully compare and aggregate results across countries.
[13]Appendix E presents the graph of these coefficients which, unsurprisingly, selects the same group of countries as in our previous exercise.

these countries, we will refer to them as the Stopping Rule World.

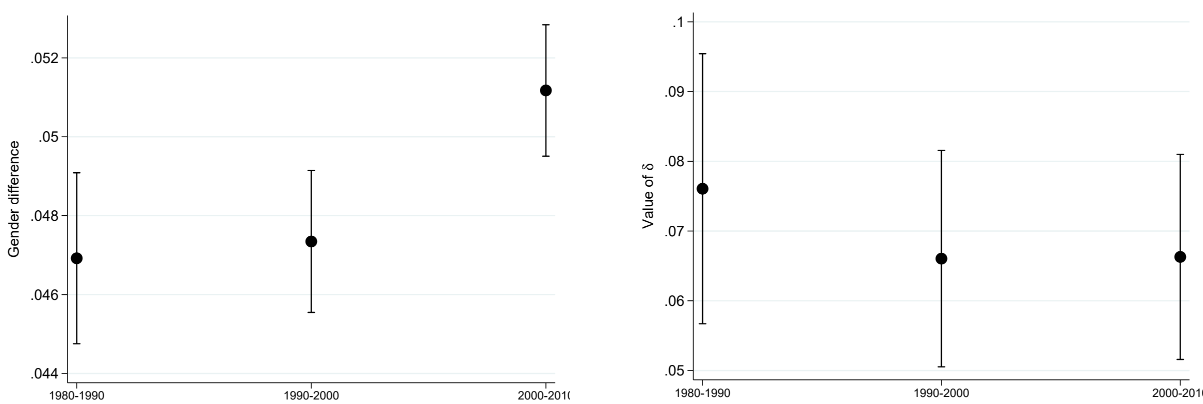
In Panel (a) of Figure 9, we present the evolution of the average gender difference in the number of younger siblings over the last three decades in the Stoping Rule World. These estimates and their confidence intervals where constructed by averaging country level differences and weigthing by the average population size of each country during the corresponding decade.[14] Difference in sibling competition has signifcantly increased for births taking place in the most recent decade.

We also run an estimation of mortality based on Equation 4, allowing the coefficient of interest to vary across decades.

$$Mortality_{im} = \alpha + \sum_{c \in C} \delta_c * Cohort\_c_{im} * Sibling\_Competition_{im} + M_m + C_{im} + \epsilon_{im} \tag{6}$$

where $Cohort\_c$ a dummy indicating if child i was born in decade c (1980s, 1990s or 2000s). Following the strategy of the previous section, we instrument $Sibling\_Competition_{im} * Cohort\_c$ by a dummy indicating if child i had twins among her younger siblings born before she was 5, interacted with $Cohort\_c$. We present the estimated δ_c coefficients in panel (b) of Figure 9, aggregated at the Stopping Rule World level. (These estimates and the corresponding confidence intervals were constructed in the same manner as sibling competition in panel (a). At the country level, the evolution of the estimated *delta* are presented in Appendix H). As expected, the causal impact of sibling competition on mortality has decreased over time.

Figure 9: The evolution of sibling competition and its effect on mortality over time in the Stopping Rule World



(a)Gender difference in sibling competition, 0-5.

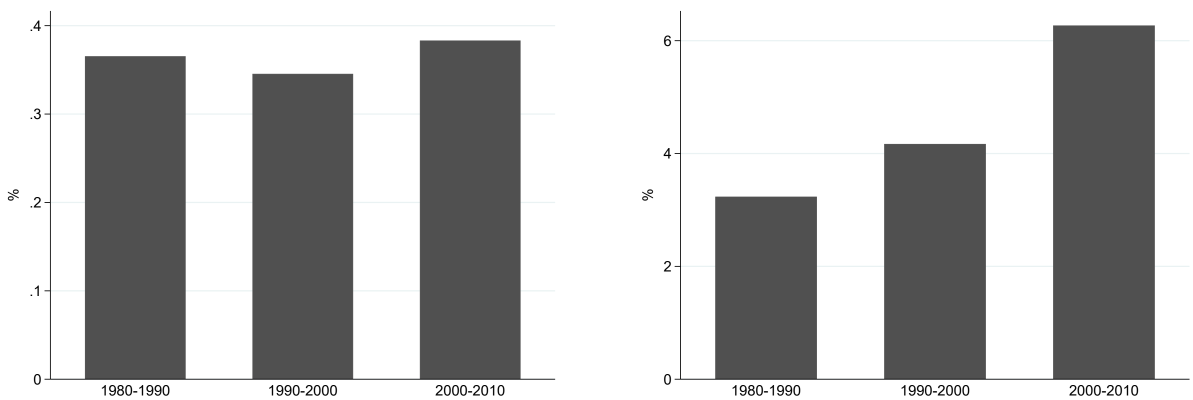
(b) Causal effect of sibling competition on mortality.

Combining these two sets of estimates, we present in Figure 10 the evolution of the probability of death due to the stopping rule. As above, these estimations were constructed by taking the average of the country level

[14]See the right-hand side graphs of Appendix H for the country level estimates of gender difference in sibling competition.

estimations, weighted by the girl population by country and decade. [15] Overall, the probability of death due to the stopping rule has ambiguously evolved: first, the fall in mortality rates associated with sibling competition is large enough to compensate for the increase in the difference in sibling competition. Then the increase in sibling competition compensate the evolution of mortality which increased the probability of death. In the right-hand side panel of the same Figure, we report, among the observed deaths, the proportion that can be attributed to the stopping rule (again taking a weigthed average across the relevant countries). The importance of our mechanism as a cause of mortality has increased over time. While overall mortality has consistently decreased, the fall in mortality associated with the stopping rule has been less pronounced than for other causes of mortality.

Figure 10: Evolution of the probability of dying and share of deaths explained by the Stopping Rule, Stopping Rule World



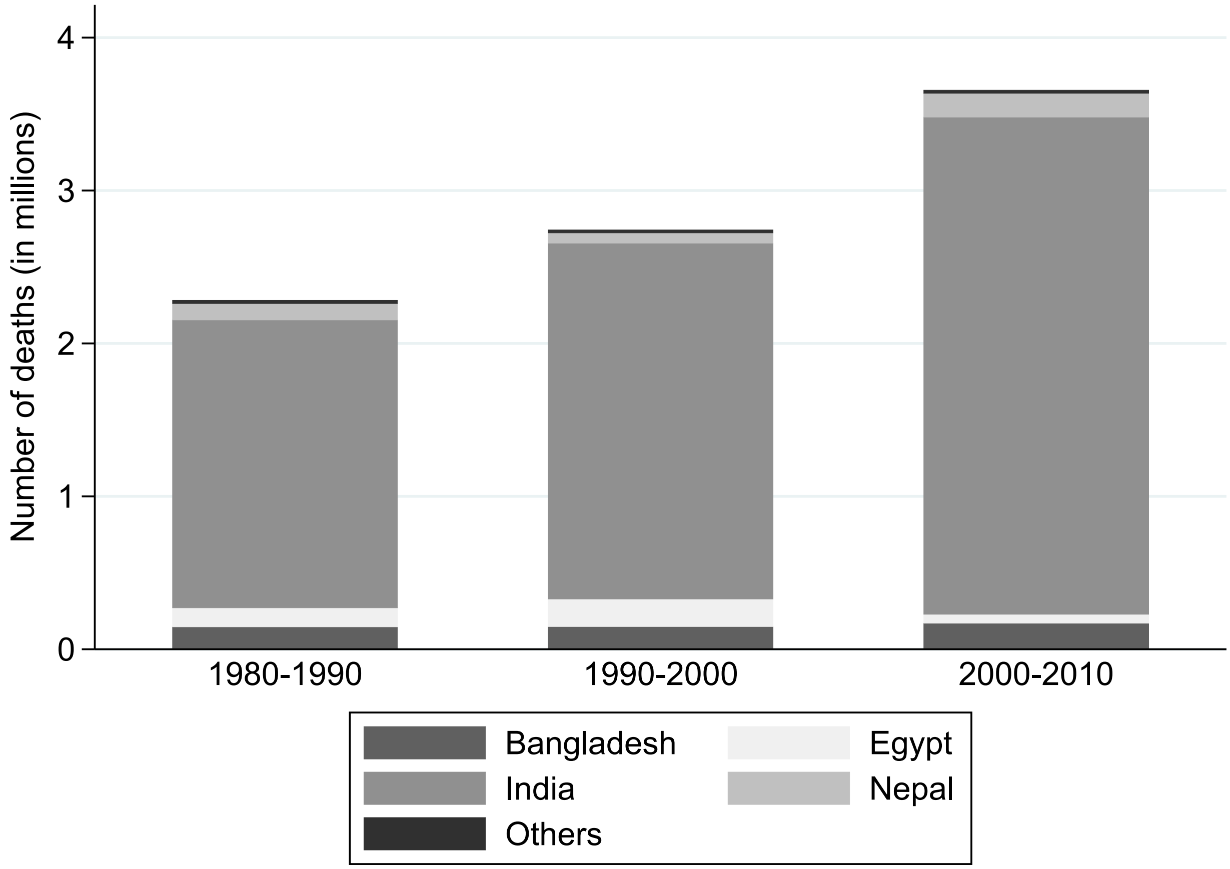
(a) Probability of dying under 5 because of the stopping rule.

(b) Share of deaths caused by the stopping rule.
.

Finally, we briefly explore this evolution at the country level. Figure 11 presents the evolution of the number of deaths due to the stopping rule and their distribution across countries. The four main contributors are Egypt, India, Bangladesh and Nepal. Given its population size, India is the country with the most deaths. Over time, the stopping rule as a cause of child mortality is increasingly concentrated in South Asia which, in the last decade, represented more than 90% of the deaths. At the country level (see Appendix H) countries evolved differently. While Egypt became negligible in the last decade, the relative importance of India increased over time. Unlike others, India indeed combines an increase in sibling competition with stable mortality coefficients.

[15] Appendix I reports the corresponding estimates at the country level.

Figure 11: Evolution of the number of deaths across countries due to the stopping rule



6 Conclusion

In many countries, young girls face higher mortality risks than boys. This disadvantage partly follows from active discrimination in the family in terms of access to health care or essential resources. In this paper, we showed that passive or implicit discrimination also plays an important role. In an environment where male children are strongly preferred, parents tend to keep having children until they reach their desired number of boys, a behaviour known as the stopping rule. As a result, girls have on average more younger siblings and therefore experience a stronger intra-household competition for scarce resources. We develop a new measure to identify countries in which the stopping rule prevails. Those countries are mainly situated in Northern Africa, Eastern Europe, Central and South Asia. Computing the number of young girls who died under the stopping rule, we find that the resulting increase in sibling competition explains up to 10% of under five girl mortality in India, 16% in Nepal and more than 20% in Armenia or Azerbadjan.

References

Abrevaya, J. (2009). Are there missing girls in the united states? evidence from birth data. *American Economic Journal: Applied Economics*, 1(2).

Anderson, S. & Ray, D. (2010). Missing women: Age and disease. *Review of Economic Studies*, 77, 1262–1300.

Angrist, J. D. & Evans, W. N. (1998). Children and their parents' labor supply: Evidence from exogenous variation in family size. *American Economic Review*, 88(3).

Angrist, Joshua D.and Lavy, V. & Schlosser, A. (2010). Multiple experiments for the causal link between the quantity and quality of children. *Journal of Labor Economics*, 28(4).

Arnold, F. (1985). Measuring the effect of sex preference on fertility: the case of korea. *Demography*, 22(2).

Arnold, F. (1997). Gender preferences for children. *Demographic and Health Surveys Comparative Studies*, 23.

Arnold, F., Choe, M. K., & Roy, T. K. (1998). Son preference, the family-building process and child mortality in india. *Population Studies*, 52(3), 301–315.

Arnold, F., Kishor, S., & Roy, T. K. (2002). Sex-selective abortions in india. *Population and Development Review*, 28(4), 759–785.

Barcellos, S. H., Carvalho, L. S., & Llenas-Muney, A. (2014). Child gender and parental investments in india: Are boys and girls treated differently? *American Economic Journal: Applied Economics*, 6(1).

Basu, D. & de Jong, R. (2010). Son targeting fertility behavior: Some consequences and determinants. *Demography*, 47(2), 521–536.

Ben-Porath, Y. & Welch, F. (1976). Do sex preferences really matter? *Quarterly Journal of Economics*, 90(2).

Black, S. E., Devereux, P. J., & Salvanes, K. G. (2005). The more the merrier? the effect of family size and birth order on children's education. *The Quarterly Journal of Economics*, 120(2), 669–700.

Chahnazarian, A. (1988). Determinants of the sex ratio at birth: Review of recent literature. *Biodemography and Social Biology*, 35(3-4), 214–235.

Clark, S. (2000). Son preference and sex composition of children: Evidence from india. *Demography*, 37(1), 95–108.

Das Gupta, M. (1987). Selective discrimination against female children in rural punjab, india. *Population and Development Review*, 13(1), 77–100.

Das Gupta, M., Zhenghua, J., Bohua, L., Zhenming, X., Chung, W., & Hwa-Ok, B. (2003). Why is son preference so persistent in east and south asia? a cross-country study of china, india and the republic of korea. *Journal of Development Studies*, 40(2), 153–187.

Dimri, A., Gille, V., & Ketz, P. (2019). Measuring sex-selective abortion: Are there repeated abortions? *Working Paper*.

Ebenstein, A. (2014). Patrilocality and missing women. *Working Paper*.

Ebenstein, A. & Leung, S. (2010). Son preference and access to social insurance: Evidence from china's rural pension program. *Population and Development Review*, 36(1), 47–70.

Haughton, J. & Haughton, D. (1998). Are simple tests of son preference useful? an evaluation using data from vietnam. *Journal of Population Economics*, 11(4), 495–516.

Hesketh, T. & Xing, Z. W. (2006). Abnormal sex ratios in human populations: Causes and consequences. *PNAS*, 103(36), 13271–13275.

Jayachandran, S. (2015). The roots of gender inequality in developing countries. *Annual Review of Economics*, 7, 63–88.

Jayachandran, S. (2017). Fertility decline and missing women. *American Economic Journal: Applied Economics*, 9(1).

Jayachandran, S. & Kuziemko, I. (2011). Why do mothers breastfeed girls less than boys? evidence and implications for child health in india. *The Quarterly Journal of Economics*, 126(3), 1485–1538.

Jayachandran, S. & Pande, R. (2017). Why are indian children so short? the role of birth order and son preference. *American Economic Review*, 107(9), 2600–2629.

Jensen, R. (2003). Equal treatment, unequal outcomes? generating sex inequality through fertility behaviour. Working Paper, Mimeo, Harvard University.

Lambert, S. & Rossi, P. (2016). Sons as widowhood insurance: Evidence from senegal. *Journal of Development Economics*, 120, 113–127.

Norling, J. (2015). A new framework for measuring heterogeneity in childbearing strategies when parents want sons and daughters. *Working Paper*.

Palloni, A. & Millman, S. (1986). Effects of inter-birth intervals and breastfeeding on infant and early childhood mortality. *Population Studies*, 40, 215–236.

Park, C. B. & Ho, N.-H. (1995). Consequences of son preference in a low-fertility society: Imbalance of the sex ratio at birth in korea. *Population and Development Review*, 21(1), 59–84.

Qian, N. (2008). Missing women and the price of tea in china: The effect of sex-specific income on sex imbalance. *Quarterly Journal of Economics*, 123(3), 1251–1285.

Qian, N., Lin, M.-J., & Liu, J.-T. (2014). More missing women, fewer dying girls: The impact of abortion on sex ratios at birth and excess female mortality in taiwan. *Journal of the European Economic Association*, 12(4), 899–926.

Ray, D. (1998). *Development Economics*. Princeton University Press.

Retherford, R. D., Choa, M. K., Thapa, S., & Gubhaju, B. B. (1989). To what extent does breastfeeding explain birth-interval effects on early childhood mortality? *Demography*, 26(3), 439–450.

Rossi, P. & Rouanet, L. (2015). Gender preferences in africa: A comparative analysis of fertility choices. *World Development*, 72, 326–345.

Sen, A. (1990). More than 100 million women are missing. *New York Review of Books*, 37(20), 61–66.

Sheps, M. C. (1963). Effects on family size and sex ratio of preferences regarding the sex of children. *Population Studies*, 17(1).

Waldron, I. (1998). *Factors Determining the Sex Ratio at Birth*. United Nations.

Williamson, N. E. (1976). *Sons or Daughters. A Cross Cultural Survey of Parental Preferences*. Sage Publications.

Yamaguchi, K. (1989). A formal theory for male-preferring stopping rules of childbearing: Sex differences in birth order and in the number of siblings. *Demography*, 26(3).

Appendix

A Proof of Proposition 1

Let us first assume that the child at rank k is a boy and consider his younger siblings. Three cases arise. In a first case, the desired number of boys is obtained before reaching the maximal number of children, which occur with probability $\sum_{j=1}^{\overline{N}-k-1} B\left(b^{*}-e-1, j\right)$ (where $B(a, b)$ is the simple binomial probability of having exactly a successes in b trials). In the second case, one needs exactly $\overline{N}$ children to reach the desired number of boys, b^{*}. This occurs with probability $\left(p \sum_{j=1}^{\overline{N}-k-1} B\left(b^{*}-e-2, j\right)\right)$: with their $\overline{N}-1$ younger children, the parents have exactly $n^{*}-1$ boys and, with probability p, their last child, at rank $\overline{N}$, is a boy. Finally, one finds parents who do not reach their desired number of boys when having $\overline{N}$ children.

Consider now a girl of the same rank k who has e older brothers. Suppose first that her next sibling is a boy. For all families that reach their desired number of boys with less than $\overline{N}$ children, this boy will have exactly the same expected number of younger siblings to that of a boy of rank k who has e older brothers. For families which, with a boy at rank k, reach a size $\overline{N}$, his expected number of younger siblings is equal to the expected number of younger siblings of a boy of rank k minus 1. In other words, the expected number of siblings of this boy of rank $k+1$, which we denote by $E\left(Y_b(k+1, e) \mid g_k\right)$ (to indicate that her sibling of rank k is a girl, g), is given by:

$$E\left(Y_b(k+1, e) \mid g_k\right)=E\left(Y_b(k, e)\right)\left(\sum_{j=1}^{\overline{N}-k-1} B\left(b^{*}-e-1, j\right)\right)+\left(E\left(Y_b(k, e)\right)-1\right)\left(1-\sum_{j=1}^{\overline{N}-k-1} B\left(b^{*}-e-1, j\right)\right)$$

$$\Longleftrightarrow E\left(Y_b(k+1, e) \mid g_k\right)=E\left(Y_b(k, e)\right)-1+\left(\sum_{j=1}^{\overline{N}-k-1} B\left(b^{*}-e-1, j\right)\right).$$

Suppose instead that her next sibling is a girl. Following the same reasoning as above, this girl, of rank $k+1$, has an expected number of younger siblings which is given by:

$$E\left(Y_g(k+1, e) \mid g_k\right)=E\left(Y_g(k, e)\right)-1+\left(\sum_{j=1}^{\overline{N}-k-1} B\left(b^{*}-e, j\right)\right).$$

As a result, the expected number of younger siblings for a girl of rank k with e older brothers, $E\left(Y_g(k, e)\right)$,is given by 1 plus expectation of the number of younger siblings of that girl's next sibling:

$$E\left(Y_g(k, e)\right)=1+p E\left(Y_b(k+1, e) \mid g_k\right)+(1-p) E\left(Y_g(k+1, e) \mid g_k\right)$$

$$= 1+p\left(E\left(Y_b\left(k,e\right)\right)-1+\left(\sum_{j=1}^{\overline{N}-k-1}B\left(b^*-e-1,j\right)\right)\right)+(1-p)\left(E\left(Y_g\left(k,e\right)\right)-1+\left(\sum_{j=1}^{\overline{N}-k-1}B\left(b^*-e,j\right)\right)\right)$$

$$= E\left(Y_b\left(k,e\right)\right)+\left(\sum_{j=1}^{\overline{N}-k-1}B\left(b^*-e-1,j\right)\right)+\frac{1-p}{p}\left(\sum_{j=1}^{\overline{N}-k-1}B\left(b^*-e,j\right)\right)$$

$$\Longrightarrow E\left(Y_g\left(k,e\right)\right)>E\left(Y_b\left(k,e\right)\right),\ \forall k<\bar{N}, e\leq b^*-1.$$

□

B List of DHS surveys

Table 2 displays the years of interview and the number of observations we use in our analysis. The *main sample* describes the datasets we use in Sections 3 & 4 while the *evolution over time sample* describes those used in Section 5.

Table 2: List of DHS surveys

	Main Sample		Evolution over time Sample	
	Year of interview	Observations	Year of interview	Observations
Afghanistan	2015	125,715	2015	89,201
Albania	2017	16,128	2009	12,699
			2017	11,590
Angola	2015	42,002	2015	25,106
Armenia	2016	8,771	2000	9,161
			2005	9,711
			2010	8,123
			2016	6,664
Azerbaijan	2006	13,565	2006	13,239
Bangladesh	2014	43,772	1994	21,193
			1996	21,697
			2000	25,532
			2007	28,732
			2011	42,427
			2014	36,263
Benin	2017	45,853	1996	14,919
			2001	17,706
			2006	55,762
			2012	41,341
			2017	24,844
Bolivia	2008	40,355	1989	10,969
			1994	16,740
			1998	24,440
			2003	42,084
			2008	39,538

Brazil	1996	25,513		
Burkina Faso	2010	56,178	1993	14,131
			1999	18,841
			2003	38,952
			2010	53,585
Burundi	2016	45,419	1987	5,689
			2010	22,274
			2016	27,207
Cambodia	2014	33,290	2000	37,031
			2005	39,549
			2010	36,153
			2014	26,613
Cameroon	2011	42,312	1998	12,675
			2004	27,717
			2011	38,584
Central Africa	1994	16,936		
Chad	2015	68,989	1997	20,828
			2004	20,111
			2015	50,004
Colombia	2015	62,593	1986	3,768
			1990	7,920
			1995	14,600
			2000	17,399
			2005	66,173
			2010	89,045
			2015	49,340
Comoros	2012	11,497	1996	6,102
			2012	9,613
Congo	2015	16,687	2015	15,875
Cote d'Ivoire	2015	28,211	1994	18,576
			1999	6,391
			2005	12,755
			2012	24,409

DR Congo	2013	59,276	2007	28,814
			2013	44,013
Dominican Republic	2013	18,167	1986	6,450
			1991	9,095
			1996	14,180
			1999	2,264
			2007	55,603
			2013	15,377
Ecuador	1987	11,835		
Egypt	2014	59,266	1988	14,965
			1992	23,146
			1995	39,710
			2000	44,954
			2005	58,173
			2008	47,619
			2014	44,952
El Salvador	1985	6,381		
Ethiopia	2016	41,392	2000	37,852
			2005	38,036
			2011	41,667
			2016	27,051
Gabon	2012	23,109	2000	14,463
			2012	20,233
Gambia	2013	26,601	2013	20,981
Ghana	2014	23,118	1988	6,583
			1993	9,431
			1998	10,950
			2003	14,047
			2008	11,680
			2014	17,367
Guatemala	2015	55,398	1987	7,310
			1995	28,363
			1999	15,459

			2015	42,626
Guinea	2018	28,887	1999	19,699
			2005	25,947
			2012	23,704
			2018	15,031
Guyana	2005	4,923	2005	4,567
Haiti	2017	27,809	1994	9,314
			2000	23,190
			2006	24,081
			2012	25,429
			2017	18,373
Honduras	2006	50,093	2006	48,021
India	2015	1,315,617	1993	165,165
			1999	217,118
			2006	246,465
			2015	1,008,253
Indonesia	2017	86,265	1987	13,782
			1991	39,330
			1994	57,867
			1997	64,236
			2002	69,247
			2007	81,369
			2012	74,126
			2017	58,485
Jordan	2017	47,040	1990	17,830
			1997	19,405
			2002	23,135
			2007	42,591
			2009	37,940
			2012	36,456
			2017	29,378
Kazakhstan	1999	8,106		
Kenya			1989	11,600

	2014	83,591	1993	15,903
			1998	18,796
			2003	20,390
			2009	22,159
			2014	64,055
Kyrgyzstan	2012	16,180	1997	6,918
			2012	13,568
Lesotho	2014	11,710	2004	13,777
			2009	14,291
			2014	8,628
Liberia	2013	30,804	1986	6,836
			2007	21,227
			2013	25,510
Madagascar	2009	48,464	1992	12,132
			1997	17,373
			2004	18,556
			2009	48,827
Malawi	2015	68,074	1992	10,735
			2000	35,599
			2004	34,155
			2010	69,482
			2015	48,724
Maldives	2017	13,922	2009	19,724
			2017	9,307
Mali	2011	33,379	1987	5,112
			1996	19,517
			2001	43,902
			2006	50,736
			2012	27,724
			2018	16,257
Mexico	1987	22,676		
Moldova	2005	9,903	2005	9,263
Morocco			1987	9,495

	2003	32,494	1992	12,773
			2003	28,999
Mozambique	2011	37,984	1997	20,764
			2003	34,692
			2011	33,686
Myanmar	2016	22,989	2016	17,052
Namibia	2013	18,090	1992	8,776
			2000	13,172
			2007	18,958
			2013	14,281
Nepal	2017	26,028	2001	25,393
			2007	25,682
			2011	24,496
			2017	18,391
Nicaragua	2001	34,157	2001	29,542
Niger	2012	44,183	1992	16,049
			1998	24,216
			2006	33,490
			2012	37,923
Nigeria	2018	127,545	1990	16,932
			2003	21,047
			2008	102,744
			2013	97,415
			2018	67,261
Pakistan	2018	50,495	1991	16,516
			2006	38,132
			2012	43,331
			2018	29,861
Paraguay	1990	15,346		
Peru	2012	47,261	1986	4,426
			1991	22,062
			1996	54,552
			2000	55,818

			2007	85,442
			2012	42,290
Philippines	2017	47,244	1993	23,559
			1998	26,582
			2003	28,369
			2008	28,126
			2013	26,373
			2017	30,511
Rwanda	2015	30,058	1992	13,432
			2000	25,206
			2005	29,284
			2010	30,973
			2015	22,079
Sao Tome & Principe	2008	7,620	2008	7,492
Senegal	2016	22,740	1986	5,679
			1993	13,941
			1997	22,288
			2005	38,115
			2011	39,362
			2016	14,044
Sierra Leone	2008	21,136	2008	20,790
South Africa	1998	22,934		
Sri Lanka	1987	17,705		
Sudan	1990	25,805		
Swaziland	2006	11,410	2006	10,847
Tajikistan	2017	21,985	2012	17,137
			2017	12,760
Tanzania	2015	37,169	1991	17,606
			1996	19,126
			1999	9,999
			2004	28,779
			2010	29,168
			2015	25,372

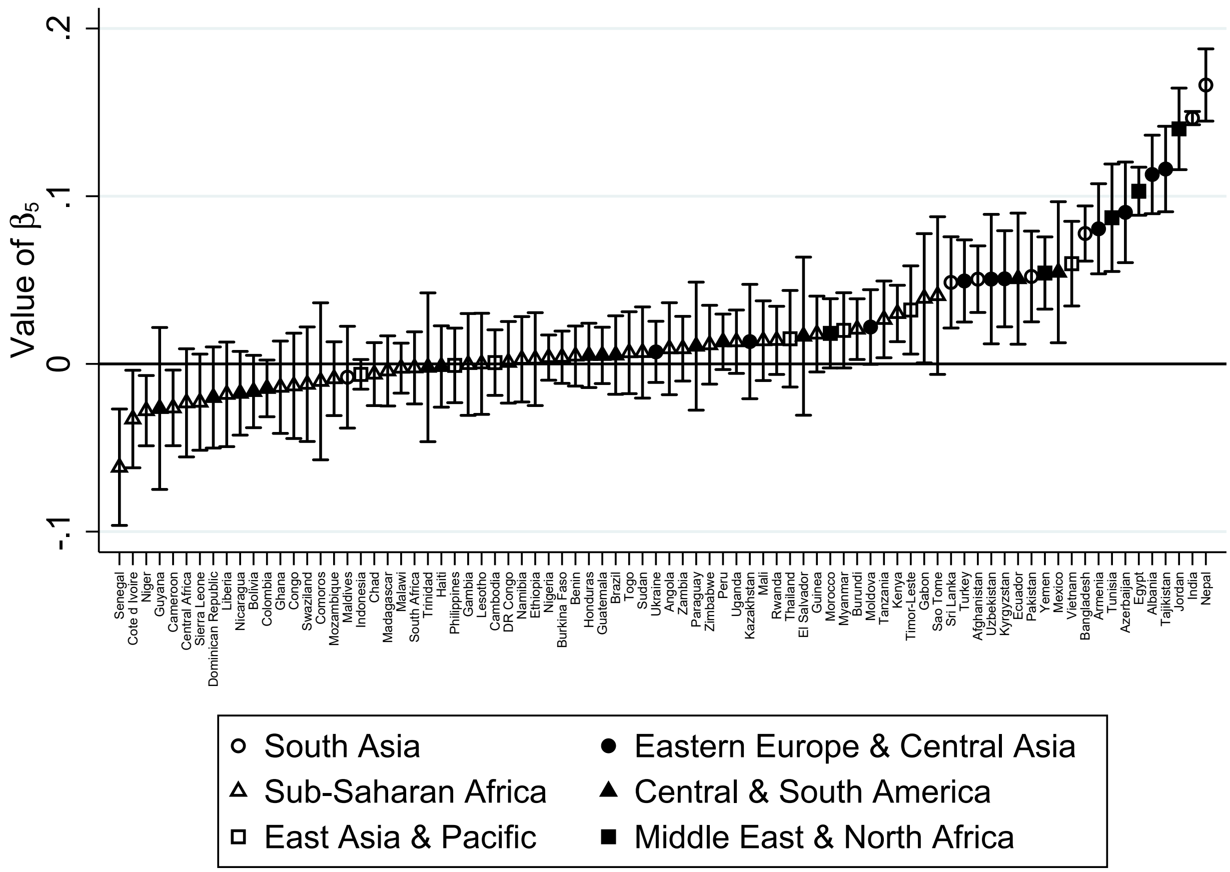
Thailand	1987	17,796		
Timor-Leste	2016	28,682	2009	35,836
			2016	18,756
Togo	2014	26,264	1988	5,345
			1998	22,073
			2014	20,466
Trinidad and Tobago	1987	7,837		
Tunisia	1988	16,463		
Turkey	2013	17,871	1993	11,576
			1998	13,443
			2004	20,345
			2008	19,270
			2013	15,089
Uganda	2016	57,906	1988	8,233
			1995	17,684
			2001	21,110
			2006	29,139
			2011	25,727
			2016	37,047
Ukraine	2007	8,007	2007	7,840
Uzbekistan	2011			
Vietnam	2002	14,383	2002	13,192
Yemen	2013	64,602	1991	18,635
			2013	51,954
Zambia	2013	49,207	1992	13,748
			1996	19,174
			2002	21,397
			2007	20,801
			2013	38,589
Zimbabwe	2015	20,791	1988	5,969
			1994	11,762
			1999	12,028
			2005	18,605

		2010	17,767
		2015	13,761

C Stopping rule: Ever born siblings instead of ever born younger siblings

Below, the β^5 using the number of ever born siblings rather that the number of ever born younger siblings, i.e., using Equation 1 instead of Equation 2

Figure 12: Differential number of younger siblings of girls at age 5, by country



Data source: DHS data, all children aged 5+ at the time of the survey.

D Relative number of younger siblings versus sex ratio of the last born

How does our manner of selecting stopping rule countries compares to the approach relying on the sex ratio of the last born? The table below shows a comparison of countries selected by both techniques. A country is selected as a stopping rule country by the sex ratio of the last born method if the average sex ratio of the last born in this country is significantly higher (for stopping rule against girls) or lower (for stopping rule against boys) than 105 boys for 100 girls. While our technique for selecting stopping rule countries uses all mothers whatever their age, Table 3 uses only the subsample of mothers aged 40+. This is to ensure comparability with the sample used to identify the gender of the last born.

There are 17 countries selected by both approaches for applying the stopping rule against girls [16], while only 3 are selected by both for applying the stopping rule against boys [17]. There are 6 countries we select as applying the stopping against girls [18] while the sex ratio of the last born does not. This is also the case for 2 countries when it comes to the stopping rule against boys [19]. The same thing happens the other way around. 11 countries are selected for applying the stopping rule against girls with the sex ratio of the last born approach, but are not with ours [20] Most of them are countries in Sub-Saharan Africa, Central or South America. This also happens for 15 countries for the stopping rule against boys [21]. Interestingly, Namibia is selected for applying the stopping rule against girls with our approach, but is selected for applying it against boys when using the sex ratio of the last born approach. Note that the selection of countries into one group or another with the sex ratio of the last born method is heavily dependant on the threshold used, while there is in fact no natural threshold.

Table 3: Comparison of selected countries

	Stopping Rule against female		Stopping Rule against male	
	Number of younger siblings	Sex of last born	Number of younger siblings	Sex of last born
Afghanistan	X			
Albania	X	X		
Angola			X	X

[16] Albania, Armenia, Azerbaijan, Bangladesh, Egypt, Gabon, India, Jordan, Kazakhstan, Kenya, Kyrgyzstan, Nepal, Pakistan,Tajikistan, Tunisia, Turkey, and Vietnam.

[17] Angola, Malawi, and Niger

[18] Afghanistan, Guinea, Moldova, Namibia, Sri Lanka, and Yemen

[19] Dominican Republic and Sudan

[20] Bolivia, Congo, Ghana, Guyana Honduras, Indonesia, Morocco, Myanmar, Thailand, Uganda, and Ukraine.

[21] Brazil, Chad, DRC, Gambia, Guatemala, Haiti, Mali, Mexico, Mozambique, Namibia, Sierra Leone, Tanzania, Togo, Trinidad & Tobago, and Zambia

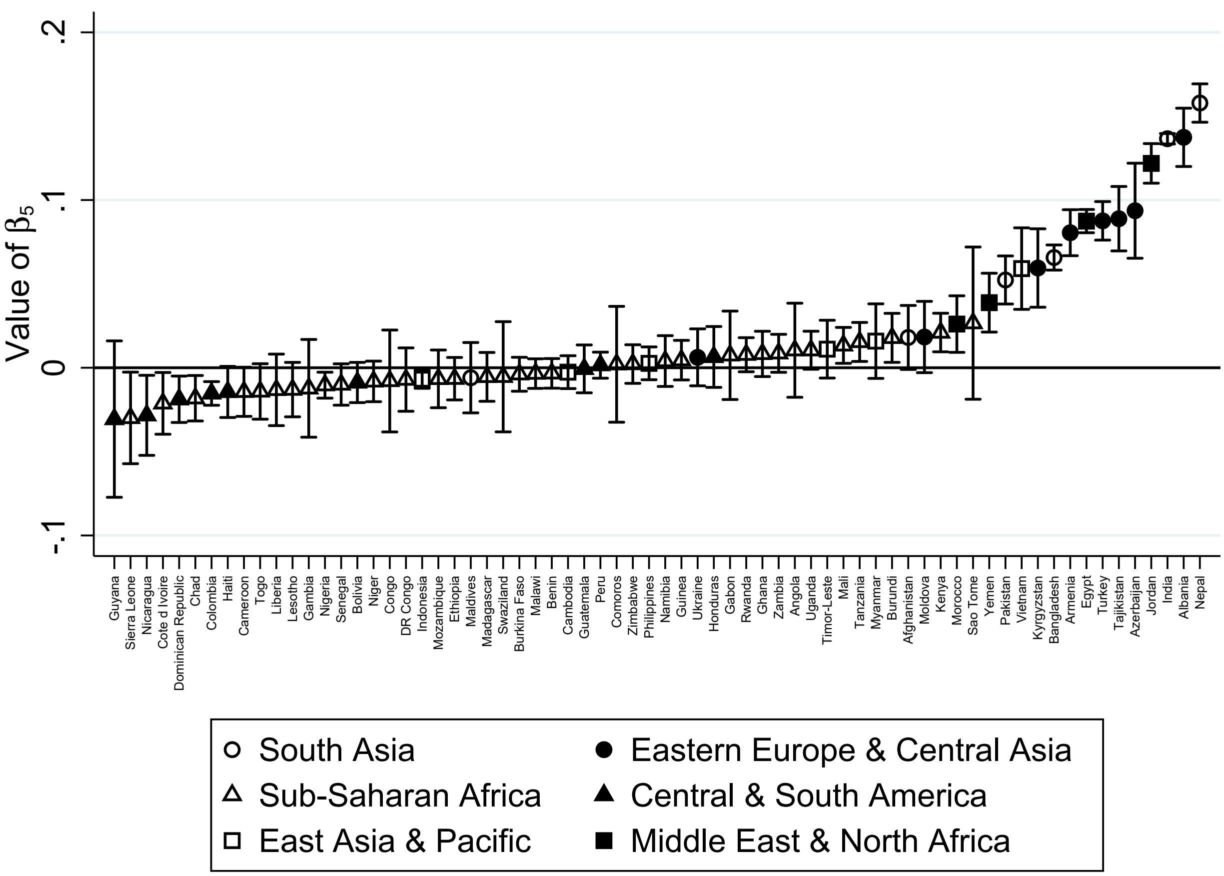
Armenia	X	X		
Azerbaijan	X	X		
Bangladesh	X	X		
Bolivia		X		
Brazil				X
Chad				X
Congo		X		
DR Congo				X
Dominican Republic			X	
Egypt	X	X		
Gabon	X	X		
Gambia				X
Ghana		X		
Guatemala				X
Guinea	X			
Guyana		X		
Haiti				X
Honduras		X		
India	X	X		
Indonesia		X		
Jordan	X	X		
Kazakhstan	X	X		
Kenya	X	X		
Kyrgyzstan	X	X		
Malawi			X	X
Mali				X
Mexico				X
Moldova	X			
Morocco		X		
Mozambique				X
Myanmar		X		
Namibia	X			X

Nepal	X	X		
Niger			X	X
Pakistan	X	X		
Sierra Leone				X
Sri Lanka	X			
Sudan			X	
Tajikistan	X	X		
Tanzania				X
Thailand		X		
Togo				X
Trinidad and Tobago				X
Tunisia	X	X		
Turkey	X	X		
Uganda		X		
Ukraine		X		
Vietnam	X	X		
Yemen	X			
Zambia				X

E Stopping rule countries: 1980-2010 sample

Figure 13 plots the β_5 coefficients for all the countries in our sample. Once again, notice the cluster of countries on the right-hand side of the Figure with high positive differences in the number of younger siblings between girls and boys. Interestingly those countries are the same selected in Subsection 3.2.2 which emphazizes the prevalence of our mechanism in those regions, i.e., South Asia, Eastern Europe, Central Asia and North Africa.

Figure 13: Differential number of younger siblings of girls at age 5, by country



Data source: DHS data, all children aged 5+ at the time of the survey.

F Descriptive Statistics

Afghanistan

	(1)	(2)	(3)
VARIABLES	mean	sd	N
year of birth	2,002	6.106	92,997
Female child	0.465	0.499	92,997
Age of mother at birth	24.29	5.928	92,997
Under 5 mortality	0.0839	0.277	92,997
Time from previous birth ¡= 18 months	0.189	0.392	92,997
Younger siblings between 0 and 5	0.783	0.494	92,997
Twin births between 0 and 5	0.0148	0.121	92,997

Albania

	(1)	(2)	(3)
VARIABLES	mean	sd	N
year of birth	2,002	6.196	13,382
Female child	0.482	0.500	13,382
Age of mother at birth	24.86	4.512	13,382
Under 5 mortality	0.0125	0.111	13,382
Time from previous birth ¡= 18 months	0.0565	0.231	13,382
Younger siblings between 0 and 5	0.277	0.353	13,382
Twin births between 0 and 5	0.00418	0.0646	13,382

Armenia

	(1)	(2)	(3)
VARIABLES	mean	sd	N
year of birth	2,000	6.726	7,061
Female child	0.475	0.499	7,061
Age of mother at birth	23.04	4.041	7,061
Under 5 mortality	0.0214	0.145	7,061
Time from previous birth ¡= 18 months	0.121	0.326	7,061
Younger siblings between 0 and 5	0.337	0.395	7,061
Twin births between 0 and 5	0.00651	0.0805	7,061

Azerbaijan

	(1)	(2)	(3)
VARIABLES	mean	sd	N
year of birth	1,991	5.853	11,268
Female child	0.477	0.499	11,268
Age of mother at birth	24.36	4.316	11,268
Under 5 mortality	0.0792	0.270	11,268
Time from previous birth ¡= 18 months	0.201	0.400	11,268
Younger siblings between 0 and 5	0.480	0.472	11,268
Twin births between 0 and 5	0.00745	0.0860	11,268

Bangladesh

	(1)	(2)	(3)
VARIABLES	mean	sd	N
year of birth	1,998	7.013	35,886
Female child	0.489	0.500	35,886
Age of mother at birth	21.51	5.342	35,886
Under 5 mortality	0.0922	0.289	35,886
Time from previous birth ¡= 18 months	0.0816	0.274	35,886
Younger siblings between 0 and 5	0.362	0.404	35,886
Twin births between 0 and 5	0.00596	0.0770	35,886

Burundi

	(1)	(2)	(3)
VARIABLES	mean	sd	N
year of birth	2,004	5.839	32,318
Female child	0.494	0.500	32,318
Age of mother at birth	25.61	5.842	32,318
Under 5 mortality	0.116	0.320	32,318
Time from previous birth ¡= 18 months	0.0967	0.296	32,318
Younger siblings between 0 and 5	0.599	0.386	32,318
Twin births between 0 and 5	0.0166	0.128	32,318

Cameroon			
	(1)	(2)	(3)
VARIABLES	mean	sd	N
year of birth	1,997	6.669	30,580
Female child	0.489	0.500	30,580
Age of mother at birth	23.55	6.081	30,580
Under 5 mortality	0.134	0.341	30,580
Time from previous birth ¡= 18 months	0.114	0.318	30,580
Younger siblings between 0 and 5	0.589	0.447	30,580
Twin births between 0 and 5	0.0270	0.162	30,580

Colombia			
	(1)	(2)	(3)
VARIABLES	mean	sd	N
year of birth	1,999	6.731	50,946
Female child	0.489	0.500	50,946
Age of mother at birth	23.06	5.477	50,946
Under 5 mortality	0.0287	0.167	50,946
Time from previous birth ¡= 18 months	0.0941	0.292	50,946
Younger siblings between 0 and 5	0.356	0.431	50,946
Twin births between 0 and 5	0.00607	0.0776	50,946

Cote d Ivoire			
	(1)	(2)	(3)
VARIABLES	mean	sd	N
year of birth	1,998	6.585	20,435
Female child	0.489	0.500	20,435
Age of mother at birth	23.63	6.160	20,435
Under 5 mortality	0.142	0.349	20,435
Time from previous birth ¡= 18 months	0.0893	0.285	20,435
Younger siblings between 0 and 5	0.530	0.419	20,435
Twin births between 0 and 5	0.0267	0.161	20,435

Ecuador			
	(1)	(2)	(3)
VARIABLES	mean	sd	N
year of birth	1,973	6.161	8,784
Female child	0.480	0.500	8,784
Age of mother at birth	23.97	5.672	8,784
Under 5 mortality	0.129	0.335	8,784
Time from previous birth ¡= 18 months	0.223	0.416	8,784
Younger siblings between 0 and 5	0.686	0.525	8,784
Twin births between 0 and 5	0.0149	0.121	8,784

Egypt			
	(1)	(2)	(3)
VARIABLES	mean	sd	N
year of birth	2,000	6.537	43,418
Female child	0.485	0.500	43,418
Age of mother at birth	24.38	5.150	43,418
Under 5 mortality	0.0484	0.215	43,418
Time from previous birth ¡= 18 months	0.111	0.314	43,418
Younger siblings between 0 and 5	0.453	0.426	43,418
Twin births between 0 and 5	0.0148	0.121	43,418

Gabon			
	(1)	(2)	(3)
VARIABLES	mean	sd	N
year of birth	1,997	6.983	17,042
Female child	0.503	0.500	17,042
Age of mother at birth	23.43	6.230	17,042
Under 5 mortality	0.0721	0.259	17,042
Time from previous birth ¡= 18 months	0.108	0.310	17,042
Younger siblings between 0 and 5	0.527	0.459	17,042
Twin births between 0 and 5	0.0233	0.151	17,042

India			
	(1)	(2)	(3)
VARIABLES	mean	sd	N
year of birth	2,000	6.590	1.056e+06
Female child	0.473	0.499	1.056e+06
Age of mother at birth	23.06	4.943	1.056e+06
Under 5 mortality	0.0667	0.250	1.056e+06
Time from previous birth ¡= 18 months	0.113	0.317	1.056e+06
Younger siblings between 0 and 5	0.475	0.446	1.056e+06
Twin births between 0 and 5	0.00629	0.0791	1.056e+06

Jordan			
	(1)	(2)	(3)
VARIABLES	mean	sd	N
year of birth	2,003	6.404	36,459
Female child	0.482	0.500	36,459
Age of mother at birth	25.74	5.432	36,459
Under 5 mortality	0.0183	0.134	36,459
Time from previous birth ¡= 18 months	0.168	0.374	36,459
Younger siblings between 0 and 5	0.585	0.481	36,459
Twin births between 0 and 5	0.0135	0.115	36,459

Kenya

VARIABLES	(1) mean	(2) sd	(3) N
year of birth	2,000	6.332	62,627
Female child	0.494	0.500	62,627
Age of mother at birth	23.74	5.833	62,627
Under 5 mortality	0.0723	0.259	62,627
Time from previous birth ¡= 18 months	0.110	0.312	62,627
Younger siblings between 0 and 5	0.545	0.451	62,627
Twin births between 0 and 5	0.0157	0.124	62,627

Kyrgyzstan

VARIABLES	(1) mean	(2) sd	(3) N
year of birth	1,997	6.362	11,817
Female child	0.488	0.500	11,817
Age of mother at birth	24.59	4.815	11,817
Under 5 mortality	0.0449	0.207	11,817
Time from previous birth ¡= 18 months	0.106	0.308	11,817
Younger siblings between 0 and 5	0.421	0.416	11,817
Twin births between 0 and 5	0.00779	0.0879	11,817

Mexico

VARIABLES	(1) mean	(2) sd	(3) N
year of birth	1,972	6.302	17,349
Female child	0.496	0.500	17,349
Age of mother at birth	24.06	5.591	17,349
Under 5 mortality	0.0926	0.290	17,349
Time from previous birth ¡= 18 months	0.238	0.426	17,349
Younger siblings between 0 and 5	0.720	0.535	17,349
Twin births between 0 and 5	0.0143	0.119	17,349

Moldova

VARIABLES	(1) mean	(2) sd	(3) N
year of birth	1,989	6.203	8,351
Female child	0.484	0.500	8,351
Age of mother at birth	23.69	4.360	8,351
Under 5 mortality	0.0375	0.190	8,351
Time from previous birth ¡= 18 months	0.0730	0.260	8,351
Younger siblings between 0 and 5	0.246	0.364	8,351
Twin births between 0 and 5	0.00515	0.0716	8,351

Nepal

VARIABLES	(1) mean	(2) sd	(3) N
year of birth	2,002	6.444	21,034
Female child	0.484	0.500	21,034
Age of mother at birth	22.89	4.845	21,034
Under 5 mortality	0.0833	0.276	21,034
Time from previous birth ¡= 18 months	0.0885	0.284	21,034
Younger siblings between 0 and 5	0.445	0.410	21,034
Twin births between 0 and 5	0.00718	0.0844	21,034

Niger

VARIABLES	(1) mean	(2) sd	(3) N
Year of birth	1,999	6.072	31,625
Female child	0.484	0.500	31,625
Age of mother at birth	23.57	6.176	31,625
Under 5 mortality	0.190	0.393	31,625
Time from previous birth ¡= 18 months	0.145	0.352	31,625
Younger siblings between 0 and 5	0.736	0.448	31,625
Twin births between 0 and 5	0.0245	0.155	31,625

Pakistan

VARIABLES	(1) mean	(2) sd	(3) N
year of birth	2,004	6.064	37,887
Female child	0.489	0.500	37,887
Age of mother at birth	24.65	5.458	37,887
Under 5 mortality	0.0751	0.264	37,887
Time from previous birth ¡= 18 months	0.186	0.389	37,887
Younger siblings between 0 and 5	0.670	0.503	37,887
Twin births between 0 and 5	0.0139	0.117	37,887

Senegal

VARIABLES	(1) mean	(2) sd	(3) N
year of birth	2,003	6.082	16,015
Female child	0.488	0.500	16,015
Age of mother at birth	24.49	6.038	16,015
Under 5 mortality	0.0996	0.299	16,015
Time from previous birth ¡= 18 months	0.0766	0.266	16,015
Younger siblings between 0 and 5	0.569	0.384	16,015
Twin births between 0 and 5	0.0207	0.142	16,015

Sierra Leone

VARIABLES	(1) mean	(2) sd	(3) N
year of birth	1,995	6.408	15,505
Female child	0.493	0.500	15,505
Age of mother at birth	22.99	6.218	15,505
Under 5 mortality	0.197	0.398	15,505
Time from previous birth ¡= 18 months	0.106	0.308	15,505
Younger siblings between 0 and 5	0.519	0.435	15,505
Twin births between 0 and 5	0.0208	0.143	15,505

Sri Lanka

VARIABLES	(1) mean	(2) sd	(3) N
year of birth	1,972	6.621	13,695
Female child	0.491	0.500	13,695
Age of mother at birth	24.76	5.300	13,695
Under 5 mortality	0.0571	0.232	13,695
Time from previous birth ¡= 18 months	0.136	0.343	13,695
Younger siblings between 0 and 5	0.562	0.466	13,695
Twin births between 0 and 5	0.0104	0.101	13,695

Tajikistan

VARIABLES	(1) mean	(2) sd	(3) N
year of birth	2,003	6.324	15,841
Female child	0.477	0.499	15,841
Age of mother at birth	24.57	4.643	15,841
Under 5 mortality	0.0496	0.217	15,841
Time from previous birth ¡= 18 months	0.126	0.332	15,841
Younger siblings between 0 and 5	0.489	0.435	15,841
Twin births between 0 and 5	0.00947	0.0969	15,841

Timor-Leste

VARIABLES	(1) mean	(2) sd	(3) N
year of birth	2,003	5.863	21,502
Female child	0.478	0.500	21,502
Age of mother at birth	25.98	5.767	21,502
Under 5 mortality	0.0495	0.217	21,502
Time from previous birth ¡= 18 months	0.126	0.332	21,502
Younger siblings between 0 and 5	0.604	0.445	21,502
Twin births between 0 and 5	0.0104	0.102	21,502

Tunisia

VARIABLES	(1) mean	(2) sd	(3) N
year of birth	1,975	6.107	11,986
Female child	0.490	0.500	11,986
Age of mother at birth	25.52	5.361	11,986
Under 5 mortality	0.118	0.323	11,986
Time from previous birth ¡= 18 months	0.213	0.409	11,986
Younger siblings between 0 and 5	0.722	0.490	11,986
Twin births between 0 and 5	0.0187	0.135	11,986

Turkey

VARIABLES	(1) mean	(2) sd	(3) N
year of birth	1,998	6.556	14,223
Female child	0.488	0.500	14,223
Age of mother at birth	24.00	4.978	14,223
Under 5 mortality	0.0509	0.220	14,223
Time from previous birth ¡= 18 months	0.148	0.355	14,223
Younger siblings between 0 and 5	0.439	0.493	14,223
Twin births between 0 and 5	0.00696	0.0831	14,223

Uganda

VARIABLES	(1) mean	(2) sd	(3) N
year of birth	2,002	6.413	42,464
Female child	0.499	0.500	42,464
Age of mother at birth	23.65	5.947	42,464
Under 5 mortality	0.115	0.319	42,464
Time from previous birth ¡= 18 months	0.122	0.327	42,464
Younger siblings between 0 and 5	0.641	0.434	42,464
Twin births between 0 and 5	0.0210	0.143	42,464

Uzbekistan

VARIABLES	(1) mean	(2) sd	(3) N
year of birth	1,982	5.888	7,306
Female child	0.490	0.500	7,306
Age of mother at birth	24.67	4.599	7,306
Under 5 mortality	0.0617	0.241	7,306
Time from previous birth ¡= 18 months	0.128	0.334	7,306
Younger siblings between 0 and 5	0.570	0.453	7,306
Twin births between 0 and 5	0.00985	0.0988	7,306

Vietnam			
	(1)	(2)	(3)
VARIABLES	mean	sd	N
year of birth	1,988	5.806	12,172
Female child	0.486	0.500	12,172
Age of mother at birth	24.95	4.768	12,172
Under 5 mortality	0.0522	0.222	12,172
Time from previous birth ¡= 18 months	0.0755	0.264	12,172
Younger siblings between 0 and 5	0.411	0.409	12,172
Twin births between 0 and 5	0.00353	0.0593	12,172

Yemen			
	(1)	(2)	(3)
VARIABLES	mean	sd	N
year of birth	2,000	6.193	48,509
Female child	0.484	0.500	48,509
Age of mother at birth	23.90	5.949	48,509
Under 5 mortality	0.0855	0.280	48,509
Time from previous birth ¡= 18 months	0.282	0.450	48,509
Younger siblings between 0 and 5	0.786	0.574	48,509
Twin births between 0 and 5	0.0109	0.104	48,509

G Results of the First and Second Stage of the IV regression (Instrument: Twins)[22]

Colomn 2 presents the detailed δ coefficients for each country in our sample. Colomn 1 reports the corresponding First Stage when instrumenting sibling competition with the presence of twins among younger siblings. All the First Stage are strongly significantat the 1% level.

Colomn 4 shows the equivalent results when using the number of younger siblings at age 5 instead our measure of siblings competition. Note that because the number of younger siblings at age 5 is, by construction, greater than the value of our measure of sibling competition, the estimated coefficient are consistantly smaller than those of colomn 2. Colomn 3 displays the corresponding First Stage.

Mortality, Instrument: Twins

	(1) FS-Competition 0-5	(2) Competition 0-5	(3) FS-Competition at 5	(4) Competition at 5
Afghanistan	0.664***	0.0280	1.034***	0.0180
	(0.0455)	(0.0369)	(0.0607)	(0.0237)
Albania	0.667***	0.0439	1.237***	0.0237
	(0.0783)	(0.0424)	(0.127)	(0.0236)
Armenia	0.754***	0.105**	1.363***	0.0583**
	(0.0889)	(0.0465)	(0.116)	(0.0278)
Azerbaijan	0.783***	0.208***	1.322***	0.123***
	(0.0853)	(0.0719)	(0.108)	(0.0428)
Bangladesh	0.596***	0.0448	1.022***	0.0261
	(0.0417)	(0.0418)	(0.0770)	(0.0248)
Burundi	0.505***	0.125***	0.878***	0.0722***
	(0.0298)	(0.0379)	(0.0437)	(0.0224)
Cameroon	0.603***	0.0382	1.043***	0.0221
	(0.0224)	(0.0250)	(0.0380)	(0.0146)

[22] Controls: Urban vs. rural, Mother's education, Year of birth of mother, Birth Spacing FE (+ interaction with gender), Year of birth FE, Age of the mother at birth FE, Birth Rank FE, Child is the first born of her sex dummy, gender dummy, Child has a twin dummy. Standard Errors clustered at mother level

Colombia	0.869***	-0.0133	1.551***	-0.00744
	(0.0518)	(0.0109)	(0.0662)	(0.00604)
CotedIvoire	0.584***	0.0968**	1.003***	0.0564**
	(0.0342)	(0.0380)	(0.0706)	(0.0229)
Ecuador	0.723***	0.0149	1.149***	0.00934
	(0.0639)	(0.0486)	(0.0989)	(0.0306)
Egypt	0.626***	0.0845***	1.187***	0.0446***
	(0.0282)	(0.0226)	(0.0463)	(0.0127)
Gabon	0.826***	0.0193	1.452***	0.0110
	(0.0677)	(0.0289)	(0.0801)	(0.0164)
India	0.654***	0.0727***	1.110***	0.0428***
	(0.00927)	(0.00850)	(0.0164)	(0.00508)
Jordan	0.615***	0.0338	1.248***	0.0166
	(0.0342)	(0.0218)	(0.0470)	(0.0110)
Kenya	0.708***	0.00530	1.309***	0.00286
	(0.0336)	(0.0191)	(0.0570)	(0.0103)
Kyrgyzstan	0.660***	0.0966	1.272***	0.0501
	(0.0580)	(0.0845)	(0.0684)	(0.0418)
Mexico	0.649***	0.0180	1.050***	0.0111
	(0.0622)	(0.0490)	(0.0972)	(0.0302)
Moldova	0.913***	0.0615	1.529***	0.0367
	(0.110)	(0.0488)	(0.197)	(0.0296)
Nepal	0.538***	0.144**	0.984***	0.0788**
	(0.0398)	(0.0624)	(0.0695)	(0.0326)
Niger	0.565***	0.0116	0.942***	0.00693
	(0.0381)	(0.0363)	(0.0607)	(0.0219)

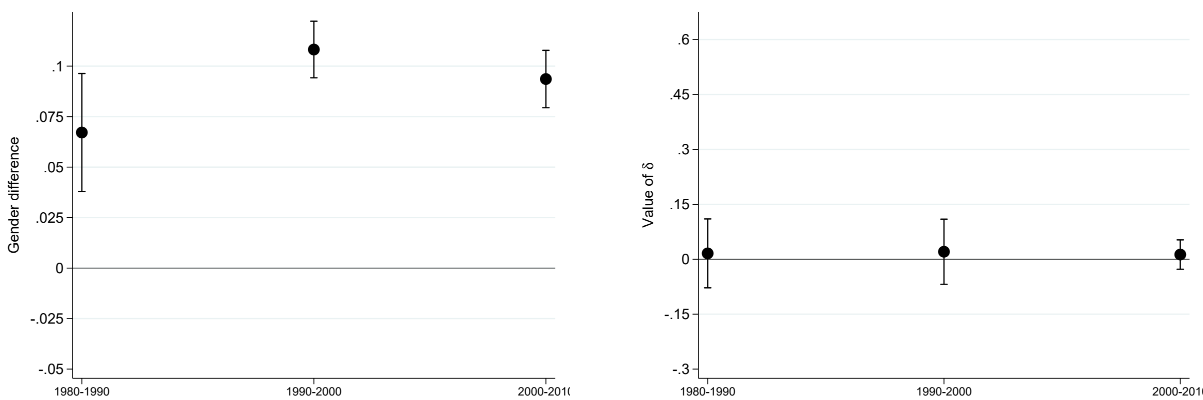
Pakistan	0.662***	0.0232	1.121***	0.0137
	(0.0464)	(0.0297)	(0.0748)	(0.0177)
Senegal	0.567***	0.0813*	1.130***	0.0409*
	(0.0311)	(0.0429)	(0.0564)	(0.0219)
SierraLeone	0.702***	0.0524	1.080***	0.0341
	(0.0613)	(0.0625)	(0.115)	(0.0415)
SriLanka	0.617***	0.0549	1.042***	0.0325
	(0.0460)	(0.0382)	(0.0812)	(0.0226)
Tajikistan	0.606***	0.0324	1.118***	0.0175
	(0.0353)	(0.0372)	(0.0767)	(0.0204)
TimorLeste	0.680***	0.0152	1.322***	0.00782
	(0.0518)	(0.0265)	(0.0842)	(0.0137)
Tunisia	0.525***	0.137***	0.894***	0.0801***
	(0.0387)	(0.0466)	(0.0687)	(0.0282)
Turkey	0.924***	0.0171	1.453***	0.0109
	(0.0759)	(0.0247)	(0.107)	(0.0159)
Uganda	0.561***	0.0537**	1.050***	0.0287**
	(0.0285)	(0.0257)	(0.0496)	(0.0139)
Uzbekistan	0.551***	0.0202	1.094***	0.0102
	(0.0576)	(0.0598)	(0.102)	(0.0301)
Vietnam	0.736***	-0.00216	1.396***	-0.00114
	(0.0473)	(0.0490)	(0.104)	(0.0258)
Yemen	0.674***	0.0292	1.134***	0.0173
	(0.0438)	(0.0258)	(0.0677)	(0.0157)

Standard errors in parentheses

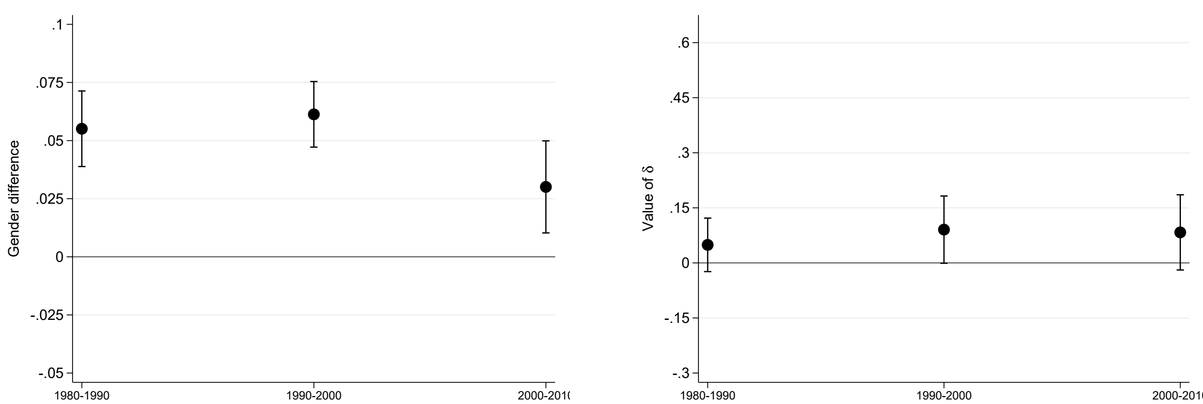
* $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$

H Evolution of country level estimates of mortality caused by sibling competitions

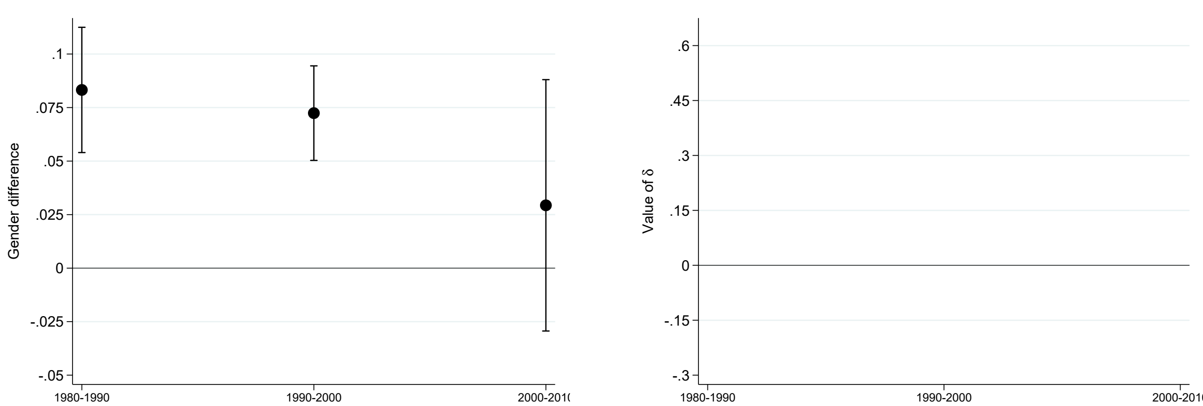
Albania



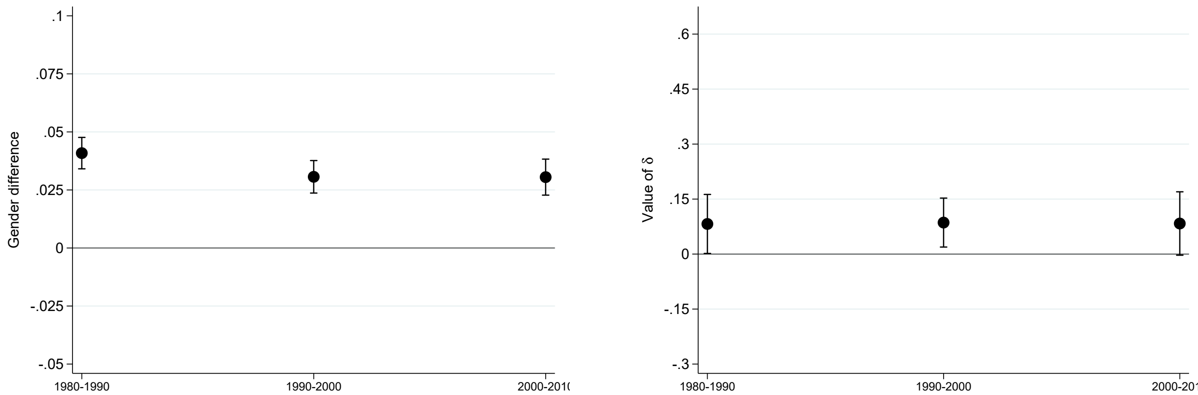
Armenia



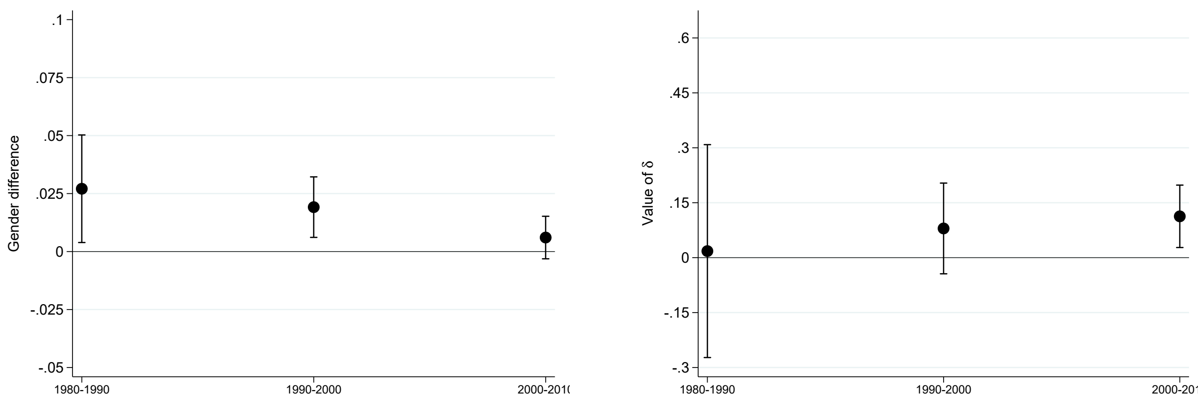
Azerbaijan



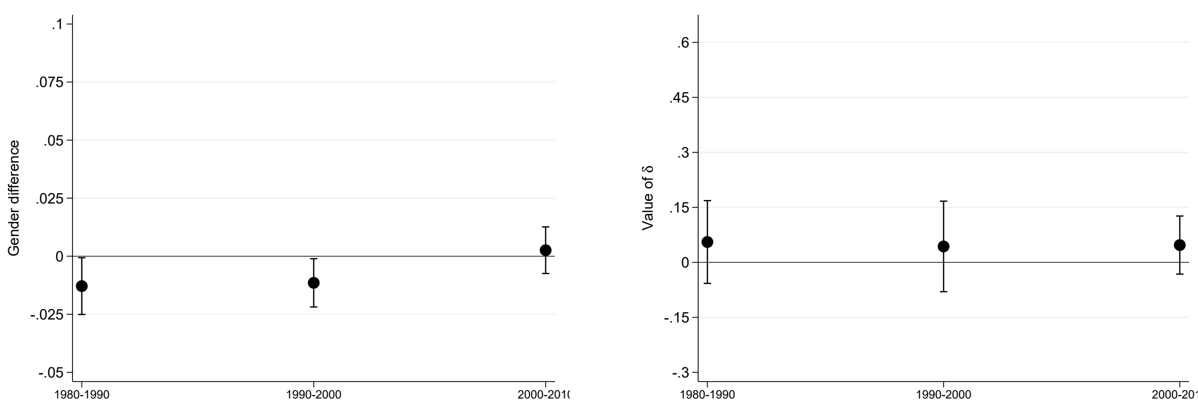
Bangladesh



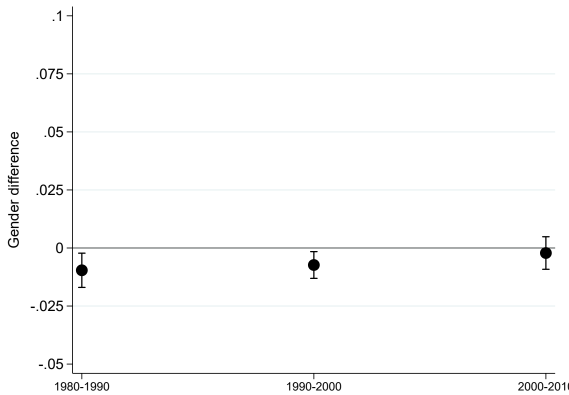
Bangladesh

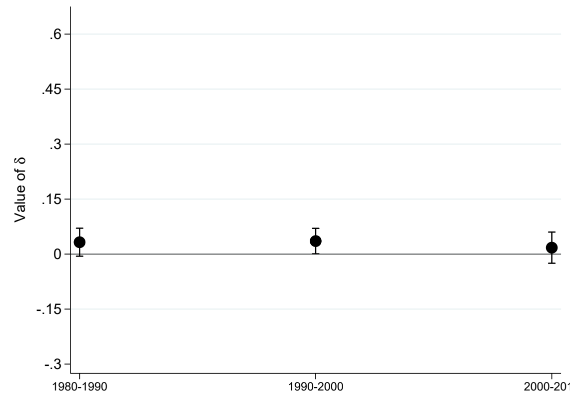


Chad

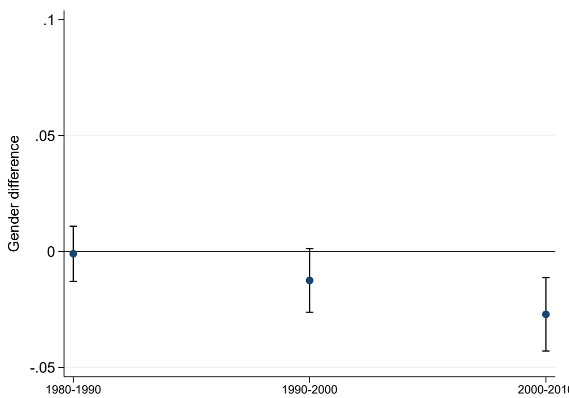


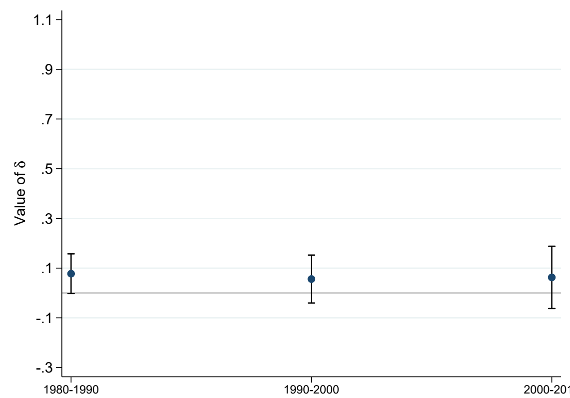
Colombia



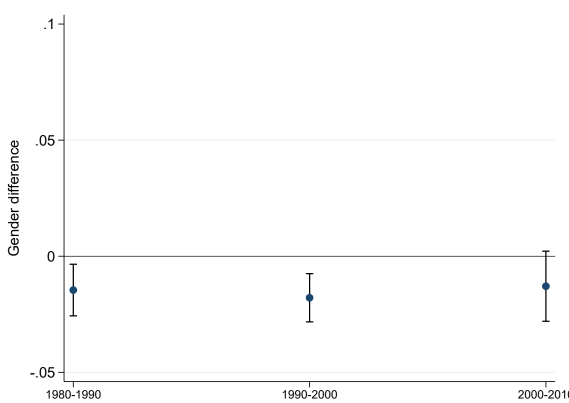


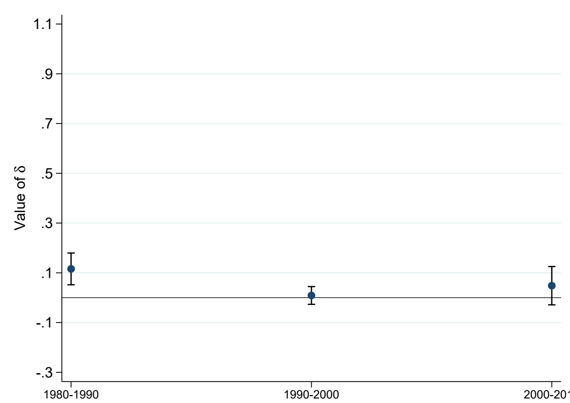
Cote d'Ivoire



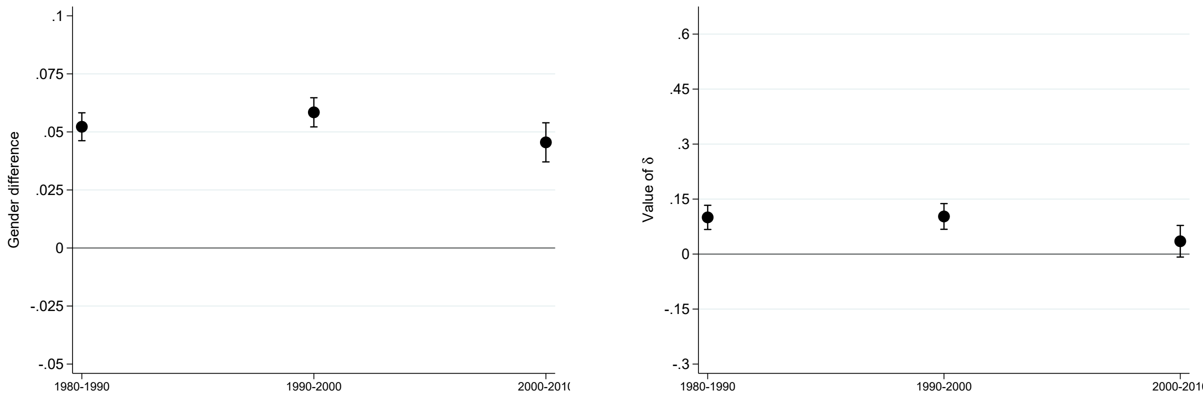


Cote d'Ivoire

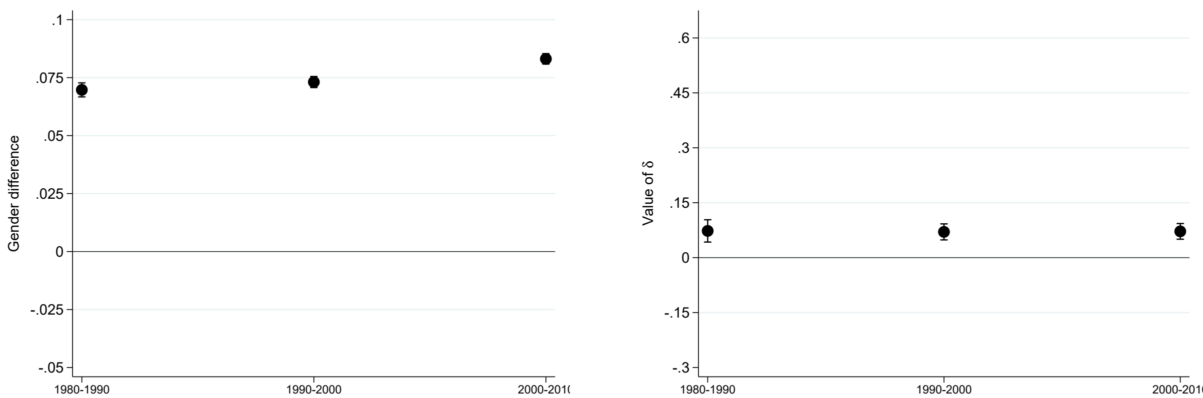




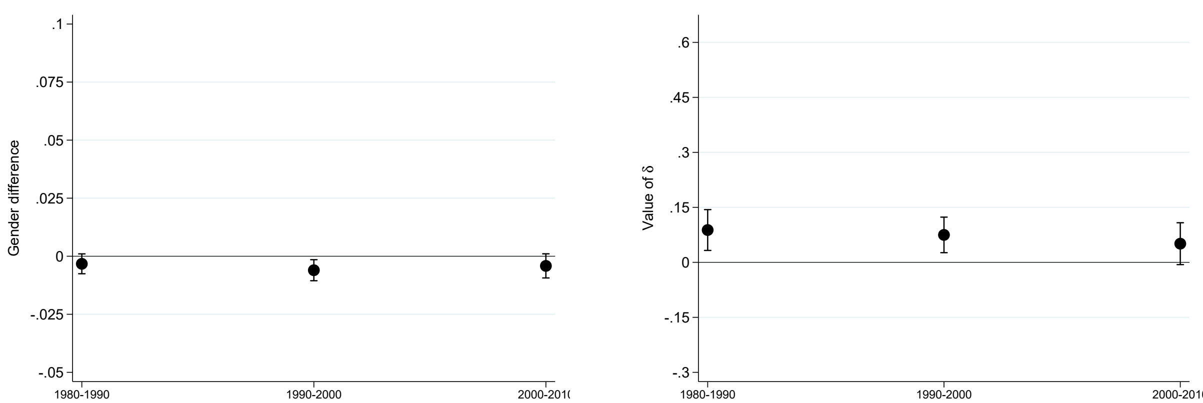
Egypt



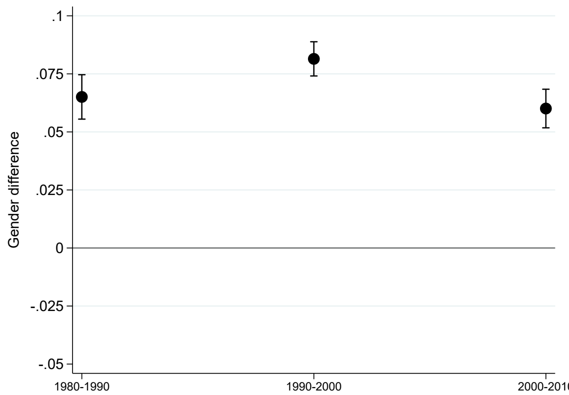
India

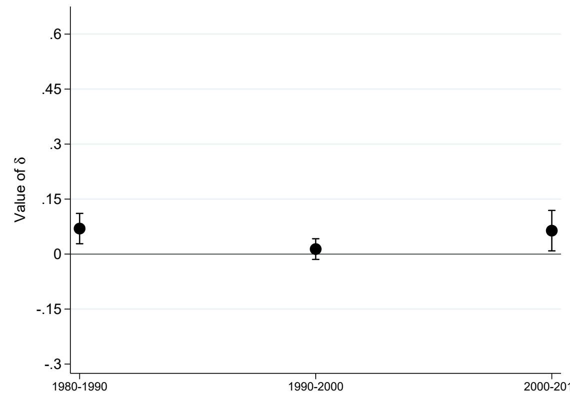


Indonesia

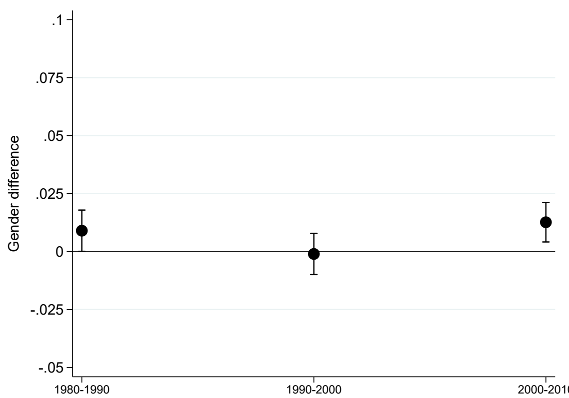


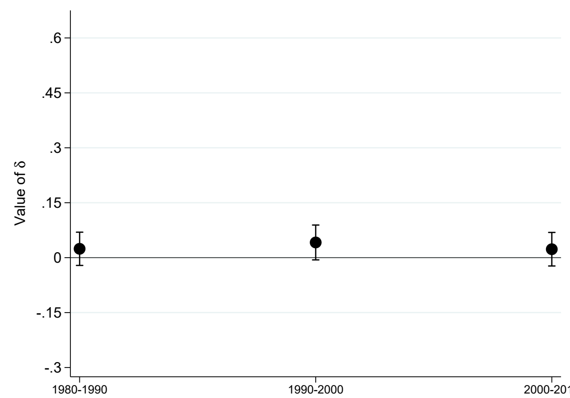
Jordan



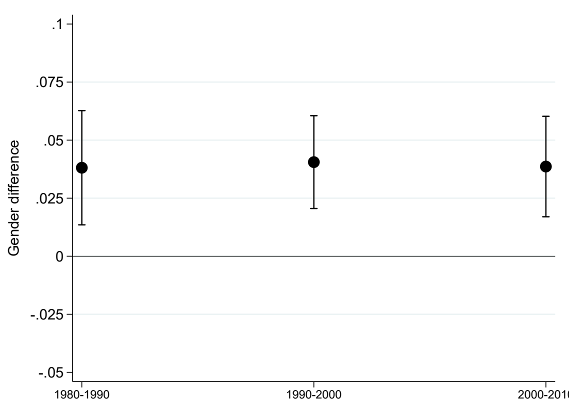


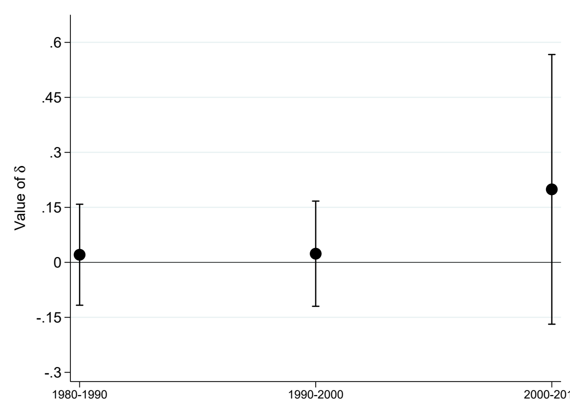
Kenya



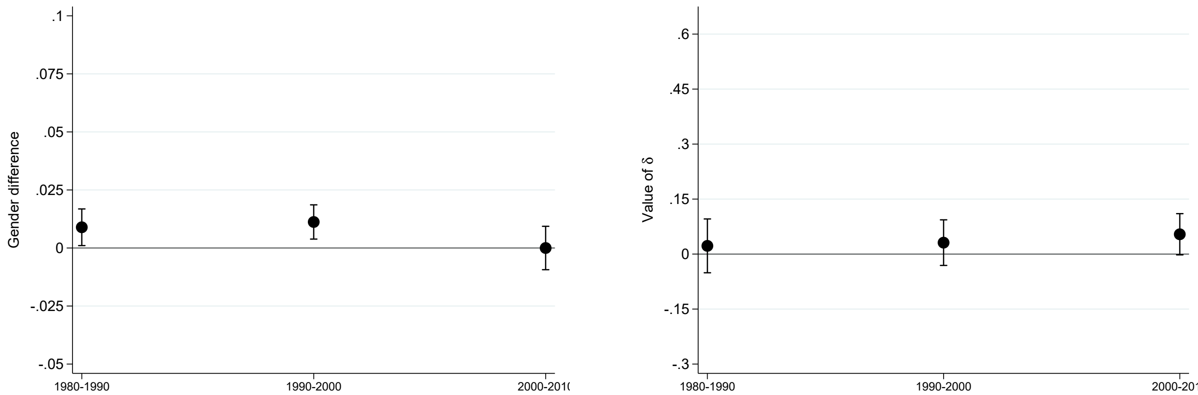


Kyrgyzstan

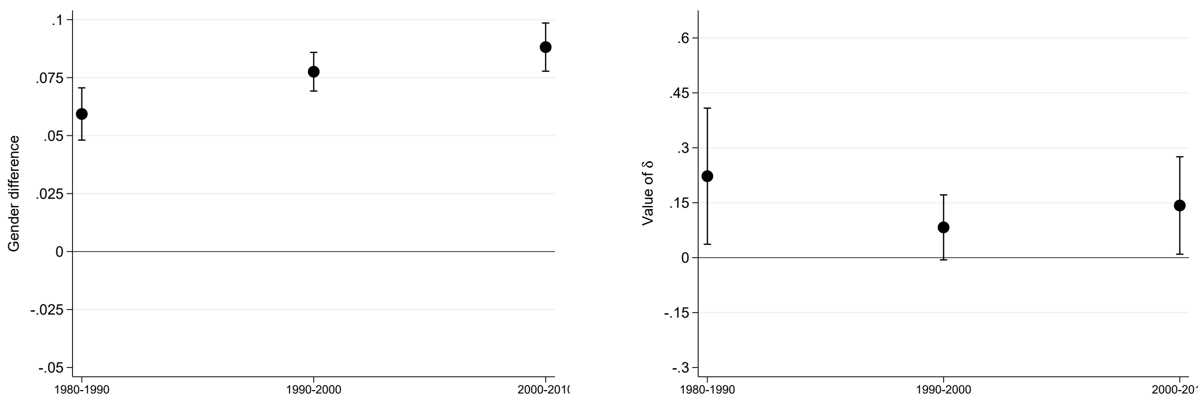




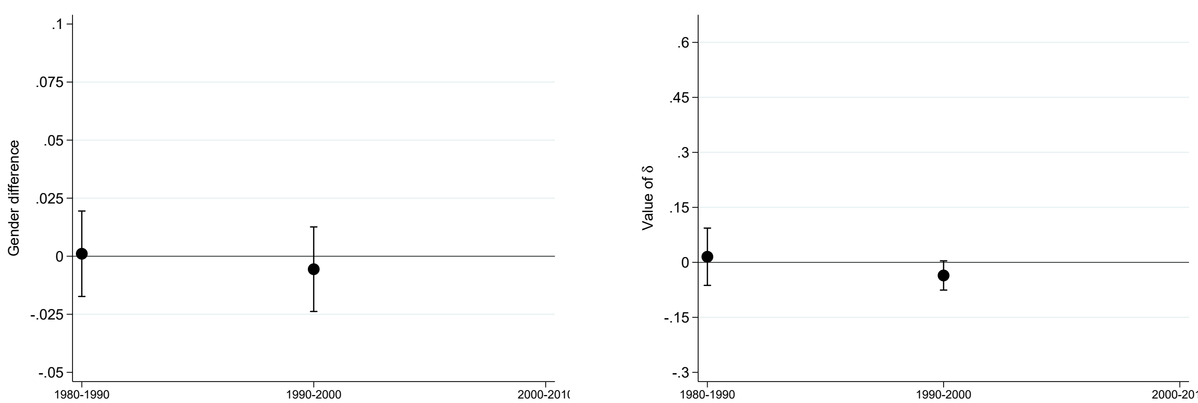
Mali



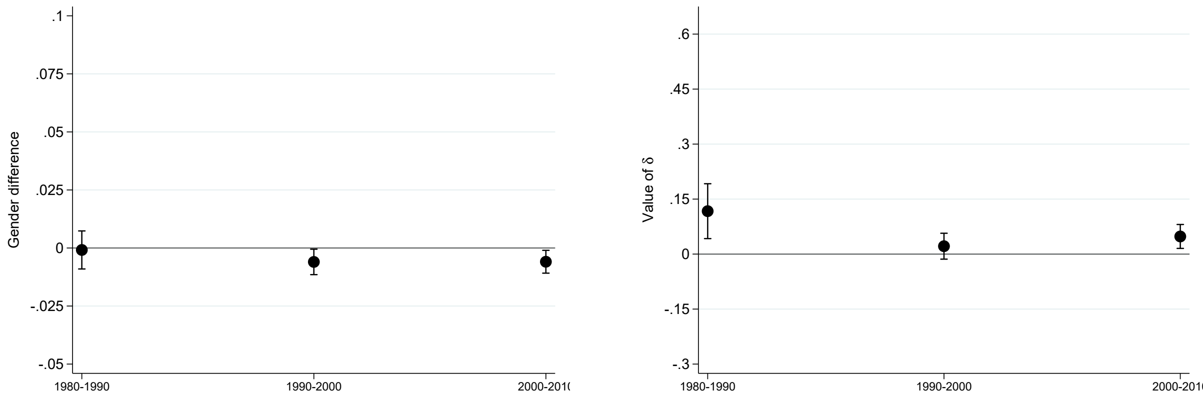
Nepal



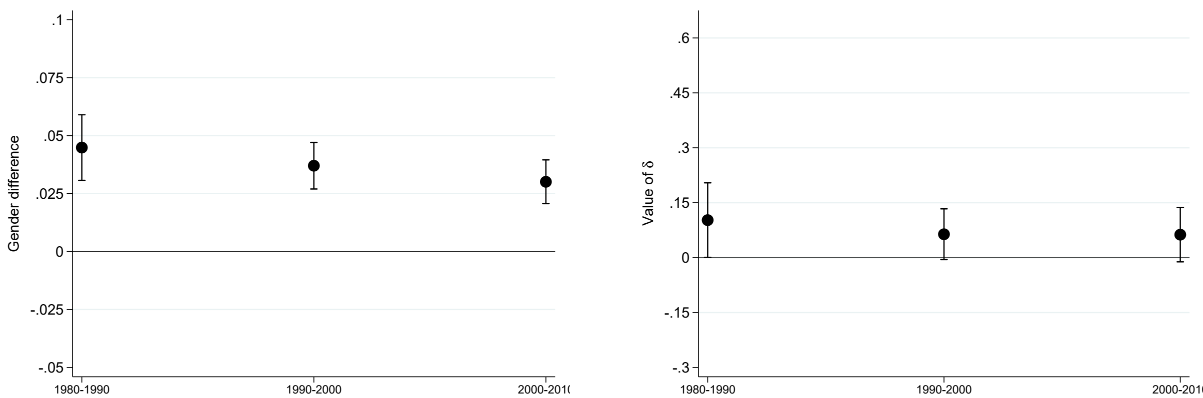
Nepal



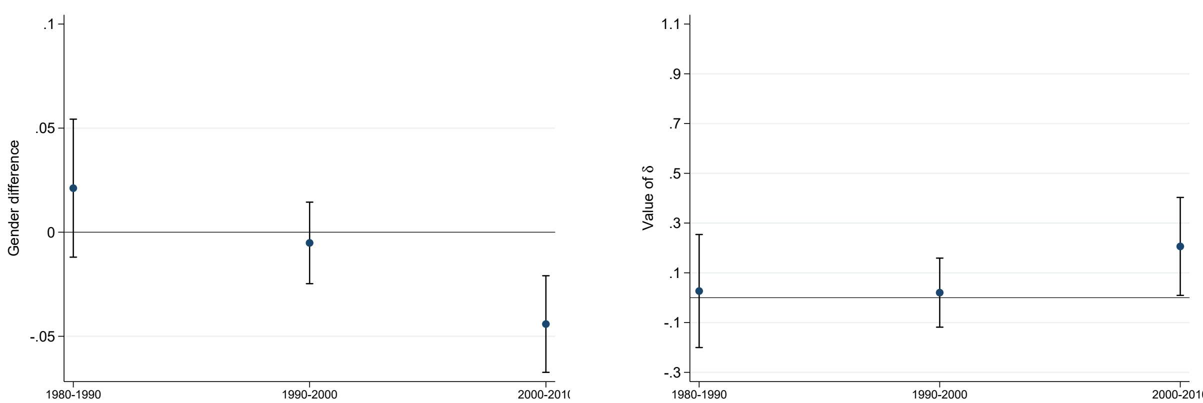
Nepal



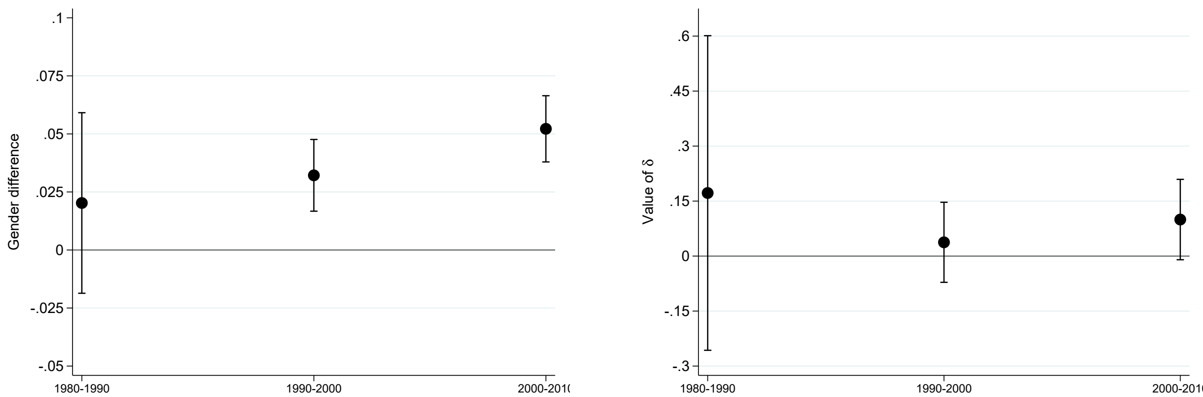
Pakistan



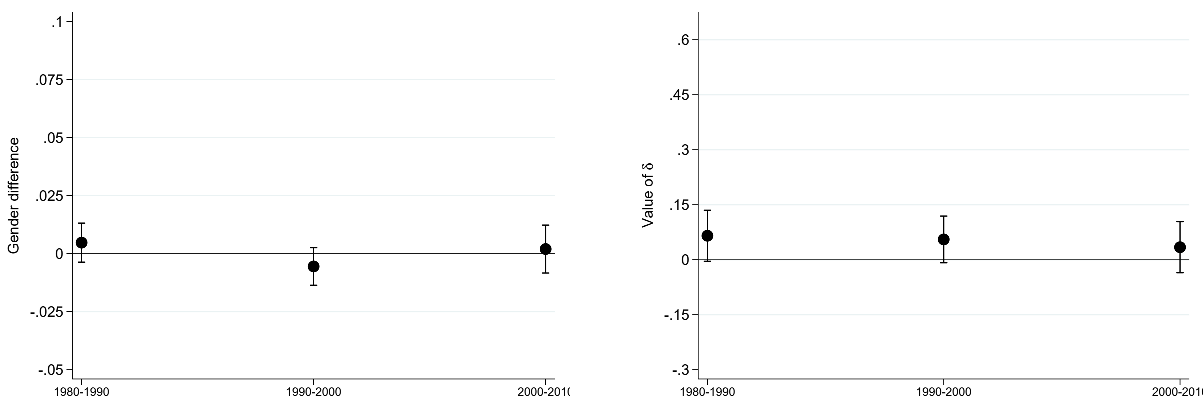
Sierra Leone



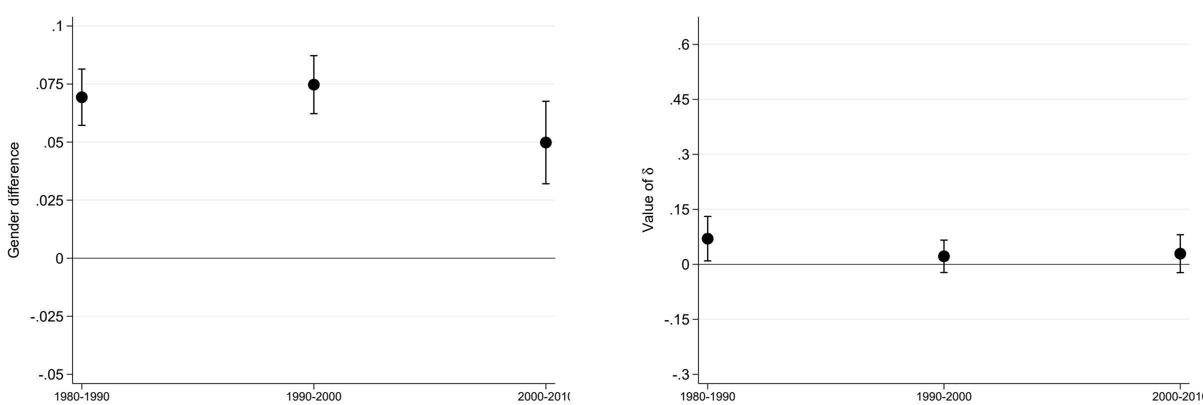
Tajikistan



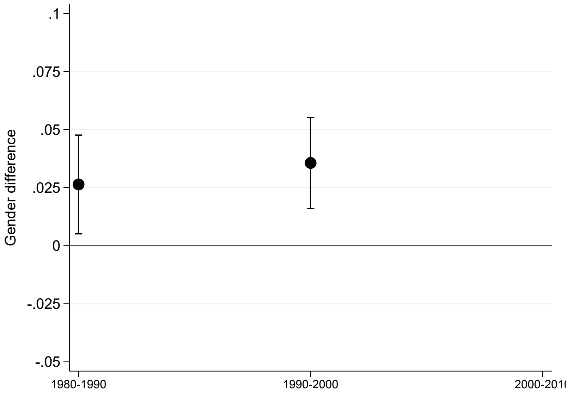
Timor-Leste

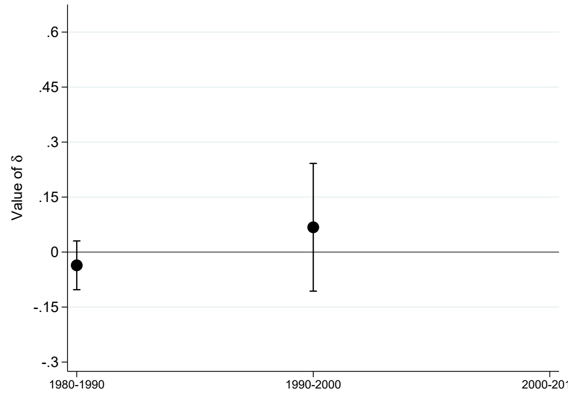


Turkey

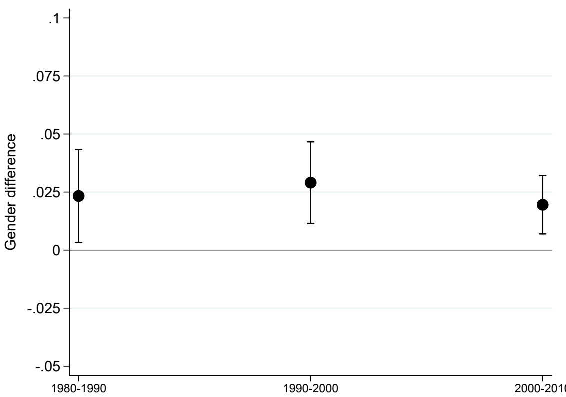


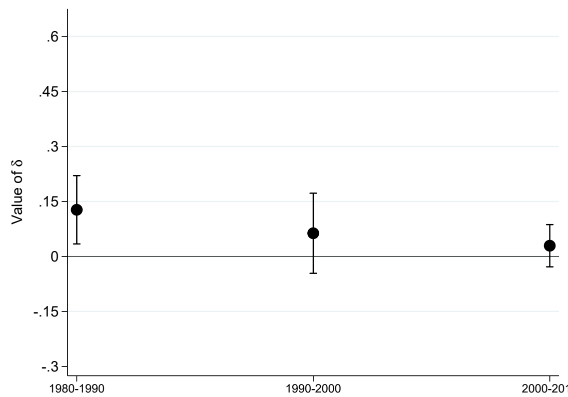
Uganda





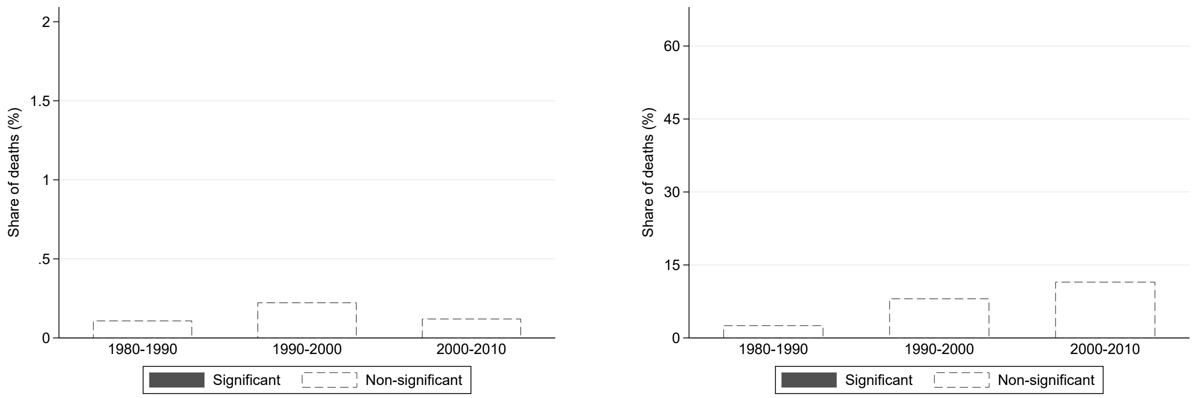
Yemen



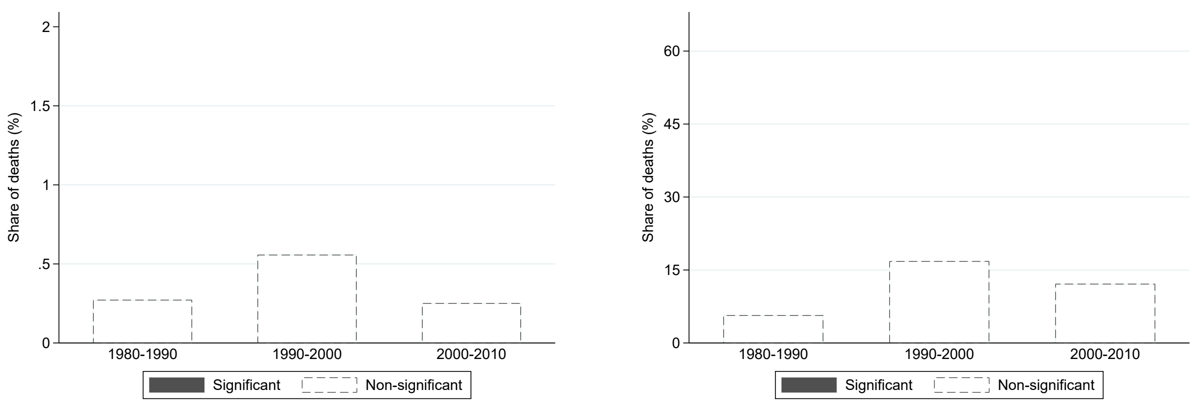


I Evolution of country level estimates of mortality caused by sibling competitions

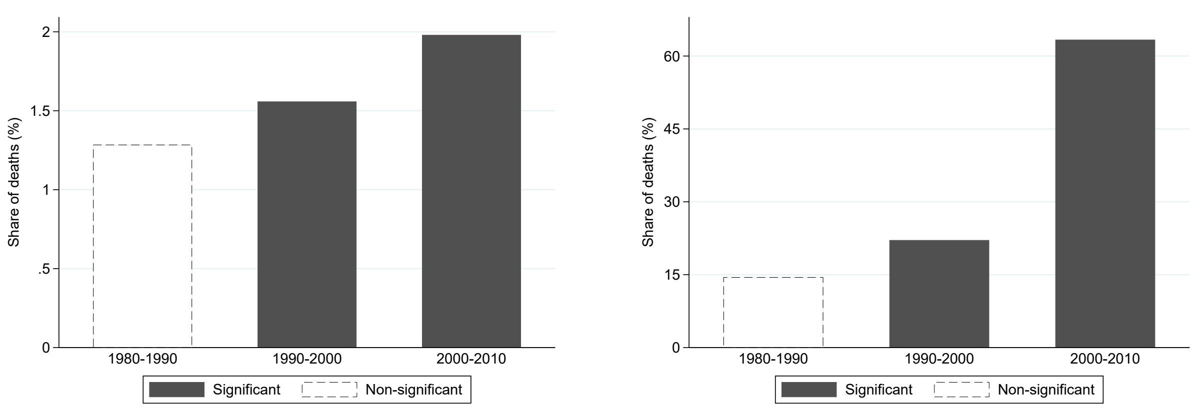
Albania



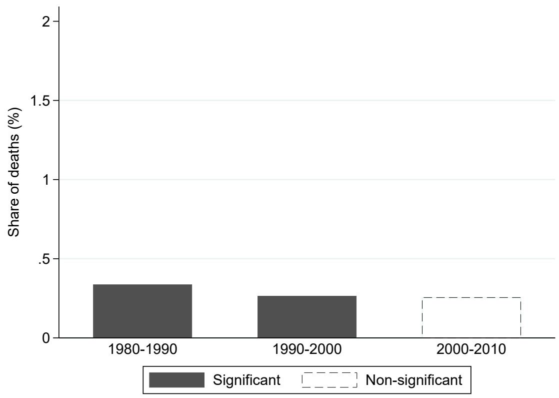
Armenia

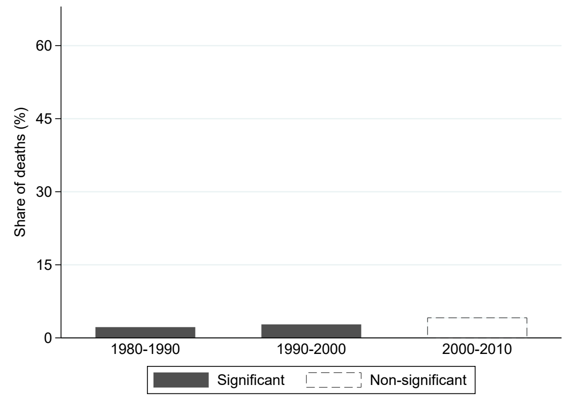


Armenia

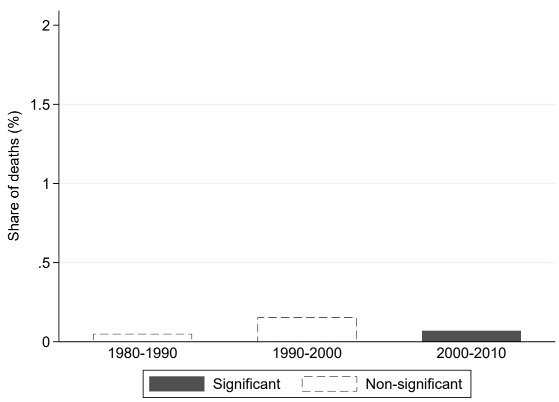


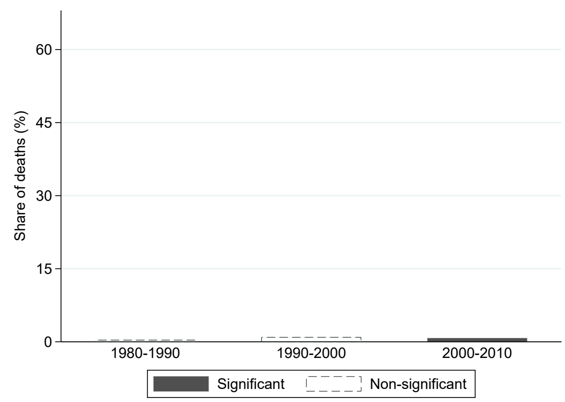
Bangladesh



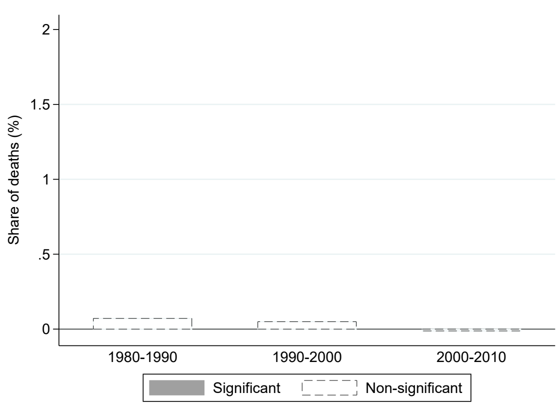


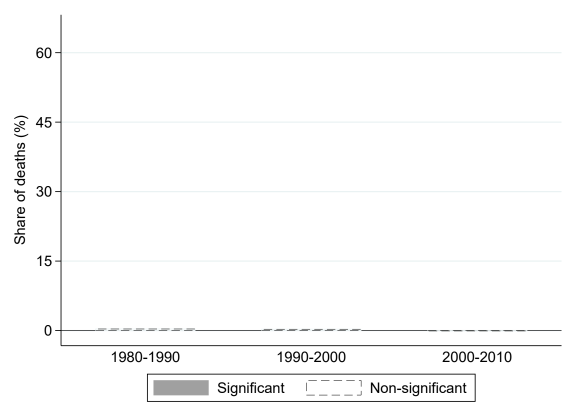
Chad



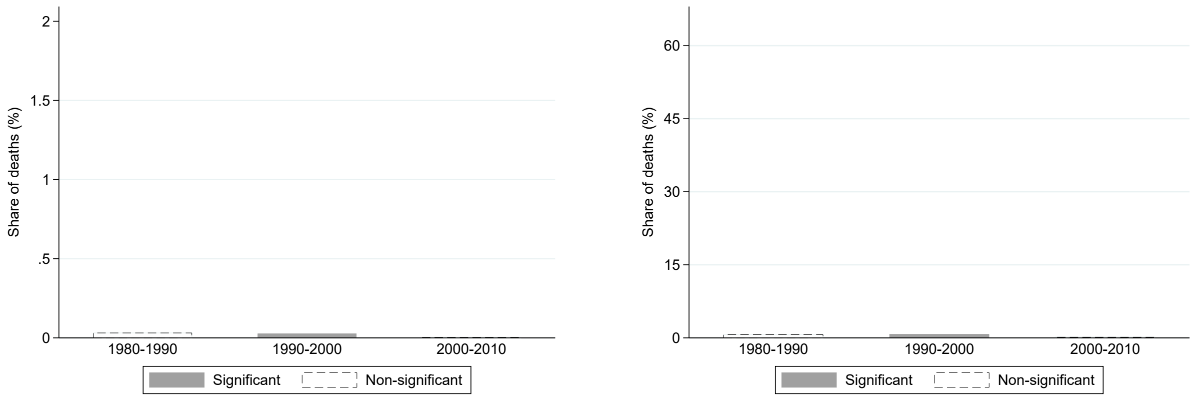


Chad



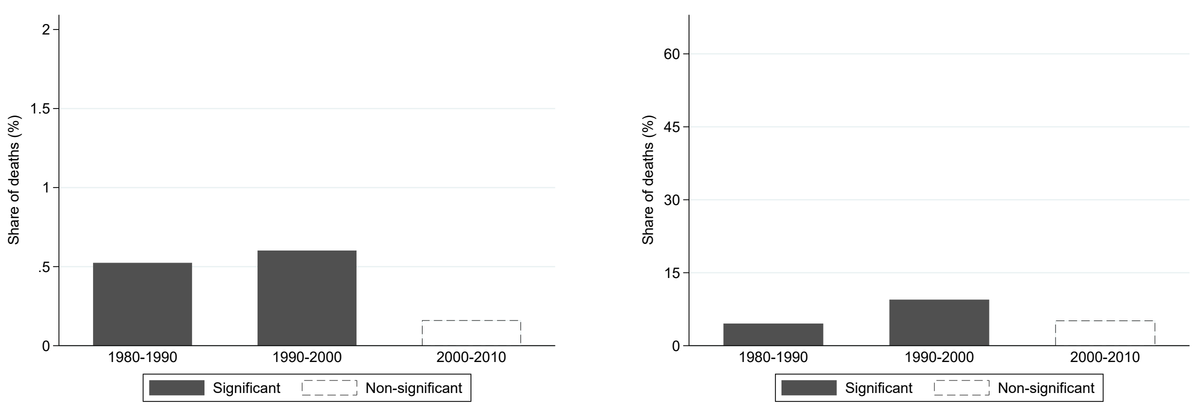


Colombia

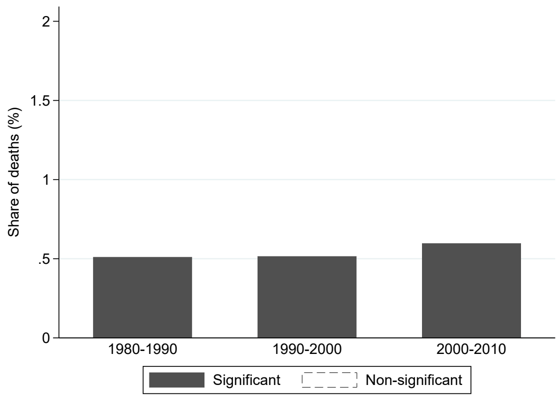


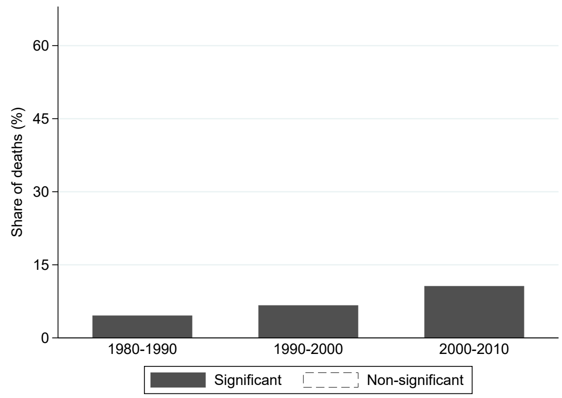
Cote d'Ivoire

Egypt

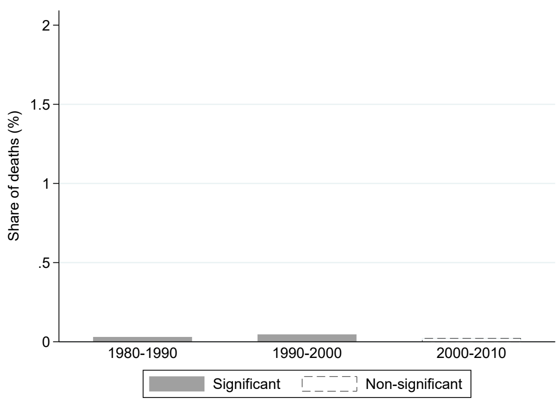


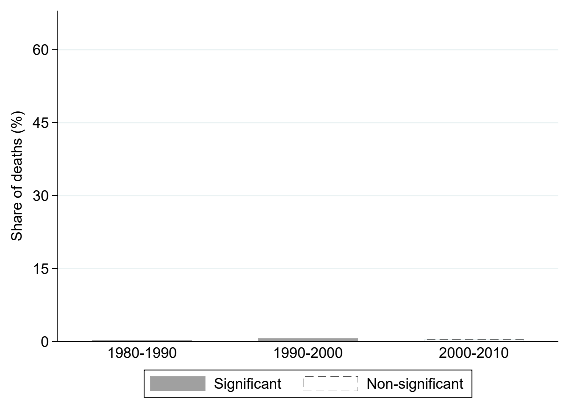
India



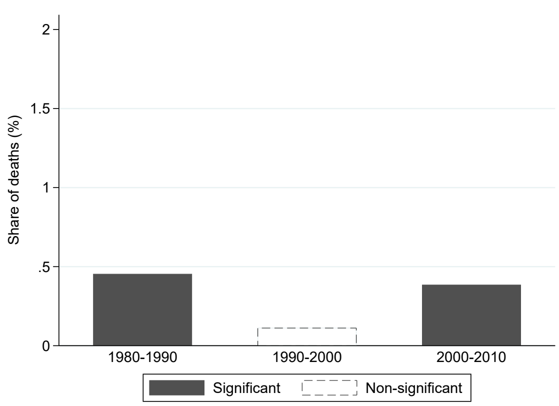


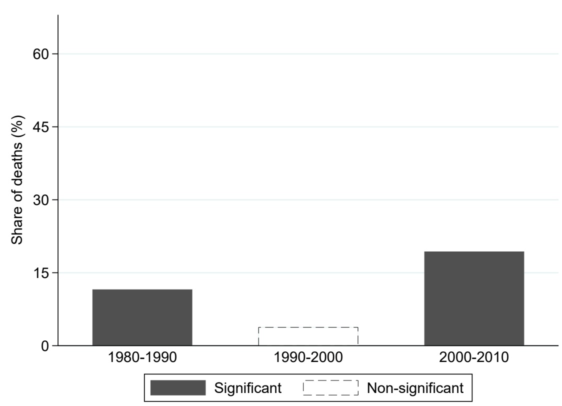
Indonesia



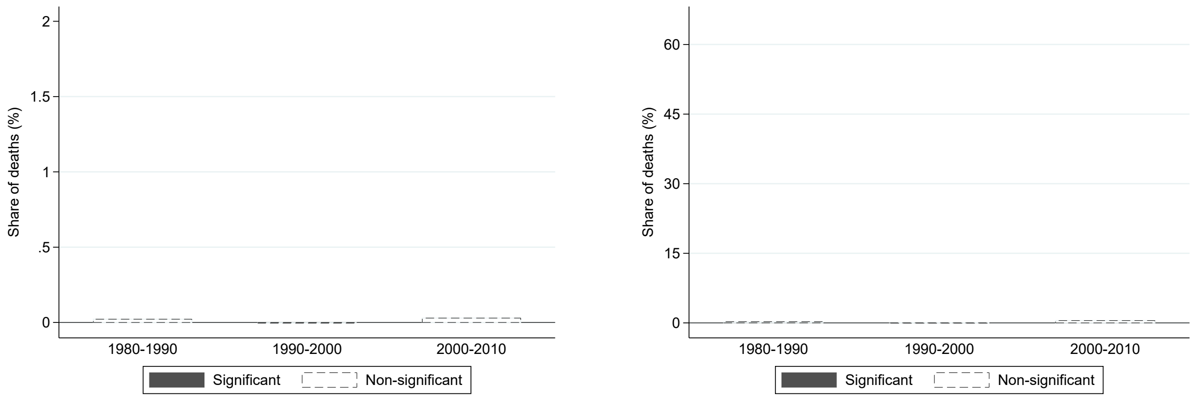


Jordan

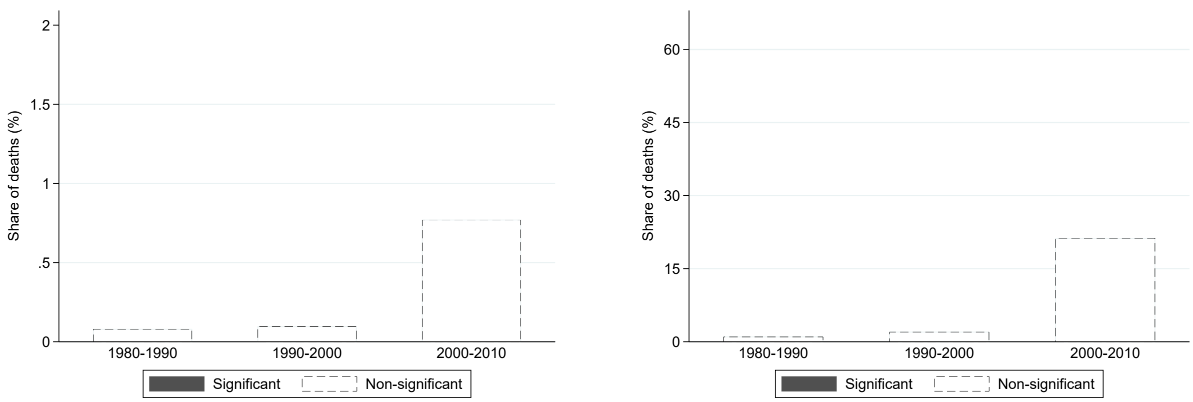




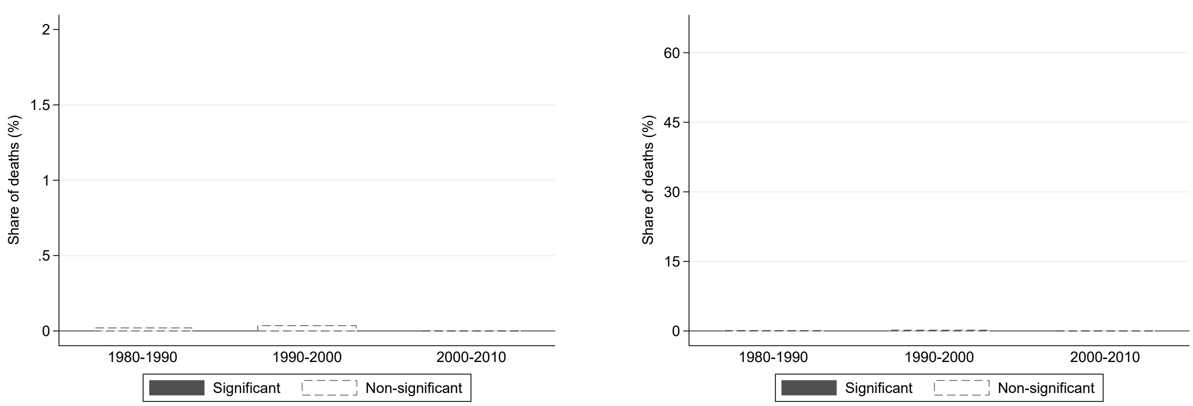
Kenya



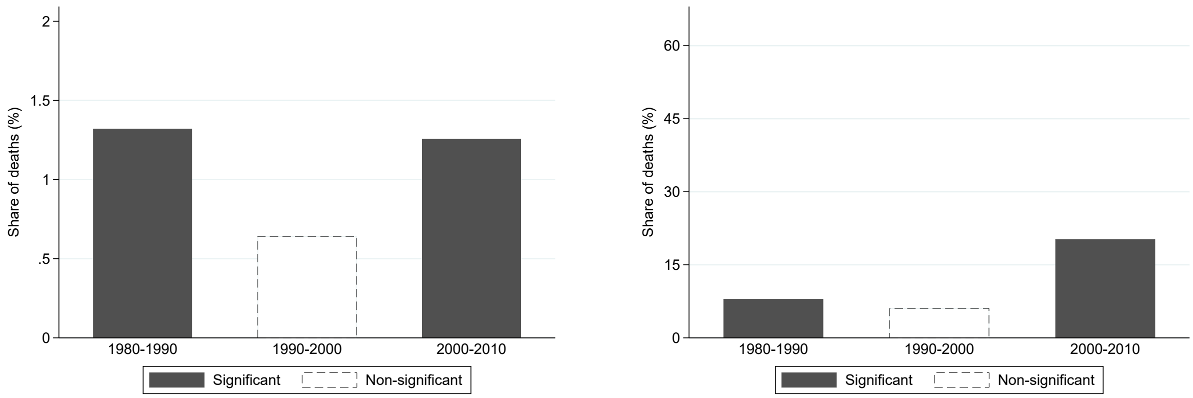
Kyrgyzstan



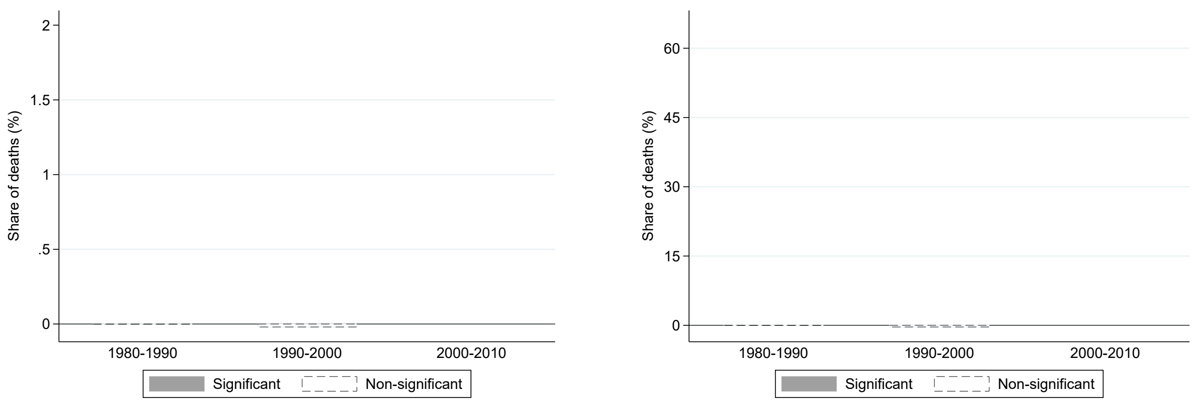
Mali



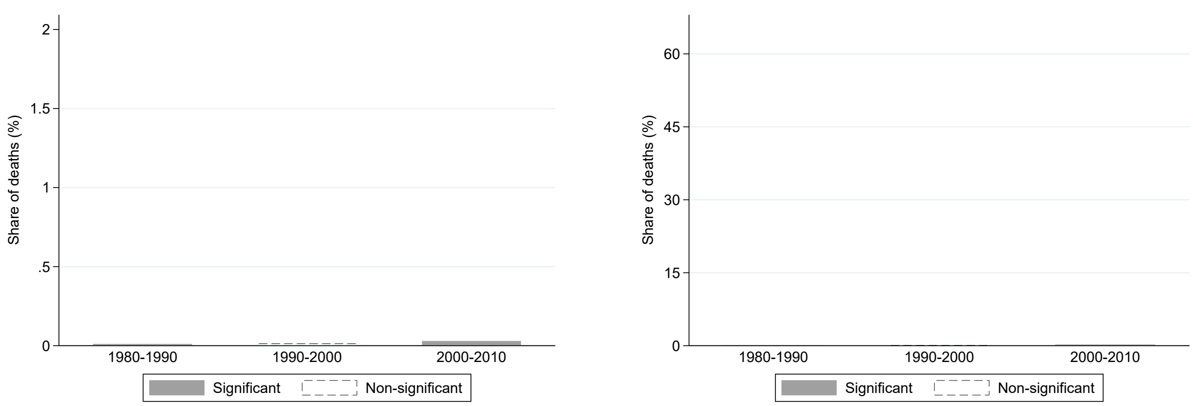
Nepal



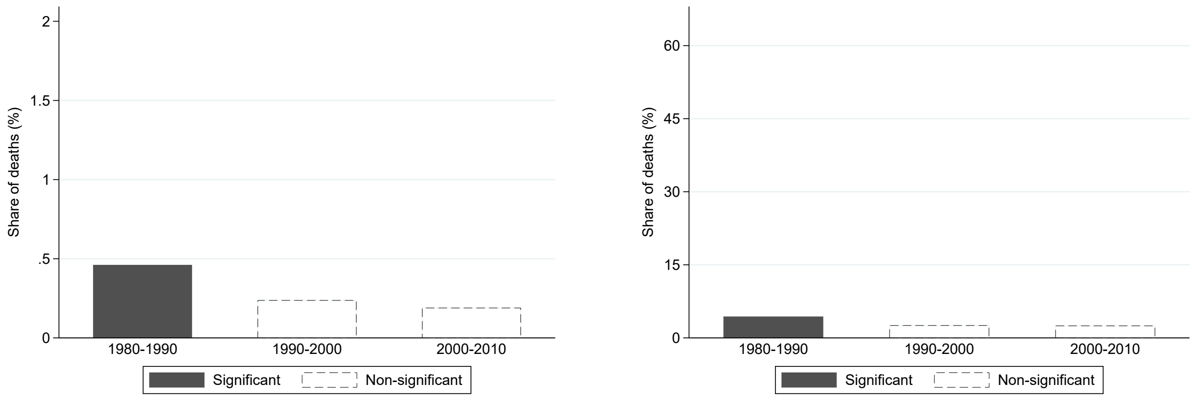
Pakistan



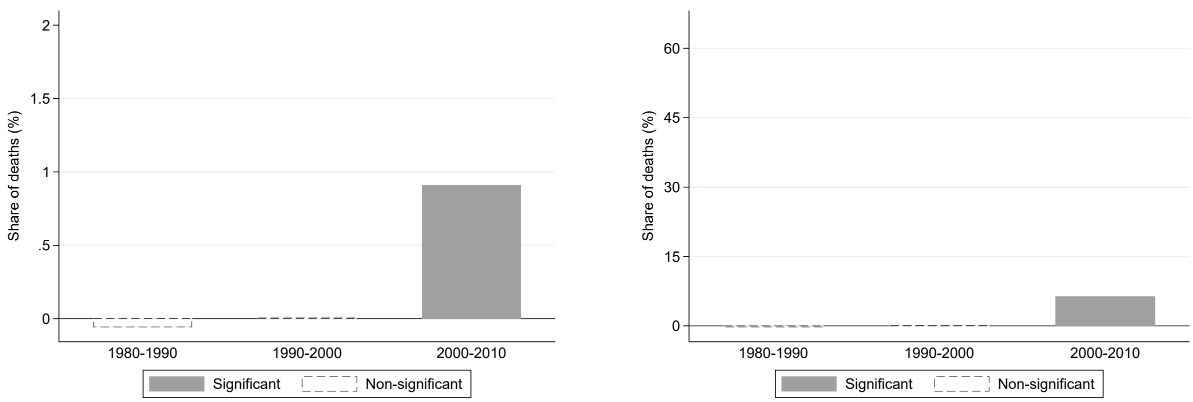
Pakistan



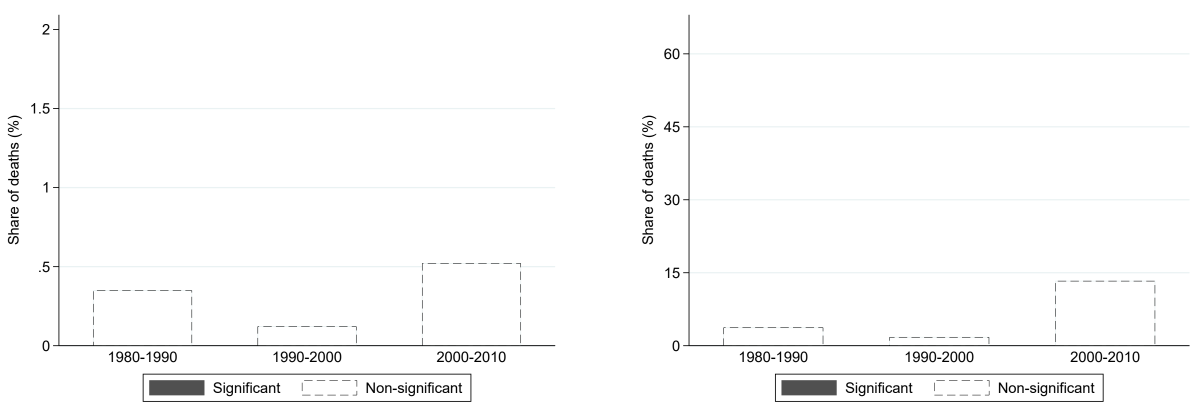
Pakistan



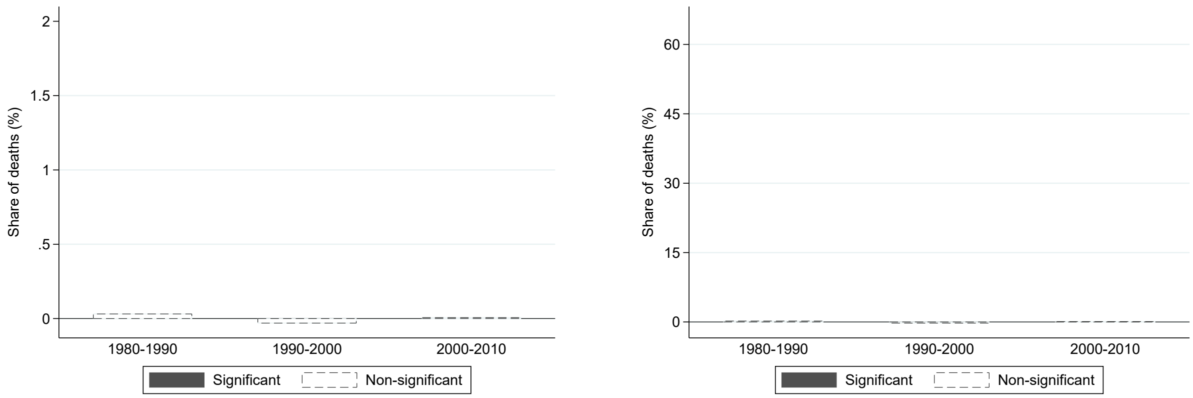
Sierra Leone



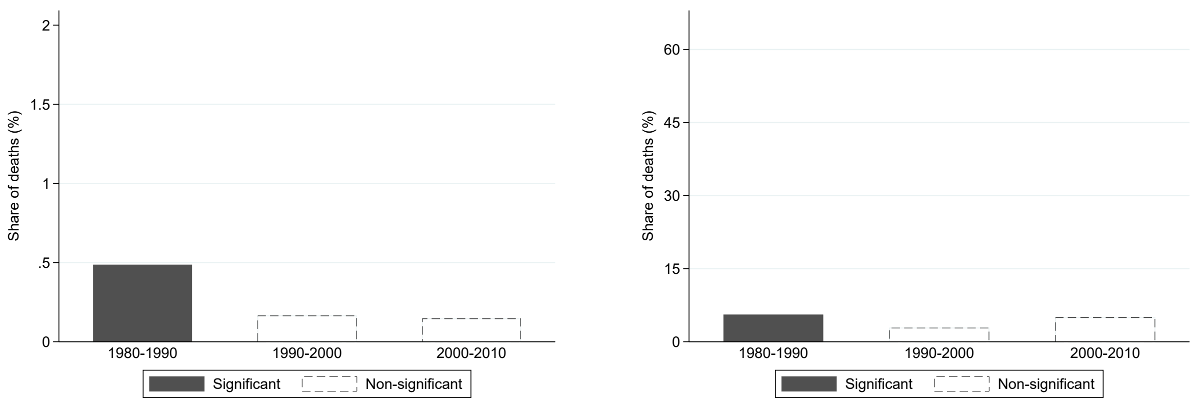
Tajikistan



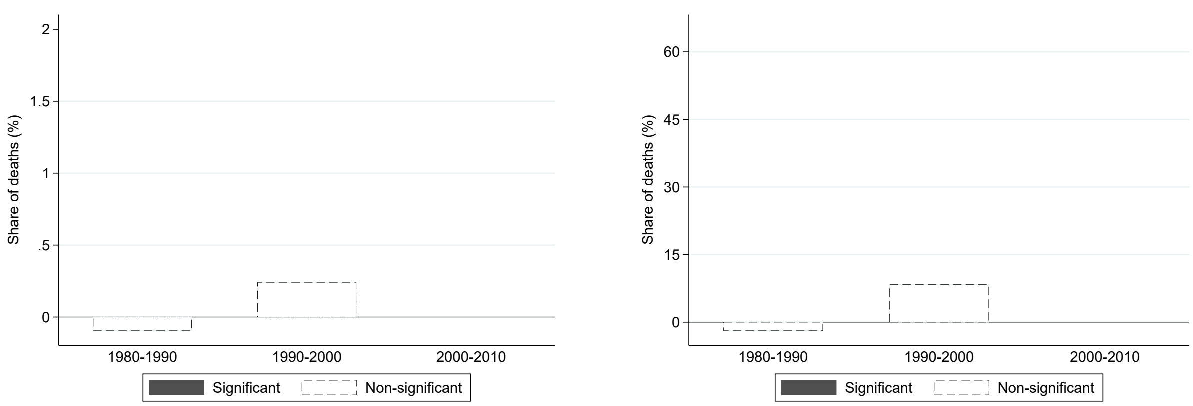
Turkey



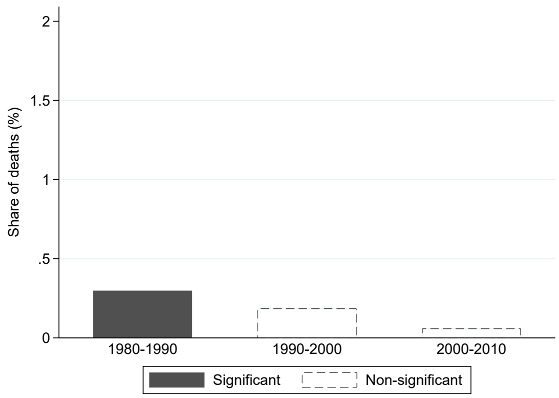
Turkey



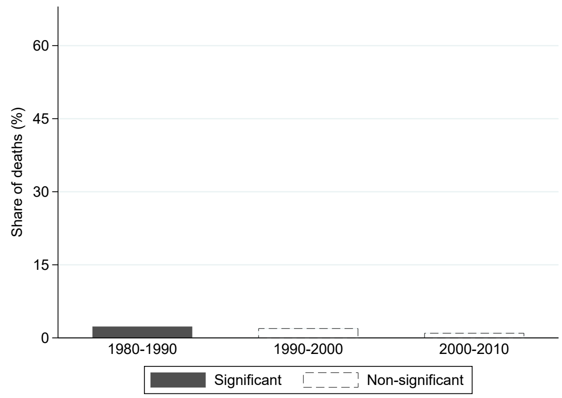
Turkey



Yemen



Share of deaths (%)
2
1.5
1
.5
0
1980-1990
1990-2000
2000-2010
Significant
Non-significant



Share of deaths (%)
60
45
30
15
0
1980-1990
1990-2000
2000-2010
Significant
Non-significant