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Child Marriage and Infant Mortality: Evidence from Ethiopia

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Child Marriage and Infant Mortality: Evidence from Ethiopia

By JORGE GARCIA-HOMBRADOS*

This study uses age discontinuities in exposure to a law that raised the legal age of marriage for women in Ethiopia to investigate the causal link between child marriage and infant mortality. Using a fuzzy regression discontinuity design, the study shows that laws banning underage marriages could be an effective strategy to tackle child marriage; and estimates that a one-year delay in women's age at cohabitation during teenage years reduces the probability of infant mortality of the first born by 3.8 percentage points. This impact is closely linked to the effect of delaying cohabitation on women's age at first birth.

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1. Introduction

In 2014, UNICEF reported that more than 700 million women worldwide first cohabited with a partner before the age of 18, with the vast majority living in developing countries (UNICEF, 2014). Considered by UNICEF a form of violence against women, decreasing the incidence of underage cohabitation, also known as child marriage[1], has become a priority for policy makers of international organizations and developing countries. In the last decades, most of the countries with a high prevalence of this practice have ratified different international agreements such as the CEDAW[2], CCMMAMRM[3] or the Maputo Protocol[4] that promote the setting and enforcement of minimum-age-of-marriage laws. Furthermore, the fight against child marriage mobilizes a large amount of resources in integrated programs and national alliances targeting the cultural, social and economic causes of this widespread practice. However, although the overall prevalence of child marriage is declining over time, its eradication is currently far from becoming a reality (Jensen and Thornton, 2003).

Using women-level data from different Asian and African countries, several studies reveal that child marriage is associated with worse levels of health, education, labor force participation, offspring mortality and participation in household decisions (Parsons et al., 2015). However, the link between child marriage and these outcomes might be driven in part by unobservable traits or by reverse causality and therefore, the statistical associations identified in *correlational* studies should not be interpreted as the causal effects of child marriage. Relying on parental anxiety for marrying off their daughters once they reach puberty, some researchers address the endogeneity in the link between child marriage and socioeconomic outcomes through using age at menarche as an instrumental variable for age at marriage. Using this approach, a handful of studies confirm the negative effects of child marriage on women's education and health (Field and Ambrus, 2008); and document the intergenerational effects on the education, health and cultural preferences of their children (Chari et al., 2017; Asadullah et al., 2016). However, although this instrumental variable approach dominates the literature on the causal effects of child marriage, evidence from

[1]UNICEF defines child marriage as the formal marriage or unmarried cohabitation before the age of 18 years. Although child marriage affects both girls and boys, the 82% of the children in the world that got married or started cohabiting with a partner before the age of 18 are girls (UNICEF, 2014).

[2]Convention on the Elimination of All Forms of Discrimination against Women, 1979.

[3]Convention on Consent to Marriage, Minimum Age for Marriage, and Registration of Marriage, 1964.

[4]Ratified in 2003.

recent medical studies showing how childhood experiences can affect age at menarche casts doubts on the assumptions behind the validity of this identification strategy[5].

With the objective of reducing the high prevalence of child marriage among Ethiopian girls, the Federal Government of Ethiopia approved in July 2000 the Revised Family Code (RFC). This law increased the legal age of marriage for women from 15 to 18 years in some regions of Ethiopia, while leaving unchanged at 18 the legal age of marriage for men. Despite the law banned underage marriage rather than underage cohabitation, the fact that unmarried cohabitation is heavily stigmatized in Ethiopia (Jones et al., 2016) makes it plausible that increasing the legal age of marriage for women lead to a delay in women's age at cohabitation. This study investigates the introduction of the RFC in some regions of Ethiopia to pursue a twofold objective. First, I investigate whether raising the legal age of marriage could help to increase women's age at cohabitation and reduce the prevalence of child marriage. Second, I investigate the causal effect of delaying women's age at cohabitation on the probability of infant mortality of the first born child, assessing also the mechanisms through which early cohabitation could affect infant mortality.

The novelty of the analysis presented in this study is based in three main contributions. First, although reducing child marriage and infant mortality are central priorities in the policy agenda of many developing countries (Weldearegawi et al., 2015) and the literature has documented the positive association between these two outcomes (see for example Raj et al. (2010)), it is unclear whether child marriage could be contributing to the high prevalence of infant mortality in many of these countries. This study addresses this gap in the literature providing the first causal estimates of the effect of child marriage on infant mortality and investigating the mechanisms driving this link. Second, this is to the best of my knowledge the first study that shows that raising the legal age of marriage could be an effective strategy to tackle child marriage and decrease the prevalence of infant mortality. Third, unlike previous studies relying on the use of age at menarche as an instrumental variable for child marriage, I address endogeneity in the link between women's age at cohabitation and socioeconomic outcomes through exploiting age discontinuities in the effective legal age of marriage faced by Ethiopian women. Using information on women's month and year of birth from the 2011 Ethiopian Demographic and Health Survey (DHS), a regression discontinuity design (RDD) is applied exploiting the fact that those women younger than 15 years when the RFC was introduced were exposed to a legal age of marriage at 18 years, while those women that were 15 years or older at the same time had the opportunity to get legally married before they turned 18 years old.

[5]The discussion in the literature is summarized in section 2.

The results show that in the five Ethiopian regions analyzed, women exposed to a legal age of marriage at 18 years are 20 percentage points less likely to be cohabiting with a partner by the age of 18 and start cohabiting significantly later than those women that have the opportunity to get legally married at the age of 15. The estimates for the causal effect of early cohabitation reveal that a one-year delay in women's age at cohabitation during teenage years decreases the probability of infant mortality of the first born by 3.8 percentage points. The results are robust to the use of different estimation techniques and bandwidths for the forcing variable. Additionally, different placebo tests rule out the possibility that the impact on infant mortality is driven by other legal dispositions included in the RFC, systematic differences between women born in different months of the year, other interventions at the national level, or over time decreases in infant mortality.

The analysis of mechanisms indicates that the impact of delaying cohabitation on the infant mortality of the first born is mainly channeled through the positive effect of delaying cohabitation on the age of women at first birth. Other possible channels such as an effect of early cohabitation on women's marriage market outcomes, participation in household decisions, education or labor force participation are explored and dismissed in the light of the evidence.

The study is structured as follows. Section 2 reviews the literature on the socioeconomic effects of child marriage. Section 3 discusses the incidence of child marriage in Ethiopia and presents the law that raised the legal age of marriage for women from 15 to 18 years in some regions of the country. Section 4 introduces the identification strategy and section 5 describes the data used in the main analysis. Section 6 presents the main results, examining their robustness to the use of alternative estimation methods, bandwidths and placebo tests. Section 7 investigates the channels through which early cohabitation could affect infant mortality. Finally, section 8 concludes the study.

2. Related Literature

The causes of child marriage have been extensively studied in anthropology and sociology. In the first economic study that aimed to model child marriage, Wahhaj (2015) enumerates three of the most commonly cited causes. First, young brides might be preferred because they are on average meeker than older ones and because they have a longer childbearing life ahead (Goody, 1990). Second, in opposition to western countries where newly married couples are expected to live without the economic support of relatives, it is very common that young couples in developing countries where the prevalence of child marriage is large are economically supported by their families, providing incentives for early marriages (Dixon, 1971). Finally, in many of these countries, the social status of the households depends strongly on the *purity* of the women of the family. In this context, families have to control the sexual behavior of the girls of the household after sexual maturation, providing parents incentives to marry their daughters as soon as possible after menarche (Moghadam, 2004). Consistent with the *purity* argument, Wahhaj (2015) explains child marriage using a marriage market model where women's *purity* is noisily observed and perceived *purity* decreases with time on the marriage market, providing households strong incentives for early marriages.

Using DHS and Multiple Indicator Cluster Surveys (MICS), different studies assess the link between child marriage and socioeconomic outcomes for women and their children. A synthesis of this literature is provided in Parsons et al. (2015). The review concludes that, overall, child marriage is associated with harmful socioeconomic outcomes for women including lower levels of participation in household decision making and worse marriage market outcomes (Jensen and Thornton, 2003; Jain and Kurz, 2007; Solanke, 2015), lower levels of labor force participation and educational attainment (Elborgh-Woytek et al., n.d.; Jensen and Thornton, 2003; Nguyen and Wodon, 2015; Wodon, Nguyen and Tsimpo, 2016), and worse maternal health (Campbell, 2002). The review also suggests that child marriage is associated with higher fertility, teenage pregnancy and lower age at first birth (Jensen and Thornton, 2003; Solanke, 2015); all of which are factors that have been linked to offspring early mortality (see for example Weldearegawi et al. (2015)). Indeed, child marriage is also associated with negative health outcomes and a higher prevalence of infant mortality for the children of these women (UNICEF, 2014; Wachs, 2008; Raj et al., 2010).

The evidence discussed so far examines the correlation between child marriage and the socioeconomic outcomes of these women and their children. This could be problematic because the statistical association between child marriage and these outcomes might be driven by reverse causality or by unobservable factors correlated with both. A few studies address empirically this problem through exploiting age at menarche as an instrumental variable for women's age at marriage. These studies show that the arrival of puberty determines the entrance in the marriage market and that a delay in the age at menarche increases significantly women's age at marriage. Using this approach, Field and Ambrus (2008), Asadullah et al. (2016) and Hicks and Hicks (2015) find that early marriage decreases educational attainment for women and antenatal health investments in Bangladesh, India and Kenya. Using the same instrumental variable, Chari et al. (2017), Sekhri and Debnath (2014), Asadullah et al. (2016) and Asadullah and Wahhaj (2016) document for India and Bangladesh that early marriage also has intergenerational effects, leading to negative impacts on the educational attainment, cultural values and health investments received by the children of women marrying young. On the other hand, Hicks and Hicks (2015) do not find any effect of early marriage on labor market outcomes, beliefs and marriage market outcomes of women in Kenya. To the best of my knowledge, this instrumental variable approach has not been used to investigate the link between child marriage and offspring mortality.

Although age at menarche is arguably an *external* and *relevant* source of variation for the age at marriage and its use as an instrumental variable for the latter can help to address endogeneity concerns, the correct identification of the effects of early marriage in these studies requires age at menarche not to be driven by unobservable factors that affect the outcome of interest. The latter assumption would be however problematic if unobservable childhood experiences such as episodes of sexual or physical abuse, that may have long-term effects on socioeconomic outcomes, have also an effect on the age at menarche, as suggested by Karapanou and Papadimitriou (2010) and Barrios et al. (2015).

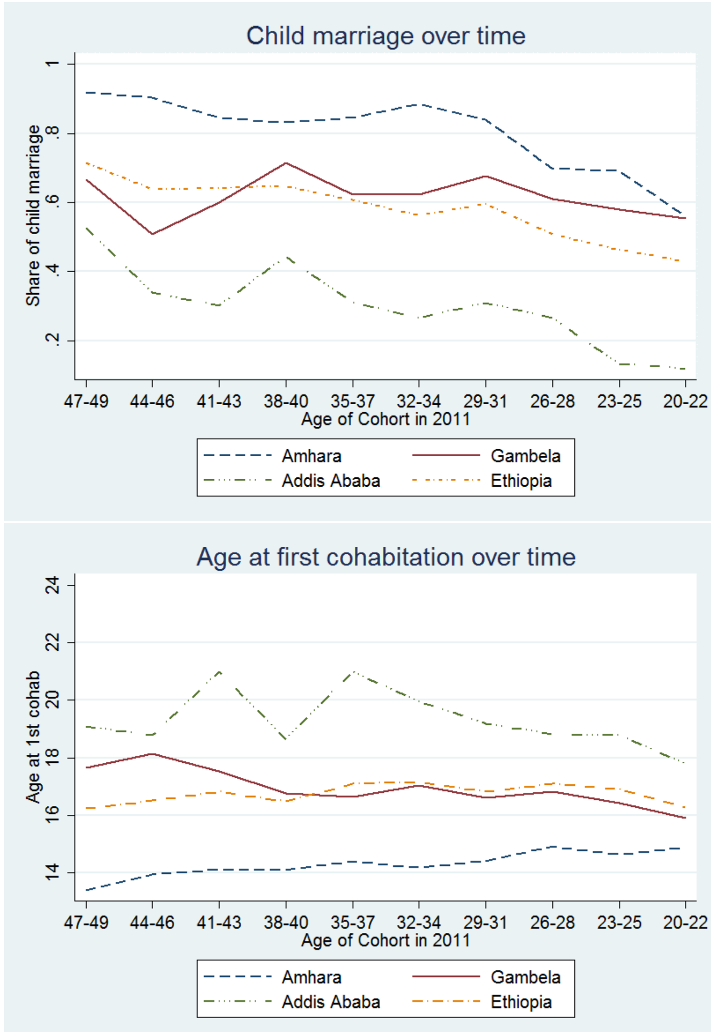
3. Child Marriage in Ethiopia and the Revised Family Code

It is estimated that 1,974,000 women in Ethiopia cohabited before the age of 18, making Ethiopia the country with the 5th largest number of child marriages in the world[6]. Although the practice of child marriage is conducted in all the country, its prevalence ranges substantially across regions and over time. Figure 1 displays the evolution of the prevalence of early cohabitation and of the mean age at first cohabitation in the country. In order to illustrate the variation across regions, the figure also displays the evolution of these indicators for a selection of three Ethiopian regions, including Addis Ababa and Amhara, the regions with the lowest and the largest prevalence of child marriage. For women aged 20-22 years old, the figure shows that the incidence of child marriage ranges between 56% in Amhara and 12% in Addis Ababa[7]. Another interesting pattern that emerges from the graph is that although during the last decades the share of women cohabiting before the age of 18 has been decreasing in Ethiopia, the trend is very different across regions and while for example the incidence of child marriage has decreased sharply in Amhara or Addis Ababa, the magnitude of the reduction is much modest in Gambela.

In the last decades, increasing age at cohabitation for women has become a priority for policy makers in Ethiopia. Following the ratification of the CEDAW, which encourages governments to set and enforce laws and programs to delay age at cohabitation, the Federal Government of

[6]Girls not Brides website. *http://www.girlsnotbrides.org/where-does-it-happen/*
[7]The causes that explain this strong regional variation are discussed in UNICEF (2015). The authors of the report highlight the existence of different social norms across the country.

FIGURE 1. CHILD MARRIAGE OVER TIME IN ETHIOPIA (DHS 2011)



Ethiopia approved the Revised Family Code (RFC) in July 2000. This law established the legal age of marriage for both men and women at 18 years.

Before the Federal Government of Ethiopia passed the RFC, the legal age of marriage for women and men was regulated by the 1960 Family Code. The latter law set a legal age for marriage of 15 years for women and of 18 years for men. Thus, while the RFC raised the legal age for marriage for women from 15 to 18 years, it left unchanged the minimum age of marriage for men at 18 years. At the same time, the RFC recognized the validity of marriages celebrated before the approval of the RFC that complied with the 1960 Family Code. Additionally, the RFC provided women authority to administer common marital property, abolished the right of husbands to forbid women to work outside home and facilitated the divorce procedure. The potential confounding effects of these additional dispositions will be examined and dismissed in section 6.6.4.

However, the approval of the RFC by the Federal Government of Ethiopia did not imply the immediate application of the law over the entire country. Under the Federal Constitution of Ethiopia, the family law is jurisdiction of the regional governments. In consequence, the approval of the RFC by the Federal Government of Ethiopia in July 2000 only implied its immediate application in the chartered cities of Addis Ababa and Dire Dawa. The application in the rest of Ethiopian regions required the approval of the regional governments. Although the enactment of the law by the Federal Government of Ethiopia paved the road, its approval by the different regional governments was not immediate (Hallward-Driemeier and Gajigo, 2015).

Figures 2 and 3 display the local linear density estimator (McCrary, 2008) of the age at first cohabitation for those Ethiopian women that were 12-14 years old and for those women aged 15-17 at the time of approval of the RFC in their region[8]. The figures show that while for the younger cohort of girls, exposed only to a legal age of marriage at 18 years, the most frequent age at first cohabitation is 18 years, the density function for the second cohort of women, exposed at least for some time after they turned 15 to a legal age of marriage at 15 years, reaches its peak at the age of 15 years. Furthermore, the figures suggest that while for the older cohort there seems to be a discontinuity in the density of women that first cohabited with a partner at the age of 15 and there is no discontinuity at 18 years, for the younger cohort the discontinuity at 15 years seems to be smaller and a new discontinuity emerges at the age of 18. In this line, figure A1 in the appendix compares in the same graph the distribution of the age at cohabitation for different cohorts of women in the analytical sample. The figure shows that the distribution shift to the right for the younger cohort of women, exposed to a legal age of marriage at 18 years.

The shift in the distribution of the age at first cohabitation for these two cohorts of women suggests that the rise in the legal age of marriage increased the mean age at cohabitation for women. Indeed, the fact that the most frequent age at cohabitation in each cohort is the legal age for marriage that they are exposed to could be indicating that the minimum ages of marriage set in the 1960 and 2000 Family Codes were to some extent enforced. On the other hand, the figures confirm that the percentage of women that cohabit with a partner before reaching the minimum age of marriage is non-negligible among women from both cohorts. Different reasons can explain why the introduction of the RFC has not eradicated child marriage. First, although the RFC bans civil, religious and customary marriages before the age of 18 years, underage unmarried cohabitation

[8]These two figures are constructed using the sample of ever cohabited women in the 2011 Ethiopian DHS that were older than 18 at the time of the survey and lived in one of the five Ethiopian regions that are included in the analysis.

FIGURE 2. AGE AT 1ST COHABITATION (COHORTS 12-14 AT RFC): DISCONTINUITIES AT 15 AND 18

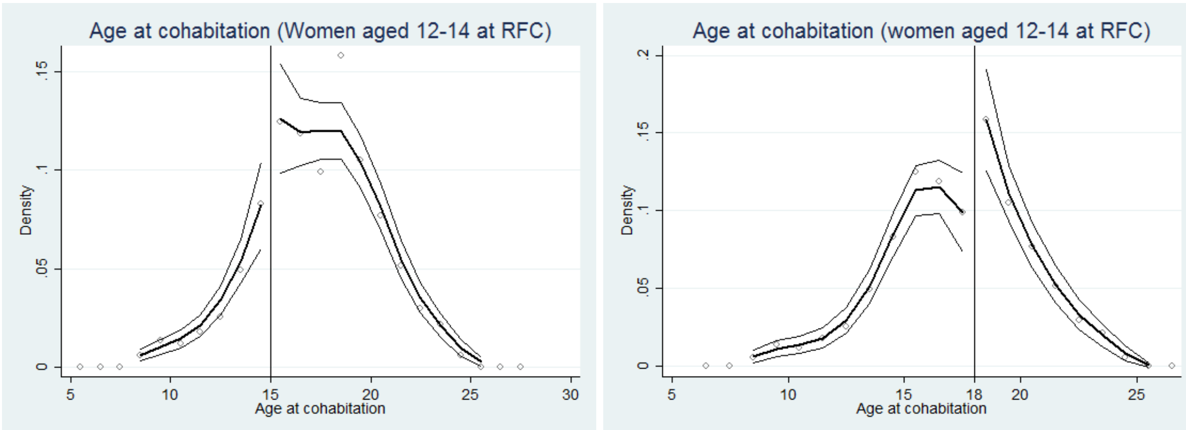
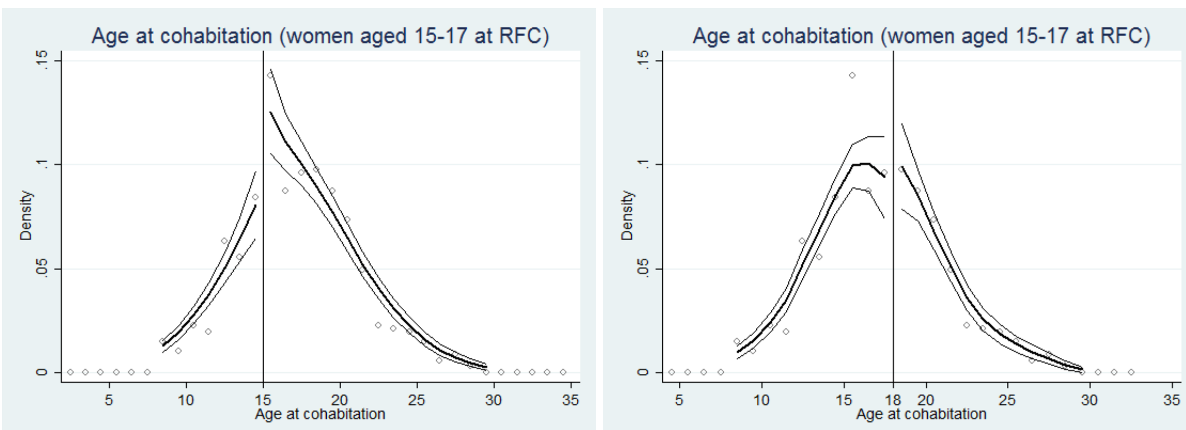


FIGURE 3. AGE AT 1ST COHABITATION (COHORTS 15-17 AT RFC): DISCONTINUITIES AT 15 AND 18



(which is one of the forms of child marriage) is not explicitly forbidden in the law. Second, although underage marriage is not permitted in the RFC, the criminal law did not sanction it until the year 2005[9]. Third, the institutional capacity to enforce the law is limited, particularly in rural areas where the presence of the state administration is narrow (Jones et al., 2016). Fourth, the law attributes the obligation to verify that both bride and groom are at least 18 years to the official or priest celebrating the wedding. However, the lack of birth and school registers for wide sectors of the population makes more difficult this checking procedure (Jones et al., 2016). For these reasons, and despite the fact that unmarried cohabitation is stigmatized in Ethiopia (Jones et al., 2016), we

[9]Unlike the introduction of the law in the 5 regions used in the sample, the introduction of child marriage as an offence in the criminal law did not generate a sharp change in the mean age at cohabitation in these regions. Therefore, this reform in the criminal code is not used for identification purposes

cannot expect the law to completely end child marriage.

Taken together, these patterns suggest that although it does not end with this practice, exposure to an effective legal age of marriage at 18 increases the mean age at first cohabitation for women. In section 6.6.1, I examine whether this change is sharp at the cut-off and statistically relevant to validate the estimations conducted in the study.

4. Identification Strategy

The RFC raised the legal age of marriage for women in Ethiopia from 15 to 18 years. This legal change generated variation in the legal age of marriage faced by women of different ages. First, those women that were younger than 15 years old when the RFC was approved in their region were only exposed to an effective legal age for marriage at 18 years. Second, those women that were older than 18 when the RFC was approved were not directly affected by the change in the legal age for marriage. Third, those women aged between 15 and 18 years old at the same time were exposed, at least for some time after their 15th birthday, to a legal age of marriage at 15 years. Thus, these women had the opportunity to marry legally before the age of 18.

The identification strategy relies on the change in the mean age at cohabitation for those women that were just above 15 years at the time of the approval of the law; and therefore had the opportunity to get legally married at the age of 15 before the introduction of the RFC, relative to those women that were just below 15 years old at the same time, and could not get legally married until the age of 18. If the mean age at cohabitation increases sharply for those women younger than 15 at the time of the approval of the RFC, the setting would be ideal for the implementation of a regression discontinuity design (RDD) using age of the women at the time of the approval of the RFC as the forcing variable.

The regression discontinuity framework used in this study has three particularities. First, the approval of the RFC hindered underage marriage for those women younger than 18 years and not yet married when the RFC was approved but did not eradicate child marriage among them[10]. In other words, exposure to a legal age of marriage at 18 did not entail that all these women in the sample started cohabiting after the age of 18. Second, the forcing variable is defined as the age of the women at the time of the rise in the legal age of marriage measured in months, with the cut-off at the age of 15 years. Those women that turned 15 in the month in which the RFC was approved were also dropped from the sample used in the analysis. The reason for this is that the DHS survey used in the analysis only collected information on the month and year of birth, making

[10]The possible causes for this are discussed in section 3.

it impossible to determine for those women that turned 15 in the same month, whether they did before or after the approval of the RFC. Third, the RFC was not applied simultaneously in every Ethiopian region. Although this variation is not exploited for identification purposes in the RDD analysis[11], the different timing in the application of the rise in the legal age for marriage across regions provides variation in the current age of women that were approximately 15 year old when the RFC was approved in their region[12]. Because in this RDD setting the causal estimates are only identified for women around the cut-off, the variation in the age of women at the cut-off makes the RDD estimates more generalizable.

In order to estimate (a) the effect of exposure to a legal age of marriage at 18 on women's age at first cohabitation, (b) the effect of exposure to a legal age of marriage at 18 on the probability of infant mortality of the first born child and (c) the effect of women's age at cohabitation on the probability of infant mortality of the first born child, I estimate the following three regressions:

$$(1) \qquad \textit{Age at Cohab.}_i = \alpha_0 + \alpha_1(\textit{Age at RFC<15}_i) + \alpha_2 F(\textit{Age at RFC}_i) + \alpha_3 X_i + \mu_i$$

$$(2) \qquad InfantMortality_i = \delta_0 + \delta_1(\textit{Age at RFC<15}_i) + \delta_2 F(\textit{Age at RFC}_i) + \delta_3 X_i + \epsilon_i$$

$$(3) \qquad InfantMortality_i = \beta_0 + \beta_1(\widehat{\textit{Age at Cohab.}_i}) + \beta_2 F(\textit{Age at RFC}_i) + \beta_3 X_i + u_i$$

where *Age at Cohab*$_i$ indicates the age of woman i at first cohabitation, $InfantMortality_i$ is a dummy variable equal to 1 if the first born child of woman i died within the first year of life, *Age at RFC<15* is a dummy variable that indicates whether the woman was younger than 15 when the RFC was approved in her region and therefore, was exposed to an effective legal age of marriage at 18 years. X is a vector of control variables indicating the region of residence, the age of the woman at the time of survey, ethnic and religion affiliation, gender of the first born and whether

[11]The sequential introduction of the RFC could be in principle exploited using a difference-in-difference (Diff-in-Diff) strategy to estimate the effect of the law on infant mortality. The Diff-in-Diff analysis would exploit across cohort and regional variation in the legal age of marriage faced by women that lived in different regions. Although less robust in terms of statistical significance, the results of this analysis are consistent with those obtained in the RDD analysis (the results are not reported in the paper). However, I believe the RDD analysis has additional benefits in terms of identification and attribution of the effect to the change in the legal age of marriage.

[12]The age of the women that were approximately 15 when the RFC was approved in their region at the time of the 2011 survey ranges from 18 in Tigray to 26 in Addis Ababa and Dire Dawa.

the woman lives in a rural area. *F(Age at RFC)* is a function of the age of the woman in months when the legal age for marriage was raised in her region. Finally, $\widehat{Age\ at\ Cohab.}_i$ is the predicted age at cohabitation for woman i estimated from equation 1.

Equation 1 is the first stage regression. The parameter α_1 measures the effect of exposure to a legal age of marriage at 18 on the age at first cohabitation, relative to women that had the possibility of getting legally married at 15. Equation 2 is the reduced form equation. The parameter δ_1 yields the effect of exposure to a legal age of marriage at 18 on the probability of infant mortality of the first born, relative to women that had the possibility of getting legally married at 15. Equation 3 is the second stage equation. It regresses infant mortality against the predicted age at cohabitation estimated in equation 1. The parameter β_1 yields the effect of a one-year delay in women's age at cohabitation with a partner during teenage years on the probability of infant mortality of the first born.

The estimation of equations 1, 2 and 3 is conducted using non-parametric local polynomial regressions based on triangular kernel functions[13]. The study follows the state of the art procedure described in Calonico, Cattaneo and Titiunik (2014) and Calonico et al. (2016) for the selection of the optimal bandwidth and for the calculation of bias-corrected RD estimates with robust variance estimator. Standard errors are clustered at the running variable level as recommended by Lee and Lemieux (2010) for RDDs based on discrete forcing variables. As a robustness check, I also estimate equations 1, 2 and 3 using (a) two alternative bandwidths equal to 0.75 and 1.5 times the optimal bandwidth and (b) conventional and bias-corrected non-parametric RD estimation procedures with conventional variance estimators. Additionally, I also estimate the equations using parametric methods and windows of 2, 3, 4, 5 and 6 years at both sides of the cut-off. Following Gelman and Imbens (2017), that discourage the use of polynomials of order 3 and above in parametric RDDs, I include polynomials of order 1 and 2 for the forcing variable, allowing also a different polynomial function at either side of the cut-off.

Both when estimated using parametric and non-parametric methods, the identification of the causal effects on infant mortality of early cohabitation and of the increase in the minimum age of marriage relies on two main conditions. The first identification assumption requires that facing an effective legal age of marriage at 18 years increases the mean age at cohabitation. In other words, if the RFC did not change sharply the mean age at cohabitation for those women at the cut-off,

[13]Following Calonico, Cattaneo and Titiunik (2014) and Calonico et al. (2016), the conventional estimates are conducted using local linear regressions and the bias-corrected estimates are conducted using local quadratic regressions

the estimated parameter β_1 in equation 3 would not be efficient, potentially leading to a problem of weak instruments (Bound, Jaeger and Baker, 1995). Equally, if the reform did not change the mean age at cohabitation for the women at the cut-off, the expected coefficient of the parameter δ_1 in equation 2 would be 0. Although the descriptive analysis presented in section 3 suggests that the rise in the minimum age for marriage led to an increase in the mean age at cohabitation, the existence of a sufficiently sharp change in the mean age at cohabitation and in the incidence of child marriage at the cut-off will be tested empirically in section 6.6.1.

The second identification assumption of the RDD is that the determinants of infant mortality unaffected by the legal change should be continuously related to the forcing variable at the cut-off. Although this condition cannot be tested for every determinant of infant mortality, I examine in section 6.6.4 the existence of discontinuities at the cut-off for some of these determinants that are unlikely affected by the legal age of marriage. If the placebo analysis shows discontinuities at the cut-off for these variables, we would need to consider the possibility that confounding factors might be driving the results. In addition to presenting the results of this placebo test, section 6.6.4 examines the robustness of the results to other identification threats and discusses the feasibility of alternative explanations for the results.

5. Data and Descriptive Statistics

The data used in the analysis is from the Ethiopian Demographic and Health Survey (DHS) conducted in late 2011. DHS have been implemented in more than 100 low- and middle-income countries across the world for more than three decades. These surveys are widely used by researchers in the fields of fertility and health, and the high quality and accuracy of the information collected is showed in Pullum (2008). Indeed, the vast majority of the studies and reports that explore empirically the incidence of child marriage rely on these surveys (Parsons et al., 2015).

Although DHS questionnaires are designed to produce comparable statistics across countries and over space, the exact questionnaire and the size and characteristics of the sample vary in every DHS. The 2011 Ethiopian DHS collected household, child, male and female level information for a sample of 16,702 households, representative at the national and regional level. The female module of the survey was applied to all women aged 15-49 living in the households sampled. This module includes questions on health, anthropometrics, demographics, fertility and women status within the household, providing information on the birth and mortality history of their children, as well as on the age at first cohabitation, which is used to measure child marriage. On the other hand, the survey provides little information on labor market outcomes and does not record the age at marriage[14]. The data on antenatal and postnatal behavior before and after each birth and children's health status are only recorded for those children born in the last 5 years, and even among these children, the share of missing values is substantive.

In total, the female module was applied to 16,515 women aged 15-49 living in the 11 Ethiopian regions. Out of these women, 8,685 live in Addis Ababa, Dire Dawa, SNNP, Tigray and Amhara, the regions that approved the RFC between 2000 and 2007 and that will be used in the analysis. The remaining 6 Ethiopian regions were excluded from the analytical sample for two reasons. First, four of these regions did not implement the rise in the legal age for marriage before 2008. Thus, even if the RFC was approved in these regions before 2011, women that were 15 when the RFC was approved would still be underage at the time of the survey and therefore, excluded from the analysis. Second, the regions of Gambela and Oromia were excluded from the analysis because despite having approved the RFC before 2008, they did not seem to enforce it in any way[15]. Thus, the inclusion of these two regions in the analysis would decrease the magnitude and significance of the parameter of interest in the first stage equation, reducing the efficiency of the parameter that yields the effect of early cohabitation on infant mortality and potentially leading to a problem of weak instruments (Bound, Jaeger and Baker, 1995).

Out of the 8,685 women living in these 5 regions, I use in the analysis the sample of 5,078 women aged 18-49 at the time of the survey, that ever cohabited with a partner[16] and gave birth to their first child more than one year before the survey[17].

Table 1 provides the descriptive statistics for the main variables used in the analysis for this sample of 5,078 women. However, it is important to remark that the regression discontinuity analysis does not use all these women to estimate the parameters of interest but only those that fall within the bandwidth used in the non-parametric analysis or the relevant window in the parametric

[14]Unmarried cohabitation is strongly stigmatized in Ethiopia and is unfrequent for most ethnic groups (Jones et al., 2016). Furthermore, even if age at cohabitation and age at marriage are two separate events, it seems reasonable to think that the former is a more relevant determinant of infant mortality and woman wellbeing.

[15]This pattern can be observed in figure A2 in the appendix. Reasons for this could be that the capacity of the state to enforce the law might be weaker in these two regions or more traditional attitudes among the citizens of them.

[16]The percentage of women that gave birth reporting never cohabited with a partner only represents the 1.31% of all the women that gave birth to at least one child in these regions. These observations are not used in the analysis.

[17]Because infant mortality is defined as mortality within the first year of life, the sample is restricted to those women that gave birth to their first child more than one year before the survey to avoid censoring in the dependent variable.

analysis. Given that the estimates yielded by the RDD are local in the sense that they are interpreted as the effects for those women that were approximately 15 when the RFC was approved in their region, the table also includes the mean of the variables for those women aged 14-15 at the time of the rise in the legal age of marriage.

TABLE 1—SUMMARY STATISTICS: WOMEN THAT EVER COHABITED AND EVER BORE A CHILD IN THE REGIONS INCLUDED IN THE STUDY.

	Aged 14-15 at RFC.					Full sample (Aged 18-49 2011)		
	N	Mean	Standard deviation	Min	Max	N	Mean	Diff (FS - 1y bw)
Women characteristics								
Age (2011)	308	23.38	2.28	18	26	5,078	32.88	9.50
Age at policy	308	14.58	0.49	14	15	5,078	25.36	10.78
Work (0/1)	308	0.33	0.47	0	1	5,077	0.36	0.03
Anemia (0/1)	292	0.20	0.40	0	1	4,824	0.17	-0.03
Years schooling (compl.)	308	2.89	3.93	0	15	5,078	2.16	-0.73
Level educ: Secondary or higher (0/1)	308	0.16	0.37	0	1	5,078	0.12	-0.04
Rural (0/1)	308	0.58	0.49	0	1	5,078	0.71	0.13
Orthodox (0/1)	308	0.57	0.50	0	1	5,078	0.59	0.02
Eth. Oromiya (0/1)	308	0.23	0.42	0	1	5,078	0.12	-0.11
Marriage market								
Age at 1st cohab	308	16.13	3.18	8	24	5,078	16.43	0.30
Child married (0/1)	308	0.66	0.47	0	1	5,078	0.65	-0.01
Currently partner (0/1)	308	0.86	0.34	0	1	5,078	0.83	-0.03
Divorced (0/1)	308	0.13	0.33	0	1	5,078	0.12	-0.01
Same partner (0/1)	308	0.73	0.45	0	1	5,076	0.62	-0.11
Empowerment index (0-2)	265	0.92	0.41	0	2	4,181	0.89	-0.03
Particip. social life decisions (0-2)	265	1.06	0.68	0	2	4,175	1.03	-0.03
Particip. health decisions (0-2)	265	1.01	0.62	0	2	4,170	0.97	-0.04
Particip. purchase decisions (0-2)	264	0.79	0.56	0	2	4,164	0.75	-0.04
Particip. husband earnings (0-2)	262	0.83	0.48	0	2	4,142	0.82	-0.01
Age difference with partner	266	6.35	4.58	-5	30	4,171	7.82	1.47
Years of schooling (partner)	307	4.12	4.59	0	16	4,999	3.53	-0.59
Wealth index (1-5)	308	3.48	1.55	1	5	5,078	3.18	-0.30
Fertility outcomes								
N children	308	1.85	1.01	1	7	5,078	4.22	2.37
Age at 1st birth	308	18.30	2.78	11	25	5,078	18.89	0.59
First born characteristics								
Years since born	308	5.00	2.73	1	13	5,078	13.96	8.96
Male (0/1)	308	0.56	0.50	0	1	5,078	0.52	-0.04
Deceased before 1st year (0/1)	308	0.09	0.28	0	1	5,078	0.10	0.01
Maternal and infant health: First born								
Months breastfeed	63	23.14	12.14	0	51	240	20.69	-2.45
Birth weight (kg)	38	3.06	0.78	1	5	204	3.12	0.06
N vaccines (1-9)	144	6.44	2.76	0	9	781	6.44	0.00
Delivery at home (0/1)	164	0.48	0.50	0	1	862	0.48	0.00
Ever antenatal visit (0/1)	107	0.62	0.49	0	1	520	0.64	0.02
Child has anemia (0/1)	128	0.30	0.46	0	1	686	0.32	0.02

Note: Descriptive statistics are provided for two different samples: (a) women aged 18-49 in the five regions of interest that ever cohabited with a partner and have given birth and (b) women aged 18-49 in the five regions of interest that were 14-15 when the RFC was approved in their region, ever cohabited with a partner and have given birth. The last column reports the difference in means between these two samples.

The table shows that the age in 2011 for the women aged 14-15 when the RFC was approved ranges between 18 (in Tigray) and 26 (in Addis Ababa and Dire Dawa). The average number of years of education among these women is very low (less than 3), highlighting that most Ethiopian women are probably out of school by the time they start cohabiting with a partner. The participation in the labor market among the women in the sample exceeds 30% and approximately 60% of these women live in rural areas. Interestingly, 13% of women aged 14-15 when the RFC was approved in their region had separated or divorced from their first cohabiting partner.

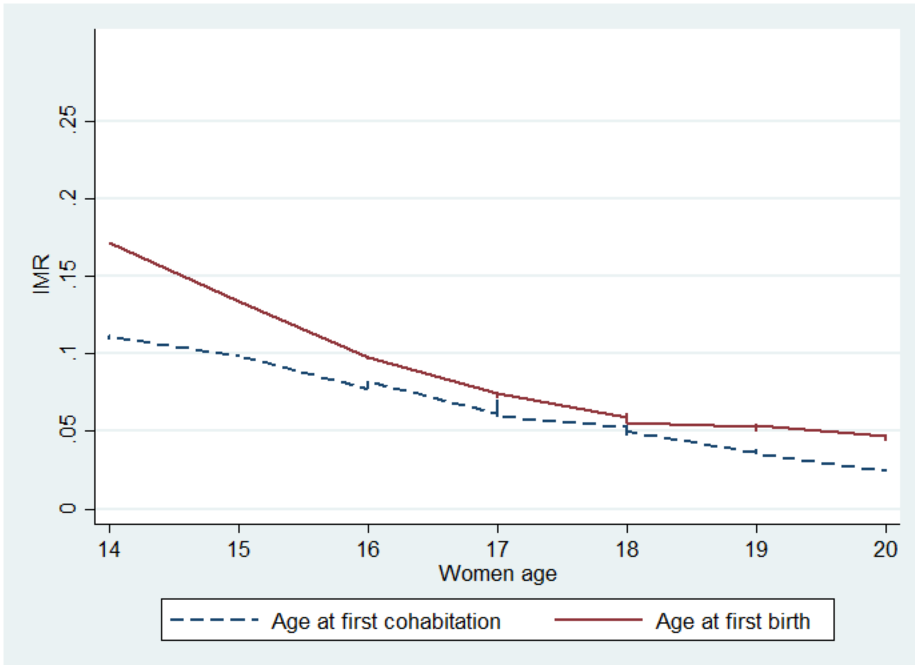
The mean age at cohabitation for the women in the full sample is approximately 16.4 years, and more than 60% of these women first cohabited with their partner before the age of 18 years. Remarkably, the incidence of child marriage falls from 70% among those women aged 15 when the RFC was approved in their region to 60% among those women aged 14 and therefore, exposed to a legal age of marriage at 18 years. Similarly, the mean age at first cohabitation increases from 15.8 to 16.6 for the same groups of women[18]. The differences in terms of age at cohabitation and prevalence of child marriage between women aged 14 and women aged 15 at the time of the RFC are statistically significant at the 5% and 10% significance levels.

In the full sample, the mean age of women at first birth is 18.9 years and this number is smaller (18.3) among those women aged 14-15 when the RFC was introduced in their region. The infant mortality rate of the first born among the women in the full sample is 10%. If we focus on the women aged 14-15 when the RFC was introduced in their region, the probability of infant mortality of the first born falls from 12% among women aged 15 when the RFC was introduced to 5% among women aged 14 at the same time and therefore, exposed to a legal age of marriage at 18 years. This difference is statistically significant at the 5% level of significance.

Figure 4 displays the statistical association between women's age at cohabitation, age at first birth and infant mortality of the first born during teenage years. The curves are estimated using LOWESS regressions for the sample of women aged 13-17 at the time of the RFC. Although these relations should not be interpreted as causal, the figure shows a strong correlation between age at first birth and infant mortality during puberty. The graph suggests that delaying age at first birth

[18]The large incidence of child marriage and the low age at cohabitation among the youngest women in the sample, which represent an important part of the women within one year from the cut-off, is mechanically driven by the way in which the sample is selected. Because the sample only includes women aged 18-49 that have given birth to their first child more than one year ago and have ever cohabited, it is very likely that the vast majority of the youngest women in the sample (e.g. aged 18 or 19) cohabited with their partner before the age of 18. The general evolution of the prevalence of child marriage across age cohorts in these regions of Ethiopia could be better observed in figure A3 in the appendix, which is constructed using the women aged 18-49 in the DHS data regardless of whether they ever cohabited or gave birth.

FIGURE 4. AGE AT COHABITATION, AGE AT BIRTH AND INFANT MORTALITY OF THE FIRST BORN (LOWESS REGRESSIONS)



from 15 to 17 is associated with a decrease in the incidence of infant mortality of the first born from approximately 13.5% to 7.5%. On the other hand, rises in the age at first birth after the age of 18 are associated with smaller reductions in the probability of infant mortality of the first born. Although the slope of the estimated function that displays the statistical association between age at cohabitation and infant mortality of the first born during early adolescence is less pronounced, the negative statistical association between these two variables is also evident in the graph.

6. Results

6.1. The Effect of the RFC on the Age at Cohabitation

The first condition for the validity of the identification strategy outlined in section 4 is the existence of a discrete change in the mean age at first cohabitation at the cut-off. The size and statistical significance of this discontinuity is yielded by the parameter α_1 in the first stage equation. Columns 1, 3 and 5 of table 2 report the estimates for this parameter using non-parametric techniques with different estimation procedures and bandwidths. The results of the preferred estimation are reported in column 3 and show that exposure to a legal age of marriage at 18 relative to the possibility of getting legally married at 15 increases women's age at cohabitation by approximately 2 years in the five Ethiopian regions included in the sample. The coefficients of the variable when the alternative bandwidths and non-parametric estimation procedures are used are also positive and in most cases, similar in magnitude. The coefficients measuring the effect of exposure to a legal age of marriage at 18 across the different non-parametric estimations are all statistically significant at the 1% and satisfy the *relevance* condition (F>10) required for the estimation of the second stage equation.

The results reported in columns 1, 3, 5, 7 and 9 of table 3 show that the rise in the age at first cohabitation for women younger than 15 when the RFC was implemented in their region remains large (0.6-2 years) and statistically significant (p-value<0.01) when equation 1 is estimated using parametric methods with several windows and different polynomials for the forcing variable. The sharp change in the mean age at cohabitation at the cut-off is also evident in figure 5. Consistently, the results reported in column 1 of table 4 show that the rise in the mean age at cohabitation is accompanied by a decrease in 20 percentage points in the incidence of child marriage at the cut-off. The results of the first stage equation are in line with the main conclusions of the descriptive analysis conducted in section 3 showing how the distribution of the age at cohabitation changes across cohorts of women exposed to a different legal age of marriage.

The evolution of the prevalence of child marriage and mean age at cohabitation across age cohorts observed in figure 5 deserves two additional comments. First, the large discontinuity at the cut-off and the polynomial behavior at the right of the cut-off are consistent with the hypothesis that through pushing women slightly over 15 when the RFC was introduced to get married before the approval of the RFC, the introduction of the RFC could have reduced the mean age at cohabitation for the cohorts of women aged just above 15 at the time of the RFC. For example, a woman slightly older than 15 years old when the RFC was approved had the possibility of getting legally married as soon as she turned 15, but if she waited some months and the RFC is approved, she would not be able to get legally married until the age of 18. This fact might push women aged 15 at the time of the RFC that were planning to get married over the next year or two to marry as soon as they turned 15. This has an important implication for the interpretation of the results: the estimates of interest in the first stage and reduced form equations measure the effect of exposure to a legal age of marriage at 18 relative to the possibility of getting married at 15, rather than to exposure to a legal age of marriage at 15.

Second, beyond the large discontinuity at the cut-off, the figure suggests that the prevalence of child marriage is larger among the younger women in the sample, exposed to a legal age of marriage at 18. Although we cannot expect child marriage to be eradicated among women exposed to a legal age of marriage at 18, the fact that the largest incidence of child marriage in the sample

TABLE 2—NON-PARAMETRIC METHODS: RFC, AGE AT FIRST COHABITATION AND INFANT MORTALITY.

	Conventional		Bias-corrected		Robust	
	(1) FS Age at 1st cohab	(2) SS/RF Infant Mortality	(3) FS Age at 1st cohab	(4) SS/RF Infant Mortality	(5) FS Age at 1st cohab	(6) SS/RF Infant Mortality
Bandwith A: Calonico et al. (2016)						
Age<15 at RFC	1.774***	-0.073**	2.055***	-0.079**	2.055***	-0.079*
	(0.321)	(0.034)	(0.321)	(0.034)	(0.371)	(0.041)
Age at 1st cohab.		-0.041**		-0.038**		-0.038*
		(0.019)		(0.019)		(0.023)
N		5078		5078		5078
N effect. obs.		581		990		990
Bandwidth		24.0		40.3		40.3
Bandwith B: 1.5 × C C T						
Age<15 at RFC	1.380***	-0.073***	1.584***	-0.082***	1.584***	-0.082***
	(0.289)	(0.027)	(0.289)	(0.027)	(0.328)	(0.031)
Age at 1st cohab.		-0.053***		-0.052***		-0.052**
		(0.019)		(0.019)		(0.022)
N		5078		5078		5078
N effect. obs.		874		1451		1451
Bandwidth		36.1		60.5		60.5
Bandwith C: 0.75 × C C T						
Age<15 at RFC	1.869***	-0.078*	2.081***	-0.082*	2.081***	-0.082
	(0.341)	(0.044)	(0.341)	(0.044)	(0.392)	(0.053)
Age at 1st cohab.		-0.042*		-0.039*		-0.039
		(0.023)		(0.023)		(0.029)
N		5078		5078		5078
N effect. obs.		453		733		733
Bandwidth		18.0		30.2		30.2

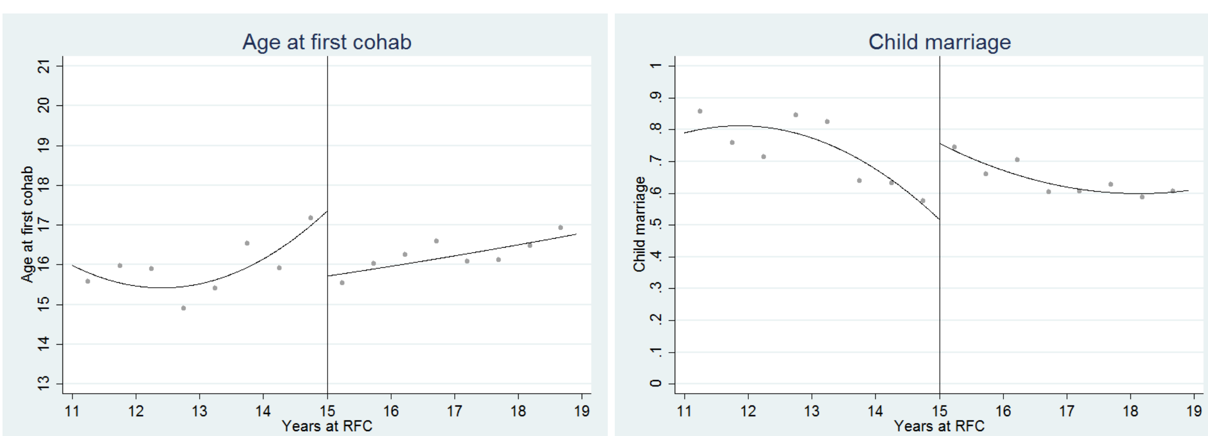
Note: Each coefficient provided in the table is estimated using a separate regression. The table reports the estimates of interest for the first stage (FS), reduced form (RF) and second stage (SS) equations using different bandwidths and the three procedures decribed in Calonico et al. (2016): conventional variance estimator, bias-corrected variance estimator and robust variance estimator. The coefficients for the variable *Age<15 at RFC* measure the effect of the RFC on the age at first cohabitation (first stage) in columns 1, 3 and 5; and the effect of the law on the prevalence of infant mortality of the first born (reduced form) in columns 2, 4 and 6. The coefficients for the variable *Age at 1st cohab* measure the effect of delaying one year the age at cohabitation during teenage years on the prevalence of infant mortality of the first born (second stage equation). The results provided for Bandwidth A are estimated using the optimal bandwidth calculated following Calonico et al. (2016). The results provided for Bandwidths B and C are estimated using 1.5× the optimal bandwidth and 0.75× the optimal bandwidth. The specifications include the following control variables: dummies for the regions of residence, the age of women at survey, ethnic and religion affiliation, gender of the first born, a rural/urban dummy variable and a non-parametric function for the age of the women at RFC. The sample size and the bandwidths used in the RF, FS and SS regressions are the same within each estimation procedure and bandwidth used. Standard errors are clustered at the forcing variable. Standard errors are in parentheses.***p<0.01;**p<0.05;**p<0.1.

TABLE 3—PARAMETRIC METHODS USING DIFFERENT TIME WINDOWS: RFC, AGE AT FIRST COHABITATION AND INFANT MORTALITY.

	2 years window (N=571)		3 years window (N=849)		4 years window (N=1164)		5 years window (N=1432)		6 years window (N=1646)	
	(1) FS Age at 1st cohab	(2) SS/RF Infant mortality	(3) FS Age at 1st cohab	(4) SS/RF Infant mortality	(5) FS Age at 1st cohab	(6) SS/RF Infant mortality	(7) FS Age at 1st cohab	(8) SS/RF Infant mortality	(9) FS Age at 1st cohab	(10) SS/RF Infant mortality
Pol. order 1										
Age<15 at RFC	1.706***	-0.078**	1.008***	-0.078**	0.920***	-0.060**	0.825***	-0.050	0.630**	-0.046*
	(0.331)	(0.034)	(0.309)	(0.030)	(0.269)	(0.027)	(0.269)	(0.024)	(0.241)	(0.024)
Age at 1st cohab		-0.046**		-0.077**		-0.065**		-0.061*		-0.073*
		(0.019)		(0.032)		(0.031)		(0.031)		(0.044)
Pol. order 2										
Age<15 at RFC	1.707***	-0.077**	1.066***	-0.069**	0.913***	-0.056**	0.860***	-0.051**	0.837***	-0.044*
	(0.332)	(0.033)	(0.312)	(0.028)	(0.276)	(0.027)	(0.266)	(0.024)	(0.251)	(0.022)
Age at 1st cohab		-0.045**		-0.065**		-0.061**		-0.059**		-0.052**
		(0.018)		(0.026)		(0.030)		(0.028)		(0.026)
Interacted pol. order 1										
Age<15 at RFC	1.707***	-0.076**	1.057***	-0.070**	0.919***	-0.053**	0.849***	-0.047**	0.769***	-0.040*
	(0.332)	(0.032)	(0.311)	(0.028)	(0.275)	(0.026)	(0.268)	(0.022)	(0.249)	(0.021)
Age at 1st cohab		-0.045**		-0.066**		-0.058**		-0.055**		-0.052*
		(0.018)		(0.026)		(0.028)		(0.027)		(0.028)
Interacted pol. order 2										
Age<15 at RFC	1.744***	-0.068	2.045***	-0.085**	1.494***	-0.091**	1.326***	-0.085***	1.141***	-0.084***
	(0.454)	(0.060)	(0.384)	(0.041)	(0.360)	(0.036)	(0.339)	(0.031)	(0.330)	(0.031)
Age at 1st cohab		-0.039		-0.041**		-0.061**		-0.064***		-0.074**
		(0.034)		(0.020)		(0.024)		(0.024)		(0.031)

Note: Each coefficient provided in the table is estimated using a separate regression. The table reports the estimates of interest for the first stage (FS), reduced form (RF) and second stage (SS) equations using the sample of women aged within different age windows around the cut-off and interacted and uninteracted polynomials of order 1 and 2 for the forcing variable. Interacted polynomials allow the running variable to follow a different polynomial at either side of the cut-off. The coefficients for the variable *Age<15 at RFC* measure the effect of the RFC on the age at first cohabitation (first stage) in columns 1, 3, 5, 7 and 9; and the effect of the law on the prevalence of infant mortality (reduced form) in columns 2, 4, 6, 8 and 10. The coefficients for the variable *Age at 1st cohab* measure the effect of delaying one year the age at cohabitation during teenage on the prevalence of infant mortality of the first born (second stage equation). The specifications include the following control variables: dummies for the regions of residence, the age of women at survey, ethnic and religion affiliation, gender of the first born, a rural/urban dummy variable and interacted and uninteracted polynomial functions of order 1 and 2 for the age of the women at RFC. The order of the polynomial function for the forcing variable used in every specification is reported above each estimation set. Standard errors are clustered at the forcing variable. Standard errors are in parentheses.***p<0.01;**p<0.05;**p<0.1.

FIGURE 5. MAIN ANALYSIS: AGE AT FIRST COHABITATION AND CHILD MARRIAGE AT THE CUT-OFF



is found for the youngest cohorts of women is apparently puzzling. To reconcile this paradox, one should take into account that the sample used in the analysis only includes women that ever cohabited with a partner and have given birth to their first child at least one year before the survey. In this scenario, the prevalence of child marriage among the youngest cohorts of women in the sample, barely aged 18 at the time of the survey, is expected to be very close to 100%. The *true* evolution of the prevalence of child marriage across age cohorts in the 5 Ethiopian regions of interest is presented in figure A3 in the appendix. Using the full sample of women aged 18-49 included in the DHS data regardless of whether they ever cohabited with a partner or have given birth, the figure shows that the prevalence of child marriage is lower among younger cohorts of women, and changes sharply at the cut off.

6.2. The Effect of the RFC on Infant Mortality

The next step is determining whether the 20 percentage points drop in the incidence of child marriage and the 2 years increase in the mean age at cohabitation at the cut-off for women exposed to a legal age of marriage at 18 affected the probability of infant mortality of the first born. This effect is yielded by the parameter δ_1 in equation 2. The estimates for this parameter using non-parametric estimations with different estimation procedures and bandwidths are reported in columns 2, 4 and 6 of table 2. The coefficients suggest that exposure to a minimum age of marriage at 18 years relative to the possibility of getting legally married at 15, decreases significantly the probability of infant mortality of the first born by 7.3-8.2 percentage points, depending on the bandwidth and estimation procedure selected. The most conservative estimate[19] is reported in column 6 and yields an effect of 7.9 percentage points, statistically significant at the 10%. Columns 2, 4, 6, 8 and 10 of table 3 provide the estimates for the parameter δ_1 using parametric techniques with different windows and polynomials functions for the forcing variable. Overall, the estimates presented in these two tables confirm the robustness of the results to the use of parametric and non-parametric methods. The discontinuity in the infant mortality rate of the first born for those women older than 15 at the time of the approval of the RFC is graphically displayed in figure 6.

The magnitude of the effect of exposure to a legal age of marriage at 18 on the infant mortality of the first born at the cut-off seems large, particularly when compared with the mean incidence of infant mortality of the first born among the women in the sample (10%). However, the interpretation of this coefficient requires a few considerations. First, the women just above the cut-off had

[19]In line with Calonico, Cattaneo and Titiunik (2014), this corresponds to the estimation conducted using the optimal bandwidth and the bias-corrected RD with robust variance estimator.

the opportunity of getting legally married as soon as they turned 15, but if they wait a few months and the RFC is approved, they would face a legal age of marriage at 18. Therefore, the parameter of interest in the reduced form equation measures the effect on infant mortality of the first born of exposure to a legal age of marriage at 18 relative to the possibility of getting legally married at the age of 15, rather than to exposure to a legal age of marriage at 15 years.

Second, the RDD estimates of the effects on infant mortality of the first born presented in this section are larger but aligned with those obtained in simple correlation analysis. For example, the results of the LOWESS analysis displayed in figure 4 reveal that delaying the age of women at first cohabitation from 15 to 17, which is the approximate change in mean age at cohabitation at the cut-off, would be associated with a decrease of approximately 4-4.5 percentage points in the incidence of infant mortality of the first born.

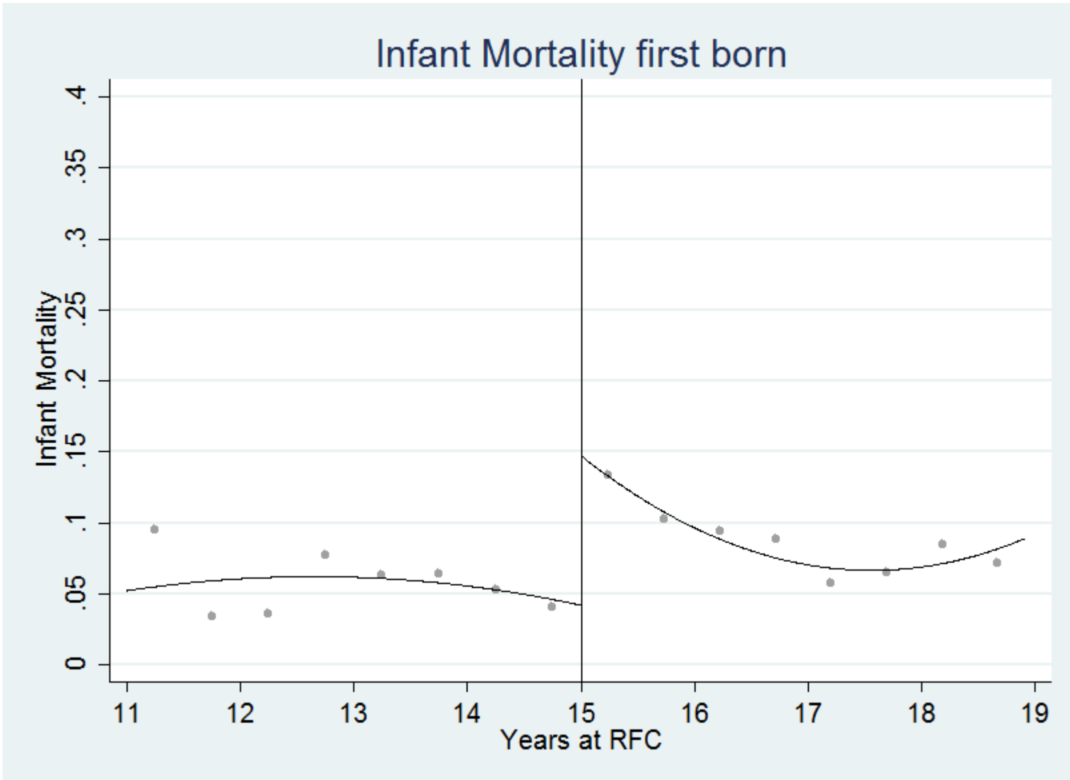
Third, the estimates of the parameter δ_1 discussed in this section should be interpreted as local treatment effects: they measure the effect on infant mortality of exposure to a minimum age of marriage at 18 for those women in the sample that were approximately 15 years old when the RFC was approved in their region. These are women that when they were surveyed in 2011, had an age ranging between 25 (Addis Ababa and Dire Dawa) and 18 years (Tigray), have already cohabited with a partner and have given birth to their first child at least one year before the survey.

6.3. *The Effect of Women's Age at Cohabitation on Infant Mortality*

The causal effect of women's age at cohabitation during teenage years on the infant mortality of the first born is yielded by the parameter β_1 in the second stage equation. The results for the non-parametric estimations are reported in columns 2, 4 and 6 of table 2 and reveal that a one-year delay in the age at first cohabitation decreases the probability of infant mortality of the first born by 3.8-5.2 percentage points, depending on the bandwidth and estimation procedure used. The most conservative estimation, reported in column 3, indicates that a one-year delay in the age of women at cohabitation decreases the probability of infant mortality of the first born by 3.8 percentage points. The effect is statistically significant at the 90% confidence level. The results for the effect of women's age at cohabitation reported in columns 2, 4, 6, 8 and 10 of table 3 confirm that the findings of the non-parametric analysis are overall robust to the use of parametric techniques with different windows and polynomial functions for the forcing variable.

When interpreting the coefficients of the second stage equation, it is important to consider that the estimated effect of early cohabitation on the infant mortality of the first born is a local average treatment effect. More specifically, the parameter of interest in the regression measures the effect

FIGURE 6. MAIN ANALYSIS: INFANT MORTALITY RATE AT THE CUT-OFF



of a one-year delay in the age at cohabitation during teenage years for those women in the sample that were approximately 15 years old when the RFC was approved in their region and delayed cohabitation because they were exposed to a legal age of marriage at 18 years. In other words, the effect of delaying marriage on infant mortality is identified for the subsample of Ethiopian women aged between 18 and 25 that ever cohabited with a partner, gave birth to their first child at least one year before the survey and either live in areas where the capacity of the institutions to enforce the law is strong or are particularly law abiding.

6.4. Robustness Checks

Using parametric and non-parametric methods, the previous section shows that in the five Ethiopian regions analyzed, exposure to a minimum age of marriage at 18 relative to the possibility of getting legally married at 15 increases significantly age at cohabitation and decreases the incidence of child marriage and infant mortality of the first born. In this section, I discuss and explore alternative explanations for the results.

Firstly, an important threat to the attribution of the effects identified on infant mortality to the rise in the legal age for marriage would be that the RFC not only raised the legal age for marriage but also set some additional provisions aiming to change the balance of power within the household through facilitating the procedure of divorce, abolishing the right of husbands to forbid women to work and providing women the right to administer the common marital property. These legal changes may have improved women economic status, empowerment, marriage market outcomes such as participation in household decisions ultimately affecting infant mortality. Although this seems a plausible possibility, the point here is that all of these norms were applied retrospectively regardless of whether women were already married or not, and therefore, they should not affect differently women aged just below and above 15 years when the RFC was approved. Nonetheless, I examine empirically whether the law affected differently labor force participation, divorce rates, participation in household decisions and other marriage market outcomes of women at both sides of the cut-off. To measure woman participation in household decisions, I use a set of questions in the DHS survey that provide information on who decides on visits to relative, household purchases, health expenditure and the administration of the money earned by the husband. Each of these variables take the value of 0 if the woman does not participate in the decision, 1 if the woman participates in the decision and 2 if the decision is taken alone by the woman. Then, I construct a self-reported empowerment index for each woman as an average score in these questions. Using each of these self-reported empowerment measures, labor force participation and divorce rates as dependent variables, I estimate the reduced form equation presented in section 4. The results reported in table 4 suggest that, overall, the new dispositions included in the RFC aiming to change the balance of power within the household did not affect differently labor force participation, divorce rates and participation in household decisions of women at both sides of the cut-off. Furthermore, the lack of effects on partner's characteristics such as age difference between partners or partner's years of education rule out the possibility that the additional legal dispositions could have affected infant mortality through allowing women to marry better husbands. In consequence, the evidence confirms that the effect on infant mortality identified in the study is not driven by these additional norms aiming to improve women's bargaining power within the household.

TABLE 4—RFC AND DIFFERENT OUTCOMES.

	(1) Child marriage	(2) Age at 1st cohabit.	(3) Paid work	(4) Empowerment index	(5) Decision relative visits	(6) Decision HH purchases	(7) Decision health	(8) Decision husband earn	(9) Divorced (Only ever cohab)	(10) Age diff partner	(11) Years school partner
Age<15 at RFC	-0.200***	2.079***	0.098	-0.055	-0.319**	0.133	-0.061	-0.063	0.018	-0.856	0.645
	(0.058)	(0.389)	(0.079)	(0.064)	(0.130)	(0.115)	(0.127)	(0.057)	(0.046)	(0.928)	(0.757)
N	5078	5078	5077	4181	4175	4164	4170	4142	5078	4171	4999
N effect. obs.	935	787	1163	1035	775	771	898	893	958	824	1268
Bandwidth	39.0	32.5	48.0	49.5	38.0	37.4	42.5	42.9	39.7	39.9	53.8

Note: Each coefficient provided in the table is estimated using a separate regression. The table reports the estimates of interest for the reduced form (RF) equation using the optimal bandwidth and the robust variance estimator described in Calonico et al. (2016). For each outcome, the reduce form is estimated separately for the sample of women that ever cohabited with a partner and have born a child, and for all the women in the data regardless of whether they ever cohabited with a partner or born a child. The coefficients for the variable *Age<15 at RFC* measure the effect of the RFC on the outcome variable analyzed. The regressions conducted include as control variables a set of dummies for the regions of residence, the age of women at survey, ethnic and religion affiliation, gender of the first born, a rural/urban dummy variable and a non-parametric function for the age of the women at RFC. Standard errors are clustered at the forcing variable. Standard errors are in parentheses.***p<0.01;**p<0.05;**p<0.1.

Secondly, I examine the existence of discontinuities in variables that are plausibly not affected by the reform including the ethnicity and religion of the women and the gender of their first born. This is an indirect empirical test for the second identification assumption discussed in section 4: the determinants of infant mortality should be continuously related to the forcing variable at the cut-off. In order to test this hypothesis, I estimate equations 1, 2 and 3 using the bias-corrected RD estimates with robust variance estimator and an optimal bandwidth calculated following Calonico, Cattaneo and Titiunik (2014) and whether the first born is male, whether the mother is Orthodox[20] or from Oromo ethnic group[21] as outcome variables. The results of these estimations are provided in columns 1-6 of table 5. The coefficients are small and largely insignificant, confirming that there is not any discontinuity in the value of these placebo variables at the cut-off. The absence of discontinuities at the cut-off for these placebo variables is also evident in figure 7.

Thirdly, I examine whether the difference in infant mortality rates of the first born among women at both sides of the cut-off could be driven by systematic differences between women born in different months of the year rather than by exposure to a different minimum age of marriage. To assess this possibility, I re-estimate equations 1, 2 and 3 setting a placebo cut-off for women older than 19, rather than 15, at the time of the approval of the RFC[22]. This exercise is equivalent to placing the cut-off as if the law was introduced exactly four years before the true date of approval (e.g. 4th of July of 1996 for Addis Ababa and Dire Dawa, etc). If the results of the study on infant mortality are driven by systematic differences between women born in different months, we would expect a discontinuity in the infant mortality rate of the first born among women born in different months every year. The results of this placebo test are reported in columns 7 and 8 of table 5 and they reveal no discontinuities for the mean age at cohabitation or the infant mortality rate at the false cut-off, confirming that the main conclusions of the study are not driven by systematic differences between women born in different months of the year.

[20]Ethiopian Orthodox is the most prevalent religion in Ethiopia.

[21]Oromo is the largest ethnic group in Ethiopia.

[22]The set of the placebo cut-off at 19 years is driven by the convenience of setting the false cut-off at a value of the forcing variable that left out of the estimation the observations around the real cut-off. Nonetheless, I have also conducted the analysis setting the placebo cut-off for women aged 16, 17 and 18 at the time of the approval of the RFC. This exercise is equivalent to placing the cut-off as if the law was introduced exactly 1, 2 and 3 years before the true date of approval. Consistently, the results of these placebo analyses, not reported in the study, show no statistically significant discontinuities in infant mortality and age at cohabitation at the false cut-offs when the bandwidth drop from the estimation the observations in the real cut-off.

TABLE 5—ROBUSTNESS CHECKS INFANT MORTALITY: PLACEBO ANALYSES.

	Placebo outcomes					
	Ethnicity		Gender		Religion	
	(1) FS Age at 1st cohab	(2) SS/RF Ethnic. Oromo	(3) FS Age at 1st cohab	(4) SS/RF Male	(5) FS Age at 1st cohab	(6) SS/RF Orthodox
Age<15 at RFC	1.984***	-0.015	2.068***	-0.039	1.990***	0.028
	(0.399)	(0.050)	(0.392)	(0.110)	(0.402)	(0.056)
Age at 1st cohab.		-0.006		-0.018		0.015
		(0.031)		(0.058)		(0.035)
N		5078		5078		5078
N effect. obs.		874		874		874
Bandwidth		36.6		36.9		36.2
	Placebo: RFC 48 months before		Placebo: Other Ethiopian regions		Control for year first born	
	(7) FS Age at 1st cohab	(8) SS/RF Infant Mortality	(9) FS Age at 1st cohab	(10) SS/RF Infant Mortality	(11) FS Age at 1st cohab	(12) Infant Mortality
Age<15 at RFC	0.167	-0.001	-0.469	-0.007	1.107***	-0.074**
	(0.386)	(0.024)	(0.377)	(0.025)	(0.286)	(0.032)
Age at 1st cohab.		-0.007		0.015		-0.066*
		(0.150)		(0.062)		(0.034)
N		5078		2398		5078
N effect. obs.		2733		1458		1286
Bandwidth		89.5		106.9		53.1

Note: Each coefficient provided in the table is estimated using a separate regression. The table reports the estimates of interest for the first stage (FS), reduced form (RF) and second stage (SS) equations using the optimal bandwidth and the robust variance estimator described in Calonico et al. (2016). The coefficients for the variable *Age<15 at RFC* measure the effect of the RFC on the age at first cohabitation (FS) and on the outcome variable analyzed (RF). The coefficients for the variable *Age at 1st cohab* measure the effect of delaying one year the age at cohabitation during teenage years on the outcome variable analyzed (SS). The sample size and the bandwidths used in the RF, FS and SS regressions are common within every outcome analyzed. Columns 1-6 report the results for placebo variables, arguably unaffected by the RFC. Columns 7 and 8 report the results of a placebo test with a false cut-off set 4 years before the approval of the RFC. Columns 9 and 10 report the results of a placebo test using the regions that did not approve the RFC before 2008: Afar, Harari and Gumuz. The regressions conducted for the estimation of the coefficients reported in columns 1 to 10 include as control variables a set of dummies for the regions of residence, the age of women at survey, ethnic and religion affiliation, gender of the first born, a rural/urban dummy variable and a non-parametric function for the age of the women at RFC. Ethnic, gender and religion are not included as control variables in the regressions where these variables are the outcome variables. Columns 11 and 12 report the results for the infant mortality including as an additional control variable a time trend for the year of birth of the infant, aiming to account for over time changes in the incidence of infant mortality. Standard errors are clustered at the forcing variable. Standard errors are in parentheses.***p<0.01;**p<0.05;**p<0.1.

FIGURE 7. PLACEBO VARIABLES: ETHNICITY, RELIGION AND GENDER

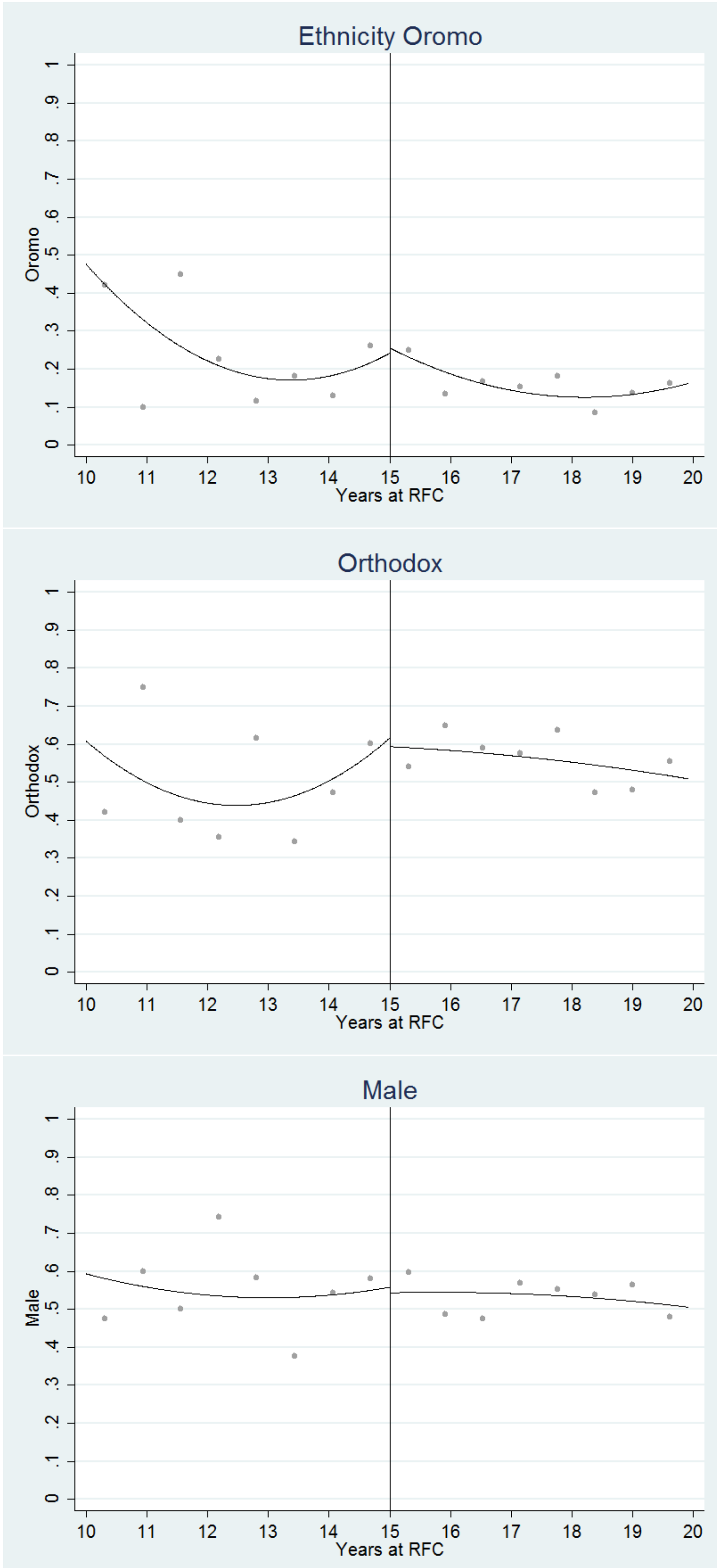
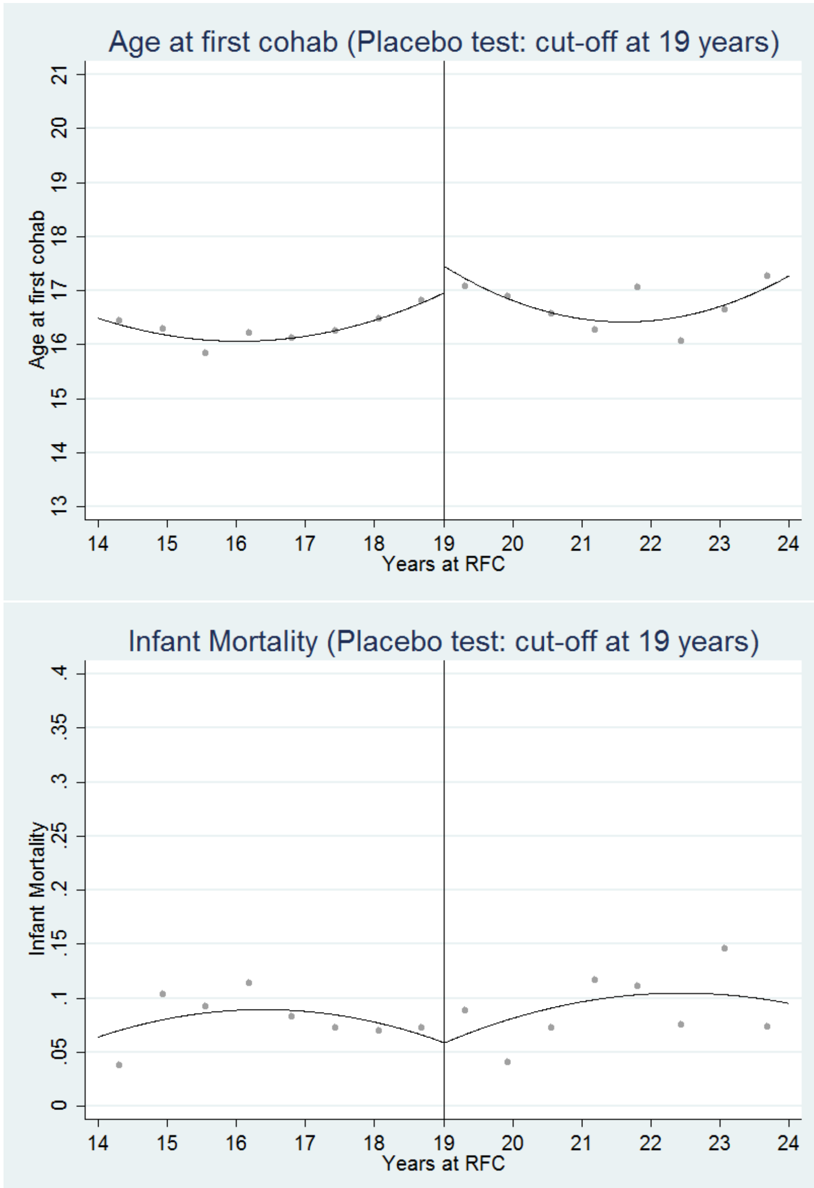


FIGURE 8. PLACEBO TEST: CUT-OFF AT 19 YEARS



Fourthly, it is also possible that the parameter could be capturing an *odd* time trend or the effect of national level policies affecting differently women at both sides of the cut-off. In order to test this hypothesis, I re-estimate the results using the Ethiopian regions of Afar, Harari and Gumuz, setting falsely the approval date of the RFC in these placebo regions in July 2000[23]. I restrict the analysis to these three regions because none of them passed the RFC before 2008[24] and therefore,

[23]The 4th of July of 2000 the Federal Government of Ethiopia approved the RFC and the law started to be applied in Addis Ababa and Dire Dawa.

[24]I exclude the region of Somali because it is unclear their exact date of introduction (Hallward-

the women used in the sample (older than 18 years at the time of the survey) were unaffected by the rise in the legal age at marriage in these regions. The results of this placebo test are reported in columns 9 and 10 of table 5 and they show that there is not any significant discontinuity in the mean age at cohabitation or in the infant mortality at the cut-off in those regions that have not approved the RFC before 2008, ruling out the possibility that the effect is driven by a national level policy affecting differently women at both sides of the cut-off.

Fifthly, since the law plausibly affected the age at first birth of women at the cut-off, it is very likely that those women aged 14 years and 11 months at the time of the approval of the RFC ended up having their first child significantly later than those aged 15 years and 1 month. In this context, the parameters δ_1 and β_1 could be only capturing over-time decreases in infant mortality unrelated with the age at first cohabitation. I investigate this possibility through re-estimating equations 1, 2 and 3 including the year of birth of the first born as a control variable. The inclusion of this variable is aiming to capture time trends in infant mortality[25]. The estimates provided in columns 10 and 11 show that the direction and significance of the parameters estimated in sections 6.6.1, 6.6.2 and 6.6.3 do not vary when the year of birth of the first child is included in the regression. The latter suggests that the reduction in the rate of infant mortality of the first born at the cut-off is not driven by over time decreases in infant mortality.

Sixthly, the main results of the study are also robust to restricting the analysis to the subsample of women that by the time of the survey were still living with their first partner. The results of this robustness check are reported in columns 1 to 4 of table A1 in the appendix and they are consistent with those obtained when using the main sample.

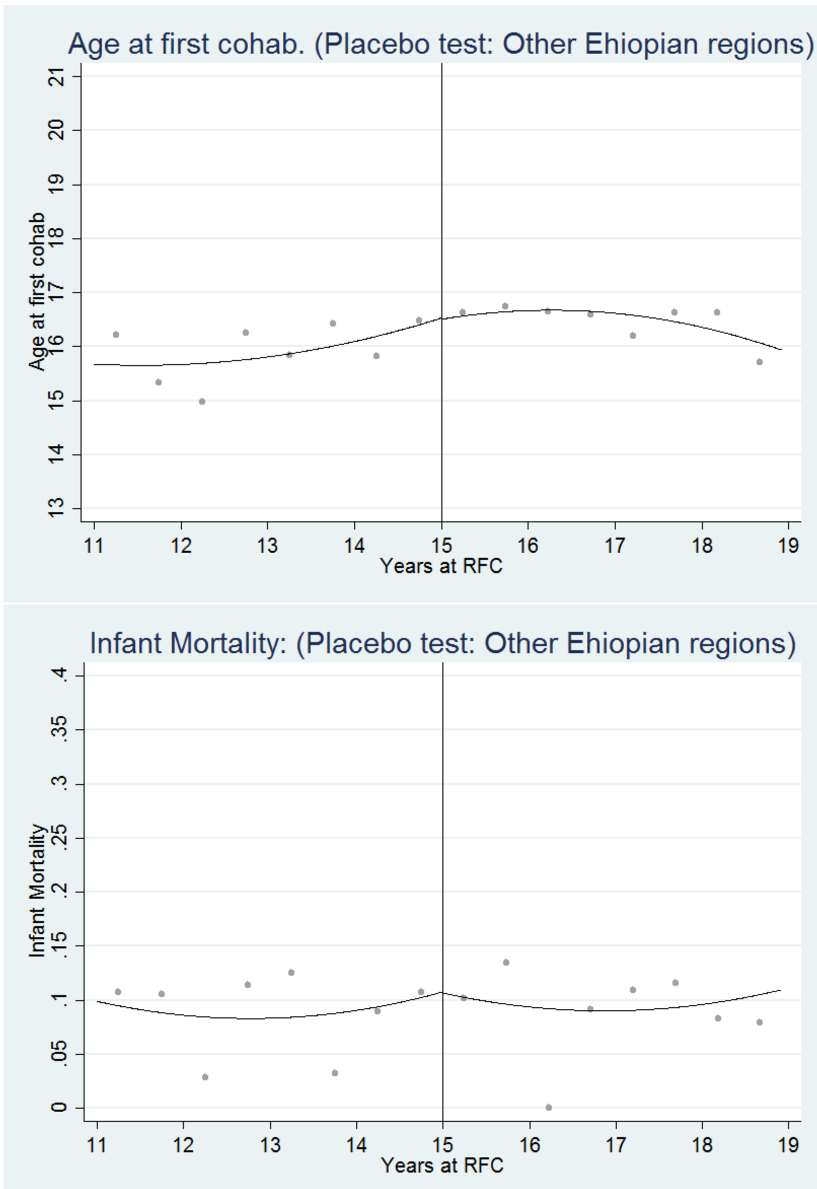
Another threat to the interpretation of the estimates could be the possible existence of selective migration. In other words, those women below the age of 15 when the legal age of marriage was raised in their region that were particularly interested in early cohabitation could have migrated to regions that did not raise the legal age of marriage. If the share of women migrating for this reason is substantial, the results might be biased by selective attrition at one side of the cut-off. Although the lack of information on women's region of origin hindered the assessment of this hypothesis, the fact that the incidence of child marriage is above 10% in every Ethiopian region even among those women exposed to a legal age of marriage at 18 years could be indicating that those women

Driemeier and Gajigo, 2015).

[25]The year of birth of the first born is not included as control variable in the main results reported in section 6 because the year of birth could be to some extent a fertility decision of the mother and therefore, plausibly affected by the age at cohabitation. Thus, including it in the main regression could lead to a *bad* control problem (Angrist and Pischke, 2008).

(or women's families) particularly interested in cohabiting before the age of 18 may not need to migrate to another region to do so. Furthermore, the lack of discontinuity in the density of the forcing variable at the cut-off examined in figure A4 in the appendix suggests that the migration of women slightly younger than 15 when the RFC was approved was not a widespread phenomenon.

FIGURE 9. PLACEBO TEST: DISCONTINUITY IN OTHER ETHIOPIAN REGIONS



I also explore the robustness of the results to the use of mortality within the first months of life as a dependent variable rather than infant mortality. For this, I re-estimate equations 1, 2 and 3 focusing on mortality of the first born within the month of birth and the following one, rather than within the first year. The estimates obtained have the same sign and statistical significance although the magnitudes are even larger than those obtained for infant mortality of the first born[26]. Therefore, the evidence confirms that the results of the study are robust to the definition of the dependent variable and suggests that most of the effect of child marriage on the mortality of the first born occurs during the very first months of life of the newborn.

Finally, the possibility of women manipulating their reported age in the survey (the base for the construction of the forcing variable) is examined conducting a McCrary test. The results of the McCrary test, graphically displayed in figure A4 in the appendix, show that the density of the age of women does not change significantly at the cut-off suggesting that women just below or above the cut-off age did not systematically misreport their age in the survey. Finally, the potential bias induced by measurement error in the reported age at cohabitation in the second stage equation is addressed through the use of women's age at RFC as an instrumental variable in the regression.

7. Mechanisms

One potential mechanism driving the effect of early cohabitation on infant mortality of the first born could be the age of the woman at first birth. Figure 4 shows the strong negative association between age at first born and infant mortality among early teenage girls in the Ethiopian data. This negative association between age at first birth and infant mortality during teenage years is well documented in the medical literature and could reflect that the body of teenage women is still not optimal for the development of a successful pregnancy and/or the effect of psychological maturity on the adoption of adequate antenatal and postnatal health behaviors (Chen et al., 2007; Olausson, Cnattingius and Haglund, 1999).

To examine this mechanism, I re-estimate equations 1, 2 and 3 using fertility outcomes, antenatal and postnatal behaviors as dependent variables in the reduced form and second stage equations.

The results on age at first birth are displayed in columns 1 and 2 of table 6 and confirm that although the coefficient is lower than 1, delaying cohabitation causally increases the age of women at first birth. In line with this *age at birth* mechanism, the results reported in columns 5 and 6 suggest that the effect of age at cohabitation on infant mortality vanishes for children born after

[26]These results are not reported in the tables provided in the study. The estimates reveal that the effect of a one-year delay in women's age at cohabitation during teenage years on the probability of mortality within the first two months of life is estimated at 5 percentage points, statistically significant at the 5% significance level. The effect of one year delay in the age at cohabitation on mortality within the month of birth, also known as neonatal mortality, is estimated at 3 percentage points, although in this case the magnitude is not statistically significant at conventional significance levels.

TABLE 6—ANALYSIS OF MECHANISMS

	Age at birth		Educ level. Secondary or higher		Infant mortality Non-first born		Years school (compl. years)		Anemia (0/1)		Work (0/1)		Empowerment index		Wealth index	
	(1) FS	(2) RF/SS	(3) FS	(4) RF/SS	(5) FS	(6) RF/SS	(7) FS	(8) RF/SS	(9) FS	(10) RF/SS	(11) FS	(12) RF/SS	(13) FS	(14) RF/SS	(15) FS	(16) RF/SS
Age<15 at RFC	2.068***	1.011***	2.053***	0.014	2.352***	0.008	2.078***	-0.731	1.809***	-0.066	2.094***	0.035	1.948***	-0.094	2.069***	-0.080
	(0.383)	(0.372)	(0.371)	(0.074)	(0.590)	(0.055)	(0.375)	(0.485)	(0.386)	(0.088)	(0.386)	(0.098)	(0.389)	(0.072)	(0.378)	(0.228)
Age at 1st cohab.		0.490**		0.008		0.004		-0.357		-0.037		0.016		-0.048		-0.037
		(0.194)		()		(0.027)		(0.308)		(0.056)		(0.054)		(0.039)		(0.129)
N		5078		5078		16373		5078		4824		5077		4181		5078
N effect. obs.		874		935		737		847		905		828		691		905
Bandwidth		36.2		38.4		32.4		35.9		39.4		34.0		33.9		37.7

	Years school partner		Age diff. partner		Months Breastfeed.		Birth weight		N Vaccin.		Birth at home		Any antenatal visits		Anemia child (0/1)	
	(17) FS	(18) RF/SS	(19) FS	(20) RF/SS	(21) FS	(22) RF/SS	(23) FS	(24) RF/SS	(25) FS	(26) RF/SS	(27) FS	(28) RF/SS	(29) FS	(30) RF/SS	(31) FS	(32) RF/SS
Age<15 at RFC	2.050***	0.783	1.898***	-0.890	2.408***	-1.230	-1.619*	-0.458	0.825	0.199	1.134***	0.011	1.119**	-0.084	0.738	-0.104
	(0.378)	(0.859)	(0.371)	(0.932)	(0.851)	(5.297)	(0.882)	(0.308)	(0.513)	(0.645)	(0.404)	(0.100)	(0.547)	(0.083)	(0.589)	(0.135)
Age at 1st cohab.		0.383		-0.467		-0.453		0.227		0.268		0.008		-0.076		-0.140
		(0.523)		(0.562)		(2.791)		(0.274)		(0.880)		(0.095)		(0.080)		(0.186)
N		4999		4171		240		204		781		862		520		686
N effect. obs.		864		900		159		105		458		623		443		506
Bandwidth		36.2		42.0		42.2		39.4		40.2		52.7		67.2		53.2

Note: Each coefficient provided in the table is estimated using a separate regression. The table reports the estimates of interest for the first stage (FS), reduced form (RF) and second stage (SS) equations using the optimal bandwidth and the robust variance estimator described in Calonico et al. (2016). The coefficients for the variable *Age<15 at RFC* measure the effect of the RFC on the age at first cohabitation (FS) and on the outcome variable analyzed (RF). The coefficients for the variable *Age at 1st cohab* measure the effect of delaying one year the age at cohabitation during teenage on the outcome variable analyzed (SS). The sample size and the bandwidths used in the RF, FS and SS regressions are common within every outcome analyzed. The regressions conducted include as control variables a set of dummies for the regions of residence, the age of women at survey, ethnic and religion affiliation, gender of the first born, a rural/urban dummy variable and a non-parametric function for the age of the women at RFC. Columns 1-20 and 27-28 report the results for the whole sample used in the main analysis. Column 21-32 conducts the analysis using the sample of women that had their first born within the last 5 years. For these cases, the survey provides information on maternal and child health outcomes. Standard errors are clustered at the forcing variable. Standard errors are in parentheses.***p<0.01;**p<0.05;**p<0.1.

the first child.

On the other hand, the effect of the age at cohabitation on the adoption of antenatal and postnatal health practices is statistically indistinguishable from 0 at conventional significance levels. The coefficient measuring the effect of women's age at cohabitation on the number of vaccines received by the children has the expected positive sign although the magnitude is small and statistically insignificant when the optimal bandwidth is used. Similarly, the coefficients for antenatal visits, months breastfeed and birth at home in the second stage equations are statistically insignificant at conventional confidence levels. Nonetheless, it is important to remark that the results on the adoption of antenatal and postnatal health behaviors should only be interpreted as suggestive because DHS data only report information on these variables when the first born is alive and was born less than 5 years ago. This could be problematic because on the one hand the sample

size used in the analysis is much smaller, reducing the statistical power of the estimations and on the other hand, the limitations in the data collection may induce a problem of sample selection bias. Indeed, using this limited sub-sample of children, I do not find any significant effect of delaying cohabitation on measures of child health such as birth weight or the prevalence of anemia (results reported in columns 24 and 32 of table 6).

However, the age at cohabitation could also affect infant mortality of the first born through other paths. For example, a younger age of the women at cohabitation may lead to lower levels of participation in household decisions (Jensen and Thornton, 2003). Since women and men have different preferences for investment in children's health (Allendorf, 2007; Majlesi, 2014), early cohabitation may lead to higher infant mortality rates. Similarly, a younger age at cohabitation may affect investments in child's health and infant mortality through constraining women's education (Field and Ambrus, 2008) and impacting labor market outcomes (Elborgh-Woytek et al., n.d.). On the other hand, given the existence of a premium for early marriage in the marriage market (Wahhaj, 2015), it is also possible that early marriage reduces infant mortality through improving women's marriage market outcomes.

To investigate these paths of impact, I estimate the effect of early cohabitation on different women outcomes including participation in household decisions, marriage market outcomes, labor force participation, health and education attainment. The results of the estimates for equations 1, 2 and 3 using these outcomes as dependent variables are reported in columns 7 to 20 of table 6. The results confirm that exposure to a legal age of marriage at 18 and delaying age at cohabitation do not seem to affect relevantly women‘s labor force participation, health, education, participation in household decisions and marriage market outcomes including age difference with partner, wealth index and partner's years of education.

Jointly, these results suggest that the effect on infant mortality of delaying age at cohabitation during teenage years for women seems to operate mainly through increasing the age at first birth. On the other hand and given the data limitations, we cannot rule out that the effect is also channeled through larger levels of antenatal and postnatal investments for girls cohabiting later. However, given the lack of effect of the age at cohabitation on self-reported empowerment, labor market outcomes or education, any potential effect of the age at cohabitation on infant mortality via antenatal and postnatal investments would probably be linked to an older age at first birth and a more mature behavior.

8. Conclusions

This study documents for the first time in the literature the causal effect of child marriage on infant mortality, shows that laws raising the legal age of marriage could be effective strategies to reduce child marriage and infant mortality and provides an alternative identification strategy that can be used to expand the analysis on the causal effect of child marriage to other outcomes and settings where similar laws have been approved. Through using a RDD strategy exploiting age discontinuities in exposure to a law that raised the legal age of marriage for Ethiopian women, the study finds that in the five Ethiopian regions analyzed, exposure to a legal age of marriage at 18 years relative to the possibility of marrying legally at 15 increases significantly the mean age at cohabitation by 2 years and decreases the probability of infant mortality of the first born child by 7.9 percentage points. The reduction in the probability of infant mortality of the first born child caused by a delay of one year in women's age at cohabitation during teenage years is estimated at 3.8 percentage points. The size of this effect is comparable to the joint effect on child mortality at the village level of moving from 0% coverage of measles, BCG, DPT, Polio and Maternal Tetanus vaccinations to 100% (McGovern and Canning, 2015)[27].

The analysis of mechanisms suggests that the strong impact of early cohabitation on infant mortality of the first born seems to be driven by the positive effect of delaying cohabitation on the age at first birth. However, due to data limitations, it is not possible to disentangle whether this *age at birth* effect is caused by purely biological reasons or by a more mature behavior raising the

[27]Using DHS data from 62 countries, McGovern and Canning (2015) estimate at 0.73 the relative risk of child mortality at the village level associated with moving from 0 percent coverage of measles, BCG, DPT, Polio and Maternal Tetanus vaccinations to 100%. At the cut-off, the relative risk of the infant mortality of the first born associated with a one-year increase in the age at cohabitation is estimated at 0.75.

adoption of antenatal and postnatal health measures. On the other hand, the analysis rules out the possibility that the effect of early cohabitation on infant mortality of the first born is driven by an effect of the former on marriage market outcomes, participation in household decisions, education or labor force participation for women. The causal link between child marriage, age at birth and infant mortality is consistent with existing evidence from studies that use correlation analysis to explore the link between early marriage, teenage pregnancy and infant mortality. For example, Raj et al. (2010) find an odds ratio of 1.5, statistically significant at the 95%, for the statistical association between infant mortality and women that marry before the age of 18 in India. In this line, using household data from Nepal, Adhikari (2003) shows that neonatal mortality rates among children of mothers 15-19 are 73% higher than among children of mothers 20-29.

Although the effects on infant mortality found in the study are large, it is important to acknowledge that the estimates yielded by the RDD employed are local and therefore, any generalization of these results to the whole Ethiopian female population or to non-first born infants should be avoided. First, the RDD approach only identifies effects for those women at the cut-off. In our case, these are women aged 18 to 26 years old at the time of the survey that ever cohabited with a partner and gave birth to their first child at least one year before the survey. Second, the causal effect of delaying cohabitation on infant mortality is only identified for the compliers. These are women that postpone their age at cohabitation because they were exposed to a different legal age at marriage, which are arguably either law abiding women or women that live in areas where the presence of the state is stronger. Third, the analysis does not find any effect of early cohabitation on the infant mortality of children born after the first child, which is consistent with the *age at birth* mechanism.

APPENDIX

FIGURE A1. AGE AT 1ST COHABITATION BY AGE COHORT.

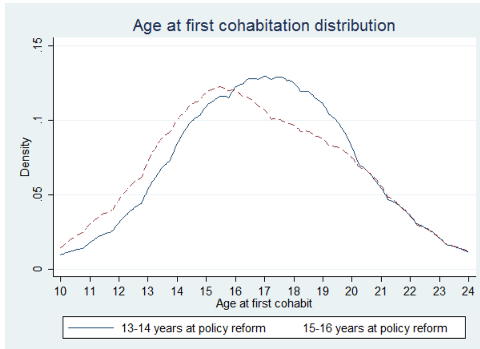


FIGURE A2. AGE AT FIRST COHABITATION AT THE CUT-OFF FOR THE ETHIOPIAN REGIONS THAT APPROVED RFC BETWEEN 2000 AND 2007

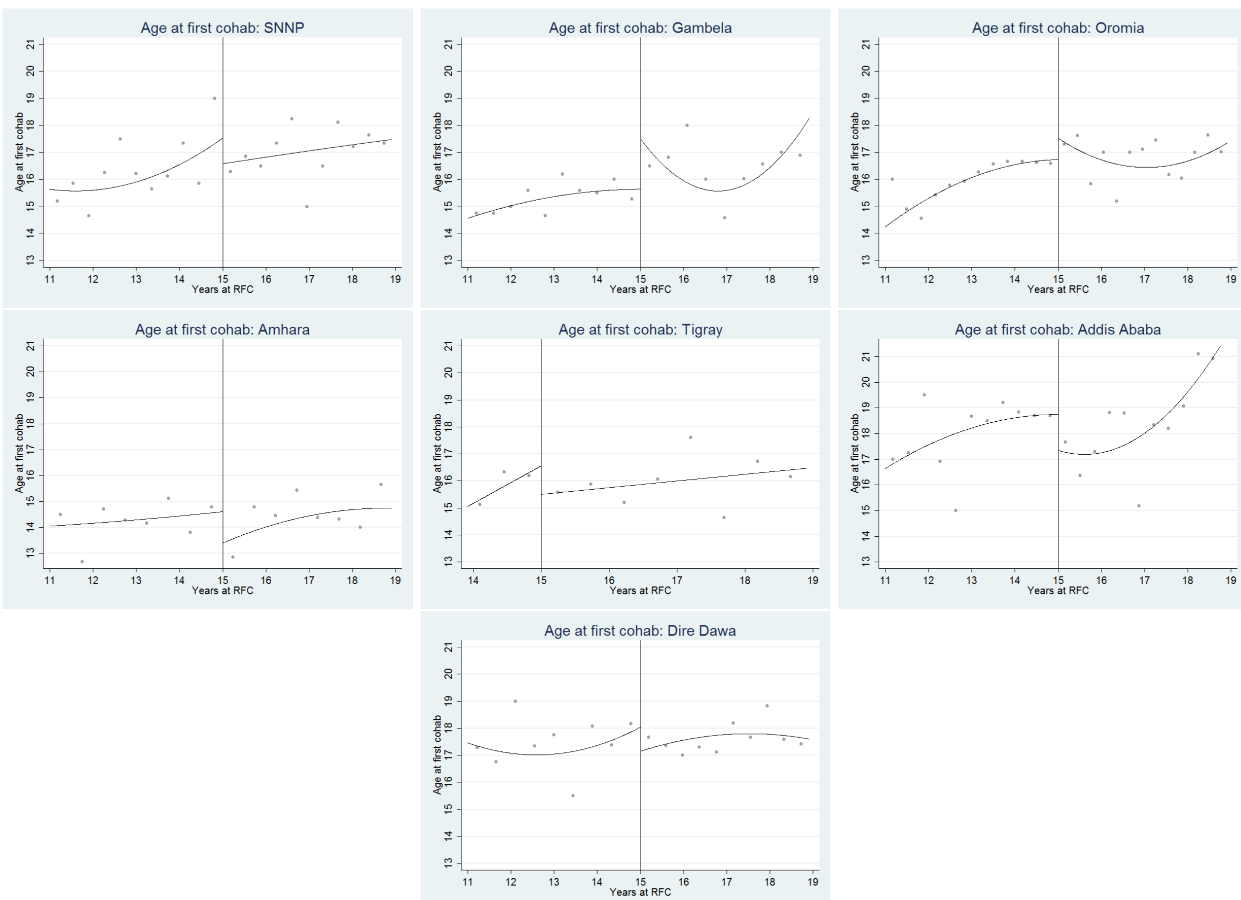


FIGURE A3. RFC AND CHILD MARRIAGE: INCLUDES ALL WOMEN (NOT ONLY THOSE THAT EVER COHABITED AND GAVE BIRTH)

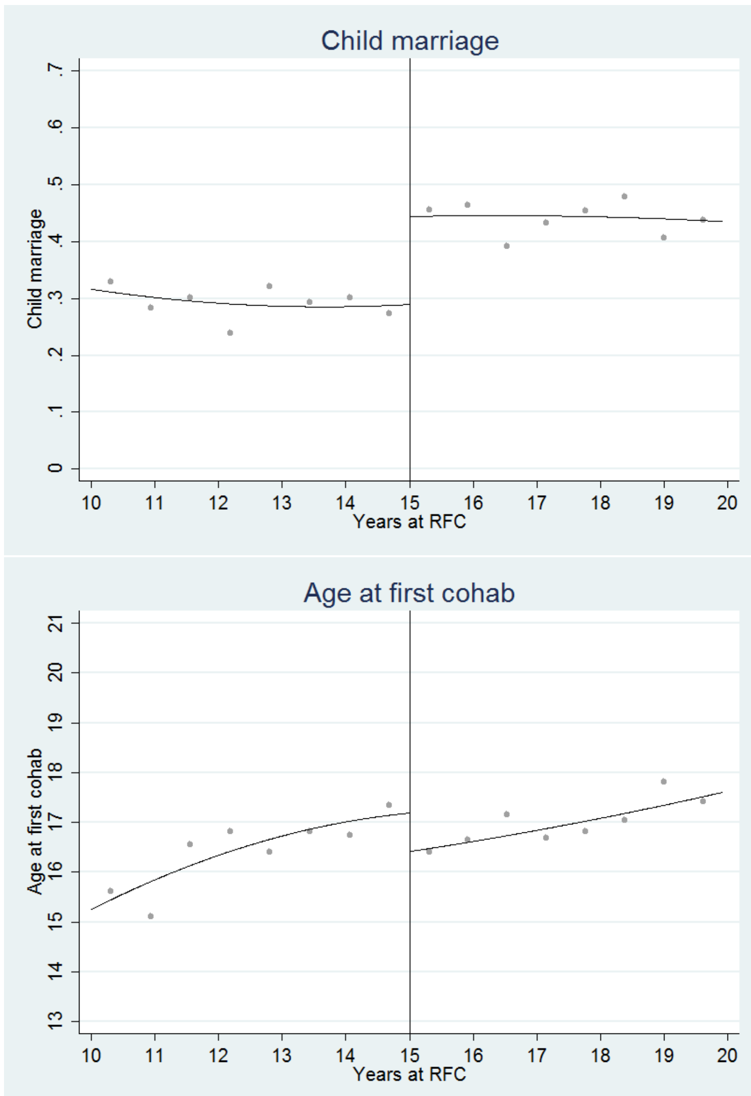
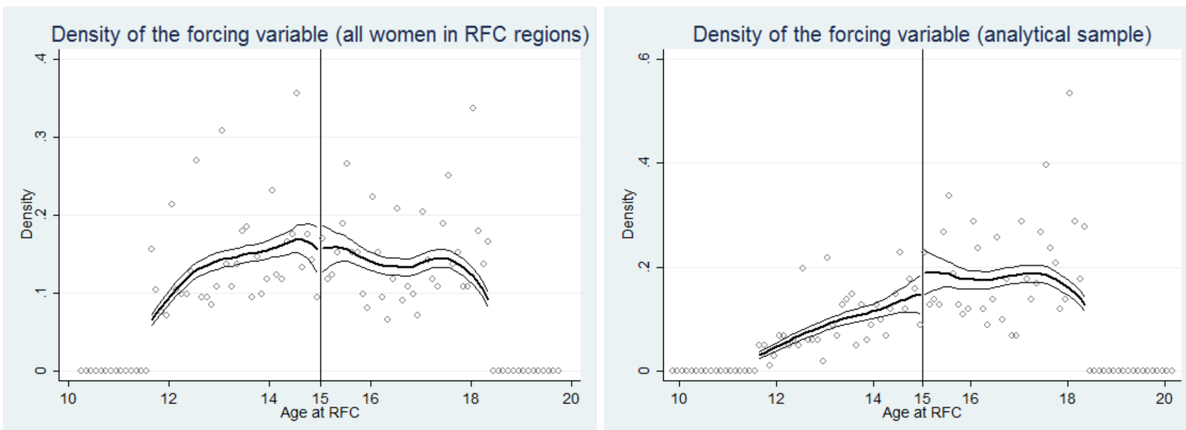


FIGURE A4. MCCRARY TEST: DENSITY OF THE FORCING VARIABLE AT THE CUT-OFF



The figures show the density function of the running variable and 95% confidence intervals following the McCrary procedure for two different samples. The first sample includes all women in the DHS data that live in the Ethiopian regions of Addis Ababa, Dire Dawa, Amhara, SNNP and Tigray. The second sample includes the women living in the same regions that ever cohabited, ever gave birth and the first born occurred at least one year before the survey. I use for the construction of the density functions those observations within a bandwidth of 40 months from the cut-off, which corresponds to the optimal bandwidth used in the main non-parametric estimations of the paper. In both cases, the results of the McCrary test show that the discontinuity in the frequency of the running variable at the cut-off is not significant at conventional confidence levels (p-value>0.1).

FIGURE A5. RFC AND WOMEN CHARACTERISTICS

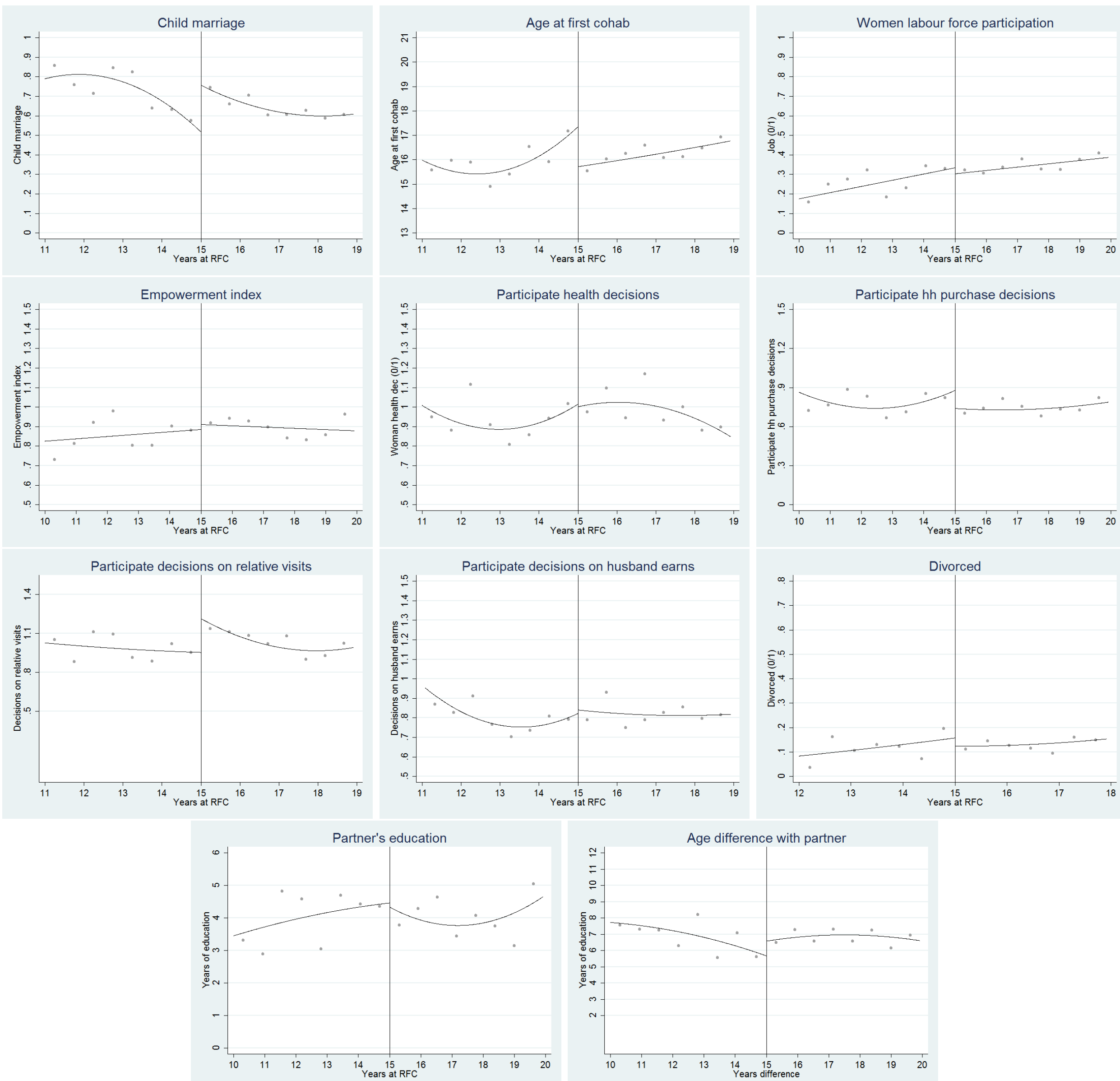


FIGURE A6. MECHANISMS: MATERNITY AND CHILD HEALTH FOR FIRST BORN

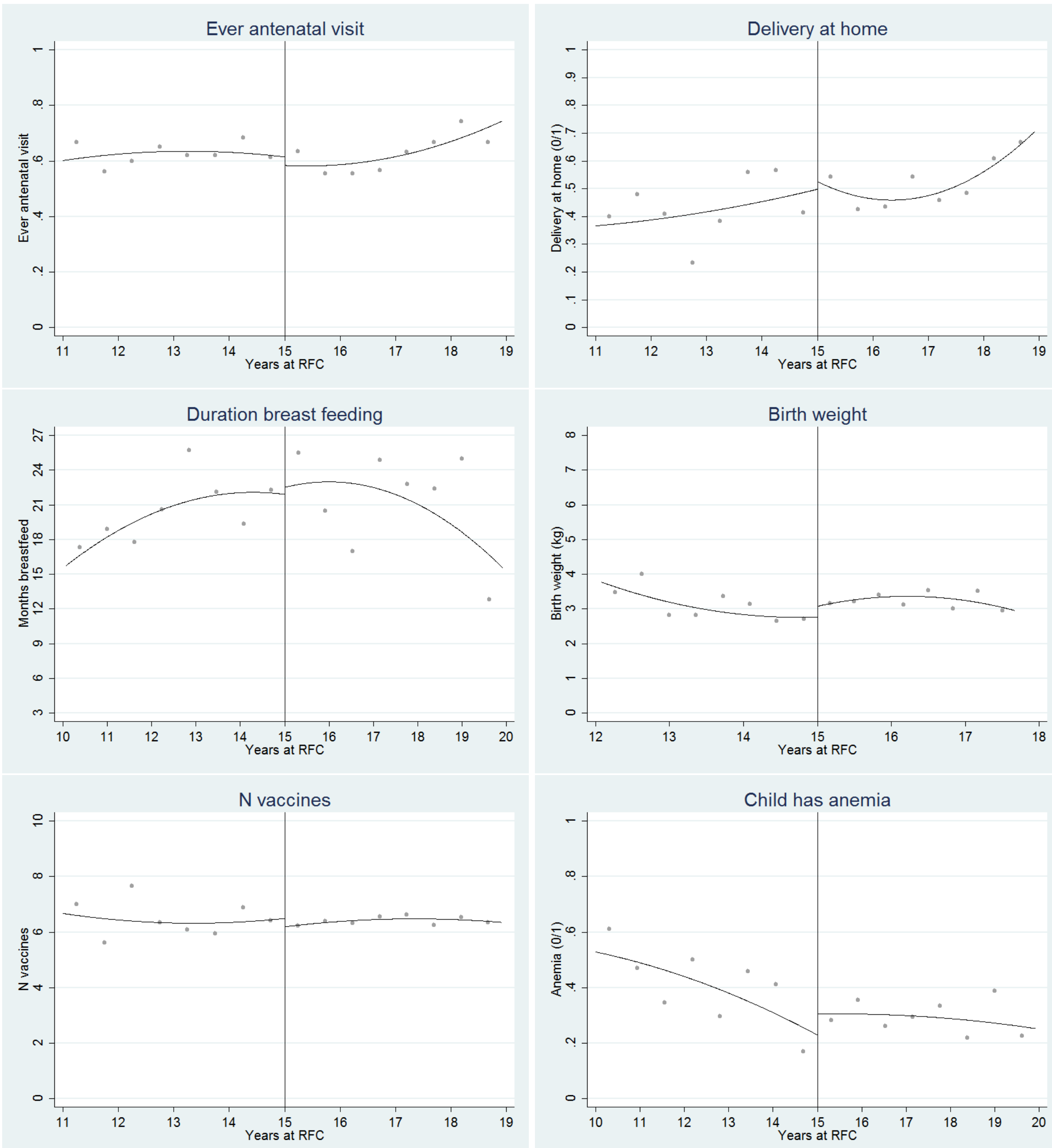


FIGURE A7. MECHANISMS: FERTILITY, MARRIAGE MARKET OUTCOMES AND OTHER WOMEN OUTCOMES

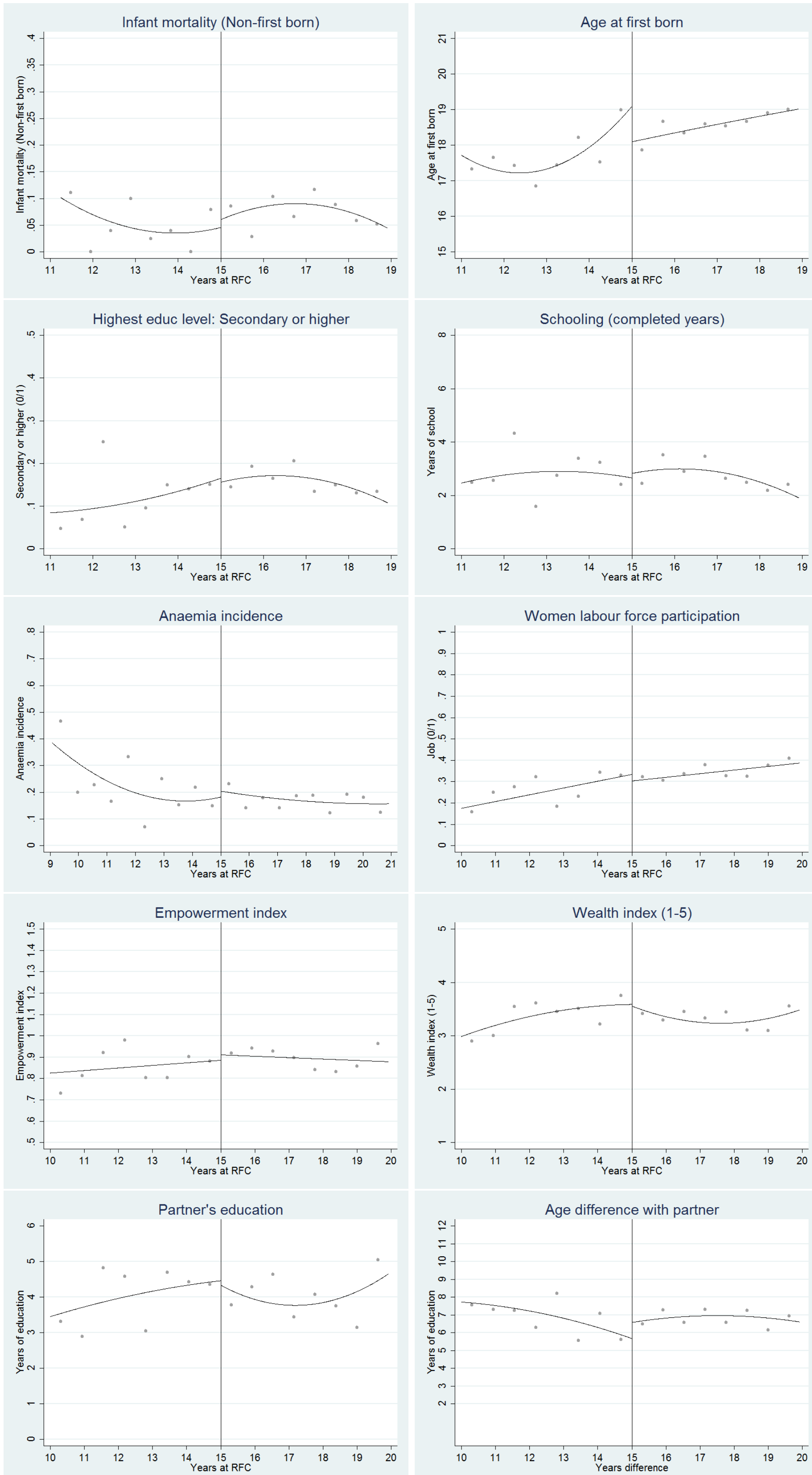


TABLE A1—ONLY WOMEN STILL COHABITING WITH FIRST COHABIT.

	Age at birth		Infant Mortality		Years school		Anaemia	
	(1) FS	(2) RF/SS	(3) FS	(4) RF/SS	(5) FS	(6) RF/SS	(7) FS	(8) RF/SS
Age<15 at RFC	1.353***	1.192***	1.333***	-0.097*	1.333***	-1.049	1.293***	-0.096
	(0.429)	(0.353)	(0.434)	(0.055)	(0.439)	(0.898)	(0.435)	(0.074)
Age at 1st cohab.		0.877***		-0.073*		-0.782		-0.075
		(0.285)		(0.042)		(0.943)		(0.077)
N		3126		3126		3126		2985
N effect. obs.		692		674		651		739
Bandwidth		39.8		38.9		37.1		43.8

	Empowerment index		Years school partner		Age Difference		Work	
	(9) FS	(10) RF/SS	(11) FS	(12) RF/SS	(13) FS	(14) RF/SS	(15) FS	(16) RF/SS
Age<15 at RFC	1.296***	-0.111	1.296***	0.001	1.282***	-0.908	1.346***	0.001
	(0.444)	(0.070)	(0.433)	(1.154)	(0.445)	(0.943)	(0.433)	(0.074)
Age at first cohabit.		-0.085		0.002		-0.681		0.002
		(0.064)		(1.049)		(0.786)		(0.066)
N		3118		3102		3112		3125
N effect. obs.		604		670		625		674
Bandwidth		35.8		38.3		36.8		38.8

Note: The analysis reported in the table is conducted using the sample of women 18-49 that ever gave birth and still cohabit with first partner. Each coefficient provided in the table is estimated using a separate regression. The table reports the estimates of interest for the first stage (FS), reduced form (RF) and second stage (SS) equations using the optimal bandwidth and the robust variance estimator described in Calonico et al. (2016). The coefficients for the variable *Age<15 at RFC* measure the effect of the RFC on the age at first cohabitation (FS) and on the outcome variable analyzed (RF). The coefficients for the variable *Age at 1st cohab* measure the effect of delaying one year the age at cohabitation during teenage on the outcome variable analyzed (SS). The sample size and the bandwidths used in the RF, FS and SS regressions are common within every outcome analyzed. The regressions conducted include as control variables a set of dummies for the regions of residence, the age of women at survey, ethnic and religion affiliation, gender of the first born, a rural/urban dummy variable and a non-parametric function for the age of the women at RFC. Standard errors are clustered at the forcing variable. Standard errors are in parentheses.***p<0.01;**p<0.05;**p<0.1.

*

REFERENCES

Adhikari, Ramesh. 2003. "Early marriage and childbearing: risks and consequences." In *Towards adulthood: exploring the sexual and reproductive health of adolescents in South Asia.* , ed. Sarah Bott, Shireen Jejeebhoy, Iqbal Shah and Chander Puri, 62–66. World Health Organization.

Allendorf, Keera. 2007. "Do Women's Land Rights Promote Empowerment and Child Health in Nepal?" *World Development*, 35(11): 1975–1988.

Angrist, Joshua D., and Jorn-Steffen Pischke. 2008. *Mostly Harmless Econometrics: An Empiricist's Companion.* Princeton University Press.

Asadullah, Mohammad, and Zaki Wahhaj. 2016. "Early Marriage, Social Networks and the Transmission of Norms." School of Economics, University of Kent Studies in Economics.

Asadullah, Niaz, Abdul Alim, Fathema Khatoon, and Nazmul Chaudhury. 2016. "Maternal Early Marriage and Cognitive Skills Development: An Intergenerational Analysis." Unpublished working paper Technical Report.

Barrios, Yasmin, Sixto Sanchez, Christina Nicolaidis, Pedro Garcia, Bizu Gelaye, Qiuyue Zhong, and Michelle Williams. 2015. "Childhood Abuse and Early Menarche among Peruvian Women." *Journal of Adolescent Health*, 56(2): 197–202.

Bound, John, David A. Jaeger, and Regina M. Baker. 1995. "Problems with Instrumental Variables Estimation When the Correlation Between the Instruments and the Endogeneous Explanatory Variable is Weak." *Journal of the American Statistical Association*, 90(430): 443–450.

Calonico, Sebastian, Matias D. Cattaneo, and Rocio Titiunik. 2014. "Robust Nonparametric Confidence Intervals for Regression Discontinuity Designs." *Econometrica*, 82: 2295–2326.

Calonico, Sebastian, Matias D. Cattaneo, Max Farrell, and Rocio Titiunik. 2016. "Regression Discontinuity Designs Using Covariates." University of Michigan.

Campbell, Jacquelyn. 2002. "Health Consequences of Intimate Partner Violence." *The Lancet*, 359(9314): 1331–1336.

Chari, Amalavoyal, Rachel Heath, Annemie Maertens, and Freeha Fatima. 2017. "The causal effect of maternal age at marriage on child wellbeing: Evidence from India." *Journal of Development Economics*, 127: 42 – 55.

Chen, Xi-Kuan, Shi Wu Wen, Nathalie Fleming, Kitaw Demissie, George Rhoads, and Mark Walker. 2007. "Teenage Pregnancies and Adverse Birth Outcomes: A Large Population Based Retrospective Study." *International Journal of Epidemiology*, 36(2): 368–373.

Dixon, Ruth. 1971. "Cross-Cultural Variations in Age at Marriage and Proportions Never Marrying." *Population Studies*, 25(2): 215–233.

Elborgh-Woytek, Katrin, Monique Newiak, Kalpana Kochhar, Stefania Fabrizio, Kangni Kpodar, Philippe Wingender, Benedict Clements, and Schwartz Gerd. n.d.. "Women, Work, and the Economy: Macroeconomic Gains from Gender Equity." 13/10, International Monetary Fund: Staff Discussion Notes.

Field, Erica, and Attila Ambrus. 2008. "Early Marriage, Age of Menarche, and Female Schooling Attainment in Bangladesh." *Journal of Political Economy*, 116(5): 881–930.

Gelman, Andrew, and Guido Imbens. 2017. "Why high-order polynomials should not be used in regression discontinuity designs." *Journal of Business & Economic Statistics*.

Goody, Jack. 1990. *The Oriental, the Ancient and the Primitive: Systems of Marriage and the Family in the Pre-Industrial Societies of Eurasia.* Cambridge University Press.

Hallward-Driemeier, Mary, and Ousman Gajigo. 2015. "Strengthening Economic Rights and Women's Occupational Choice: The Impact of Reforming Ethiopia's Family Law." *World Development*, 70: 260–273.

Hicks, Joan, and Daniel Hicks. 2015. "Lucky Late Bloomers? The Consequences of Early Marriage for Women in Western Kenya." Unpublished working paper Technical Report.

Jain, Saranga, and Kathleen Kurz. 2007. "New Insights on Preventing Child Marriage: A Global Analysis of Factors and Programs." International Center for Research on Women.

Jensen, Robert, and Rebecca Thornton. 2003. "Early Female Marriage in the Developing World." *Gender and Development*, 11(2): 9–19.

Jones, Nicola, Bekele Tefera, Guday Emirie, Bethelihem Gebre, Kiros Berhanu, Elizabeth Presler-Marshall, David Walker, Taveeshi Gupta, and Georgia Plank. 2016. "One Size Does Not Fit All: The Patterning and Drivers of Child Marriage in Ethiopia's Hotspot Districts." UNICEF and ODI.

Karapanou, Olga, and Anastasios Papadimitriou. 2010. "Determinants of Menarche." *Reproductive Biology and Endocrinology*, 8(115): 197–202.

Lee, David S., and Thomas Lemieux. 2010. "Regression Discontinuity Designs in Economics." *Journal of Economic Literature*, 48(2): 281–355.

Majlesi, Kaveh. 2014. "Labor Market Opportunities and Women's Decision Making Power within Households." Lund University, Department of Economics Working Papers 2014:4.

McCrary, Justin. 2008. "Manipulation of the running variable in the regression discontinuity design: A density test." *Journal of Econometrics*, 142(2): 698–714.

McGovern, Mark E., and David Canning. 2015. "Vaccination and all-cause child mortality from 1985 to 2011: global evidence from the Demographic and Health Surveys." *Journal of American Epidemiology*, 182(9): 791–798.

Moghadam, Valentine. 2004. "Patriarchy in Transition: Women and the Changing Family in the Middle East." *Journal of Comparative Family Studies*, 35(2): 137–162.

Nguyen, Minh, and Quentin Wodon. 2015. *Impact of Early Marriage on Literacy and Educatio Attainment in Africa in Child Marriage and Education in Sub-Saharan Africa.* World Bank.

Olausson, Petra, Sven Cnattingius, and Bengt Haglund. 1999. "Teenage Pregnancies and Risk of Late Fetal Death and Infant Mortality." *British Journal of Obstetric Gynaecology*, 106(2): 116–121.

Parsons, Jennifer, Jeffrey Edmeades, Aslihan Kes, Suzanne Petroni, Maggie Sexton, and Quentin Wodon. 2015. "Economic Impacts of Child Marriage: A Review of the Literature." *The Review of Faith & International Affairs*, 13(3): 12–22.

Pullum, Thomas W. 2008. "An Assessment of the Quality of Data on Health and Nutrition in the DHS Surveys 1993-2003." USAID.

Raj, Anita, Niranjan Saggurti, Michael Winter, Alan Labonte, Michele R Decker, Donta Balaiah, and Jay G Silverman. 2010. "The effect of maternal child marriage on morbidity and mortality of children under 5 in India: cross sectional study of a nationally representative sample." *BMJ*, 340.

Sekhri, Sheetal, and Sisir Debnath. 2014. "Intergenerational Consequences of Early Age Marriages of Girls: Effect on Children's Human Capital." *Journal of Development Studies*, 50(12): 1670–1686.

Solanke, Bola. 2015. "Marriage Age, Fertility Behaviour and Women's Empowerment in Nigeria." Obafemi Awolowo University.

UNICEF. 2014. “Ending Child Marriage: Progress and Prospects.” United Nations Chidren’s Fund.

UNICEF. 2015. “Surprising Trends in Child Marriage in Ethiopia.” United Nations Chidren’s Fund.

Wachs, Theodore. 2008. “Mechanisms Linking Parental Education and Stunting.” *The Lancet*, 371(9609).

Wahhaj, Zaki. 2015. “A Theory of Child Marriage.” School of Economics, University of Kent Studies in Economics 1520.

Weldearegawi, Berhe, Yohannes Melaku, Semaw Abera, Yemane Ashebir, Fisaha Haile, Afework Mulugeta, Frehiwot Eshetu, and Mark Spigt. 2015. “Infant Mortality and Causes of Infant Deaths in Rural Ethiopia: a Population-Based Cohorts of 3684 Births.” *BMEC Public Health*, 15(770).

Wodon, Quentin, Minh Nguyen, and Clarence Tsimpo. 2016. “Child Marriage, Education, and Agency.” *Feminist Economics*, 22(1): 54–79.