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Child Marriage and Infant Mortality: Evidence from Ethiopia

Jorge García Hombrados
University of Sussex
j.garcia-hombrados@sussex.ac.uk

Abstract: This study uses age discontinuities in the degree of exposure to a law that raised the legal age of marriage for women from 15 to 18 years in some regions of Ethiopia to provide the first evidence on (a) the beneficial effects on child marriage and infant mortality of laws that ban underage marriage and on (b) the causal effect of delaying women's age at cohabitation on infant mortality using a fuzzy regression discontinuity design. The results show that although the introduction of the law did not end child marriage among Ethiopian women, it had large effects on the incidence of child marriage and on the probability of infant mortality of the first born child. Besides, the results suggest that a one-year delay in women's age at cohabitation during teenage years decreases the incidence of infant mortality of the first born by 3.8 percentage points. The size of this effect is comparable to the joint impact on child mortality of measles, BCG, DPT, Polio and Maternal Tetanus vaccinations. This effect on infant mortality seems to be closely linked to the impact of delaying cohabitation on the age of women at first birth.

JEL classification: O1, K0.

Key words: child marriage, infant mortality, family economics

1 Motivation

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 early 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 programmes 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, labour force participation, mortality and participation in household decisions (Jensen and Thornton, 2003; Nguyen and Wodon, 2015; Wodon et al., 2015; Elborgh-Woytek et al., 2013; Wachs, 2008; UNICEF, 2014; Elborgh-Woytek et al., 2007; Solanke, 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 these 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; and document the intergenerational effects on the education, health and cultural preferences of their children (Field and Ambrus, 2008; Chari et al., 2017; Sekhri and Debnath, 2014; Hicks and Hicks, 2015; Asadullah et al., 2016; Asadullah and Wahhaj, 2016). However, although this instrumental variable approach dominates the literature on the causal effects of child marriage, it is not without its critiques. The potential limitations of this approach are discussed in section 2.

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

[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.

(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. Remarkably, exposure to a legal age of marriage at 18 relative to the possibility of getting legally married at the age of 15 decreased by 20 percentage points the incidence of child marriage among the women in the sample and increased by approximately 2 years the mean age at cohabitation for these women. However, the 2011 DHS survey shows that the share of Ethiopian women cohabiting before the age of 18 remains above 20% even among those women effectively exposed to a legal age of marriage at 18. One possible cause for this is the lack of capacity of Ethiopian institutions to enforce the minimum age of marriage established in the law. The fact that the RFC banned underage marriage but not underage cohabitation could also help to explain why, even in the presence of a strong social stigma associated with unmarried cohabitation (Jones et al., 2016), setting the minimum age of marriage at 18 years has not eradicated underage cohabitation among Ethiopian women.

This study investigates the impact of the approval of the RFC in some regions of Ethiopia to pursue a twofold objective. First, I estimate the effect for women of exposure to a legal age of marriage at 18 on the probability of infant mortality of the first born child. Second, I assess 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 banked on three main contributions to the literature. First, this is to the best of my knowledge the first study that provides evidence on the socioeconomic consequences of increasing the minimum age of marriage for women. Second, the study focuses on the link between child marriage and infant mortality, a key development outcome that has been ignored by previous causal studies assessing the consequences of child marriage. Third, unlike previous studies relying on the use of age at menarche as an instrumental variable for child marriage, I address endogeneity between women's age at cohabitation and socioeconomic outcomes through exploiting age discontinuities in the effective legal age of marriage faced by Ethiopian women. 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 equal or older than 15 years at the same time had the opportunity to get legally married before they turned 18 years old. This methodological approach can be also used to expand the analysis on the effects of child marriage to other outcomes of interest and to other settings where similar laws have been approved.

The results show that the probability of infant mortality of the first born is 7.9 percentage points lower among women exposed to a legal age of marriage at 18 than among women that had the opportunity to get married at the age of 15 before the approval of the law. They also suggest that even if they do not completely end child marriage, laws setting the minimum age of marriage for women at 18 years can be effective policies to reduce infant mortality. 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, bandwidths and windows for the forcing variable; and to different placebo tests ruling out the possibility that the impact on infant mortality is driven by systematic differences between women born in different months of the year, other interventions at the national level, other legal dispositions included in the RFC or over time decreases in infant mortality.

The analysis of mechanisms indicates that the impact of raising the legal age of marriage on the infant mortality of the first born is mainly channelled through the positive effect of delaying cohabitation on the age of women at first birth. On the other hand, the analysis suggests that the effect of early cohabitation on infant mortality is not driven by any effect of the former on women's marriage market outcomes, participation in household decisions, education or labour force participation.

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 thought 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. 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 behaviour 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) develops a theoretical model aiming to explain child marriage in developing countries. The latter paper sets 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; Elborgh-Woytek et al., 2007; Solanke, 2015), lower labour force participation and educational attainment (Elborgh-Woytek et al., 2013; Field and Ambrus, 2008; Jensen and Thornton, 2003; Nguyen and Wodon, 2015; Wodon et al., 2015), and worse maternal health (Field and Ambrus, 2008; 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). Besides, the authors of the review find that child marriage seems to be also associated with negative outcomes for the children of these women, including lower educational attainment and worse health (UNICEF, 2014; Wachs, 2008). Although not included in the review, Raj et al. (2010) and Adhikari (2003) assess empirically the statistical association between women's age at marriage, age at first birth and mortality of young children. While the former paper provides evidence of a large and positive statistical association between infant mortality and marriage before the age of 18 in India, the latter finds that being mother before the age of 20 is strongly associated with larger levels of neonatal mortality in Nepal.

With the exception of Field and Ambrus (2008), the studies reviewed in Parsons et al. (2015) examine 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 using age at menarche as an instrumental variable for age at marriage. These studies exploit the quasi-random variation generated by a delay in the age at menarche as a source of exogenous variation for age at marriage for women. They argue that the arrival of puberty determines the entrance in the marriage market and they prove empirically that a delay in the age at menarche increases significantly 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. Furthermore, 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 labour market outcomes, beliefs and marriage market outcomes of women in

Kenya; suggesting that the statistical association between child marriage and these outcomes found in studies conducting correlation analysis could be driven by reverse causality or omitted variables bias. To the best of my knowledge, this instrumental variable approach has not been used to investigate the link between child marriage and infant mortality.

The correct identification of the effect of early marriage in these studies relies on the assumption that conditional on height, socioeconomic background and location, age at menarche does not affect women's socioeconomic outcomes other than through delaying age at cohabitation. Furthermore, to be a valid instrumental variable, age at menarche should not be driven by unobservable factors that affect the outcome of interest. Although Field and Ambrus (2008) rule out different mechanisms through which age at menarche could affect the outcomes of interest, medical studies suggest that some childhood experiences such as stressful family environment or sexual abuse that may have long-term effects on socioeconomic outcomes seem to bring forward the age at menarche (Karapanou and Papadimitriou, 2010; Barrios et al., 2015). Under the latter hypothesis, the instrumental variable used might not satisfy the exclusion restriction and the estimates in these studies should be interpreted with caution.

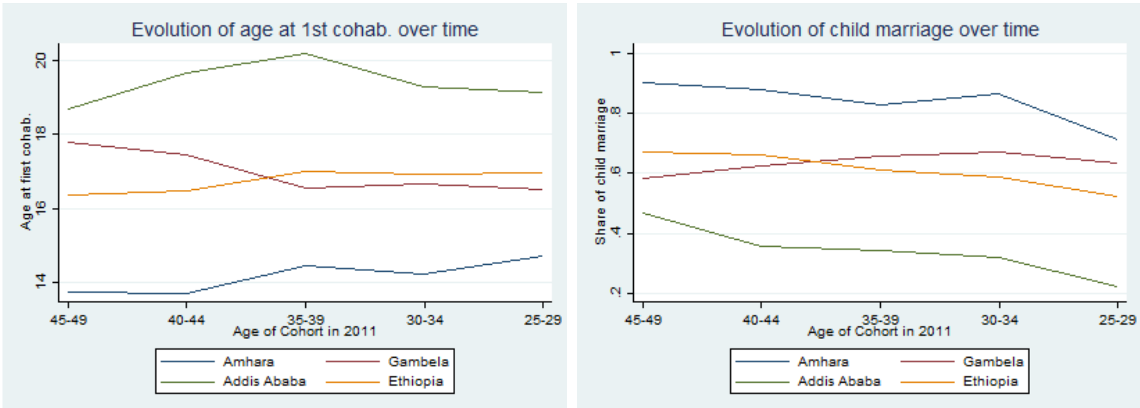
3 Child Marriage in Ethiopia and the Revised Family Code

The 2011 Demographic and Health Survey reveals that 41% of women aged 20-24 in Ethiopia first cohabited with a partner before the age of 18. The total number of women married as children in the country is estimated at 1,974,000, making Ethiopia the country with the 5th largest number of child marriages in the world[5].

Ethiopia is a federal state formed by 11 regions, which are very diverse in terms of ethnicity and religion. Although the practice of child marriage is conducted in all the country, its prevalence ranges substantially across regions. Figure 1 displays the evolution of the prevalence of early cohabitation and of the mean age at first cohabitation in Ethiopia. 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 in Ethiopia. For women aged 25-29 years old, the figure shows that the incidence of child marriage ranges between 70% in Amhara and 20% in Addis Ababa. 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, it has slightly increased in Gambela.

[5] Girls not Brides website.http://www.girlsnotbrides.org/where-does-it-happen/

Figure 1: Child marriage over time in Ethiopia (DHS 2011)



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 programmes to prevent early marriage and delay age at cohabitation, the Federal Government of 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. 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. Finally, the RFC recognized the validity of marriages celebrated before the approval of the RFC that complied with the 1960 Family Code.

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 approved after the fall of Mengistu's government, 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 show the density of the distribution of the age at first cohabitation for those 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[6]. 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 14 in appendix C compares in the same graph the distribution of the age at cohabitation for the cohort of women aged 13-14 and the cohort aged 15-16 when the RFC was introduced. 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.

Figure 2: Age at 1st cohabitation (cohorts 15-17 at RFC): Discontinuities at 15 and 18

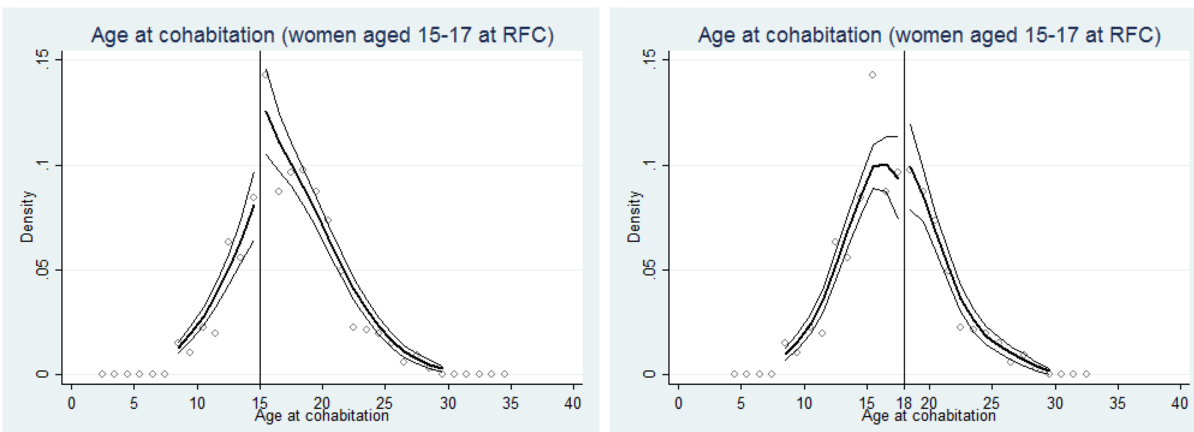
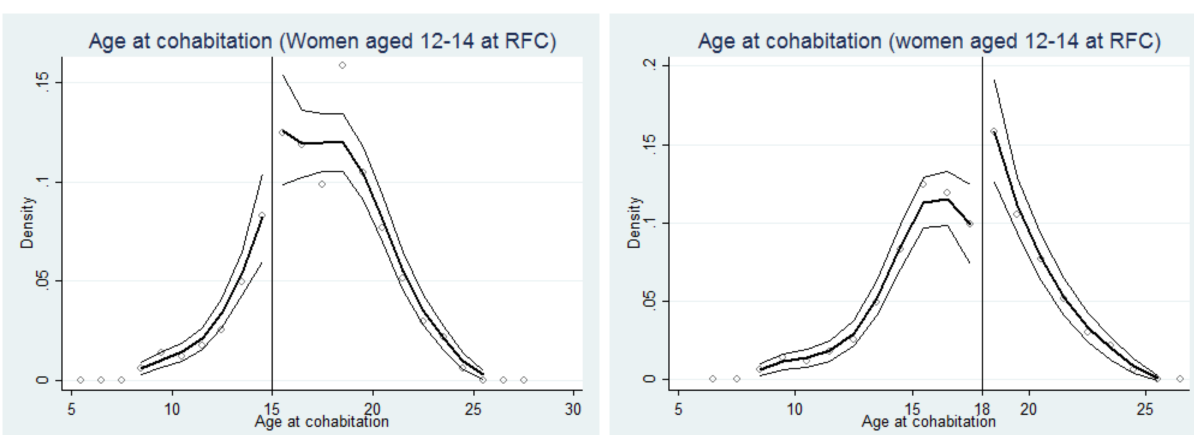


Figure 3: Age at 1st cohabitation (cohorts 12-14 at RFC): Discontinuities at 15 and 18



[6] These two figures only include the sample of Ethiopian women used in the main analysis of the study. The selection of women and regions that are used in this analysis is discussed in section 5.

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 cohabitation 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. 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 unmarried cohabitation is stigmatized in Ethiopia (Jones et al., 2016), we 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 6.1, 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 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 exploits the sharp reduction in the mean age at cohabitation with a partner for those women older than 15 years at the time of the approval of the RFC. At the extreme, the estimates rely on the change in the mean age at cohabitation for those women that were 15 years and 1 month 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 14 years and 11 months at the same time, and could not get legally married until the age 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 child marriage for those women younger than 18 years and not yet married when the RFC was approved but did not eradicate it among them[7]. In other words, exposure to a legal age of marriage at 18 does not fully determine in the sample whether a woman 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. I dropped from the sample used in the analysis those women that turned 15 the month in which the RFC was approved. The reason for this is that the DHS survey used in the analysis only collected information on the month and year of birth, making it impossible to determine for those women that turned 15 the same month, whether they did before or after the approval of the RFC. In this context, the quality of the birth data collected plays a crucial role enabling the use of the age at RFC as the forcing variable. At this point, it is important to mention that DHS surveys are widely used by researchers in the field of fertility and the high quality and accuracy of the information collected is showed in Pullum (2008)[8]. Third, the RFC was not applied simultaneously in every Ethiopian region. Although this variation is not exploited for identification purposes, the different timing in the application of the rise in the legal age for marriage across regions provides more variation in the current age of women that were approximately 15 year old when the RFC was approved in their region, making the results 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:

$$\mathit{Age\ at\ Cohab.}_i = \alpha_0 + \alpha_1(\mathit{Age\ at\ RFC{<}15}_i) + \alpha_3 F(\mathit{Age\ at\ RFC}_i) + \alpha_4 X_i + \mu_i \tag{4.1}$$

$$InfantMortality_i = \delta_0 + \delta_1(\mathit{Age\ at\ RFC{<}15}_i) + \delta_2 F(\mathit{Age\ at\ RFC}_i) + \delta_3 X_i + \epsilon_i \tag{4.2}$$

$$InfantMortality_i = \beta_0 + \beta_1(\widehat{\mathit{Age\ at\ Cohab.}_i}) + \beta_2 F(\mathit{Age\ at\ RFC}_i) + \beta_3 X_i + u_i \tag{4.3}$$

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

[8]Furthermore, unlike in other surveys conducted in sub-Saharan Africa, the distribution of the birth data collected for the DHS in Ethiopia does not show spikes in the age of surveyed individuals ended in 0 or 5.

where $InfantMortality_i$ 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 including the region of residence, the age of the woman at the time of survey, ethnic and religion affiliation, gender of the first born and a dummy variable indicating whether 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, $Age\ \widehat{at\ Cohab.}_i$ is the predicted age at cohabitation for woman i estimated from equation 4.1.

Equation 4.1 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 4.2 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 4.3 is the second stage equation. It regresses infant mortality against the predicted age at cohabitation estimated in equation 4.1. 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 4.1, 4.2 and 4.3 is conducted using non-parametric local polynomial regressions based on triangular kernel functions. The study follows the procedure described in Calonico et al. (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. As a robustness check, I also estimate equations 4.1, 4.2 and 4.3 using (a) two alternative bandwidths equal to 0.75 and 1.5 times the optimal bandwidth, (b) conventional and bias-corrected non-parametric RD estimation procedures with conventional variance estimators, and (c) parametric methods with spline polynomials of order 1 to 4 for the forcing variable and windows of 2, 3, 4 and 5 years at both sides 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, the estimated parameter β_1 in equation 4.3 would not be efficient, potentially leading to a problem of weak instrument (Bound et al., 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 4.2 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 6.1.

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 6.4 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 6.4 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 and they have been used in numerous studies on health and fertility in developing countries. 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 the questionnaires are mostly the same in DHS across the world aiming to produce comparable statistics, the exact questionnaire and the size and characteristics of the sample vary in every DHS. The 2011 Ethiopian DHS collected household, 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 labour market outcomes and does not record the age at marriage. The questions on maternal health only target the last child of the women and data on children's health status are only recorded for those children aged 0-5 at the time of the survey.

In total, the female survey 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 analysis for two reasons. First, four of these regions did not implement the rise in the legal age for marriage before 2008. Therefore, 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. 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[9]. 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.

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, that ever cohabited with a partner[10] and gave birth to their first child more than one year ago[11]. 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 analysis. Given that the estimates yielded by 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.

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 labour 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. If we focus on the sample of individuals aged 14-15 when the RFC was introduced in their region, the incidence of child marriage falls from 70% among those women aged 15 at the time of the RFC 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[12].

[9] This pattern can be observed in figure 17 in appendix C.

[10] The percentage of women that gave birth without ever cohabiting with a partner only represents the 1.1% of all the women that gave birth to at least one child in these regions.

[11] 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 ago to avoid censoring in the dependent variable.

[12] 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

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
Anaemia (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
Rural (0/1)	308	0.58	0.49	0	1	5,078	0.71	0.13
Muslim (0/1)	308	0.31	0.46	0	1	5,078	0.21	-0.10
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
Interval cohab. first child (months)	308	27.59	26.59	0	158	5,078	31.57	3.98
First born characteristics								
Years since born	308	4.96	2.75	1	12	5,078	13.94	8.98
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								
N antenatal visits	107	3.39	3.49	0	12	520	3.41	0.02
Delivery at home (0/1)	164	0.48	0.50	0	1	862	0.48	0.00
N vaccines (1-9)	144	6.44	2.76	0	9	781	6.44	0.00
Postnatal check (0/1)	66	0.06	0.24	0	1	312	0.06	0.00
Months breastfeed	63	23.14	12.14	0	51	240	20.69	-2.45

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 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 evolution of the prevalence of child marriage across age cohorts in Ethiopia could be better observed in figure 12 in appendix C, which is constructed using the women aged 18-49 in the DHS data regardless of whether they ever cohabited or gave birth.

In the full sample, the mean age of women at first birth is 18.9 years and the average interval between cohabitation and first birth is 31.6 months. The mean age at birth and the mean interval between cohabitation and first birth are smaller (18.3 and 27.6) 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 the women aged 15 when the RFC was introduced to 5% among the women aged 14 at the same time and therefore, exposed to a legal age of marriage at 18 years.

Figure 4: Age at cohabitation, age at birth and infant mortality of the first born (LOWESS regressions)

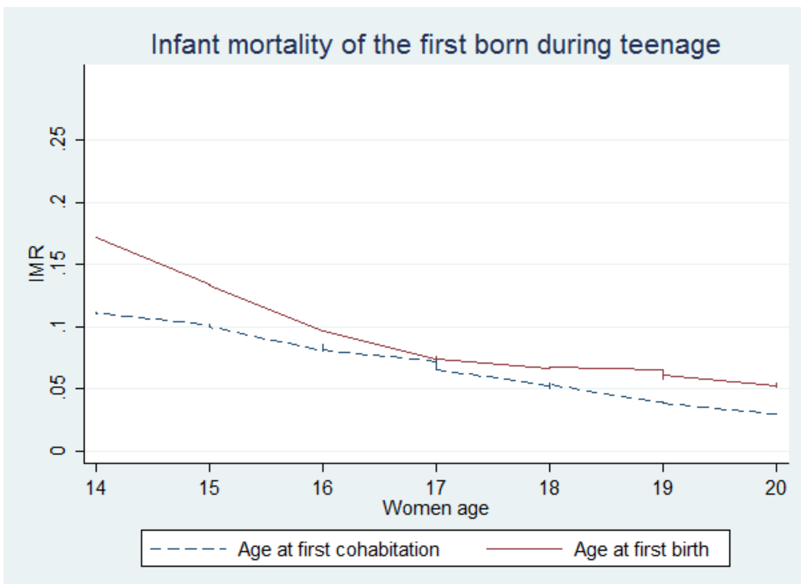


Figure 4 displays the statistical association between women's age at cohabitation, age at first birth and infant mortality of the first born during teenage years. The statistical relation between the variables is estimated using LOWESS regressions for the sample of women used in the analysis. Although 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 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 17 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 for the estimation of (a) the effect of exposure to a legal age of marriage at 18 on the probability of infant mortality of the first born and of (b) the effect of the age at cohabitation on infant mortality of the first born, 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. The coefficients of the variable when the alternative bandwidths and non-parametric estimation procedures are used are also large and positive. Overall, the coefficients measuring the effect of exposure to a legal age of marriage at 18 across the different non-parametric estimations are 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 and 7 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 (0.8-1.8 years) and statistically significant when equation 4.1 is estimated using parametric methods with several windows and spline 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 5 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 behaviour 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 2-3 years to marry as soon as they turned 15. This fact 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.

Figure 5: Main analysis: Age at first cohabitation at the cut-off

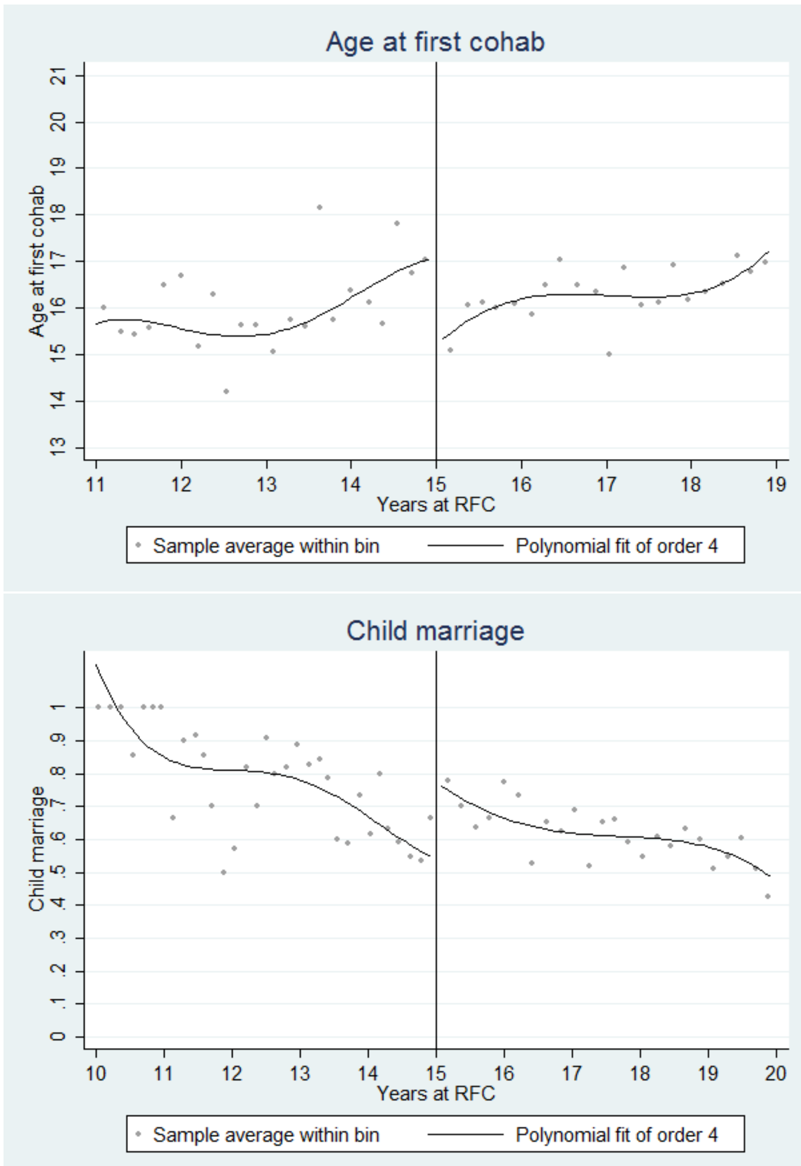


Table 2: Non-parametric methods: RFC, age at first cohabitation and infant mortality.

	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*
	(0.000)	(0.034)	(0.000)	(0.021)	(0.000)	(0.051)
Age at 1st cohab.		-0.041**		-0.038**		-0.038*
		(0.031)		(0.044)		(0.095)
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***
	(0.000)	(0.007)	(0.000)	(0.002)	(0.000)	(0.008)
Age at 1st cohab.		-0.053***		-0.052***		-0.052**
		(0.004)		(0.005)		(0.016)
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
	(0.000)	(0.074)	(0.000)	(0.059)	(0.000)	(0.122)
Age at 1st cohab.		-0.042*		-0.039*		-0.039
		(0.068)		(0.086)		(0.172)
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 (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 (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. P-values 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= 1,164)		5 years window (N= 1,432)	
	(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
Spline Pol. order 1								
Age<15 at RFC	1.707***	-0.076**	1.057***	-0.076**	0.919***	-0.053**	0.849***	-0.047**
	(0.000)	(0.022)	(0.001)	(0.022)	(0.001)	(0.041)	(0.002)	(0.040)
Age at 1st cohab		-0.045**		-0.045**		-0.058**		-0.055**
		(0.012)		(0.012)		(0.041)		(0.041)
Spline Pol. order 2								
Age<15 at RFC	1.552***	-0.080*	1.565***	-0.070**	1.335***	-0.068**	1.159***	-0.058*
	(0.000)	(0.054)	(0.000)	(0.035)	(0.000)	(0.029)	(0.000)	(0.051)
Age at 1st cohab		-0.052**		-0.045**		-0.051**		-0.050**
		(0.040)		(0.025)		(0.021)		(0.039)
Spline Pol. order 3								
Age<15 at RFC	1.552***	-0.080*	1.552***	-0.097**	1.321***	-0.069**	1.106***	-0.058*
	(0.000)	(0.054)	(0.000)	(0.018)	(0.000)	(0.036)	(0.001)	(0.064)
Age at 1st cohab		-0.052**		-0.057***		-0.052**		-0.052*
		(0.040)		(0.010)		(0.032)		(0.055)
Spline Pol. order 4								
Age<15 at RFC	1.552***	-0.080*	1.552***	-0.097**	1.782***	-0.100*	1.360***	-0.082**
	(0.000)	(0.054)	(0.000)	(0.018)	(0.003)	(0.061)	(0.002)	(0.032)
Age at 1st cohab		-0.052**		-0.057***		-0.056**		-0.060**
		(0.040)		(0.010)		(0.043)		(0.030)

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 spline polynomials of order 1 to 4 for the forcing variable. 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 and 7; and the effect of the law on the prevalence of infant mortality (reduced form) in columns 2, 4, 6 and 8. 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 (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 a spline polynomial function of order 1 to 4 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. P-values are in parentheses.***p<0.01;**p<0.05;*p<0.1.

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 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 Ethiopia is presented in figure 12 in appendix C. 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 changed dramatically at the cut off.

6.2 The Effect of the RFC on Infant Mortality

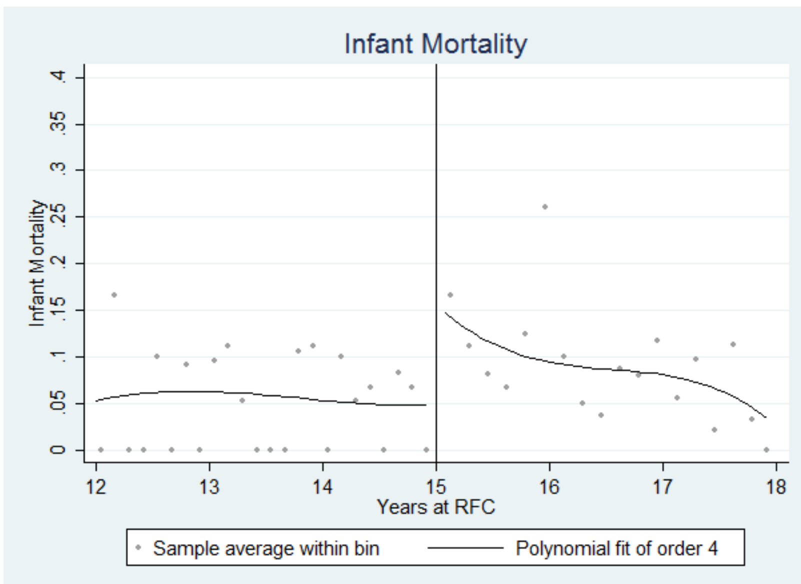
The next step is determining whether the 20 percentage points drop in the incidence of child marriage and the 2 years increase in the mean age at cohabitation at the cut-off for women exposed to a legal age of marriage at 18 affected the probability of infant mortality of the first born. This effect is yielded by the parameter δ_1 in equation 4.2. 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 preferred estimate is reported in column 6 and yields an effect of 7.9 percentage points, statistically significant at the 10%. Columns 2, 4, 6 and 8 of table 3 provide the estimates for the parameter δ_1 using parametric techniques with different windows and spline polynomials of order 1 to 4 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 with different bandwidths, windows and polynomial functions for the forcing variable. 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, women just above the cut-off had 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. The main implication for this, discussed in the next to last paragraph in section 6.1, is that 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 3.5-4 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 more than one year before the survey.

Figure 6: Main analysis: Infant mortality rate at the cut-off



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 preferred 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 and 8 of table 3 confirm that the findings of the non-parametric analysis are robust to the use of parametric techniques with different windows and spline polynomials of order 1 to 4 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 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.

6.4 Robustness Checks

Using parametric and non-parametric methods, the previous section shows that 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. These results are robust to the use of non-parametric methods with different bandwidths and estimation procedures; and parametric methods with different windows and spline polynomials for the forcing variable. In this section, I discuss and explore alternative explanations for the results.

Firstly, 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 4.1, 4.2 and 4.3 using the bias-corrected RD estimates with robust variance estimator and an optimal bandwidth calculated following Calonico et al. (2014) and whether the first born is male, whether the mother is Muslim or from Oromo ethnic group[13] as outcome variables. The results of these estimations are provided in columns 1-6 of table 4. 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.

Secondly, 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 4.1, 4.2 and 4.3 setting a placebo cut-off for women older than 19, rather than 15, at the time of the approval of the RFC[14]. This exercise is equivalent to placing the cut-off as if the law was introduced exactly

[13]Oromo is the most prevalent ethnic group in Ethiopia.

[14]The set of the placebo cut-off at 19 years is driven by the convenience of setting the false cut-off at a value

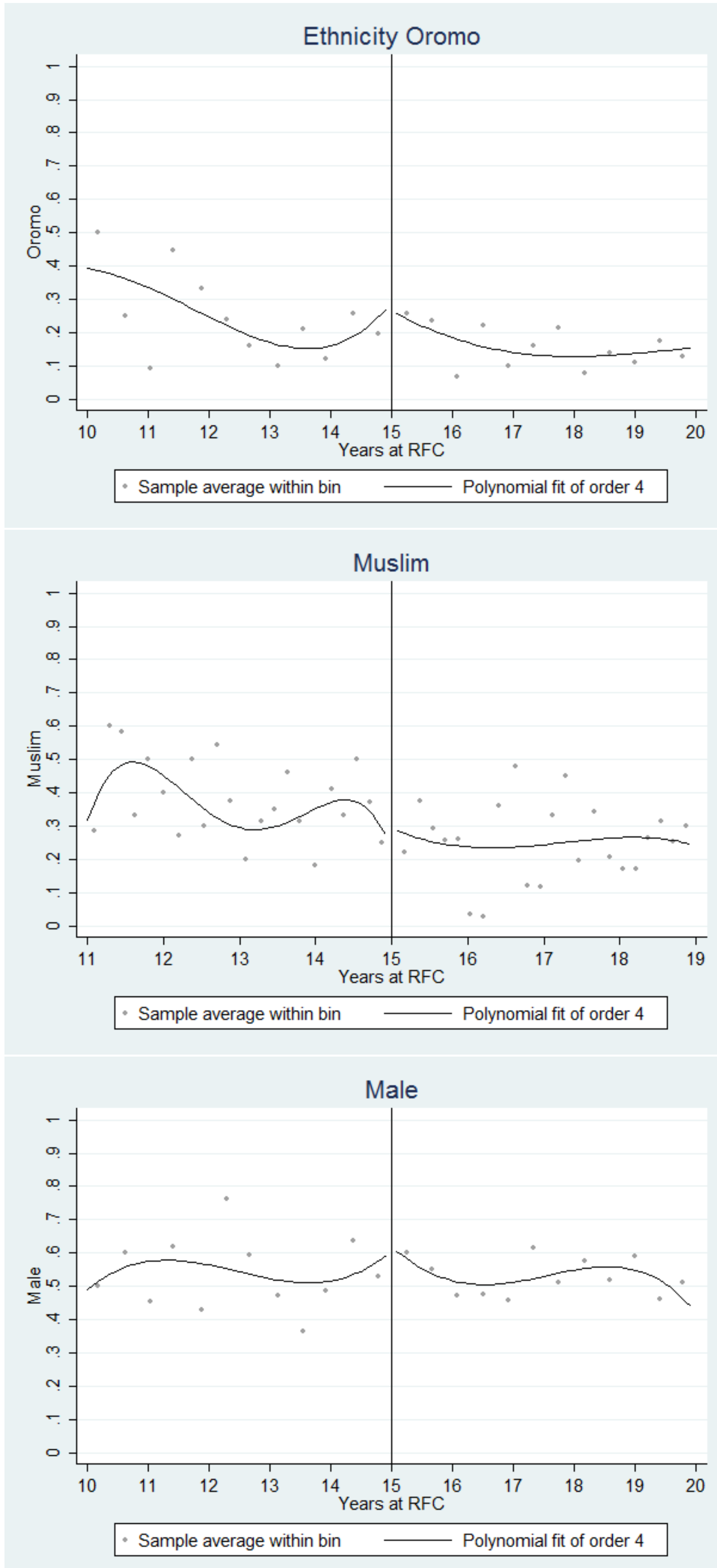
Table 4: Robustness checks Infant Mortality: Placebo analyses.

	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 Muslim
Age<15 at RFC	1.984***	-0.015	2.068***	-0.039	1.999***	0.037
	(0.000)	(0.762)	(0.000)	(0.727)	(0.000)	(0.598)
Age at 1st cohab.		-0.006		-0.018		0.017
		(0.834)		(0.755)		(0.648)
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**
	(0.666)	(0.961)	(0.213)	(0.789)	(0.000)	(0.023)
Age at 1st cohab.		-0.007		0.015		-0.066*
		(0.963)		(0.810)		(0.053)
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 by the time of the survey: Affar, 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. P-values are in parentheses.***p<0.01;**p<0.05;*p<0.1.

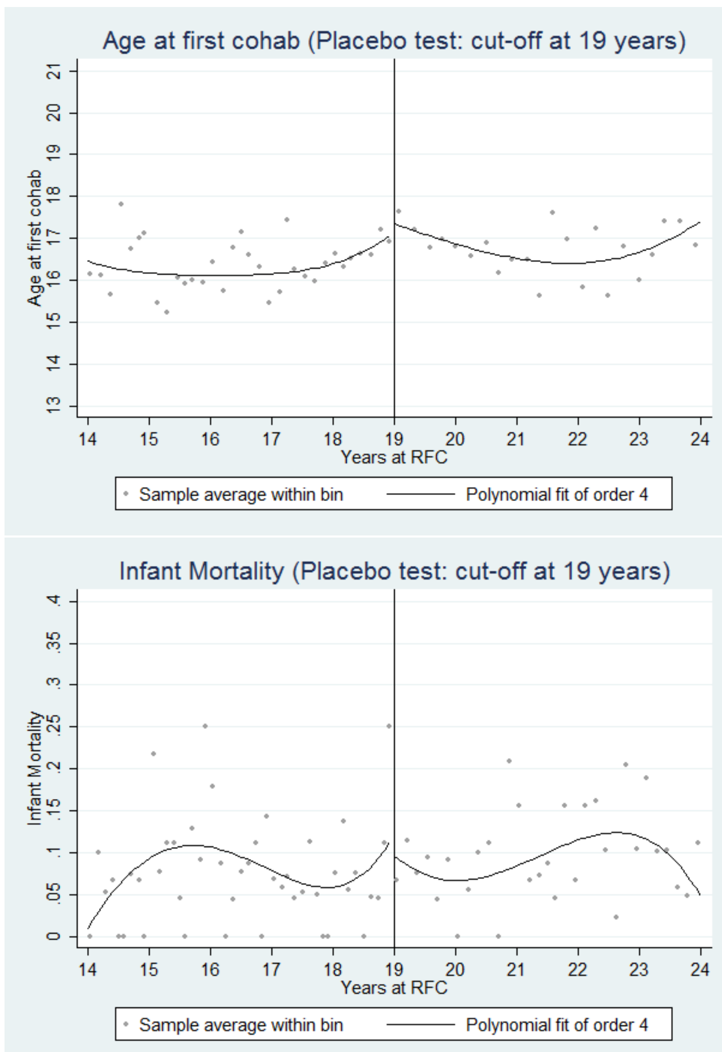
of the forcing variable that left out of the estimation the observations around the real cut-off. Nonetheless, I have also conducted the analysis setting the placebo cut-off for women aged 16, 17 and 18 at the time of the approval of the RFC. Consistently, the results of these placebo analyses, not reported in the study, show no discontinuities in infant mortality and age at cohabitation at the false cut-off.

Figure 7: Placebo variables: ethnicity, religion and gender



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 4 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.

Figure 8: Placebo test: Cut-off at 19 years



Thirdly, 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 the possibility that the main results of the study are driven by time trends in infant mortality or by a national level intervention affecting differently the cohorts of women at both sides of the cut-off as follows. First, I re-estimate the results using the Ethiopian regions of Affar, Harari and Gumuz, setting falsely the approval date of the RFC in these placebo regions in July 2000[15]. I restrict the analysis to these three regions because none of them passed the RFC before the implementation of the 2011 DHS survey (Hallward-Driemeier and Gajigo, 2015). The results of this placebo test are reported in columns 9 and 10 of table 4 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. Second, I re-estimate equations 4.1, 4.2 and 4.3 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[16]. The estimates provided in columns 10 and 11 show that the direction and significance of the parameters estimated in sections 6.1, 6.2 and 6.3 do not vary when the year of birth of the first child is included in the regression. The results of these two analyses suggest 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 or by a national level policy affecting differently women at both sides of the cut-off.

Fourthly, 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 3 and 4 of table 7 in appendix D.

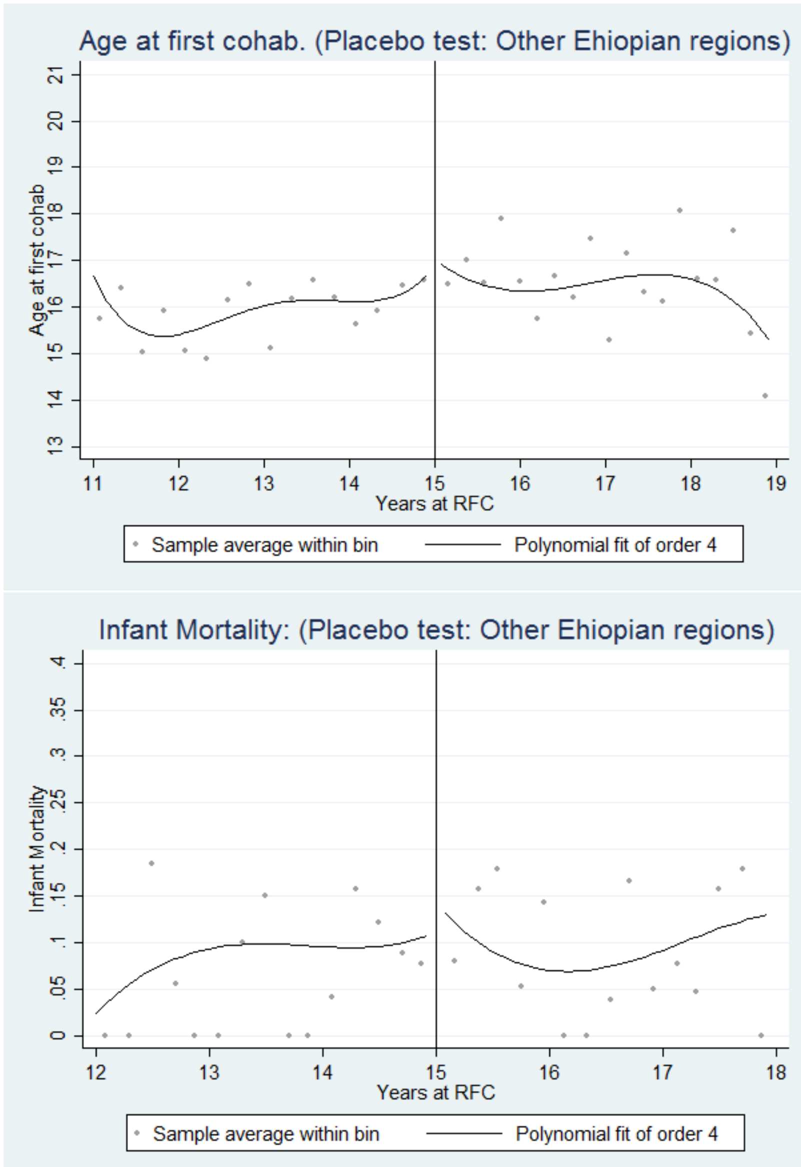
One threat to the interpretation of the estimates could be the possible existence of selective migration. In other words, those women below the age of 15 when the legal age of marriage was raised in their region that were particularly interested in early cohabitation could have migrated to regions that 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 (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 evident in figure 11 in appendix B suggests that the migration of women slightly younger

[15] 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.

[16] 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 would lead to a *bad* control problem (**?**).

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 neonatal mortality as a dependent variable rather than infant mortality. For this, I re-estimate equations 4.1, 4.2 and 4.3 focusing on mortality of the first born within the first month of life, rather than within the first year. The estimates obtained for neonatal mortality of the first born have the same sign and statistical significance although the magnitudes are even larger than those obtained for infant mortality of the first born[17]. Therefore, the evidence confirms that the results of the

[17]These results are not reported in the tables provided in the study. The estimates reveal that exposure to a legal age of marriage at 18 years relative to the possibility of getting married at 15 reduces the incidence of

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 first month of life of the newborn.

One 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 and 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 labour force participation, divorce rates and participation in household decisions 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 relative visits, 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, labour force participation and divorce rates as dependent variables, I estimate equation 4.2. The results reported in table 5 suggest that, overall, the new dispositions included in the RFC aiming to change the balance of power within the household did not affect differently labour force participation, divorce rates and participation in household decisions of women at both sides of the cut-off. In consequence, the evidence suggests 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.

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. Figure 11 in appendix B shows that the density of the age of women changes smoothly 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.

neonatal mortality of the first born at the cut-off by 9.0 percentage points. The effect of a one-year delay in women's age at cohabitation during teenage years on the probability of neonatal mortality of the first born is estimated at 4.5 percentage points.

Table 5: 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)
Age<15 at RFC	-0.200***	2.094***	0.098	-0.054	-0.318**	0.132	-0.059	-0.063	0.018
	(0.000)	(0.000)	(0.211)	(0.395)	(0.014)	(0.247)	(0.643)	(0.275)	(0.686)
N	5078	5078	5077	4181	4175	4164	4170	4142	5078
N effect. obs.	958	812	1176	1035	801	771	898	893	958
Bandwidth	39.2	33.1	48.1	49.3	38.0	37.4	42.2	42.9	39.5

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). 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. P-values are in parentheses.***p<0.01;**p<0.05;*p<0.1.

7 Mechanisms

One potential mechanism driving the effect of early cohabitation on infant mortality of the first born could be the age of women 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 behaviours (Olausson et al., 2007, 1999).

To examine this mechanism, I re-estimate equations 4.1, 4.2 and 4.3 using fertility outcomes, antenatal and postnatal behaviours as dependent variables in the reduced form and second stage equations. The results on fertility outcomes are displayed in columns 1 to 4 of table 6 and confirm that although delaying cohabitation causally reduces the number of months between cohabitation and first birth, it also increases significantly the age of women at first birth. In line with this *age at birth* mechanism, the results reported in columns 7 and 8 suggest that the effect of age at cohabitation on infant mortality vanishes for infants given birth after the first born.

On the other hand, the effect of the age at cohabitation on the adoption of antenatal and postnatal health practices is mixed. The results reported in columns 17 to 26 indicate that delaying cohabitation increases significantly the probability of conducting a postnatal check. The coefficients measuring the effect of women's age at cohabitation on the probability of giving birth at home and on the duration of breastfeeding have the expected sign (negative the former and positive the latter) although the magnitudes are small and statistically insignificant when the optimal bandwidths are used. However, the coefficients for vaccinations and antenatal visits in the second stage equations have an unexpected negative sign, although statistically insignificant at conventional confidence levels. Nonetheless, it is important to remark that the results on the adoption of antenatal and postnatal health behaviours 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. For some of these variables, the information is only available if the child is also the last birth of the women before the survey. 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.

Nonetheless, 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 labour market outcomes (Elborgh-Woytek et al., 2013). 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 impacts infant mortality through affecting 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, labour force participation and education attainment. The results of the estimates for equations 4.1, 4.2 and 4.3 using these outcomes as dependent variables are reported in columns 5 to 6, 9 to 16 and 27 to 28 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 labour force participation for women, participation in household decisions, years of education 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 channelled 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, labour market outcomes or education, any potential effect of the age at cohabitation on infant mortality via antenatal and postnatal investments would probably

Table 6: Analysis of mechanisms

	Age at birth		Months: F. Born -F. Cohab.		Years school (compl. years)		Infant mortality Non-first born		Empowerment index		Years school partner		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
Age<15 at RFC	2.070***	0.996***	2.062***	-10.612*	2.064***	-0.706	2.329***	0.011	1.964***	-0.095	2.043***	0.783	2.058***	-0.086
	(0.000)	(0.007)	(0.000)	(0.067)	(0.000)	(0.145)	(0.000)	(0.846)	(0.000)	(0.192)	(0.000)	(0.356)	(0.000)	(0.705)
Age at 1st cohab.		0.482**		-5.148*		-0.346		0.005		-0.048		0.384		-0.041
		(0.013)		(0.056)		(0.264)		(0.860)		(0.219)		(0.461)		(0.754)
N		5078		5078		5078		16373		4181		4999		5078
N effect. obs.		905		935		905		737		691		864		958
Bandwidth		37.4		38.3		37.1		32.6		33.4		36.5		39.9

	Age diff. partner		Months Breastfeed.		Postnatal check		N Vaccin.		Birth at home		N antenatal visits		Work	
	(15) FS	(16) RF/SS	(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
Age<15 at RFC	1.904***	-0.848	2.682***	-0.023	2.035***	0.268**	0.633	-0.173	1.187***	-0.015	1.049*	-1.016*	2.097***	0.035
	(0.000)	(0.377)	(0.001)	(0.997)	(0.001)	(0.026)	(0.176)	(0.768)	(0.006)	(0.880)	(0.092)	(0.054)	(0.000)	(0.711)
Age at 1st cohab.		-0.440		0.014		0.133*		-0.271		-0.015		-0.981		0.016
		(0.444)		(0.996)		(0.080)		(0.766)		(0.872)		(0.157)		(0.763)
N		4171		240		312		781		862		520		5077
N effect. obs.		853		146		204		594		584		376		828
Bandwidth		40.4		36.6		43.7		56.0		49.0		50.2		34.5

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. 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 1-16 and 27-28 report the results for the whole sample used in the main analysis. The information on age difference with partner and on partner's years of education refer to the last partner. Column 17-26 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 health outcomes. Standard errors are clustered at the forcing variable. P-values are in parentheses.***p<0.01;**p<0.05;*p<0.1.

be linked to an older age at first birth and a more mature behaviour.

8 Conclusions

This study uses a regression discontinuity design to assess the impact for women of exposure to a legal age of marriage at 18 years on the incidence of child marriage, mean age at cohabitation and infant mortality of the first born. Then, the study estimates the causal effect of women's age at cohabitation during teenage years on the probability of infant mortality of the first born. The methodological design exploits age discontinuities in exposure to a law that raised the legal age of marriage for women from 15 to 18 years in some regions of Ethiopia.

The RDD estimates suggest that 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 incidence of child marriage by approximately 20 percentage points. Besides, the analysis reveals that increasing the minimum age of marriage in the law has beneficial effects on infant mortality. In the preferred estimation, exposure to a legal age of marriage at 18 relative to the possibility of getting legally married at the age of 15 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. These results are robust to the use of different estimation methods, alternative bandwidths, windows and up to 4th order spline polynomials for the forcing variable. Different placebo tests are also conducted to rule out the possibility that the effects are driven by over time reductions in infant mortality, systematic differences between women born in different months of the year, other legal dispositions included in the RFC or by national level policies affecting differently women at both sides of the cut-off. There is also no evidence of discontinuities in variables plausibly unaffected by the change in the minimum age of marriage neither of manipulation in the forcing variable at the cut-off. Finally, the results are robust to the use of neonatal mortality as dependent variable, suggesting that most of the effect of child marriage on infant mortality occurs during the first month of life.

The effects of delaying cohabitation on infant mortality are aligned 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. The effect of one-year delay in women's age at first cohabitation on the probability of infant mortality of the first born found in this study 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)[18].

However, 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 population of Ethiopian women or to non-first born children should be avoided. Indeed, the analysis does not find any effect of early cohabitation on the infant mortality of children that are not the first born.

The analysis of mechanisms suggests that the strong effect 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 behaviour raising the 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 labour force participation for women.

This study contributes to the thin literature that documents the negative effects of child marriage, providing also a credible methodological alternative to previous studies relying on instrumental variable that can be used to expand the analysis to other outcomes and settings. Besides, although the effects estimated should be interpreted as local, the study shows for the first time that, even if imperfectly enforced, laws increasing the legal age of marriage can contribute to reducing infant mortality. Finally, the study provides the first causal evidence on the beneficial effect on infant mortality of delaying women's age at cohabitation.

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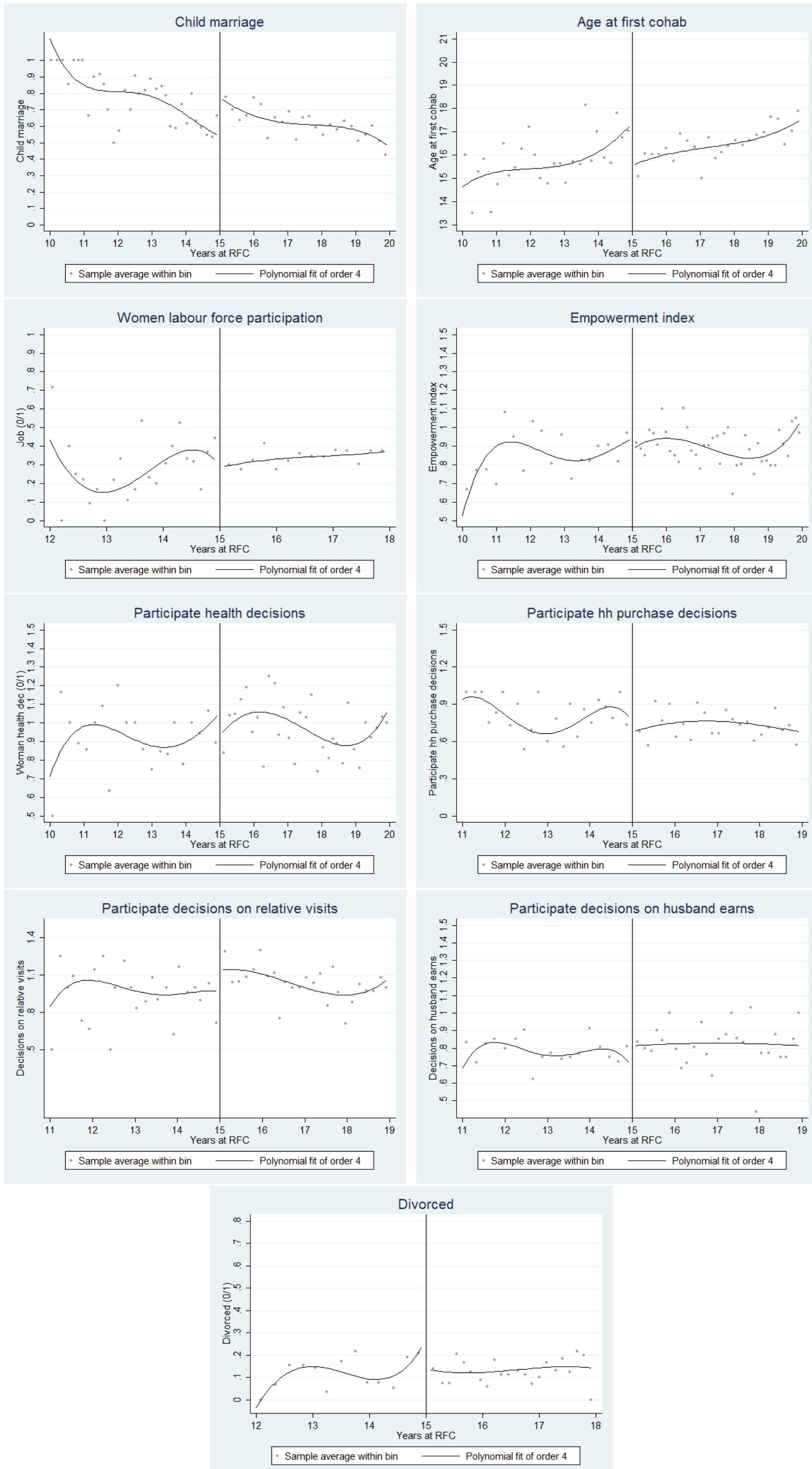
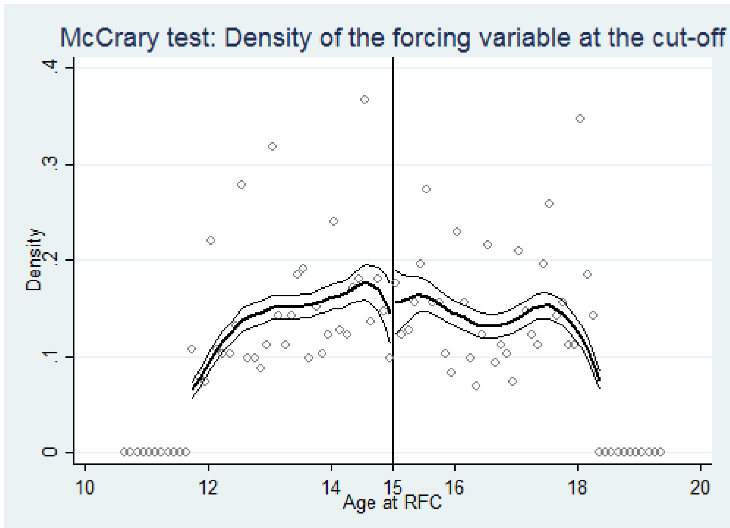
A RFC and women characteristics

Figure 10: RFC and women characteristics



B McCrary Test: Density of the Forcing Variable at the Cut-Off

Figure 11: Density of the forcing variable at the cut-off.



C Additional Graphs

Figure 12: RFC and child marriage: Includes all women (not only those that ever cohabit and gave birth)

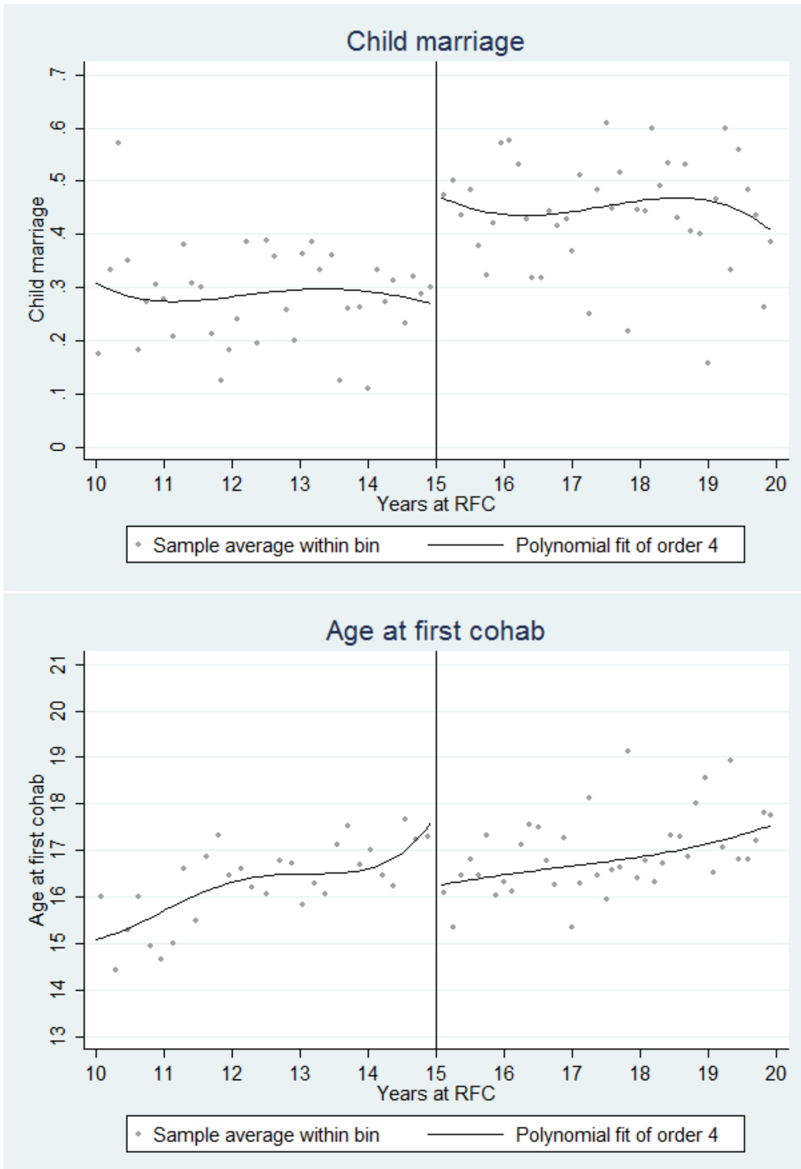


Figure 13: Age at first cohabitation and infant mortality of first born: Full sample of women (18-49)

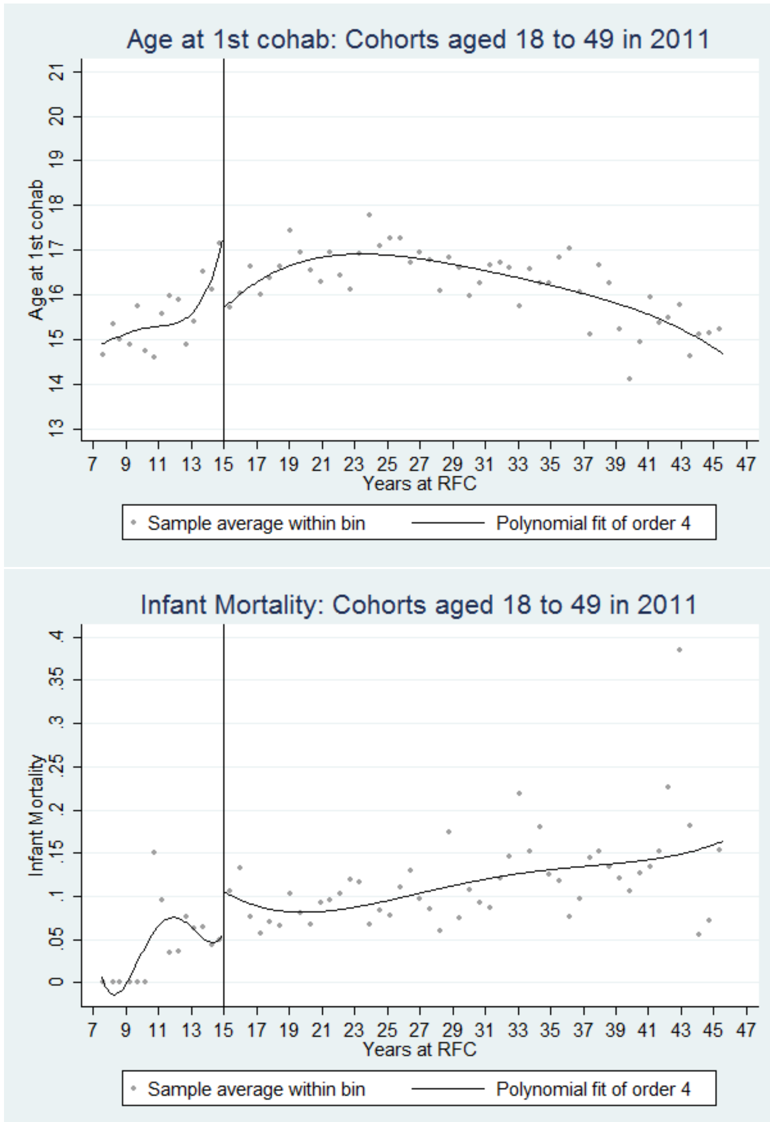


Figure 14: Age at 1st cohabitation by age cohort.

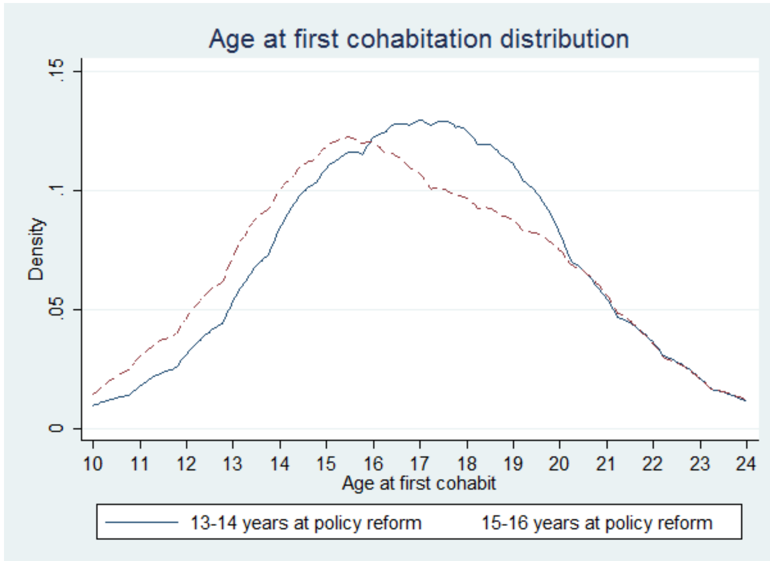


Figure 15: Mechanisms: Fertility, marriage market outcomes and other women outcomes

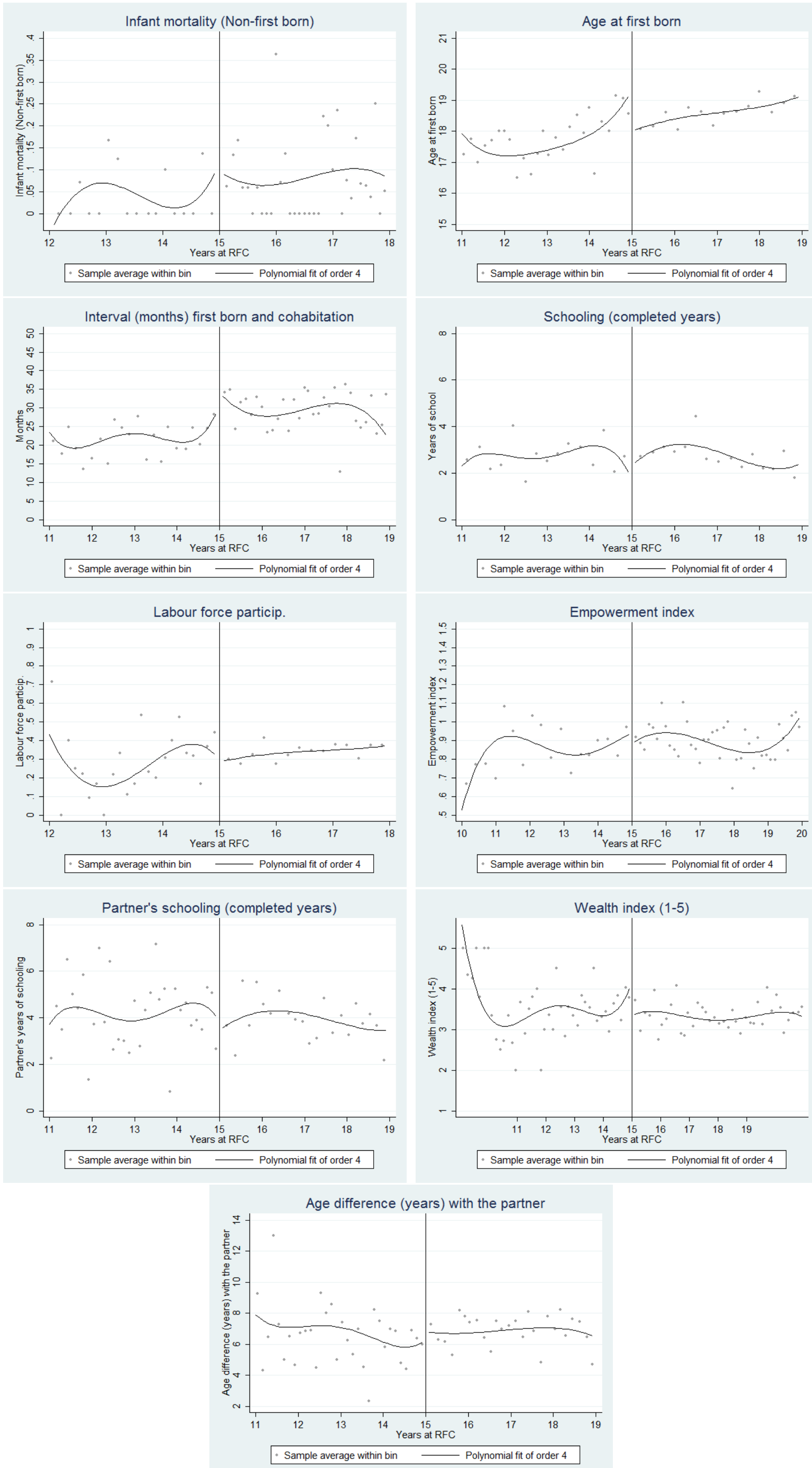


Figure 16: Mechanisms: Maternity and child health for first born

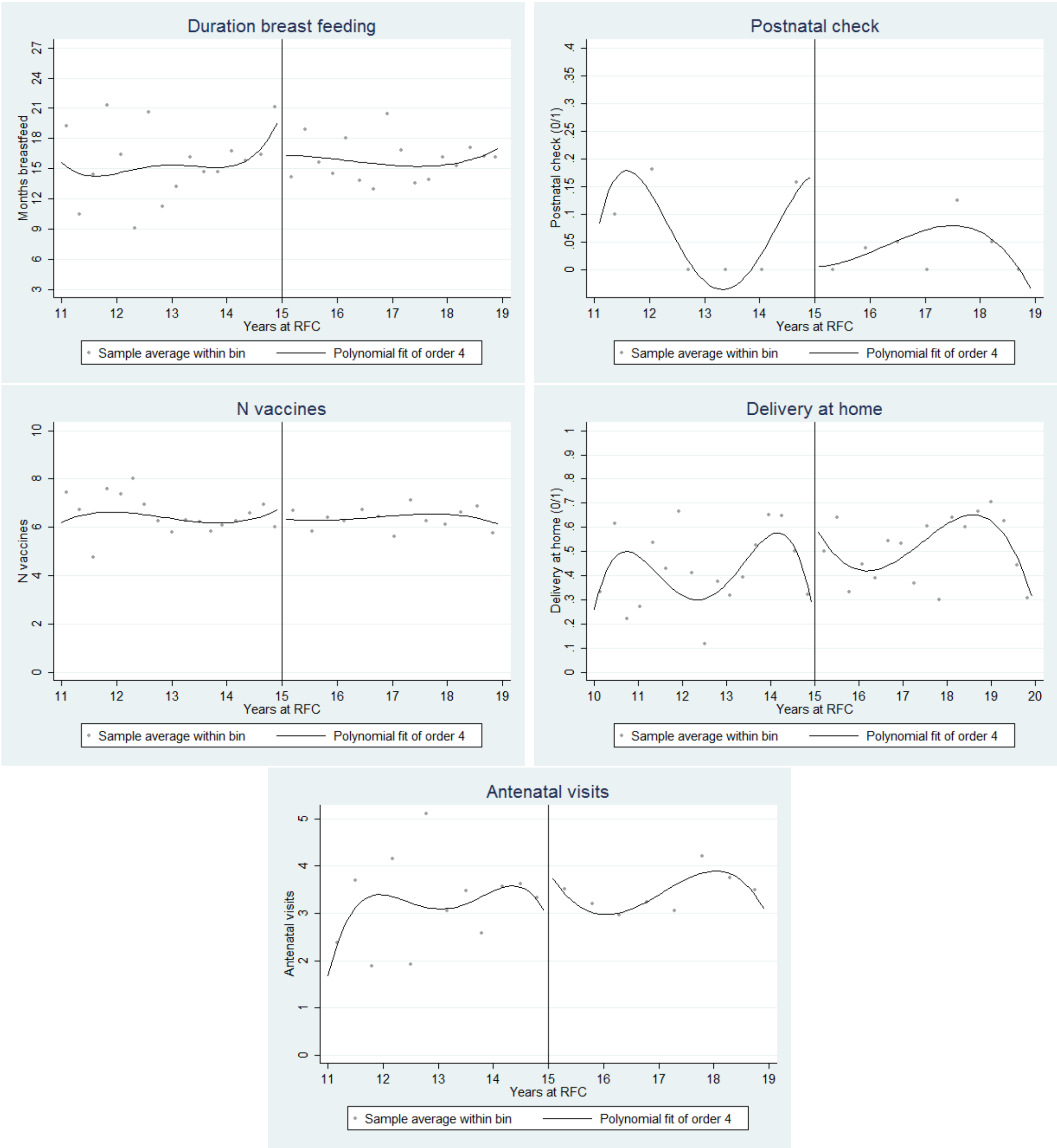
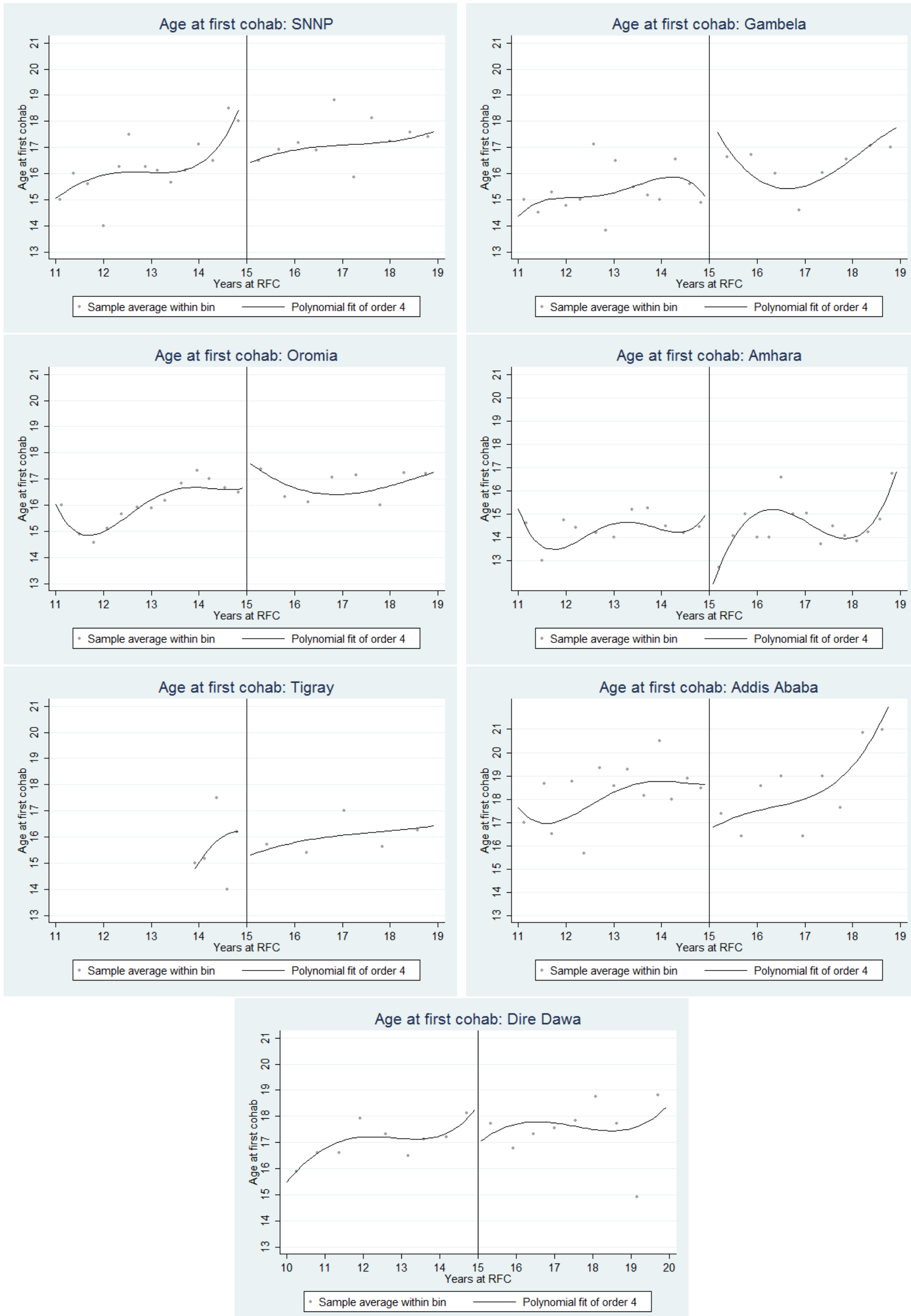


Figure 17: Age at first cohabitation at the cut-off for the Ethiopian regions that approved RFC between 2000 and 2007



D Sample of Women Still Cohabiting with the First Partner

Table 7: 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	1.233***	1.204***	1.291***	-0.097*	1.324***	-1.020	1.268***	-0.094
	(0.005)	(0.001)	(0.003)	(0.085)	(0.002)	(0.260)	(0.003)	(0.228)
Age at 1st cohab.		0.977***		-0.076*		-0.766		-0.075
		(0.003)		(0.082)		(0.420)		(0.351)
N		3126		3126		3126		2985
N effect. obs.		595		651		651		739
Bandwidth		34.8		37.9		37.1		43.0

	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	1.270***	-0.116*	1.286***	-0.028	1.336***	-0.964	1.258***	-0.014
	(0.004)	(0.098)	(0.003)	(0.981)	(0.002)	(0.293)	(0.004)	(0.861)
Age at first cohabit.		-0.091		-0.022		-0.692		-0.011
		(0.156)		(0.984)		(0.349)		(0.874)
N		3118		3102		3112		3125
N effect. obs.		593		647		691		626
Bandwidth		34.4		37.9		39.2		36.5

Note: The analysis reported in the table is conducted using the sample of women 18-49 that ever bore a child 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 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. 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. P-values are in parentheses.***p<0.01;**p<0.05;*p<0.1.