Effects of Motherhood Timing, Breastmilk Substitutes and Education on the Duration of Breastfeeding: Evidence from Egypt[1]

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Forthcoming in the *World Development*

[1] Acknowledgements: We thank Mehtabul Azam, Omar Dahi, Bidisha Lahiri, John Winters, seminar participants at the Oklahoma State University in 2018, and the two anonymous referees of the journal for their comments and suggestions on earlier drafts of this paper.
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Abstract

Breastfeeding has significant health and human capital effects on both mothers and infants. However, breastfeeding rates vary significantly within and across countries as societal, political, economic and cultural factors along with individual choices shape the breastfeeding practices. Using data from the Egyptian Demographic and Health Surveys, this study examines the effects of first motherhood timing, availability of breastmilk substitutes, and mothers' education levels on breastfeeding duration in a major developing country, Egypt. The empirical analysis, which corrects for the estimation errors that plagued previous research, shows that delaying the first motherhood timing and increasing the availability of infant formulas have statistically significant negative effects on breastfeeding duration. Furthermore, breastfeeding duration is found to be decreasing in mothers' education levels.

Keywords: Breastfeeding duration; Motherhood timing; Breastmilk substitutes; Education

JEL Codes: I12, J13, O12

1 Introduction

Early childhood development is a significant predictor of future human capital development. Particularly, parental investments during early childhood are shown to influence short-, medium-, and long-term health and cognitive outcomes with strong hysteresis effects (Cunha and Heckman, 2008; Cunha et al., 2010). In this paper we focus on one type of parental investment during early childhood, breastfeeding, and examine its various economic determinants. Recent studies have shown that breastfeeding has significant benefits for infants, including decreasing diarrhea, asthma, respiratory infections, ear infections, sudden infant death syndrome (SIDS), childhood obesity, gastrointestinal diseases, lymphoma, leukemia, Hodgkin's disease and chronic digestive diseases (American Academy of Pediatrics, 2018; Anderson et al., 1999; Modrek et al., 2017; Stanley et al., 2007; Thompson et al., 2017; Vestergaard et al., 1999; Victora et al., 2016). Breastfeeding also has significant and positive long term effects on infants' cognitive skills, which determine educational outcomes, and knowledge and skill accumulation during both childhood and adult life (Boucher et al., 2017; Cunha et al., 2010; Currie and Almond, 2011; Elwood et al., 2005; Gomez-Sanchiz et al., 2004; Heckman et al., 2006; Richards et al., 2002). Furthermore, breastfeeding has significant health benefits for the mothers, such as reducing the likelihood of breast and ovarian cancers, type 2 diabetes, earlier return to pre-pregnancy weight, and increased child spacing (American Academy of Pediatrics, 2018; Victora et al., 2016).

The benefits of breastfeeding are further multiplied in developing countries where there are significant problems with access to postnatal health care for mothers and infants. Furthermore, limited access to clean water, sanitation services and early childhood vaccines further increases the health risks to infants and mothers. Children in developing countries are also at a further disadvantage because of institutional and physical barriers to early childhood development including access to education infrastructure. Due to its substantial benefits, both the World Health Organization (WHO) and the United Nations International Children's Emergency Fund (UNICEF) recommend exclusive breastfeeding for at least the first six months of life, and continued breastfeeding until at least two years of age. Similarly, many national governments and all major international organizations have launched initiatives to encourage breastfeeding.[1] Recent studies all suggest significant heath benefits from breastfeeding beyond six months, with many suggesting the effects becoming stronger as the duration increases (Boucher et al., 2017; Modrek et al., 2017; Thompson et al., 2017).

Nevertheless, despite the overwhelming evidence showing that breastfeeding provides long-term benefits to both infants and mothers, the duration of breastfeeding significantly varies across countries (McFadden et al., 2017). For example, in the U.S., less than 25% of infants are exclusively breastfed during the first six months after birth, and globally the same rate is only 43% (CDC, 2018; WHO and UNICEF, 2016). Likewise, in the UK, the rate of exclusive breastfeeding at four months is just 7% (Brady, 2012). It is estimated that 823,000 deaths among children aged under five years, and 20,000 deaths due to breast cancer could be prevented globally each year if there was universal breastfeeding (Victora et al., 2016). Furthermore, the economic cost of not breastfeeding is estimated to be close to 0.5% of the world GDP (Rollins et al., 2016).

The existing evidence regarding the determinants of breastfeeding suggests that structural factors play a pivotal role. First, inequalities in health care affect women's decisions to breastfeed. In low-income communities and countries, limited access to health care and doctor consultations hampers the information flow to mothers regarding the benefits of breastfeeding, therefore inhibits breastfeeding practices.[2] Furthermore, sociocultural and economic factors, such as societal norms, the job market status, education and income levels, appear to be important predictors of a woman's decision to breastfeed (Bue and Priebe, 2018; McFadden et al., 2017; Rollins et al., 2016).

Among the economic determinants of breastfeeding practices, Roe et al. (1999) and Chetterji and Frick (2005) report that the opportunity cost of time spent breastfeeding increases after a mother returns to work, and thus these papers predict a clustering of breastfeeding termination immediately prior to returning to work. Numerous studies have also shown that post-birth maternal employment is associated with a shorter breastfeeding duration (Fein and Roe, 1998; Gielen et al., 1991; Kurinji et al., 1989; Visness and Kennedy, 1997). Furthermore, Lindberg (1996) shows that mothers who work part-time breastfeed longer than those employed full time. It is also possible that mothers with more education and higher incomes choose to breastfeed less because of their social status and values, availability of breastmilk substitutes, greater labor market pressure and higher opportunity costs on women's time (Victora et al., 2016). However, breastfeeding initiation and duration may increase among women with higher incomes because of them being more informed about its benefits. This trend in breastfeeding duration is particularly obvious in developed countries where breastfeeding initiation and duration increase with education (Logan et al., 2016). The existing evidence also suggests that breastfeeding rates are inversely related to income levels with higher rates observed in lower than higher income countries (Victora et al., 2016). Additionally, we observe a significant difference between rural and urban areas with the urban areas having a shorter breastfeeding duration (Adair et al., 1993). Among the societal drivers of breastfeeding behavior, gender-biased preferences, particularly in favor of sons, are shown to affect breastfeeding practices, as breastfeeding reduces the likelihood of future pregnancies (Chakravarty, 2015; Jayachandran and Kuziemko, 2011).

What is missing in this literature, however, is the effect of the timing of the first birth, which is expected to play a significant role in women's breastfeeding practices.[3] Globally, we observe an increase in the average age of mothers at the first birth, which can have significant effects on breastfeeding practices, and therefore on the health and human capital development of both infants and mothers (Bongaarts et al., 2017; OECD, 2018). Holding everything else constant, we expect that the increasing age of first-time mothers will have a negative effect on the duration of breastfeeding. We suggest that there are two major drivers of this negative relationship, which are biological constraints and labor force participation.

First, women face the biological constraint of declining fecundity by age (Boukje et al., 1991). Therefore, women who delay their first childbearing could lower the risk of future infertility before achieving their desired number of children by reducing the gap between successive children (i.e. child-spacing). One way to reduce the gap between two successive children is to shorten the breastfeeding duration because breastfeeding inhibits postnatal fertility and acts as a biological constraint, working as an effective natural birth control method (i.e. lactation amenorrhea method) (Jayachandran and Kuziemko, 2011; Wijden et al., 2003).[4]

Second, a large body of literature has already established that delaying motherhood timing has significant positive effects on human capital accumulation and labor market outcomes for women, including life time earnings and wages (Bratti and Cavalli, 2014; Buckles, 2008; Cristia, 2008; Herr, 2016; Hotz et al., 1997; Miller, 2011). Miller (2011), for example, shows that the wage premium in the US from delaying the timing of first motherhood by one additional year is approximately 3% to 10%. Similarly, women who breastfeed their children for at least six months are found to experience significant decreases in their lifetime earnings compared to mothers who do not breastfeed or breastfeed for shorter duration (Rippey and Noonan, 2012). Thus, women face trade-offs between motherhood timing and lifetime earnings.[5] We expect that these tradeoffs play a greater role in countries where female labor force participation is lower.

However, we know little about the effects of delaying the motherhood age on breastfeeding duration. Thus, we contribute to the literature by examining the effects of the timing of the first child on the duration of breastfeeding. Furthermore, we examine other major determinants of breastfeeding duration, including the availability of breastmilk substitutes and mothers' education levels. Notably, the effect of breastmilk substitutes on breastfeeding practices is a major concern for health organizations. Increasing evidence suggests a negative association between the increasing availability and accessibility of breastmilk substitutes and the duration of breastfeeding (Adair et al., 1993; Brady, 2012). Infant formula production is a profitable business for food companies and despite the efforts exerted by the WHO, its use is increasing (Piwoz and Huffman, 2015).[6] For example, in the UK, breastmilk substitute companies spend ten times more on marketing than the

Department of Health spends on promoting breastfeeding (Brady, 2012). The increasing use of breastmilk substitutes across the globe is a major concern as there are significant differences between breastmilk and its formula based substitutes in terms of nutritional content and health effects both for infants and mothers (Carignan et al., 2015; Hanson et al., 2016). There is also an increased risk of negative health effects from breastmilk substitutes in areas where access to clean water is limited, including both developed and developing countries (Piwoz and Huffman, 2015).

In the empirical analysis, we use the seven waves of Egyptian Demographic and Health Surveys (EDHS) with over 50,000 observations covering the period of 1988-2014. The empirical results, based on 2SLS, Nevo and Rosen (2012) bounding approach and the hazard model, suggest that delaying the first motherhood timing adversely affects the decision of when to wean a child, and this effect is strongest for the first child and decreases with birth order. We also find that the availability of breastmilk substitutes has a significantly negative effect on the duration of breastfeeding. Furthermore, more educated women are found to breastfeed for shorter duration. Similarly, women who live in urban areas or have a greater number of living children breastfeed for shorter duration. Confirming previous studies, we further show that girls are breastfed for a shorter period than boys. These results are robust to a variety of estimation techniques and a rich battery of sensitivity tests.

The remainder of the paper is organized as follows. Section two presents the empirical methodology including the data and estimation methods. Section three discusses the empirical results, followed by sensitivity analysis in section four. Section five concludes.

2 Empirical Methodology

2.1 Data

The empirical analysis is based on the seven most recent waves of the EDHS, which were carried out in 1988, 1992, 1995, 2000, 2005, 2008, and 2014. The EDHS is a nationally representative household survey of all women aged 15-49 years, and includes information about the breastfeeding duration of all children born over the past five years to each woman.[7] Additionally, the survey collects information regarding the fertility history, birth order, the sex composition of siblings, desired fertility, child health and mortality, education and labor market status of the mother and standard demographic and household characteristics.

Tables 1 and 2 present the summary statistics of the variables included in the empirical analysis. Additional descriptive statistics are provided in the Appendix. We find that the average observed breastfeeding duration is 13.473 months, which includes both children actively breastfed and children already weaned at the time of the survey. In Table 1 we also compare the demographic characteristics of mothers in five different age groups based on their first motherhood timing. Consistent with earlier studies, we find that younger mothers aged 13-17 marry early, have more children, have a larger ideal family preference, have a longer gap between successive children, and are more likely to live in rural areas. In contrast, mothers who do not have their first child before age 18 have more years of schooling and, as shown in Table 2, better labor market outcomes. We also find that mothers aged 13-17 breastfeed around two months longer than those mothers who are older. A similar pattern is observed throughout the complete duration of breastfeeding.

<Insert Tables 1-2 Here >

Figure 1 shows that the duration of breastfeeding and the age of the mother at first birth are negatively associated for the first child and all children, which is likely driven by biological constraints and labor market choices, as previously discussed. Figure 2 shows that the probability of infecundity in mothers who are 18 years or older is more than twice higher than that in mothers who are younger than 18.[8] As shown in the left panel of Figure 2, the first motherhood timing and infecundity are positively correlated, and there is a discrete increase in the probability of infecundity after the mid-thirties. In Table 1, we also see that irrespective of their age at first birth, mothers achieve their desired number of children, as suggested by the closing of the gap between the ideal and actual number of children.[9] Therefore, we expect the timing gap between successive children to be shorter among mothers who delay their first childbearing compared to that among mothers who have children when they are younger than 18. Indeed, as shown in the right panel of Figure 2, the average gap between two successive children and the mothers' age at the first birth are negatively correlated, and there is a discrete increase in the average time gap at age 25.

<Insert Figures 1-2 Here >

In Figure 3, we show that the age of the mothers at the first birth is positively correlated with their years of schooling and labor force participation. On the left panel, we see that mothers who had their first child between the ages of 13 and 20 have less than four years of schooling, and as shown in the the right panel, the labor force participation of this group of mothers is below 10%. We also note that both years of schooling and labor force participation monotonically increase among women who had their first motherhood timing at ages 21-30. In Table 1 we also see that the mothers who delayed their first childbirth until after 21 years of age completed three-to-eight additional years of schooling. In Table 2 we provide further descriptive statistics on the health by age and work related characteristics, broken down by the age of the first birth. We find that less than 9% of mothers aged 13-17 at the first birth work outside the home, and just 1% of these mothers work in skilled occupations. In contrast, on average, 27.36% of mothers aged 22-45 at first birth work outside the home, and 78% of these mothers work in skilled occupations.[10] As shown in Table 2, natural fecundity declines with age, making delayed childbearing riskier.

< Insert Figure 3 Here >

This sharp contrast in labor market outcomes between the two groups of mothers based on education and age at first birth implies that the opportunity cost of childbearing is much lower for less-educated mothers who have a child at an earlier age. This relationship also explains why mothers aged 13-17 have more children, a longer time gap between successive children, and a longer breastfeeding duration than mothers who delay their childbearing. To examine this effect more carefully, in Table 3, we present the results from a probit model, which is used to estimate the probability of labor market outcomes for three groups of women, including women who work outside the home (columns (3)-(4)), work full time (columns (5)-(6)), and work in skilled jobs (columns (7)-(8)), according to the age at the first pregnancy.[11] Our baseline group consists of women who were pregnant for the first time at age 13-18. The results suggest that all three groups of women who were pregnant for the first time at an older age have a much higher probability of participating in the labor force, working full time, and working in a skilled occupation. Furthermore, these effects get stronger as the age at the first pregnancy increases. In columns (1)-(2) we also confirm that the probability of miscarriage increases by age, especially after the mid-20s.[12]

< Insert Table 3 Here >

2.2 Econometric Specification

We test our hypotheses using the following linear regression model in Eq. (1), which is similar to Jayachandran and Kuziemko (2011):

$$y_{mi} = \beta_0 + \beta_1 A_m + X'_{mt}\beta + \alpha_i + \gamma_t + \delta_{at} + u_m \tag{1}$$

where y_{mi} is the number of months a mother (m) breastfeeds child i. We estimate Eq. (1) first for the first child and then for the average breastfeeding duration of all children. A_m is the mother's age at the first birth, and the coefficient β_1 measures the impact of delaying the timing of first motherhood by one additional year on the duration of breastfeeding. X is a vector of covariates. α_i is a vector of regional fixed effects, γ_t is the childbirth-year fixed effects, and δ_{at} is the child's age-in-months fixed effects. In total we include 50 age-in-month, 35 childbirth-year and 22 region fixed effects.[13] u_m is the unobserved error term.

X includes the following control variables:

$MilkSubst_t$, is the share of breastmilk substitutes in total merchandise imports of Egypt, standardized by the mean and standard deviation so that it is between zero and one.[14] The share of breastmilk substitutes in merchandise imports increased significantly in Egypt after the 1980s, rising from 0.03% during the 1960s and 1970s to 0.07% during the 1980s, and 0.11% between 1990 and 2014. We expect that the increasing availability of breastmilk substitutes has a detrimental effect on the duration of breastfeeding, which is a concern increasingly voiced by the WHO and UNICEF. The increasing marketing efforts by breastmilk substitute companies are also expected to negatively affect breastfeeding duration (Adair et al., 1993; Brady, 2012).[15]

$Education - Primary_m$, $Education - Secondary_m$, and $Education - High_m$ are three control variables used to examine the effect of the mother's education level. These variables are measured as dummy variables equalling one if the mother has primary, secondary or higher education. The missing category is mothers with less than primary education. We expect that women with more education breastfeed for a shorter duration as the opportunity cost on their time increases through work-family tradeoff. The income level is also positively correlated with the education level. However, the increasing education level may also encourage women to breastfeed longer because of the increasing awareness of the benefits of breastfeeding.

$MotherWorking_m$ is a dummy variable equaling one if the mother works outside the home at the time of the survey. We expect that this variable has a negative effect on breastfeeding duration because of the work-family tradeoff.[16]

Gender is a dummy variable equaling one if the gender of the child being breastfed is female. Because of male-biased gender preferences in Egypt, as discussed by Chakravarty (2015), we expect girls to be breastfed for a shorter duration than boys.

Religion is a dummy variable equaling one (zero) if the mother is Muslim (Christian), and controls for unobserved cultural, societal or religious practices that may affect the breastfeeding duration.

HouseholdHead is a dummy variable equaling one if the mother is the head of the household. Increasing demand on a mother's time as the head of a household is expected to lower the breastfeeding duration.

$Education\!-\!Primary_f$, $Education\!-\!Secondary_f$, and $Education\!-\!High_f$ are three control variables used to examine the effect of the father's education level. They are measured by dummy variables equaling one if the father has primary, secondary or higher education. The missing category is fathers with less than primary education. We expect increasing education level of fathers to be negatively correlated with breastfeeding duration if it positively affects the mother's likelihood to work or have access to breastmilk substitutes. This variable may also increase the breastfeeding duration because of an increased awareness of the benefits of breastfeeding. However, father's education level is likely to be correlated with the mother's education level and employment status and, therefore, may be captured by other control variables.

TotalChildren is the total number of living children born to the respondents and controls for the effect of family size on breastfeeding duration. The demand for women's time increases as the family size grows because of domestic work and the attending needs of the other children, and therefore, we expect that this variable has a negative effect on the breastfeeding duration.

Rural is a dummy variable equaling one (zero) if the respondent lived in a rural (urban) area. We expect that mothers residing in the countryside breastfeed longer than those living in urban areas because of economic and societal factors.

2.3 Estimation method

The simplest strategy to estimate equation (1) is to use OLS. However, OLS estimates of β_1 are likely biased because the relationship between the age at the first birth and the duration of breastfeeding may be driven by selection, either due to unobserved heterogeneity related to health conditions, or delays associated with unobserved determinants of labor market outcomes. Thus, these two types of nonrandom selection likely cause $cov(A_m, u_m) \leq 0$, leading to a downward bias in OLS estimates. Therefore, to consistently estimate Eq. (1) we rely on the 2SLS method, and include the following two instruments: changes in the length of compulsory education in Egypt, and the regional variation in predicted age-specific first birth growth rates.

The obvious challenge in estimating the effects of delaying motherhood on the duration of breastfeeding is identifying plausible exogenous variation in the first birth timing. The first birth timing is likely correlated with the mother's unobserved health conditions and other unobserved factors, such as career goals or productivity, which also affect the duration of breastfeeding. We use the changes in the length of compulsory schooling in Egypt in 1981, 1988, and 1999 as our primary IV to explain the effects of first motherhood timing on breastfeeding duration. Exogenous changes in the length of compulsory education is a widely used and reliable IV for motherhood timing, changes in women's fertility, and education decisions. The education system in Egypt is centralized and based on a 6+3+3 system with six years of primary, three years of secondary and three years of high school education. Until 1980, compulsory education in Egypt included only the first tier with six years of education. In 1981, lawmakers increased compulsory education to nine years, covering secondary education. However, beginning in 1988, the length of primary schooling was reduced from six to five years, which shortened compulsory education from nine to eight years. This change affected all individuals born on or after October 1, 1977. Yet, this policy was subsequently reversed in 1999, restoring primary education to six years and raising the length of compulsory schooling to nine years. These three changes in 1980, 1988, and 1999 allow us to implement a perfect identification strategy, as these changes occurred independent of women's breastfeeding decisions but are directly correlated with the timing of the first birth.

Because of these policy changes, all individuals born before October 1, 1977, or after 1988 attended one more year of schooling than those born afterwards. Similarly, individuals born before 1969 all attended three fewer years of schooling than those born afterwards. It is very unlikely that parents could have predicted these policy changes and adjusted the timing of their children's birthdays. Therefore, we can use these cutoff dates as a natural experiment to compare the first birth timing of mothers before and after 1970, 1977 and 1988. Here, the underlying assumption is that the length of compulsory education directly affects the timing of marriage and women's family planning, which is consistent with the findings of recent studies using educational policy changes as an IV to explain the effects of education on fertility (Ali and Gurmu, 2018; Black et al., 2008; Dincer et al., 2014; Huang et al., 2020).[17]

Additionally, we employ a second IV, which is the regional variations in the mothers' ages at first birth. This approach identifies how regional variations in the growth rates of first birth by age affect the decisions of individual mothers regarding the timing of their first childbearing. For a given age at the first birth, we predict the regional growth rates by the canonical product of the regional shares of first births and the national growth rates of first births. Therefore, our cross-sectional variations in the predicted change in age-specific first birth are derived from the regional demographic compositions of women by age.[18] Formally, we construct this IV ($\Delta\widehat{ABG}_{art}$) as follows:

$$\Delta\widehat{ABG}_{art} = \sum_{r=1}^{R} \varphi_{art} \times \Delta\widetilde{ABG}_{a\cancel{r}t} \tag{2}$$

where the share of first births for age group a in region r at period t is calculated as $\varphi_{art} = N_{art}/\sum_{a} N_{art}$ and N_a is the number of mothers who have their first birth at age a. The notation $\cancel{r}$ in $\Delta\widetilde{ABG}_{a\cancel{r}t}$ indicates that we exclude region r to calculate the national growth rate of first birth for age group a at period t.

In Eq. (2), we use the following three sources of variation to predict the regional age-specific first birth growth rates: (i) the age of mother at the first birth, (ii) geographical region, and (iii) time. Since we have 37 first motherhood timings, 22 regions, and 32 birth years, $\Delta\widehat{ABG}_{art}$ can take 26,048 different values.[19] This measure of predicted regional age-specific first birth growth rates in period t can serve as an instrument because it is mainly determined by the regional shares of age-specific first births. Therefore, we expect this measure to be correlated with the long-run trend of first motherhood timings but uncorrelated with the unobserved characteristics of individual mothers. However, the regional shares of first birth by age might be correlated with the unobserved component of breastfeeding duration. If this is true, our instrument will fail the exogeneity condition because the error terms in equation (1) are correlated with the region-specific shares of first births. To address this potential bias, we use a one-period lag, which is equivalent to a five-year lag of the regional shares of age-specific first births to avoid any potential contemporary correlation between the error terms and regional shares.

In addition to the 2SLS method, we also employ the Nevo and Rosen (2012) bounding approach, which helps to address any concern regarding the validity of our IVs. The Nevo and Rosen (2012) bounding approach provides a range for a consistent estimator of β when an instrument is imperfect.[20] To apply this procedure, the primary restriction our second IV needs to satisfy is that it is less endogenous than the endogenous regressor of age at first birth. The intuition behind Nevo and Rosen (2012) is that β_z^{IV} has a downward bias because of a possible negative correlation between z_m and u_m. In this setting, the Nevo and Rosen (2012) approach provides two-sided bounds of the effect of age at first birth on the duration of breastfeeding that is valid under weaker assumptions than those used in previous studies. More importantly, this sharp lower bound directly accounts for any unobserved factors that may lead to a negative correlation between the IVs and the duration of breastfeeding.

In our benchmark estimations we rely on 2SLS and the Neovo and Rosen bounding approach. However, as a third method, we use a hazard model to estimate Eq. (1). Previous studies have mostly relied on OLS and hazard model estimates to examine the determinants of breastfeeding duration. Since many children in the sample are still nursing, the regression analysis estimates the effect of delaying the first birth on the observed duration rather than the completed breastfeeding duration. Using a hazard estimation method, we find that the average complete duration of breastfeeding in Egypt is 20.766 months (Table 1). In contrast, the observed duration of breastfeeding is 13.473 months (Table 1). This discrepancy is because in the simple OLS regression model we treat the observed value of the breastfeeding duration as the end point of the breastfeeding duration. As expected, this is not true for many children, and we have a measurement error in our dependent variable. Because this measurement error is most likely correlated with the mother's age at the first birth, the OLS estimator of the first motherhood timing is biased. To address this issue, in all regression estimations we add age-month fixed effects. In addition, following Jayachandran and Kuziemko (2011), we use the Cox proportional hazard model to account for the censoring of completed breastfeeding duration. The main advantage of the hazard model over the simple regression model is that the hazard model does not impose any conditions on the baseline hazard function.[21] The failure event in the Cox model is defined as a child being weaned. A positive coefficient implies that delaying the first birth is associated with a higher probability of a child being weaned.

3 Results

3.1 Effects of First Motherhood Timing on the First Child

Table 4 reports the baseline results of the effects of the mother's age at first birth on the duration of breastfeeding of the first child. For comparison, we report the results obtained from the OLS, 2SLS, Nevo and Rosen, and hazard estimations in columns (1)-(6). The OLS estimate shown in column (1) suggests that a one-year delay in the first motherhood timing is associated with a 0.353 month decrease in the duration of breastfeeding. However, as previously discussed, this estimate is likely biased downwards because of the negative correlation between motherhood timing and the unobserved error, u_m, in equation (1). Column (2) reports the results of the 2SLS estimation and suggests that increasing the age of the first motherhood timing by one year has a significantly negative effect on the duration of breastfeeding by approximately 0.839 months. To put this number in perspective, using sample averages for women with different socio-economic backgrounds, we find that the negative effect of delaying the age of first birth on breastfeeding duration is around two months larger for working mothers compared to non-working mothers. When looking at women with secondary and higher education vs. those with primary education or less, the effect increases to 2.3 months (using beta estimates in column (2) in Table 4).[22]

< Insert Table 4 Here >

In columns (3)-(4) we report the Nevo and Rosen (2012) two-sided bounding estimates. The lower bound estimate is -1.128 and the upper bound estimate is -0.502, suggesting that each additional year increase in the age at the first birth decreases the breastfeeding duration at a minimum of two and a maximum of around four weeks. Columns (5)-(6) report the results of the Cox hazard model and confirm the findings shown in columns (2)-(4).[23] The coefficient estimate shown in column (5), i.e. 0.06, implies that delaying the first motherhood timing by one additional year increases the probability of a child being weaned earlier by 6%. In column (6), we use the first stage predicted value of the age at the first birth instead of the actual age at the first birth and again find that the hazard coefficient is positive, suggesting that delaying the first childbearing is positively correlated with a higher probability of the first child being weaned earlier. This result supports the findings presented in columns (1)-(4), showing that delaying the first motherhood timing has a negative impact on the breastfeeding duration. Overall, the results are quite robust to different estimation methods in columns (1)-(6).

In Table (4) we also find that increasing availability of breastmilk substitutes in Egypt has an economically and statistically significant negative impact on the duration of breastfeeding. This finding is also consistent with the findings of earlier research (Adair et al., 1993; Brady, 2012). According to the point estimates from column (2), a one standard deviation increase in the availability of breastmilk substitutes decreases the breastfeeding duration by over one month (-1.145). Overall, breastmilk substitutes are found to have an economically more significant negative effect than the age of the mothers at the first birth on breastfeeding duration in Egypt.

Furthermore, consistent with the modernization theory, we find that the education level of the mother significantly affects breastfeeding practices. We find that mothers with primary and secondary education, compromising 52.81% of the sample, breastfeed for shorter durations than mothers without any education, which stands in contrast to evidence from developed country experiences. The effect of higher education is not statistically significant, except for in the hazard model, as shown in columns (5)-(6). We should note that only 8% of mothers in the sample have higher education.

In addition, we find that the total number of living children born to the respondents has a significantly negative effect on the breastfeeding duration, causing a reduction of over one month. We find a similar effect among women living in urban as opposed to rural areas with a shorter breastfeeding duration. We should note that these results are consistent across different estimation methods in columns (1)-(6). Our results also confirm the male-biased breastfeeding practices in Egypt, as shown by the significantly negative coefficient of the gender variable, and support the findings of Chakravarty (2015) for Egypt and Jayachandran and Kuziemko (2011) for India. Finally, we find that religion has a significant effect on the breastfeeding duration, which is found to be higher among Muslim mothers due to other unobserved characteristics.

3.2 Effects of First Motherhood Timing on All Children

In this section we examine how the timing of the first birth affects the average breastfeeding duration for all children rather than only the first child.[24] Delaying the first motherhood timing affects the duration of breastfeeding of all subsequent children in two different ways. First, delaying the first birth by one additional year directly delays the motherhood timing of all subsequent births by at least one year. Therefore, such a delay affects the duration of breastfeeding of all children because of declining fecundity by age and the negative impact of a longer duration of breastfeeding on future fertility. Second, delaying the first birth by one additional year can delay all subsequent birth timings by changing the optimal timing of all subsequent births. Therefore, delaying the first motherhood timing can indirectly affect the duration of breastfeeding of all children through subsequent birth timings.

For this estimation, we introduce two additional variables to control for the effects of the ideal family size as described in Jayachandran and Kuziemko (2011). First, Δ *Ideal* is the difference between a child's birth order and the mother's desired number of children, as reported in the EDHS. The second variable is a dummy variable, $\Delta\ Ideal > 0$, which is equal to one if $\Delta\ Ideal$ is greater than zero. Thus, this variable indicates whether a child was born after the mother had already achieved her desired number of children. We expect that breastfeeding decreases in birth order as a mother approaches her ideal fertility if she reaches the ideal number after a certain age when the biological constraints are stronger. Alternatively, increasing the family size may also increase the opportunity cost of breastfeeding because of the growing demand on the mother's time. However, after the ideal number of children is reached, the duration of breastfeeding may increase to avoid any new pregnancies.[25]

Table 5 presents the results from estimating Eq. (1), but in this analysis our dependent variable is the average breastfeeding duration of all children. Similar to Table 4, Columns (1)-(6) show the results from the OLS, 2SLS, Nevo and Rosen, and hazard estimations. As in Table 4, Column (2) is our benchmark estimation. The findings are consistent with the previous results and confirm the significantly negative effect of first time motherhood timing on breastfeeding duration. We find that a one-year delay of the first birth decreases the duration of breastfeeding by 0.252 months among all children. As expected, these results further suggest that the total impacts of the first motherhood timing on breastfeeding duration decreases with subsequent births. Columns (3)-(4) also suggest that this effect ranges between 0.283 and 1.419 months.

< Insert Table 5 Here >

The remaining control variables show similar signs and significance levels as shown in Table 4. Briefly, we find that increasing the availability of breastmilk substitutes has a significantly negative effect on the breastfeeding duration. We also find that mothers with primary and secondary education, and those with daughters or those living in urban areas appear to breastfeed each child less on average. Women married to men with higher education appear to breastfeed for a shorter duration. Increasing the number of children born to the respondents also has a negative effect. Finally, we find that $\Delta\ Ideal$ has a negative effect and Δ $Ideal > 0$ has a positive effect on the breastfeeding duration, suggesting that the breastfeeding duration decreases after the ideal family size is reached.

4 Robustness Checks

We check the robustness of our findings using a rich battery of sensitivity tests. In all robustness tests we report only the 2SLS estimates of the motherhood timing, breastmilk substitutes and the mothers' education levels but provide full results for other control variables, and also a discussion of the estimation methodology in the online Appendix.

First, the 2SLS and Nevo and Rosen results depend on the validity of the IVs used. Therefore, in addition to the two IVs, which are the exogenous changes in the length of compulsory schooling in Egypt and average regional growth rates of age at the first birth, we include information regarding mothers with miscarriages. Previous studies have used miscarriage as an IV because of its delaying effect on the first birth and the motherhood timing (Ashcraft et al., 2013; Bratti and Cavalli, 2014; Hotz et al., 1997; Miller, 2011). Conceptually, this IV approach compares the duration of breastfeeding between two groups of women who did and did not have a miscarriage before the first birth. Because a miscarriage delays the motherhood timing of only one group, its exogenous variations can arguably be used as an IV to identify the effects of age at first birth on the duration of breastfeeding. Here, the main assumption is that the first miscarriage should not affect the duration of breastfeeding, except for through the age at first birth. This identifying assumption may not hold if miscarriages are correlated with any unobserved health or labor market variables, and thus are not randomly distributed (Ashcraft et al., 2013).[26] Furthermore, only 5.8% of the mothers in the EDHS reported having a miscarriage (Table 2), limiting the use of this variable as an IV explaining the variation in the motherhood timing across women in Egypt. There is also the problem of underreporting of miscarriages in the surveys because of some observable and unobservable characteristics of the mothers.[27] Notwithstanding these limitations, columns (1)-(2) in Table 6 show the results using miscarriage as an additional IV for the first child and all children and confirm our earlier findings.

< Insert Table 6 Here >

Next, to control for any possible nonrandom effect of the birth-month of children, we include twelve birth-month fixed effects. The results are presented in columns (3)-(4) of Table 6 and support our earlier findings. Furthermore, as discussed in Section 2.3, we used five-year lags for the measurement of the regional first birth IV variable in the benchmark estimates. For robustness, we repeated these regressions using 10-year lags instead of five-year lags and reported the results in columns (5)-(6) of Table 6. The results are similar to those reported before. Finally, to test the sensitivity of the regional IV to age-group selection, in Table 7 we experimented with different age-group criteria including the following age at first birth groups. Group 1 includes the <17, 17-19, 20-21, 22-24, and >24 age at first birth clusters and < 1985, 1985-1989, 1990-1994, 1995-1999, 2000-2004, 2005-2009, and > 2009 child birth year clusters. Group 2 includes the <18, 18-19, 20-21, 22-25, >25 age at first birth clusters and < 1985, 1985-1988, 1989-1992, 1993-1996, 1997-2000, 2001-2004, 2005-2008, and > 2008 child birth year clusters. Groups 3 includes the <18, 18-19, 20-21, 22-25, >25 age at first birth clusters and < 1985, 1985-1990, 1991-1996, 1997-2002, 2003-2008, and > 2008 child birth year clusters. Group 4 includes the <17, 17-19, 20-21, 22-24, >24 age at first birth clusters and < 1985, 1985-1990, 1991-1996, 1997-2002, 2003-2008, and > 2008 child birth year clusters. The results are reported in Table 7 and further support our earlier findings.

< Insert Table 7 Here >

As additional robustness tests, we included the following control variables in Eq. (1): children's birth order, children's birth order squared, average age of all births, average gap between all subsequent births, the mother's desire to have more children, having an unwanted pregnancy, and the mother's occupation type.[28] Although these variables are correlated with other control variables, they may still affect breastfeeding duration. The results are available in the Appendix and remain very similar to the previously reported results.

5 Conclusion

This paper examined the determinants of breastfeeding duration in Egypt while focusing on the timing of birth, availability of breastmilk substitutes, and mothers' education levels. In contrast to previous studies on the subject, our estimation approach directly addresses the possible correlation between the age at the first birth and women's unobserved health conditions. To this end, we use the exogenous changes in the number of years of compulsory education in Egypt in 1981, 1989 and 1999 along with regional growth rates of age at first birth as IVs for first motherhood timing. Using six nationally representative EDHS samples between 1988 and 2014 and based on 2SLS, the Nevo and Rosen (2012) bounding approach, and hazard model, we demonstrate that delaying the first birth has a significantly negative effect on the breastfeeding duration of the first child and all subsequent children.

We further show that increasing the availability of breastmilk substitutes in Egypt has a negative effect on the breastfeeding duration at economically and statistically significant levels. This finding is consistent with the findings reported by other scholars such as Brady (2012) who argued that increasing availability and accessibility of infant formulas along with aggressive marketing efforts of infant formula producers had a detrimental effect on breastfeeding in developing countries. We also find that increasing education levels of mothers, increasing urbanization, having a daughter, and increasing the ideal family size all have significantly negative effects on breastfeeding duration.

Existing studies suggest that breastfeeding beyond 6 months has significant health effects. Thompson et al. (2017), for example find that protective benefits of breastfeeding against SIDS increase with the duration of breastfeeding and remain significant beyond 6 months, and in fact become even more significant beyond six months in some country samples. Furthermore, Modrek et al. (2017), using a continuous measure of breastfeeding duration, find that for every extra week that the child is breastfed, the likelihood of childhood obesity at age 2 declines by 0.82%. Boucher et al. (2017) also find that breastfeeding (any or exclusive) has a significantly positive effect on cognitive development and a significantly negative effect on autistic traits. After breaking down their continuous measure of duration of any breastfeeding into sub-categories, Boucher et al. (2017) also report that, compared to children who are breastfed 0-2 months, children who are breastfed for 6-12 and beyond 12 months enjoyed a significantly larger positive effect on their cognitive development and a lower risk of autistic traits. In fact, the observed effects disappeared for children who are breastfed for 2-6 months and became significant only for 2-6, and 6-12 months of breastfeeding duration. Furthermore, the predicted effects are shown to grow stronger with the increased duration of breastfeeding and are found to be the highest for those beyond 12 months of any breastfeeding.

Given the overwhelming evidence supporting the short-, medium-, and long-term health and cognitive development benefits of increased duration of breastfeeding, therefore, our findings have significant policy implications. First, increasing the availability and accessibility of pre- and postnatal health care is vital for ensuring that women have complete information regarding the benefits of breastfeeding. Second, given that the negative labor market effects of childbearing are increasingly among of the primary reasons for delaying the first child bearing age, policy makers should find remedies to address the negative effects of having children on women's job market outcomes and lifetime earnings. This problem is particularly important in countries, such as Egypt, where women already suffer from low labor market participation (i.e. 15% during the period analyzed). There is an ongoing debate in the literature about ways in which policy makers and legislators can help establish an enabling environment for women to continue breastfeeding without having to worry about the opportunity costs in terms of their labor market outcomes. Possible interventions include legislation, increased enforcement of the laws, providing financing, public opinion campaigns, social mobilization, and increasing public awareness about health and economic costs of sub-optimal breastfeeding. Other policy actions include increasing public awareness to reduce societal pressures and stigma about breastfeeding in public spaces.[29]

Third, from a public health vs. business perspective, there is tradeoff between breastmilk and infant formula. The lobbying and marketing power of breastmilk substitute producers have reshaped the breastfeeding practices of mothers worldwide.[30] In addition, health care providers are targeted by the food industry to increase the early adoption of infant formulas as a substitute for breastmilk. Given the significantly negative effect of breastmilk substitute imports on breastfeeding duration, governments and the wider public need to be aware of the difference between profit maximization and the public health aspects of the availability of breastmilk substitutes. Overall, there needs to be an increasing understanding by the public that the promotion of breastfeeding is a societal responsibility with significant benefits to the children, women, and the wider society. The solutions to sub-optimal breastfeeding practices and efforts to provide an enabling environment for breastfeeding, therefore, also need to be multifaceted and should take into account individual attitudes, public health systems, societal attitudes, workplace and labor market responses, and the public health priorities of policy makers.

Future research could explore whether the findings of this research are applicable to other countries, both developed and developing. Notably, the average breastfeeding duration in Egypt is relatively high, and therefore longitudinal studies are needed to determine the long-term child health effects of delaying the first motherhood timing. However, it is reasonable to assume that the health effects could be significantly larger in countries, such as the U.S., where the average breastfeeding duration is on the lower end of the spectrum. Furthermore, the predicted effects in this study are for each year of delay, and therefore, these effects compound over time. The same is true for the effects of breastmilk substitutes. In addition, we expect that the health effects of the increasing use of breastmilk substitutes are heterogeneous depending on the income level of the mothers. In particular, we expect that the negative effects are compounded when the mothers lack access to safe drinking water for preparing the milk substitutes. Risks of childhood diseases due to reduced breastfeeding are also expected to be higher among lower-income households. We plan to examine these differential effects in future research.

Notes

[1]See, UNDP (2013a,b); WHO and UNICEF (2016). For a review, see Bue and Priebe (2018).

[2]For example, these inequalities play a key role in explaining breastfeeding practices in the US where a high level of variance exists in exclusive breastfeeding during the first six months with Mississippi having the lowest rate at 13% and Vermont having the highest rate at 38% (CDC, 2018).

[3]We should note that there is substantial research on the determinants of the timing of first birth, including political, societal, cultural and economic factors. See, for example, Amin and Behrman (2014); Balley (1989); Becker (1960); Becker and Lewis (1973); Cochrane et al. (1990); Lavy and Zablotsky (2015).

[4]Numerous studies and national health surveys, including those from Egypt, have shown that women commonly use breastfeeding as a traditional birth control method, particularly in lower-income countries and communities where affordable birth control methods and contraceptives are lacking (EDHS, 2015). In fact, (EDHS, 2015) suggests that prolonged breastfeeding is the most commonly recognized traditional method of family planning in Egypt (72%). For further discussion on lactation amenorrhea method, see Hatcher et al. (2011) and Wijden et al. (2003).

[5]While breast pumps could offer time flexibility to working mothers, due to their financial cost and demand on women's time, they are no less easier to adopt.

[6]In 1981 the WHO International Code of Marketing Breast Milk Substitutes was approved almost unanimously with the U.S. casting the only negative vote. The Code was proposed in response to the finding that the increasing marketing of infant formulas in developing countries contributed to increasing infant mortality and other health problems. This code prohibits any marketing or gifts of infant formulas to mothers, or bribing health care workers (Brady, 2012).

[7]The duration of breastfeeding is limited to 60 months in all EDHS waves except for in the year 1992, in which the cutoff was 50 months. Chakravarty (2015) reports that approximately

99% children are weaned by 36 months in Egypt, and therefore, this cutoff is not a binding constraint. Furthermore, because data on "exclusive" breastfeeding is not available in EDHS, our results are based duration of "any" breastfeeding.

[8]Infecundity is defined as when a woman fails to achieve pregnancy over twelve successive months of trying.

[9]The total number of children exceeds the desired number of children among mothers who are younger than 18. In addition, in our sample, we have a significant fraction of mothers who still desire to have more children in the future. Thus, if we exclude these mothers, the difference between the actual and desired number of children is statistically insignificant.

[10]In the sample, 6,116 mothers aged 22-45 at the first birth work outside the home, and of these mothers, 4,801 mothers work in skilled occupations.

[11]The probit model results show the marginal effects, which are estimated using a mother-level sample where each observation represents a unique mother.

[12]We define the outcome variable, miscarriage, as to whether a mother had a miscarriage before the first birth. In the event of first birth, we include both live and still births (in the sample, 0.51% of the first births are stillbirth).

[13]Age-in-month fixed effects help address right-censoring of breastfeeding duration Jayachandran and Kuziemko (2011).

[14]The variable $MilkSubst$ is measured at the four-digit SITC level (0488) using Comtrade data reported by the MIT Center for Economic Complexity.

[15]The breastmilk substitute data is available only at the country level. A better measure, when available, is to replace it with a variable that captures the use of breastmilk substitutes at the individual level from the health surveys.

[16]We should note that this variable indicates only whether the mother worked at the time of the survey and not when she had the child.

[17]We assume that schooling starts at the age of 6. Thus, based on the mothers' birth years and number of compulsory schooling years, we calculate the age cutoffs, which is equal to six plus the years of compulsory schooling.

[18] For similar uses, see Blanchard and Katz (1992) and Klepinger et al. (1999).

[19] In the sample, we have a limited number of observations to predict the growth rates of first birth in each region for $32(birthyears) \times 37(motherhoodtiming)$ different cells. Therefore, we clustered mothers into five groups based on their first motherhood timing over a total of seven time periods. The five groups consist of mothers aged 13-17, 18-19, 20-21, 22-25, and 26-45. We divide the 32 years uniformly by a five-year time window for each period except for the first period.

[20] Nevo and Rosen (2012) define an instrument imperfect when the strict exogeneity assumption of an instrument is not satisfied. We provide a detailed description of the assumptions and the setup of this procedure in the online Appendix.

[21] The details of the hazard model estimation are provided in the Appendix.

[22] The sample average of age at first birth is 22.735 for working and 20.360 for non-working women. It is 22.278 for those with secondary and higher education, and 19.496 for those with primary education or less.

[23] To correct for the endogeneity of motherhood timing, we use the predicted value of motherhood timing from the first stage regression. In the Appendix we report the results with the observed timing variable, which are very close to those reported here.

[24] Note that in "all children", we include the first child because some mothers only have one child. Therefore, selecting only those mothers who have more than one child could potentially create a sample selection problem.

[25] We repeat the regression analysis shown in Table 4 and include these two additional variables, and find almost identical results. These results are reported in the online Appendix. However, we should note only 2.46% of observations of this dummy variable for the first birth take the value one as the remaining respondents reported having an ideal family size greater than one.

[26] For example, if unhealthy women have on average more miscarriages than healthy women, the variable miscarriage is not completely exogenous.

[27] In the probit analysis shown in Table 3 columns (1-2), we find that the mother's age at the

first pregnancy and first miscarriage are positively correlated. Therefore, the first miscarriage is likely to be nonrandom and is expected to be negatively correlated with the unobserved error u_m. See Appendix for further discussion.

[28]Based on the DHS occupation classifications, we use the following 10 occupation categories: (i) professional, technical or managerial; (ii) clerical; (iii) sales; (iv) agricultural; (v) self-employed agricultural; (vi) household, domestic and services; (vii) household domestic; (viii) services; (ix) unskilled manual; and (x) skilled manual.

[29]Rollins et al. (2016) provide an in depth discussion of various enabling ways for increasing breastfeeding practices.

[30]For example, in 2018, the US showed strong opposition to a WHO proposal, calling governments to protect, promote and support breastfeeding. The official reason for the US was to prevent the stigmatization of breastmilk substitutes and infant formulas (Jacos, 2018).

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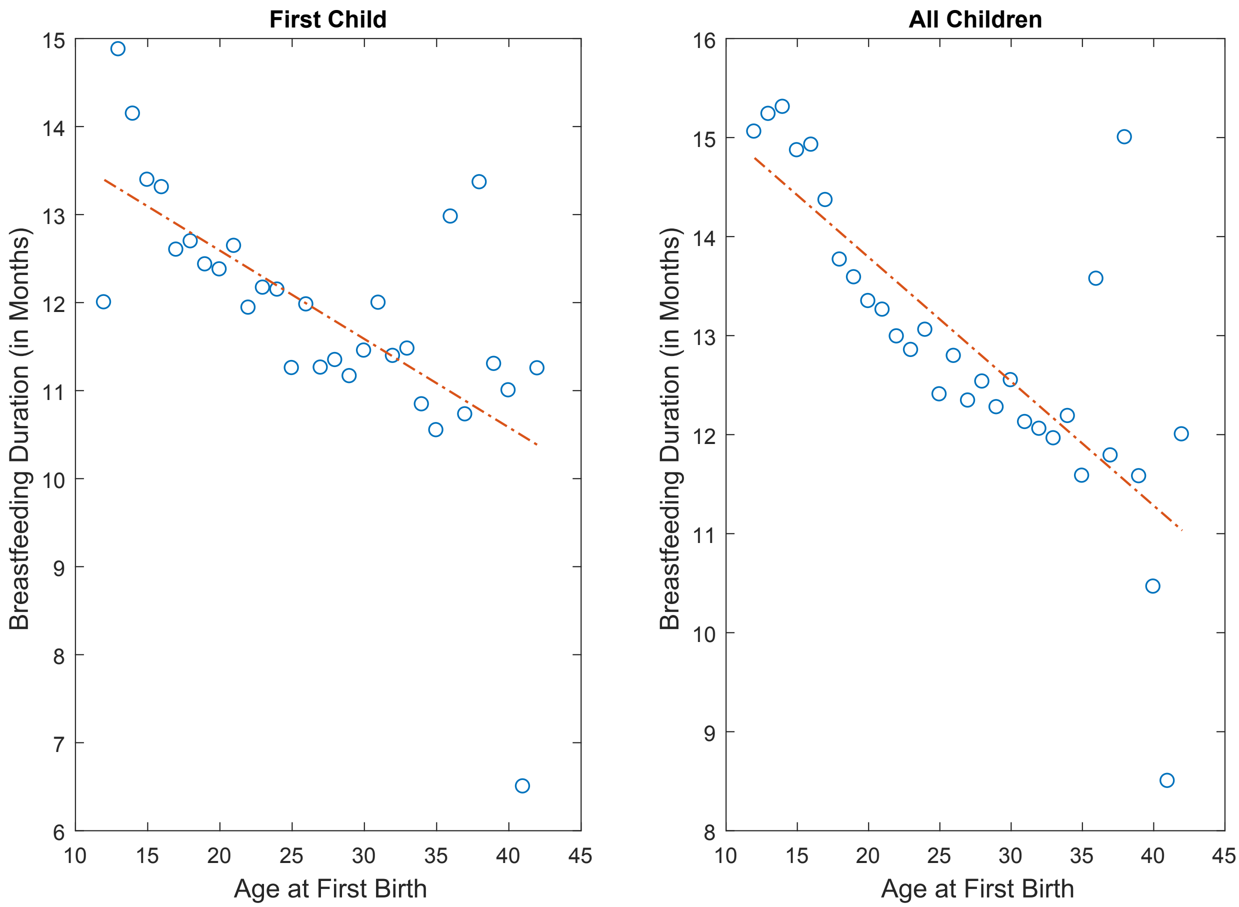
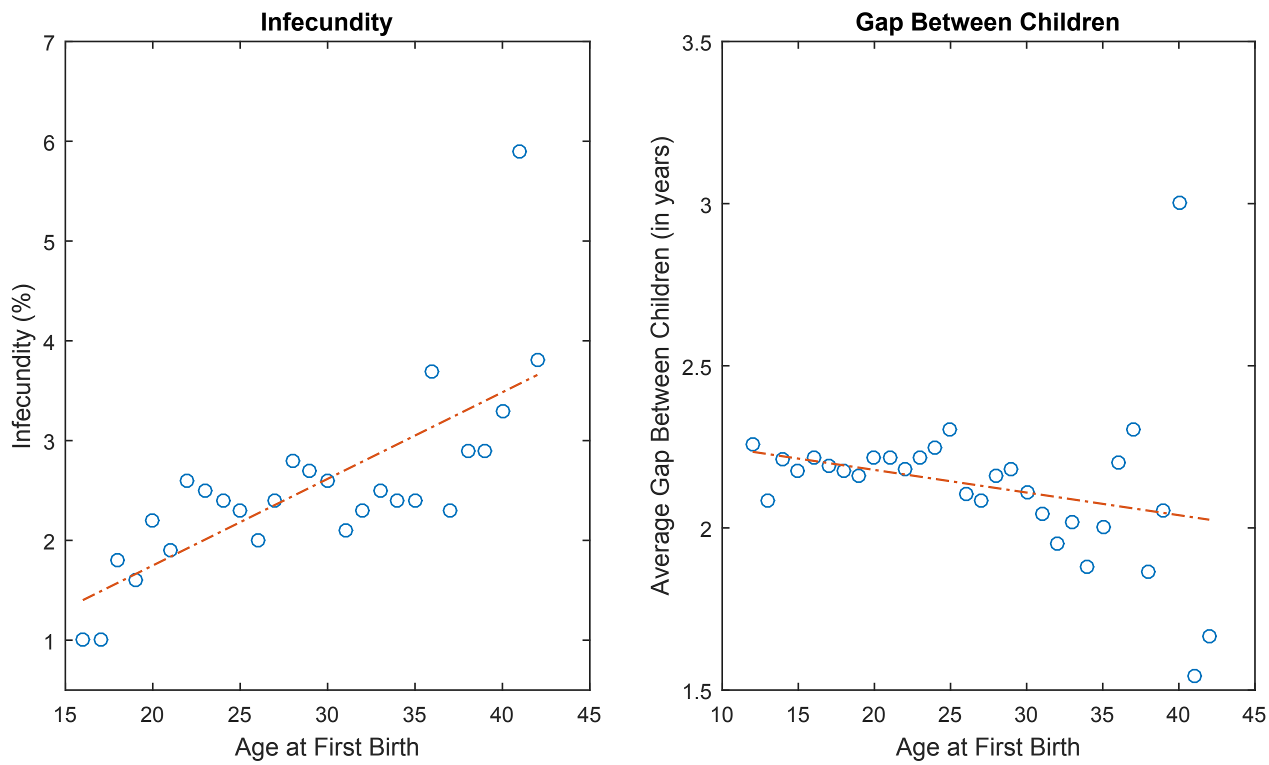
Figure 1: The Relationship Between the Duration of Breastfeeding and First Motherhood Timing



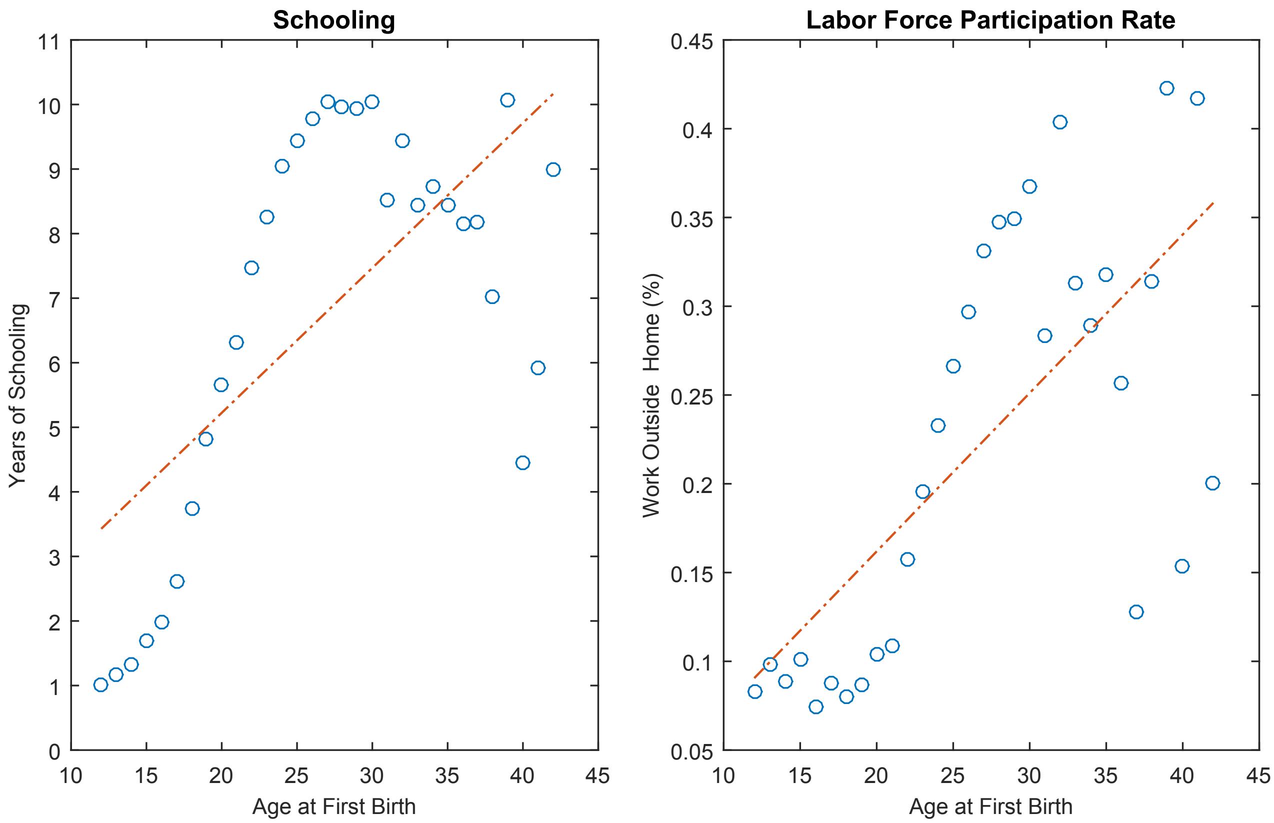
Notes: Each point represents the average observed duration of breastfeeding among mothers who had their first birth between the ages of 13 to 43. We limit the first motherhood timings at 43 because of limited observations beyond this age. The dashed lines in both panels represent the best linear fit.

Figure 2: The Relationship Among Infecundity, Average Gap Between Children and First Motherhood Timing



Notes: $Infecundity$ is defined as when a woman fails to achieve pregnancy over twelve successive months of attempts. $Gapbetweenchildren$ is the average number of years between all successive children. The dashed lines in both panels represent the best linear fit.

Figure 3: The Relationship Among Schooling, Labor Force Participation Rate and First Motherhood Timings



Notes: The left panel shows the average years of schooling based on the mother's age at the first birth. The right panel shows the labor force participation status of the mothers who had their first motherhood timings between the ages of 13-43. This variable is defined as a percentage of women who work outside the home at a given age at the first birth. The dashed lines in both panels represent the best linear fit.

Table 1: Summary Statistics (Mean and Standard Deviation)

	All	Age at First Birth				
		13 to 17	18 to 21	22 to 25	26 to 32	33 to 45
Observed Duration of Breastfeeding	13.473	14.734	13.492	12.861	12.494	12.098
(in months)	(8.474)	(8.789)	(8.448)	(8.189)	(8.303)	(8.299)
Completed Duration of Breastfeeding	20.766	21.785	21.054	20.090	19.413	18.474
(in months)	(12.006)	(11.619)	(11.968)	(12.111)	(12.358)	(12.614)
Gap Between Children	2.190	2.191	2.191	2.224	2.109	2.020
(in years)	(0.823)	(0.796)	(0.828)	(0.839)	(0.822)	(0.779)
Number of Children	3.133	4.059	3.165	2.665	2.370	1.911
	(1.838)	(2.089)	(1.830)	(1.483)	(1.248)	(0.952)
Ideal Number of Children	3.001	3.317	3.027	2.858	2.728	2.627
	(1.325)	(1.524)	(1.306)	(1.206)	(1.176)	(1.181)
Mothers' Schooling	5.838	2.108	5.118	8.375	9.823	8.296
(in years)	(5.935)	(3.419)	(5.294)	(6.169)	(6.465)	(6.464)
Fathers' Schooling	7.453	4.591	6.937	9.411	10.395	8.937
(in years)	(5.779)	(4.864)	(5.413)	(5.698)	(5.957)	(6.189)
Age at First Marriage	19.005	14.644	17.856	21.414	25.522	30.230
	(3.968)	(1.418)	(1.646)	(2.073)	(3.009)	(5.756)
Rural	0.621	0.778	0.676	0.511	0.370	0.371
(%)	(0.485)	(0.416)	(0.468)	(0.500)	(0.483)	(0.483)
Muslim	0.798	0.810	0.800	0.789	0.783	0.777
(%)	(0.402)	(0.392)	(0.400)	(0.408)	(0.412)	(0.416)
No of Observations	70,295	14,365	29,882	17,711	7,427	682

Notes: The table reports the means and standard deviations (in parentheses) of the EDHS samples for 1988, 1992, 1995, 2000, 2004, 2008 and 2014. The completed duration of breastfeeding is estimated via a hazard model. "Gap Between Children" is the average number of years between births, and "Ideal Number of Children" is the average ideal family size given by the respondents in the surveys. "Mothers' schooling" and "Fathers' schooling" refer to the average number of schooling years completed by the mothers and fathers in the surveys, respectively.

Table 2: Descriptive Statistics of the Health by Age and Work Status by First Motherhood Timing

		Age				
	All	13 to 17	18 to 21	22 to 25	26 to 32	33 to 45
Health Related Variables of Mothers						
Miscarriage	0.058	0.055	0.056	0.061	0.064	0.047
	(0.234)	(0.229)	(0.230)	(0.240)	(0.245)	(0.212)
Infecundity	0.026	0.009	0.019	0.024	0.024	0.031
	(0.159)	(0.096)	(0.138)	(0.153)	(0.154)	(0.173)
No of Observations	70,295	435	6,917	17,423	27,427	17,728
		Age at First Birth				
	All	13 to 17	18 to 21	22 to 25	26 to 32	33 to 45
Work-Related Variables of the Mothers						
Skilled occupation	0.086	0.010	0.036	0.150	0.270	0.242
(%)	(0.281)	(0.099)	(0.187)	(0.357)	(0.444)	(0.429)
Working	0.147	0.087	0.095	0.203	0.328	0.290
(%)	(0.354)	(0.282)	(0.293)	(0.403)	(0.470)	(0.454)
Working full time	0.099	0.043	0.057	0.149	0.244	0.249
(%)	(0.298)	(0.203)	(0.231)	(0.356)	(0.430)	(0.433)
No of Observations	70,295	14,365	29,882	17,711	7,427	682

Notes: The table reports the means and standard deviations (in parentheses) of the DHS samples in Egypt reported in Table 1. "Miscarriage" denotes the first pregnancy loss before the first birth. "Infecundity" is defined as when a woman fails to achieve pregnancy over twelve successive months of attempts. "Skilled occupation", "Working", and "Working full time" refer to the percentage of mothers in the sample who have a skilled occupation, work outside the home, and work full time outside the home, respectively.

Table 3: Probit Model Estimates of Health and Labor Market Outcomes by Age at First Pregnancy

	Miscarriage		Work Outside		Work Full Time		Work in Skilled Jobs	
	Number of Children = 1	Number of Children ≥ 2	Number of Children = 1	Number of Children ≥ 2	No of Children Children = 1	Number of Children ≥ 2	Number of Children = 1	Number of Children ≥ 2
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
First Pregnancy Age 19-21	0.012	0.005**	0.007	0.008	0.011	0.011**	0.028***	0.021***
	(0.009)	(0.003)	(0.011)	(0.005)	(0.010)	(0.005)	(0.010)	(0.003)
First Pregnancy Age 22-25	0.020*	0.017***	0.075***	0.079***	0.072***	0.062***	0.086***	0.064***
	(0.011)	(0.004)	(0.014)	(0.007)	(0.013)	(0.006)	(0.014)	(0.005)
First Pregnancy Age 26-45	0.258***	0.348***	0.170***	0.100***	0.157***	0.090***	0.177***	0.084***
	(0.020)	(0.011)	(0.019)	(0.007)	(0.019)	(0.007)	(0.022)	(0.006)
All Other Controls	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Age-month FE	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Birth-year FE	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Region FE	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Pseudo R^2	0.133	0.357	0.187	0.129	0.216	0.163	0.332	0.343
No of Observations	7,496	36,196	7,489	36,199	7,113	32,090	7,475	36,199

Notes: The table reports the marginal effects from a probit model on the health and labor market outcome variables in the four vertical panels. The dependent variables are "Miscarriage", "Work Outside", "Work Full Time", and "Work in Skilled Jobs". These estimates are obtained from a mother-level sample where each observation represents a unique mother. The binary outcome variable, miscarriage, denotes whether a mother had a miscarriage before the first birth. The baseline group consists of mothers who became first time pregnant in between ages 13 and 18. The set of control variables includes dummy variables for mothers' four education category variables (less than primary, primary, secondary, and higher education) and other demographic variables, including the children's gender, mothers' religion, dummy for female head of household, fathers' four education category variables, mothers' weight during the survey, children's weight at birth, mothers' body mass index, number of children living with the mothers, and a rural area dummy. All regressions include 50 age-month of birth, 32 birth-year and 22 region fixed effects. Robust standard errors are in parenthesis.

Table 4: Determinants of the Breastfeeding Duration of the First Child

	OLS	2SLS	IIV Bound		Hazard Model	
	(1)	(2)	(3)	(4)	(5)	(6)
$AgeFirstBirth$	-0.353***	-0.839***	-1.128***	-0.502***	0.062***	0.124***
	(0.050)	(0.305)	(0.422)	(0.093)	(0.007)	(0.033)
$MilkSubst$	-1.172***	-1.145***	-1.129***	-1.164***	0.627***	0.619***
	(0.099)	(0.101)	(0.103)	(0.099)	(0.021)	(0.023)
$Education-Primary_m$	-0.592**	-0.529**	-0.502**	-0.562**	0.098**	0.092**
	(0.249)	(0.250)	(0.254)	(0.249)	(0.038)	(0.039)
$Education-Secondary_m$	-0.408**	-0.288	-0.220	-0.367*	0.141***	0.127***
	(0.193)	(0.207)	(0.222)	(0.194)	(0.030)	(0.031)
$Education-High_m$	-0.008	0.232	0.377	0.064	0.225***	0.177***
	(0.293)	(0.330)	(0.366)	(0.296)	(0.048)	(0.054)
$MotherWorking$	0.217	0.282	0.322	0.235	-0.030	-0.032
	(0.219)	(0.226)	(0.232)	(0.220)	(0.032)	(0.032)
$Gender$	-0.366***	-0.372***	-0.378***	-0.365***	0.117***	0.114***
	(0.134)	(0.135)	(0.136)	(0.134)	(0.021)	(0.021)
$Religion$	0.651**	0.625**	0.614*	0.637**	-0.143***	-0.154***
	(0.307)	(0.310)	(0.313)	(0.308)	(0.047)	(0.047)
$HouseholdHead$	0.200	0.187	0.185	0.190	-0.028	-0.027
	(0.285)	(0.287)	(0.289)	(0.285)	(0.041)	(0.041)
$Education-Primary_f$	0.142	0.155	0.167	0.140	0.068*	0.067*
	(0.263)	(0.264)	(0.266)	(0.262)	(0.037)	(0.038)
$Education-Secondary_f$	0.304	0.324	0.339	0.307	-0.004	-0.006
	(0.229)	(0.232)	(0.233)	(0.229)	(0.036)	(0.037)
$Education-High_f$	-0.368	-0.268	-0.212	-0.334	0.039	0.023
	(0.308)	(0.316)	(0.322)	(0.308)	(0.049)	(0.050)
$TotalChildren$	-1.026***	-1.168***	-1.255***	-1.066***	0.435***	0.530***
	(0.188)	(0.213)	(0.230)	(0.190)	(0.020)	(0.043)
$Rural$	0.614***	0.532***	0.478**	0.594***	-0.059**	-0.050*
	(0.183)	(0.190)	(0.198)	(0.184)	(0.029)	(0.030)
Age-month FE	Yes	Yes	Yes	Yes	No	No
Birth-year FE	Yes	Yes	Yes	Yes	Yes	Yes
Region FE	Yes	Yes	Yes	Yes	Yes	Yes
Adjusted R^2	0.334	0.328	0.317	0.335		
No of Observations	13,488	13,248	13,248	13,248	12,161	11,941

Notes: The dependent variable is the duration of breastfeeding of the first child in months. $AgeFirstBirth$ is the mother's age at the first birth. $MilkSubst$ is the percentage share of breastmilk substitute imports of the total imports standardized by the mean and standard deviation. $Education-Primary_m$, $Education-Secondary_m$, and $Education-High_m$ are dummy variables equaling one if the mother has a primary, secondary, or higher education, respectively. $MotherWorking$ is a dummy variable equaling one if the mother works outside the home. $Gender$ is a dummy variable equaling one if the gender of the first birth child is female. $Religion$ is a dummy variable equaling one (zero) if the mother is a Muslim (Christian). $HouseholdHead$ is a dummy variable equaling one if the mother is the head of the household. $Education-Primary_f$, $Education-Secondary_f$ and $Education-High_f$ are dummy variables equaling one if the father has a primary, secondary, or higher education, respectively. $TotalChildren$ is the total number of living children given birth by the mother. $Rural$ is a dummy variable equaling one if the mother lives in a rural area. All regressions include 50 age-month of birth, 32 birth-year and 22 region fixed effects. Robust standard errors are shown in parenthesis and clustered by the mother. $*$, $**$, and $***$ refer to significance at the 10%, 5%, and 1% levels.

Table 5: Determinants of the Breastfeeding Duration of All Children

	OLS	2SLS	IIV Bound		Hazard Model	
	(1)	(2)	(3)	(4)	(5)	(6)
$AgeFirstBirth$	-0.190***	-0.252***	-1.419***	-0.283***	0.035***	0.044***
	(0.018)	(0.090)	(0.320)	(0.029)	(0.002)	(0.012)
$\Delta\ Ideal$	-0.180***	-0.197***	-0.462***	-0.205***	-0.023***	-0.020***
	(0.039)	(0.044)	(0.084)	(0.040)	(0.005)	(0.005)
$\Delta\ Ideal > 0$	0.294***	0.244*	-0.749**	0.217*	-0.042***	-0.039**
	(0.110)	(0.135)	(0.301)	(0.113)	(0.015)	(0.018)
$MilkSubst$	-0.706***	-0.710***	-0.755***	-0.711***	0.517***	0.516***
	(0.063)	(0.063)	(0.068)	(0.063)	(0.014)	(0.014)
$Education - Primary_m$	-0.445***	-0.442***	-0.376**	-0.440***	0.079***	0.075***
	(0.130)	(0.131)	(0.147)	(0.131)	(0.018)	(0.018)
$Education - Secondary_m$	-0.396***	-0.364***	0.301	-0.347***	0.141***	0.140***
	(0.114)	(0.124)	(0.222)	(0.115)	(0.017)	(0.018)
$Education - Higher_m$	-0.164	-0.065	1.567***	-0.020	0.175***	0.161***
	(0.191)	(0.224)	(0.494)	(0.193)	(0.029)	(0.035)
$MotherWorking$	-0.025	-0.003	0.354**	0.006	0.005	0.003
	(0.116)	(0.119)	(0.164)	(0.116)	(0.016)	(0.016)
$Gender$	-0.405***	-0.401***	-0.364***	-0.400***	0.126***	0.124***
	(0.074)	(0.074)	(0.081)	(0.074)	(0.011)	(0.011)
$Religion$	0.480**	0.469**	0.309	0.465**	-0.105***	-0.105***
	(0.202)	(0.203)	(0.228)	(0.203)	(0.030)	(0.031)
$HouseholdHead$	0.307*	0.310*	0.397*	0.312*	-0.052**	-0.052**
	(0.185)	(0.185)	(0.211)	(0.186)	(0.025)	(0.025)
$Education - Primary_f$	0.108	0.108	0.134	0.109	0.026	0.030*
	(0.132)	(0.132)	(0.145)	(0.132)	(0.017)	(0.017)
$Education - Secondary_f$	-0.091	-0.079	0.140	-0.073	0.036**	0.045***
	(0.125)	(0.126)	(0.155)	(0.125)	(0.017)	(0.017)
$Education - High_f$	-0.374**	-0.356**	0.144	-0.342*	0.085***	0.088***
	(0.174)	(0.179)	(0.240)	(0.175)	(0.024)	(0.025)
$TotalChildren$	-0.085	-0.176	-1.935***	-0.223***	0.093***	0.107***
	(0.056)	(0.147)	(0.484)	(0.064)	(0.007)	(0.018)
$Rural$	0.782***	0.775***	0.501***	0.768***	-0.089***	-0.091***
	(0.112)	(0.114)	(0.145)	(0.112)	(0.016)	(0.016)
Age-month FE	Yes	Yes	Yes	Yes	Yes	Yes
Birth-year FE	Yes	Yes	Yes	Yes	Yes	Yes
Region FE	Yes	Yes	Yes	Yes	Yes	Yes
Adjusted R^2	0.389	0.390	0.271	0.389		
No of Observations	44,951	44,005	44,005	44,005	41,183	40,295

Notes: The dependent variable is the average duration of breastfeeding of all children in months. $\Delta\ Ideal$ is the difference between a child's birth order and the mother's desired number of children. $\Delta\ Ideal > 0$ is a dummy variable equaling one if $\Delta\ Ideal$ is greater than zero. For the other variables, refer to Table 4. All regressions include 50 age-month of birth, 32 birth-year and 22 region fixed effects. Robust standard errors are shown in parenthesis and clustered by the mother. $*$, $**$ and $***$ refer to significance at the 10%, 5%, and 1% levels.

Table 6: Robustness to IV Selection and Fixed Effects

	Miscarriage IV		Birth Month Fixed Effects		10-year Lags for $\Delta\widehat{ABG}$	
	First Child	All Children	First Child	All Children	First Child	All Children
	(1)	(2)	(3)	(4)	(5)	(6)
Age at First Birth	-0.878***	-0.269***	-0.843***	-0.234***	-0.670***	-0.257***
	(0.302)	(0.073)	(0.305)	(0.077)	(0.187)	(0.046)
$MilkSubst$	-1.143***	-0.684***	-1.142***	-0.690***	-1.157***	-0.693***
	(0.101)	(0.063)	(0.101)	(0.063)	(0.100)	(0.063)
$Education - Primary_m$	-0.526**	-0.362***	-0.524**	-0.360***	-0.489*	-0.366***
	(0.250)	(0.123)	(0.250)	(0.123)	(0.256)	(0.125)
$Education - Secondary_m$	-0.279	-0.321***	-0.291	-0.349***	-0.264	-0.306***
	(0.207)	(0.120)	(0.207)	(0.121)	(0.203)	(0.113)
$Education - High_m$	0.252	0.051	0.223	-0.016	0.131	0.019
	(0.329)	(0.223)	(0.329)	(0.225)	(0.310)	(0.200)
Age-month FE	Yes	Yes	Yes	Yes	Yes	Yes
Birth-year FE	Yes	Yes	Yes	Yes	Yes	Yes
Region FE	Yes	Yes	Yes	Yes	Yes	Yes
Adjusted R^2	0.327	0.384	0.328	0.384	0.349	0.408
No of Observations	13,248	49,268	13,248	49,268	12,350	44,737

Notes: The dependent variable is the duration of breastfeeding of the first child and all children in months. Miscarriage shown in columns (1)-(2) includes miscarriage information as an additional IV. Birth-month fixed effects shown in columns (3)-(4) include 12 birth-month fixed effects. 10-year lags $\Delta\widehat{ABG}$ were used rather than five year lags. For the definitions of the other variable, refer to Table 4. All regressions include 50 age-month of birth, 32 birth-year and 22 region fixed effects. Robust standard errors are shown in parenthesis and clustered by the mother. $*$, $**$, and $***$ refer to significance at the 10%, 5%, and 1% levels.

Table 7: Testing the Validity of the Group Clustering for the Regional Growth of First Birth IV

	Group 1		Group 2		Group 3		Group 4	
	First Child	All Children	First Child	All Children	First Child	All Children	First Child	All Children
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
$AgeFirstBirth$	-0.983***	-0.256**	-0.785***	-0.260***	-0.902***	-0.345***	-0.874***	-0.311***
	(0.346)	(0.091)	(0.282)	(0.073)	(0.275)	(0.066)	(0.302)	(0.069)
$MilkSubst$	-1.137***	-0.688***	-1.147***	-0.683***	-1.142***	-0.685***	-1.144***	-0.684***
	(0.102)	(0.063)	(0.101)	(0.063)	(0.101)	(0.063)	(0.101)	(0.063)
$Education - Primary_m$	-0.526**	-0.382***	-0.546**	-0.364***	-0.513**	-0.358***	-0.537**	-0.371***
	(0.252)	(0.123)	(0.251)	(0.122)	(0.250)	(0.123)	(0.251)	(0.123)
$Education - Secondary_m$	-0.258	-0.338***	-0.307	-0.327***	-0.273	-0.273**	-0.284	-0.295**
	(0.213)	(0.123)	(0.206)	(0.119)	(0.204)	(0.117)	(0.207)	(0.118)
$Education - High_m$	0.303	0.009	0.201	0.028	0.261	0.174	0.248	0.112
	(0.343)	(0.237)	(0.324)	(0.221)	(0.325)	(0.214)	(0.332)	(0.218)
Age-month FE	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Birth-year FE	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Region FE	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Adjusted R^2	0.323	0.384	0.329	0.383	0.326	0.382	0.327	0.383
No of Observations	13,241	49,102	13,371	49,817	13,172	48,956	13,166	48,911

Notes: The dependent variable is the duration of breastfeeding of the first child and all children in months. Groups refer to different ages at the first birth and child-birth-year cutoff points. Group 1 includes the <17, 17-19, 20-21, 22-24, and >24 age at first birth clusters and < 1985, 1985-1989, 1990-1994, 1995-1999, 2000-2004, 2005-2009, and > 2009 child birth year clusters. Group 2 includes the <18, 18-19, 20-21, 22-25, and >25 age at first birth clusters and < 1985, 1985-1988, 1989-1992, 1993-1996, 1997-2000, 2001-2004, 2005-2008, and > 2008 child birth year clusters. Groups 3 includes the <18, 18-19, 20-21, 22-25, >25 age at first birth clusters and < 1985, 1985-1990, 1991-1996, 1997-2002, 2003-2008, and > 2008 child birth year clusters. Group 4 includes the <17, 17-19, 20-21, 22-24, and >24 age at first birth clusters and < 1985, 1985-1990, 1991-1996, 1997-2002, 2003-2008, and > 2008 child birth year clusters. For the definitions of the other variables, refer to Table 4. All regressions include 50 age-month of birth, 32 birth-year and 22 region fixed effects. Robust standard errors are shown in parenthesis and clustered by the mother. $*$, $**$, and $***$ refer to significance at the 10%, 5%, and 1% levels.