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When Birth Spacing Does and Does Not Matter for Child Survival: An International Comparison using the DHS

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ABSTRACT

A large body of research has found an association between short birth intervals and the risk of infant mortality in developing countries, but recent work from highly developed countries has called these claims into question, arguing that previous estimates have been biased by a failure to adequately control for unobserved heterogeneity. This study addresses this issue by estimating within-family models on a sample of 4.5 million births from 77 countries at various levels of development. We show that even after controlling for unobserved maternal heterogeneity, intervals less than 24 months substantially increase the probability of infant death, and this relationship is present in all countries in our analysis. We do show, however, that the importance of birth intervals as a determinant of infant mortality varies inversely with maternal education. Finally, we demonstrate that the mortality-reducing effects of longer birth intervals are strong at low levels of development but decline steadily towards zero as populations become healthier and wealthier. These findings offer a clear way to reconcile previous research showing that birth intervals are important for infant mortality in low-income countries, but much less consequential in high-income settings.

Keywords: Birth spacing; infant mortality; developing countries; DHS; maternal depletion; infection

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Introduction

The World Health Organization (WHO) has identified birth interval length as a critical determinant of child mortality risks and has recommended women to space their births between three and five years apart in order to reduce health risks to both children and mothers (WHO 2007). This recommendation is based on the findings that short birth intervals (i.e. shorter than 24 months) and long intervals (i.e. longer than 60 months) are associated with an elevated risk of infant death (Agustín Conde-Agudelo et al. 2012; Hobcraft et al. 1985; Rutstein 2005). The relationship between short birth intervals, in particular, and mortality has been remarkably consistent, having been demonstrated repeatedly in a variety of developmental contexts across time and space (Becher et al. 2004; Cleland and Sathar 1984; Curtis et al. 1993; Miller et al. 1992; Millman and Cooksey 1987; Nault et al. 1990; Palloni and Millman 1986; Pebley et al. 1991; Ronsmans 1996; Whitworth and Stephenson 2002). However, despite the large body of literature supporting these longstanding conclusions, recent evidence to the contrary has called the very existence of the relationship between birth interval length and perinatal outcomes into question (Ball et al. 2014; Class et al. 2017; Hanley et al. 2017). In order to properly identify whether or not birth intervals are in fact an important determinant of perinatal outcomes, it is necessary to confront two significant shortcomings in the current body of literature: a failure to address potential estimation bias from unmeasured confounding and a dearth of international comparisons.

Much of the previous literature on the relationship between birth intervals and infant mortality has not adequately addressed the issue of residual confounding by unobservable characteristics. Endogeneity is always a concern when studying the effects of fertility behavior on children's outcomes (see e.g. Angrist and Evans 1998; Angrist et al. 2010; Rosenzweig and

Wolpin 1980), and this is no different when studying the effects of birth spacing. Unobserved maternal heterogeneity can easily bias estimates of fertility's effects on child health. For example, if certain mothers are less likely to breastfeed, they may be simultaneously more likely to have shorter birth intervals and higher risks of infant mortality. The importance of this issue has recently come to the fore as several studies of mothers in rich countries have shown that, after accounting for unobserved compositional differences between women, birth intervals seem to be inconsequential for children's perinatal outcomes (Ball et al. 2014; Class et al. 2017; Hanley et al. 2017). As a result of this research in high-income settings, it has been questioned whether birth intervals really matter for perinatal outcomes at all (Klebanoff 2017). At the same time, recent research on low-income populations has shown, even after adjusting for unobserved maternal heterogeneity, that birth intervals are still highly consequential for infant mortality in Bangladesh (Molitoris, 2018), and that in historical populations the effects of birth intervals on infant mortality weakened as public health conditions improved over time (Molitoris, 2017).

Because the extant literature is largely comprised of case studies, it has been difficult to determine the extent to which differences between findings have been due to differences in methodologies, sample selection procedures, or contextual factors. The primary goal of this paper is therefore to investigate how the relationship between birth intervals and infant mortality varies across developmental contexts while applying uniform methods that can minimize residual confounding from unobserved heterogeneity. The benefit of a standardized comparative approach is that it will allow us to shed light on both the average effects of birth interval length on infant mortality and also whether the importance of birth intervals varies according to contextual conditions. An international comparison may help us to reconcile the apparently discrepant findings in the literature and provide benchmarks for knowing when increasing birth spacing may or may not be a relevant intervention for reducing infant mortality.

Our study will address the abovementioned issues by using data from 77 countries and over 200 waves of the Demographic and Health Surveys. First, we will account for the probable endogenous relationship between birth spacing and infant mortality by estimating within-family linear probability models. These models can account for unobservable maternal factors, such as maternal health or shared frailty, which may be correlated with both interval length and infant mortality risks. Second, we will then show how the relationship between birth intervals and infant mortality risks varies both within and between populations in order to identify whether specific groups of mothers are driving any observed association. Finally, we will link our estimates of birth intervals' effects on infant mortality to several macro-level indicators of development in order to understand the conditions under which birth intervals are more or less important for child survival.

Birth Intervals and Adverse Outcomes: Mechanisms and Findings

A detailed description of the theoretical mechanisms linking preceding birth intervals to children's outcomes may be found elsewhere (Agustín Conde-Agudelo et al. 2012), but we will briefly outline some of the leading explanations in the literature as they are crucial for understanding why short birth intervals may be detrimental in some contexts, but not in others. It is worth noting that these mechanisms are not mutually exclusive. They include: maternal depletion, infection transmission, and sibling competition.

The maternal depletion hypothesis argues that shorter birth intervals mean that women do not fully physically recuperate from the previous pregnancy, which subsequently results in suboptimal fetal development and a higher risk of mortality for the child born following the short interval (Winkvist et al. 1992). In a context in which food shortage is chronic, continuous, and sustained, a woman's body is known to prioritize its own wellbeing over that of the fetus in distributing energy and nutrients (Ellison 2003; Peacock 1991). Such a physiological response is thought to preserve a woman's potential for future reproduction as well as for lactation. While research continues into what, specifically, is depleted by one pregnancy and not sufficiently restored by the next (e.g. fat, micronutrients, muscle mass), some facts are well understood; for example, short birth intervals lead to folate (vitamin B_9) depletion, which is critical for the growth and development of the fetus (Greenberg et al. 2011).

Infection transmission is the second mechanism that may link birth intervals to infant mortality risks. The horizontal transmission hypothesis holds that closely spaced births will place the younger of the siblings at a greater risk of mortality (Boerma and Bicego 1992). The reasons for this are that the younger sibling will be exposed to a similar set of diseases as the older sibling while also having a less developed immune system, which will increase the ease of transmission from older to younger siblings. The weaker immune system of the latter can also increase the lethality of infectious diseases. There is also some evidence that for certain infectious childhood diseases, like measles, secondary infections acquired by an index child from their older sibling tend to have significantly higher case fatality rates (Aaby et al. 1986; Aaby et al. 1984; Garenne and Aaby 1990).

The final mechanism linking intervals to mortality is sibling competition, which implies that closely spaced children are more likely to compete for the same resources, such as parental time and investment. Generally, competition for most resources would not be so much a result of the interval length *per se* but as a result of an increase in family size, leading to a decrease in parental attention and investment in the first years of life for the index child. However, direct competition for one resource, breastmilk, would be directly related to the length of a birth interval. Some evidence from developing countries suggests that breastfeeding-pregnancy overlap is not uncommon (Boerma and Bicego 1992; Ramachandran 2002), and may result in a lower quality and quantity of breastmilk for the child born following the interval, leading to diminished neonatal growth (Marquis et al. 2002; Marquis et al. 2003).

Until now, we have only discussed mechanisms that would explain why *shorter* preceding birth intervals may cause adverse perinatal outcomes. This focus has been intentional, as the literature on the topic has been overwhelming in showing that shorter intervals are associated with higher rates of mortality, stillbirth, low birth weight, and other poor outcomes. But it is worth mentioning that there is also a smaller literature showing that *long* intervals (i.e. longer than 60 months) are also disproportionately associated with higher risks of adverse perinatal outcomes (Agustín Conde-Agudelo et al. 2006; Zhu et al. 1999). Why longer intervals would be detrimental has not yet been firmly established, but one explanation, 'maternal regression', states that the longer a woman goes without conceiving a subsequent child, the more her physiology (and consequently her perinatal outcomes) resembles that of a primigravid woman (Zhu et al. 1999). Evidence from countries across the developmental spectrum has suggested that women who give birth following long birth intervals experience similar risks for pre-eclampsia and eclampsia as women having their first birth (Agustin Conde-Agudelo and Belizán 2000; Skjærven et al. 2002). Nevertheless, it is important to recognize that the exposure to intervals beyond 60 months is much smaller than the exposure to intervals shorter than, say, 24 months. In developing countries, approximately 25% of births occur within 24 months of the preceding birth, while only about 6% of births occur after 60 months (Rutstein 2005). Short birth intervals therefore are a considerably greater risk in most populations.

That short interbirth intervals are predictive of higher mortality risks is a consistent finding in the literature, but it is not universal. Some recent studies of high-income populations in Sweden, Canada, and Australia have found that, when controlling for unobserved maternal heterogeneity via sibling fixed-effects, short birth intervals did not lead to higher risks of adverse outcomes (Ball et al. 2014; Class et al. 2017; Hanley et al. 2017), suggesting that the apparent relationship between interval length and children's outcomes may be attributable to the non-random distribution of birth intervals across mothers. Nevertheless, other recent research accounting for unobserved maternal heterogeneity has found quite different results. Two studies of poor, high-mortality populations, 19th century Stockholm, Sweden, and contemporary Bangladesh, have shown that shorter birth intervals increased the risk of neonatal, post-neonatal, and child mortality (Molitoris 2017, 2018). Furthermore, the latter two studies presented results that may explain the discrepancy in findings mentioned above. First, the effects of birth interval length on mortality risks decreased over time as the overall level of mortality declined in Sweden (Molitoris 2017). Second, even within a high-mortality context, the size of the effects of interval length on mortality varied inversely with the educational level of the mother (Molitoris 2018).

Taken together, we believe all of these findings may fit into the same picture. Given the mechanisms outlined earlier in this section, one should expect that as economic and epidemiological conditions improve, short birth intervals should become a less important predictor of infant mortality. Maternal depletion, infection transmission, and resource competition should all become relatively less important as the general nutrition and health of the population improves, thereby making birth intervals a weaker determinant of infant mortality until, eventually, they are virtually irrelevant. In order to examine if this is indeed the case, we will apply uniform statistical methods that can account for unobserved heterogeneity to data from a variety of low- to middle-income contexts, and explicitly examine whether the association varies across their respective levels of development.

Data

Demographic and Health Surveys

FIGURE 1 HERE

This study uses data on 77 countries and 207 waves of the Demographic and Health Surveys (DHS) (see appendix table 1A for list of included countries and their respective numbers of cases). The DHS is a household survey, with a separate survey for women aged 15-49. The household response rates in the 207 surveys used in this study range from 83.8-99.9%, with a mean of 97.5% and standard deviation of 2.45%, while the response rates for the woman's questionnaire ranges from 77.0-99.6%, with a mean of 93.6% and standard deviation of 3.92%. Our analyses are based upon the self-reported fertility histories of each woman surveyed. The outcome of interest in this study is infant mortality, defined as mortality between birth and 12 months. We have restricted the pooled data in several ways for our analysis. First, only children born at parities two or higher are included in the analysis, as firstborns have an undefined birth interval. Second, index children born as a set of a multiple birth (e.g. twin, triplet, etc.) were excluded. Third, children with unusually long birth intervals (greater than 10 years) were excluded from the analysis as intervals of this length are highly unusual; 99% of birth intervals are ‘closed’ within 10 years in the data. Fourth, index children must have come from mothers with three or more children. This restriction is necessary as the within-family approach we will adopt later requires at least two birth intervals (i.e. three births) per woman. In total, the final analytical sample included approximately 4.56 million births to over 1.15 million women. Of these children, approximately 370,000 died in the first year of life.

The distribution of birth intervals across the 77 countries is shown in figure 2. The mean birth interval was nearly 35 months (median = 29 months) with a standard deviation of 25 months. The distributions observed here followed typical distributions of birth spacing and were mostly similar across populations. While the majority of populations conformed to the average distribution of intervals, some exceptional populations had unusually large shares of children born after very short birth intervals. For example, 15% of children in Yemen were born following an interval less than 12 months, and nearly a quarter of our countries had more than 50% of their birth intervals below 24-27 months. Such a high prevalence of short intervals is not necessarily indicative of data problems. There are certainly pronounced regional differences in spacing patterns across the developing world (Casterline and Odden 2016). Nevertheless, our results rely heavily on the reliability of the birth histories and in order to be certain that they are not biased due to misreporting of births we will later include two robustness checks in our analysis. First, we will stratify our statistical models by United Nations sub-region to account for differences in the accuracy of birth histories across regions of the world (see Schoumaker 2014). Second, we will omit births occurring more than 10 years before the interview year because of the tendency for women to displace older births when recounting their birth histories (Potter 1977). A 10-year cut-off is adopted because a technical investigation of the DHS has shown that the displacement of births tends to be quite low (about 2% or less) within that time frame (Pullum and Becker 2014). Furthermore, reducing the reference period any further would limit our ability to conduct a within family analysis, as it would force any women included in the sensitivity analysis to have more closely spaced births.

FIGURE 2 HERE

World Bank Indicators of Development

In order to understand how the effects of birth intervals on infant mortality vary according to level of development we link various indicators of development by country-year. We sourced the data from the World Development Indicators database, which is the primary World Bank collection of development indicators, compiled from officially recognized international sources. It presents the most current and accurate global development data available, and includes national, regional and global estimates. In our analyses we particularly focus on GNI per capita, the national infant mortality rate, and e_0, estimated life expectancy at birth.

Methods

To analyze the effects of birth spacing on infant mortality, we estimate the following linear probability model:

$$Y_{ij} = S_{ij}\beta_{1,ij} + \mathbf{X}_{ij}\beta_{k,ij} + \theta_j + \varepsilon_{ij} \tag{1}$$

The dependent variable, Y, is binary and indicates whether or not child i of mother j died in the first year of life. Our main independent variable, S, is the length of the preceding interbirth interval (i.e. the time between the birth of the older adjacent sibling and the birth of the index child). We have treated it as a continuous variable with a quartic functional form in order to account for the well-known nonlinear relationship between interval length and mortality risks (Hobcraft et al. 1985; Rutstein 2005). Because a major goal of this paper is to provide comparable estimates across many populations, we have adopted parsimonious models that control for basic demographic characteristics that may vary across siblings. The controls, **X**, include the sex of the index child, (centered) birth year, survival status of the previous child, and birth order. Summary statistics of the model's covariates may be found in table 1.

TABLE 1 HERE

Most previous studies on this topic in low-income countries have not dealt with the fact that birth interval length is most probably an endogenous regressor when studying its effects on infant health. This is because interval length may be correlated with a host of characteristics that may be unobserved, such as maternal breastfeeding preferences or health behaviors, and may themselves influence the probability of infant mortality. Recent work has called attention to the importance of accounting for unobserved factors that may bias estimates of the effect of birth spacing on child outcomes (Ball et al. 2014; Barclay and Kolk 2017; DaVanzo et al. 2008; Molitoris 2017, 2018). We therefore partition the error term into a mother-specific component, θ, and an individual-specific component, ε, by subtracting the within-mother means of all variables from their observed values. This essentially allows us to estimate within-family models by controlling for sibling fixed effects (FE). Thus, our models compare children born to the same mother, which means our results should not be driven by unobserved, time-invariant differences between mothers that correlate with interval length, such as religious affiliation, ethnicity, or, insofar as it is time-invariant, socioeconomic status, amongst other factors.

The within-family approach is not without limitations, however. First, we are not able to control for any source of endogeneity that emerges as a result of *time-varying* unobserved heterogeneity. With that in mind, our modelling strategy does, however, offer a more robust control strategy than has generally been applied. Second, the within-family approach necessarily restricts our analysis sample to only include women with three or more births. Because we are studying high-fertility populations, the problem this restriction poses for the generalizability of the findings is not severe. In our sample, 23% of parous women had a completed family size of two or fewer children, a low share compared to that of a post-transitional population, like modern Sweden, for example, where the same restriction would prohibit us from drawing inference from about 70% of the parous female population. Considering this issue from the child's perspective, the limitations to generalizability seem to be even less severe as one-child sibling groups do not contribute any observations to the universe of birth intervals, while two-child sibling groups contribute only one birth interval. In contrast, a three-child sibling group contributes twice as many birth intervals to the universe of birth intervals as a two-child group, and a four-child group three times as many, and so on. Given the high fertility in our data, we calculate that by focusing on sibling groups with at least three children we include 91.5% of the measurable birth intervals in the surveys.

In our analysis, we will first compare the between-family estimates (OLS) to the within-family estimates (FE) using the pooled sample of 77 countries and 207 waves to identify whether or not the relationship between preceding interval length and infant mortality persists after minimizing residual confounding from maternal heterogeneity. We will then proceed to stratify the sample by UN sub-region and maternal education in order to identify if the relationship varies between or within populations. This exercise can be valuable, as it can highlight whether the aggregated patterns are being driven by a few exceptional parts of the world, and reveal whether infant mortality is more vulnerable to birth spacing in some groups than others. Based on the theoretical mechanisms described previously, we would, in fact, suspect that children born to women with less education would be more vulnerable to infection or resource scarcity than those born to more highly educated women. Recent evidence from Bangladesh has indeed shown this to be the case (Molitoris 2018), and it is important to identify if this finding is generalizable to the rest of the world, as more precise targeting of vulnerable groups by family planning programs may be required in order to offset recent funding cuts to international aid organizations (Bingenheimer and Skuster 2017; Starrs 2017).

After estimating these models, we then adopt a comparative perspective. Once again using the within-family approach, we estimate the effects of birth intervals on mortality for each country-cohort combination in the pooled DHS sample. In practice, we estimate separate models for each country and include an interaction term between the preceding birth interval and the birth year of the index child. We then estimate the effect of increasing the interval from 12 to 24 months on infant mortality for each birth cohort with at least 30 observations in each country. The estimates are then linked to World Bank data to understand whether the effects of birth intervals vary according to the level of development, proxied using data on the life expectancy at birth, infant mortality rate, and gross national income (GNI) per capita for each country-birth cohort combination. These indicators were chosen because they are among the most consistently available pieces of information across countries and years, and because they serve as good general proxies for social and economic development.

Results

Controlling for Unobserved Heterogeneity

To begin our analysis, we first estimate the model described in equation (1) with and without controlling for sibling fixed effects. To facilitate the discussion of the results, we present the results graphically as predicted probabilities, but the full output of the models may be found in the appendix. Figure 3 shows the predicted probabilities of infant mortality by the length of the preceding birth interval for the between-family (OLS) and within-family (FE) models. The probabilities were estimated while holding all other variables at their means. Both the between- and within-family models provided fairly similar estimates on the effects of short birth intervals, pointing towards the same substantive conclusions: when intervals are shorter than about 24 months, increasing the length of the birth interval reduces the probability of infant mortality substantially. The only significant difference between the estimated effects emerged at longer birth intervals. The estimates from the between-family models suggest that the risk of infant mortality plateaus once intervals reach between about 36 and 48 months in length. The within-family estimates, on the other hand, diverged at this point. They showed that the probability of infant mortality actually continued to decline as intervals became longer, albeit at a much slower pace. In other words, the marginal benefit of increasing a birth interval when the interval was already greater than about 36 months was fairly small, whereas increasing the length of an interval shorter than 36 months would be highly beneficial in terms of infant mortality risks. It is also worth highlighting here that, in spite of the WHO recommendation for optimal spacing between three and five years, we find no evidence of an *increase* in mortality risks as birth intervals get longer.

FIGURE 3 HERE

Identifying Regional and Socioeconomic Variation

As described earlier, the models were then stratified according to 13 UN sub-regions and the mother's highest level of education in order to explore heterogeneity in the relationship between birth intervals and infant mortality risks. Figure 4 shows the estimates from the models stratified by UN sub-region. Regardless of region, birth intervals less than about 24 months were uniformly associated with a significantly higher risk of infant mortality. When we compare regions in terms of the percentage change in infant mortality associated with increasing birth intervals from 12 to 24 months in length, the smallest relative improvements in mortality were seen by the populations of Western, Middle, and Eastern Africa. In those populations, increasing birth intervals from 12 to 24 months was associated with about a 30% reduction in infant mortality risks. This is clear from the figure, which shows a more gradual decline in mortality risks as intervals grow longer in those populations. On the other hand, the populations with the largest relative decrease in infant mortality for the same increase in spacing were those in Western and Central Asia, Northern Africa, Central America, all of which showed an expected reduction of about 50% when increasing intervals from 12 to 24 months. Put differently, it appears that in some regions, the 'optimal' spacing for maximizing child health is considerably longer than in others. In regions like South and Eastern Europe or the Americas, the benefits of increasing birth intervals beyond even 24 to 36 months seem negligible. Once reaching intervals of that length, the mortality risk more or less remains constant. Yet in Eastern and Western Africa as well as Southern Asia, there appears to be a nearly linear negative relationship between birth interval length and infant mortality risks. That is, the longer the interval, the better the chances of infant survival.

FIGURE 4 HERE

Interestingly, the variation just described above in the regional comparison also resembles the variation we observed in the comparison between educational groups (see figure 5). Among women of any level of education, children born after intervals shorter than 24 months have an elevated risk of infant mortality. Yet the size of the mortality penalty for children born following shorter intervals varied inversely with a woman's level of education. Children born to women with no education had a probability of dying between 0.07 and 0.18 if they were born following an interval of 12 to 36 months. These probabilities declined as maternal education increased. Among women with a tertiary education, children born following the same interval lengths had probabilities of dying between 0.02 and 0.06. Just as in the regional comparison, the point of diminishing returns to further spacing differed across educational groups. For women with no education, we can see the same pattern that characterized some of the least developed regions on earth: a nearly linear negative relationship between interval length and the probability of infant mortality. Children born to women with at least a primary education had a different pattern, in which the probability of dying declines until intervals reach about 36 months in length, after which the mortality risk plateaus. This pattern is also evident for women with secondary and tertiary education, with the only difference being the point at which mortality risks flatten out; at higher levels of education, the risks plateau at shorter interval lengths. It is worth noting here that stratifying our models by the education of the mother will necessarily mean that the underlying populations being represented by each model are changing, which may partially explain why women with the patterns of women with low education resemble the least developed regions. For example, when we consider the group of women with tertiary education, they will be disproportionally drawn from more developed regions, where the relationship between spacing and mortality may be less dramatic. Although we do estimate within-family models that can implicitly control for regional differences, it is important to recognize that stratifying the model does change the analysis samples underlying characteristics, making it difficult to be certain the extent to which the changing nature of the relationship is due to differences between educational groups and differences between regions.

FIGURE 5 HERE

Comparing the Effects of Spacing across Levels of Development

The final part of our analysis will compare how the effects of spacing vary across stages of demographic and economic development. To do this, we have estimated similar FE linear probability models as in equation (1) but have included an interaction term between the length of the preceding birth interval and the birth year of the index child. These models were estimated separately for each country. We then estimated the effect of increasing a birth interval from 12 to 24 months in length in each birth cohort of each country. We check the sensitivity of these results by estimating the effects of increasing intervals from 18 to 30 and 24 to 36 months as well and the substantive conclusions remain the same (see appendix). This procedure effectively allows us to generate over 3000 data points that can be plotted against the three aforementioned development indicators: the infant mortality rate, life expectancy at birth for both sexes, and the GNI per capita (Figure 6). It is important to note that the estimated effects have been scaled to reflect a percentage change in the respective probabilities of dying before age one (${}_{1}q_{0}$) in each country-cohort combination in order to allow for comparison across years and populations. The vertical axis can therefore be interpreted as the expected percentage change in the probability of dying if a birth interval increased from 12 to 24 months in a specific country and cohort. All plots were then fitted with a cubic spline trend.

FIGURE 6 HERE

Panel (a) first plots the effects against the infant mortality rate (IMR). At levels of IMR over about 100 infant deaths per 1,000 live births, increasing a birth interval from 12 to 24 months would reduce the probability of dying before age one by about 50%, on average. Such large effects are persistent until the IMR falls well below 100. After the IMR reaches about 80 deaths per 1,000, the protective effect of increasing a birth interval from 12 to 24 months begins to weaken. At levels of IMR around 30 per 1,000 and lower, the effect of increasing intervals from 12 to 24 months becomes statistically indistinguishable from zero.

Panel (b) tells a very similar story. At low levels of life expectancy at birth (e_0), the marginal effect of increasing intervals would reduce the probability of dying by around 50%. As the general level of mortality declines in the population and e_0 rises, we again see a weakening of the importance of birth intervals for infant survival. In this instance, when e_0 increases beyond about 55 years, extending short birth intervals will decrease the probability of infant death less and less. Eventually, as e_0 reaches just over 70 years, the relationship becomes statistically indistinguishable from zero. Thus panels (a) and (b) both reveal a pattern in which the mortality-reducing effects of birth spacing are only visible in populations with a high level of mortality. Once mortality has decreased significantly, we can no longer identify any effect of birth spacing on infant mortality risks.

In the final plot, panel (c), we view the effects of longer intervals against the level of wealth in a population using information on GNI per capita. Here, we have taken the natural logarithm of GNI per capita and marked the World Bank thresholds for low, low-middle, and upper-middle income countries to facilitate interpretation. At low levels of GNI, the effects of increasing birth intervals are clearly the largest. Again, they are on average at about 50%. Relatively small improvements in GNI appear to be associated with reductions in the importance of birth interval length for infant mortality. Within the low-income category, small improvements in wealth bring the estimated effect of longer intervals from a 50% reduction to a 30% reduction. After this point, however, further increases in national wealth do not seem to change the nature of the relationship between birth intervals and mortality. Countries in the low-middle income and upper-middle income categories have, on average, the same marginal benefit from increasing intervals from 12 to 24 months as those at the upper limit of the low-income category. It should be pointed out, however, that these categories are much more sparsely populated than the low-income category. As a result, any conclusions that we are able to draw regarding these groups must be more tentative.

Supplementary Analyses

In addition to our main results, we have conducted several supplementary analyses to both further explore heterogeneity in our findings and check the robustness of our results (see appendix). First, we stratify the models by a woman's children ever born (CEB) and also index children's birth cohorts. The goal of these exercises is to identify how applicable our findings are for small versus large families and also to identify if the patterns observed until now are driven exclusively by older birth cohorts, or are still an ongoing phenomenon. Then, we restrict our analysis to a subsample of births occurring within ten years preceding the survey in order to account for the possible displacement or omission of births from women's self-reported birth histories (Potter 1977). It has been shown that the displacement of births is about 2% or less within that time frame in the DHS (Pullum and Becker 2014). Next, we again estimate our models using two different subsamples of the data, one which included only even-parity births and one which included only odd-parity births from families of five or more. We examine these subsamples because our analysis included a control for the death of the preceding child, which, in a within-family framework, allows the death of a single child to contribute to the variance of both the dependent variable and independent variable. To be sure that this is not affecting our results in unanticipated ways, we re-estimated the models on a subsample of children whose deaths cannot themselves enter into to the estimation as both dependent and independent variables. In other words, the indicator for the death of a previous child in our model will never represent the death of one of the siblings included in the analysis. The focus on families with five children or more is simply for comparability between the two subsamples, as the within-family framework requires at least two observations per family with a defined preceding interval, and the first odd-numbered parities that meet that criterion are parities three and five. Finally, we checked the robustness of our comparative results by estimating the effects of increasing a birth interval at different interval lengths. The original analysis (see figure 6) only compared the effects of increasing an interval from 12 to 24 months across populations. In this supplementary analysis, we perform the same procedure but comparing the effects of increasing an interval from 18 to 30 and 24 to 36 months on infant mortality risks.

When stratifying our models by a woman's total children ever born (CEB), the results show that as a woman's CEB increases, the relationship between interval length and infant mortality becomes more apparent. In families of all sizes we could see a negative relationship between interval length and mortality risks, but the differences were smallest in three child families. As family size grew larger, the characteristic shape of the relationship between intervals and infant mortality emerged. In all family sizes above three, when intervals were shorter than 24 months, significant improvements in mortality risks could be gained by increasing spacing. From 36 months and above, we again could see the diminishing returns to lengthening intervals even further. The only exception to this pattern was among mothers with three CEB. In those families, there was a more or less linear decline in mortality risks as intervals grew longer.

Stratifying the analysis by birth cohort generated similar results to those found in the main analysis. Regardless of period of birth, we again see the characteristic pattern of high mortality following intervals shorter than 24 months. The difference between the cohorts was that in earlier ones, we saw a virtually linear decline in mortality as intervals grew longer whereas in the later cohorts, we see evidence of a plateauing of the mortality risk after intervals reach about 36 months.

We then estimated the models for our three subsamples: (1) index children born within ten years preceding the survey, (2) children from five-child families or larger born at even parities, and (3) children from five-child families or larger born at odd parities. The estimates and substantive findings based on these sub-samples were generally in line with the main findings. The one difference that emerged in the subsample of children born within the ten years preceding the survey was that we no longer found a continued decline in the probability of dying as intervals grew longer. Instead, when intervals reached about 48 months in length, there was a virtual flattening of the mortality risk at any interval length thereafter. Based on the results presented in all analyses so far, including the stratified and subsample analyses, we can conclude that there are certainly diminishing returns to increased spacing beyond 36 to 48 months. The question that remains is whether or not the marginal benefit of further spacing at longer interval lengths is still present but small or if it actually becomes zero.

Finally, turning to the comparative analysis, the substantive findings of the robustness checks were similar with the original analysis, though there were some differences. When comparing the marginal effects across levels of GNI, the pattern was virtually identical regardless of which reference interval length was used. When plotting against infant mortality and life expectancy, the results were slightly different than those of the original analysis. We do not find a complete disappearance of the mortality-reducing effect of increasing birth intervals at higher levels of life expectancy or lower levels of IMR, although we do see a substantial weakening. It is important to keep in mind, however, that all of the findings in this paper have indicated that the substantial changes in mortality risks due to changing birth interval lengths have been almost exclusively driven by intervals less than two years in length. In other words, the main mortality-reducing effect of increasing birth intervals applies to those children born less than two years after their older sibling, and based on the previously discussed mechanisms, it is this specifically that should be expected to change according to the context.

Discussion

There are several important findings in this study. First, we have shown that the relationship between birth interval length and infant mortality persists even after applying a within-family methodology to account for unobserved heterogeneity between mothers. We found that the probability of dying increases greatly as intervals fall below 24 to 36 months and this pattern was highly consistent across regions of the world. Second, we find no evidence that intervals longer than 60 months are associated with an elevated probability of dying. On the contrary, the evidence presented here suggests that the probability of dying either plateaus or continues to decline, albeit at a slower pace, as birth intervals get longer. Finally, and most significantly, the results from our international comparison show that the importance of birth spacing as a determinant of infant mortality declines at more advanced levels of development. These findings have a number of important implications.

First, this was the first international comparative paper to apply a within-family approach to analyze the effects of birth spacing on infant mortality risks. In contrast to recent studies using the same approach to analyze populations from more developed countries, our study finds that birth spacing indeed does have significant implications for infant survival. Because we have adopted a within-design, this pattern cannot be explained by unobserved heterogeneity between mothers. Regardless of how we stratified the analysis or limited the sample, the same pattern remained: short preceding birth intervals, especially those shorter than 24 months, were significantly associated with a higher probability of infant mortality.

Second, our results only partially support the WHO recommendation for spacing births between three and five years apart. We showed that the largest improvements in the probability of survival consistently come from increasing spacing until at least 36 months. Where our findings differ from the current recommendation is that we find little evidence that *longer* birth intervals will be detrimental for infant mortality. In most of our analyses, the probability of infant mortality either stagnates once intervals reach about 36 to 48 months in length or even continues to decline as intervals grow longer. In some of the UN sub-regions, we find evidence of a reversal in mortality risks followed by a continued decline, but these estimated increases are often statistically indistinguishable or so slight as to be of little practical significance. Furthermore, the sole instance in which a reversal of mortality risks was evident, Western Asia, the reversal occurred already by 30 months. Thus, while our results certainly support the idea of diminishing returns to longer spacing for mitigating infant mortality risks, they do not consistently support any ‘upper bound’ for safe spacing.

Perhaps the most significant implications of our findings, however, come from the comparative portion of our analysis. We showed that as the level of development increases, as measured by infant mortality levels, life expectancy at birth and, to a lesser extent, GNI per capita, the average beneficial effect of increasing a birth interval from 12 to 24 months approaches zero. This finding was entirely consistent with the variation we observed within populations which showed that birth intervals became less consequential for infant mortality as maternal education increased. These findings are consistent with recent work on other high-mortality populations. A recent study of 19th century Sweden showed that the effects of birth intervals on mortality declined as overall mortality levels declined (Molitoris 2017), and a study of neonatal mortality risks in Bangladesh showed that the effects of spacing on mortality were weaker as maternal education increased (Molitoris 2018). Although we found that the marginal effect of birth interval length on infant mortality declined at lower levels of aggregate mortality, it varied less by the level of national wealth. There was a significant mortality-reducing effect of lengthening birth intervals at the absolute lowest levels of GNI per capita, and the effect only weakened slightly at higher levels of national wealth. However, it is important to note that we had relatively few populations in our sample that would be considered above the World Bank's "Low Income" category, so our conclusions regarding the relationship between the wealth of a society and the effects of spacing on mortality must be more tentative.

Finally, because we have shown that the strength of the relationship between birth interval length and infant mortality declines and disappears as mortality falls, the comparative results here help to reconcile the differences in findings reported elsewhere. Recent research using data from high-income, low-mortality populations such as Australia, Sweden, and Canada had cast doubt on the importance of interpregnancy intervals for perinatal outcomes such as preterm birth and low birth weight (Ball et al. 2014; Class et al. 2017; Hanley et al. 2017). These studies had also applied the same sibling fixed effects approach used in this study in order to account for unobserved heterogeneity. Consequently, it was unclear if the discrepant findings in those studies was due to differences in methodologies, data, or context. Based on our comparative findings, it seems to be the latter. The null results from highly developed contexts are entirely consistent with the patterns observed in low-income contexts. As development progresses, birth intervals become less significant for child health. Considering the causal mechanisms involved with this relationship, it would indeed be a surprise to find that birth intervals are significant for infant survival in contexts where infant mortality is extremely rare. In such populations, the average level of nutrition is high and the burden of infectious diseases is low. Furthermore, a wide availability of both ante- and postnatal medical interventions can save many vulnerable young lives. In poor, less-healthy populations, however, where childhood stunting and wasting may be common, infectious disease is prevalent and access to any modern medical care may be limited, infant mortality may be more sensitive to all inputs, including factors such as birth spacing.

Our study does have some important limitations to consider. Because our estimates are based on within-family models, we have been able to show that the relationship between birth intervals and infant mortality in low-income contexts is not attributable to time invariant compositional differences between women. Nevertheless, our approach cannot remove the influence of *time-varying* unobserved heterogeneity that can be correlated with both interval length and infant mortality risks, which has the potential to bias our estimates. Examples of unobserved time-varying factors might include negative shocks to maternal health or socioeconomic resources in the household. We should also acknowledge that the DHS data are based upon self-reported fertility histories, and there will undoubtedly be a certain degree of measurement error in our data (Pullum and Becker 2014; Schoumaker 2014). However, we have checked the sensitivity of our results to misreporting of births, and it appears that the only difference in the findings was that in the restricted sample the estimated probability of dying plateaued after intervals surpassed about 48 months in length instead of continuing to decline. In addition, it is worth noting that the development indicators that we draw from the World Bank refer to the national level of infant mortality, life expectancy at birth, and GNI per capita, and may not closely correspond to the local conditions that the respondents to the survey actually experienced and which might be more important determinants of infant mortality risk.

Nevertheless, we feel that our study also has important strengths. This study is the first to apply a methodology that can account for unobserved heterogeneity in a comparative framework to identify the effects of birth spacing on infant mortality. In doing so, we have confirmed many of the findings of previous research, while also uncovering new details that can help revise general recommendations for birth spacing practices. By adopting a comparative approach, we feel that our study has also helped to reconcile some of the supposed inconsistencies in the current body of literature.

The findings presented here also offer several promising paths forward for future research. First, future research ought to focus more explicitly on identifying the causal mechanisms connecting birth interval length to infant mortality. While our study has sought to identify if the relationship between birth spacing and mortality holds when adopting a robust control strategy, it was beyond its scope to identify which mechanisms facilitate this relationship. In order to explicitly identify the relative importance of the three mechanisms linking short intervals to infant mortality, longitudinal data that includes detailed information on factors such as biomarkers, household spending, and medical information would be required. Second, more comparative work that also includes wealthier populations would help to fill in the gaps regarding why birth intervals seem to matter a great deal in low-income contexts, but much less in high-income contexts. In the present study we were limited by the fact there were no populations that would be considered even moderately wealthy in our dataset. Third, it would be worthwhile to investigate if the relationships between birth spacing and other outcomes are similarly moderated by the level of population health or other development indicators. While short intervals are not associated with higher mortality rates in healthier populations, they may be associated with poor perinatal outcomes, such as low birth weight or preterm birth. Further comparative research may therefore help us to understand the conditions under which birth intervals matter for child health, and those under which they do not.

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Tables and Figures

Table 1. Distribution of covariates used in models.

	N	%	IMR	Mean Interval	St. Dev. Interval
Infant Deaths	369,227	8.1			
Preceding Interval:					
<12	147,128	3.2	210.1	0.84	0.08
12-14	265,978	5.8	156.5	1.09	0.07
15-17	280,031	6.1	124.6	1.34	0.07
18-20	344,059	7.5	104.7	1.59	0.07
21-23	474,945	10.4	92.9	1.84	0.07
24-26	533,548	11.7	81.5	2.08	0.07
27-29	423,655	9.3	75.5	2.33	0.07
30-32	338,117	7.4	67.5	2.58	0.07
33-35	301,998	6.6	59.5	2.83	0.07
36-38	265,019	5.8	53.3	3.08	0.07
39-41	195,262	4.3	51.7	3.33	0.07
42-44	149,155	3.3	50.4	3.58	0.07
45-47	128,310	2.8	44.9	3.83	0.07
48-50	112,670	2.5	41.1	4.08	0.07
51-53	86,132	1.9	41.0	4.33	0.07
54-56	69,504	1.5	40.9	4.58	0.07
57-59	62,972	1.4	38.7	4.83	0.07
60-62	57,423	1.3	37.6	5.08	0.07
63-65	44,476	1.0	38.6	5.33	0.07
66-68	37,336	0.8	37.2	5.58	0.07
69-71	34,062	0.8	36.9	5.83	0.07
72-74	31,416	0.7	36.3	6.08	0.07
75-77	25,078	0.6	36.6	6.33	0.07
78-80	20,942	0.5	36.5	6.58	0.07
81-83	19,617	0.4	38.2	6.83	0.07
84+	115,295	2.5	37.2	8.11	0.83
Sex:					
Male	2,328,349	51.0	85.5	2.73	1.50
Female	2,235,779	49.0	76.1	2.74	1.50
Survival Status of Previously born Sibling:					
Alive	3,868,540	84.8	63.7	2.82	1.51
Died	695,588	15.2	176.5	2.27	1.32
Birth Order					
2	1,140,772	25.0	86.1	2.58	1.40
3	1,149,562	25.2	71.0	2.85	1.60

4	803,513	17.6	75.4	2.81	1.54
5	548,504	12.0	80.4	2.78	1.50
6	368,638	8.1	85.2	2.73	1.46
7	239,333	5.2	90.2	2.70	1.42
8+	313,806	6.9	100.9	2.60	1.35
Maternal education:					
No education	2,094,677	45.9	101.1	2.63	1.36
Primary	1,635,935	35.9	73.2	2.75	1.52
Secondary	713,587	15.6	47.2	2.93	1.70
Tertiary	118,939	2.61	32.1	3.08	1.87
Missing/Unknown	990	0.02	82.8	2.66	1.47
UN Sub-region:					
Caribbean	166,187	3.6	61.55	2.64	1.57
Central America	155,316	3.4	57.66	2.64	1.52
Central Asia	36,549	0.8	56.58	2.87	1.73
Eastern Africa	782,653	17.2	88.55	2.73	1.37
Middle Africa	208,658	4.6	79.34	2.76	1.42
Northern Africa	289,526	6.3	81.02	2.66	1.55
South America	533,459	11.7	68.58	2.83	1.73
South-eastern Asia	502,336	11.0	72.75	2.88	1.68
Southern Africa	66,759	1.5	61.15	3.26	1.76
Southern and Eastern Europe	689,361	15.1	85.10	2.62	1.39
Southern Asia	8,948	0.2	44.14	3.22	1.85
Western Africa	939,886	20.6	98.16	2.75	1.35
Western Asia	184,490	4.0	54.91	2.41	1.47

	N	Mean	Standard Deviation	Min	Max
Birth year	4,564,128	1990.67	10.48	1952	2014
Preceding Interval	4,564,128	2.73	1.50	0.50	9.92

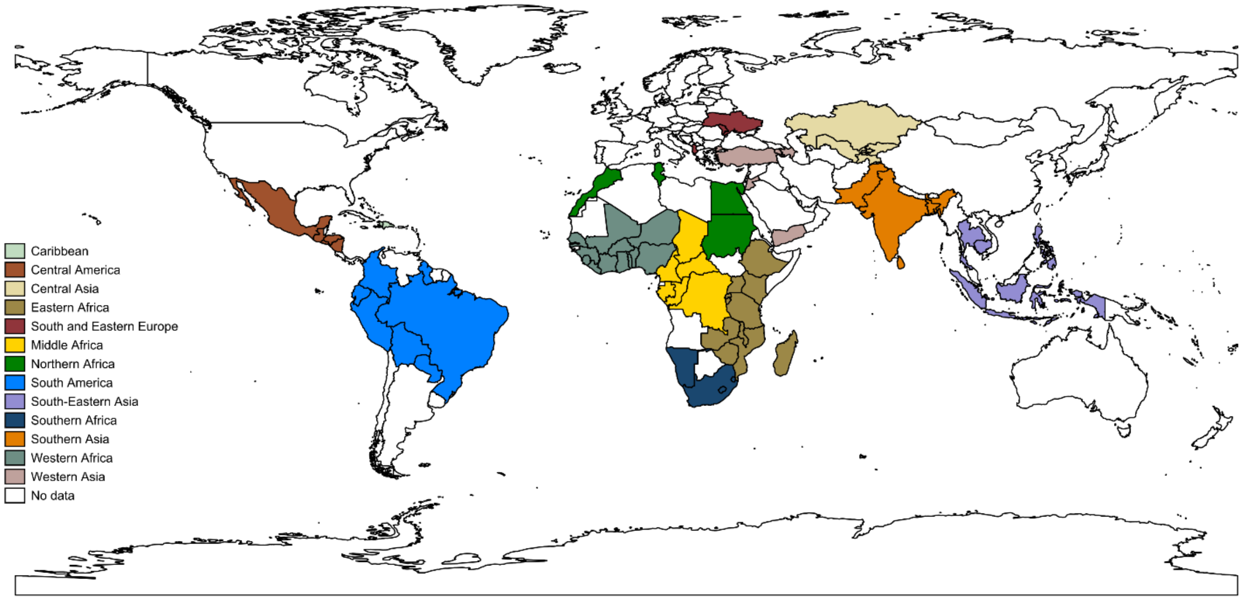


Figure 1. Map of countries included in analysis grouped into UN sub-regions.

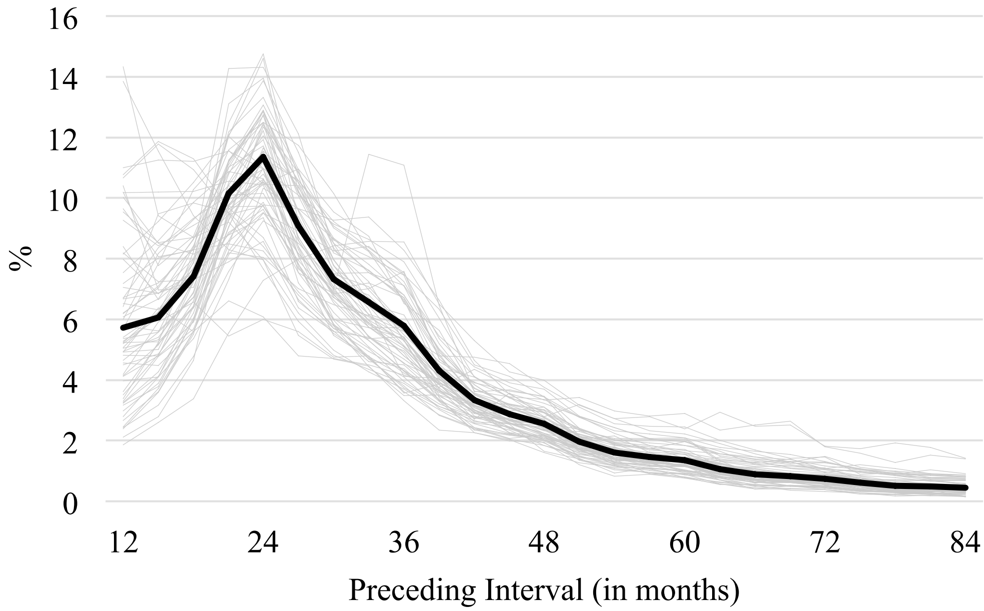


Figure 2. Distribution of preceding birth intervals (in months) in DHS countries.
Note: Bold line indicates the average distribution of all countries.

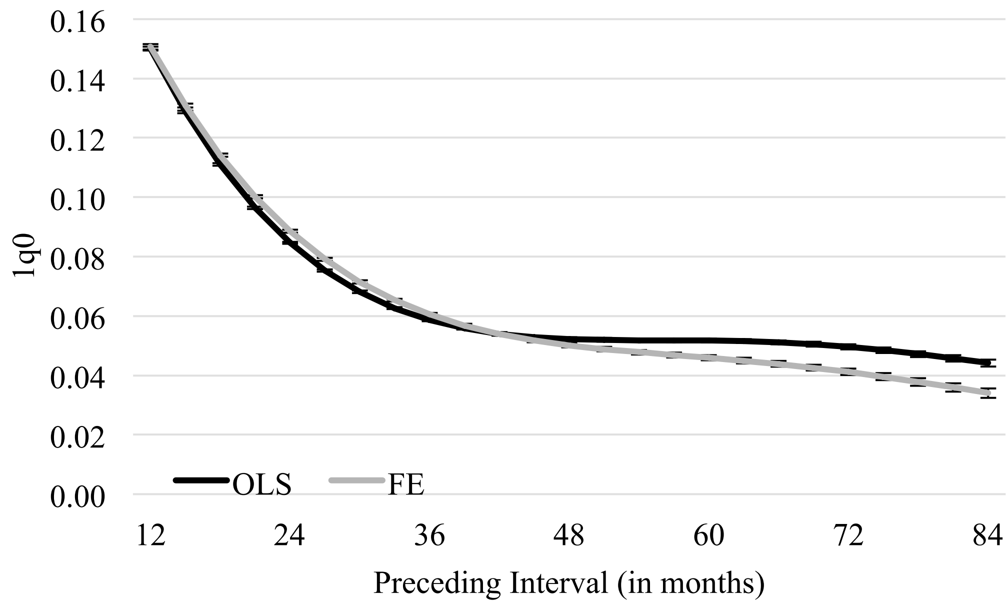
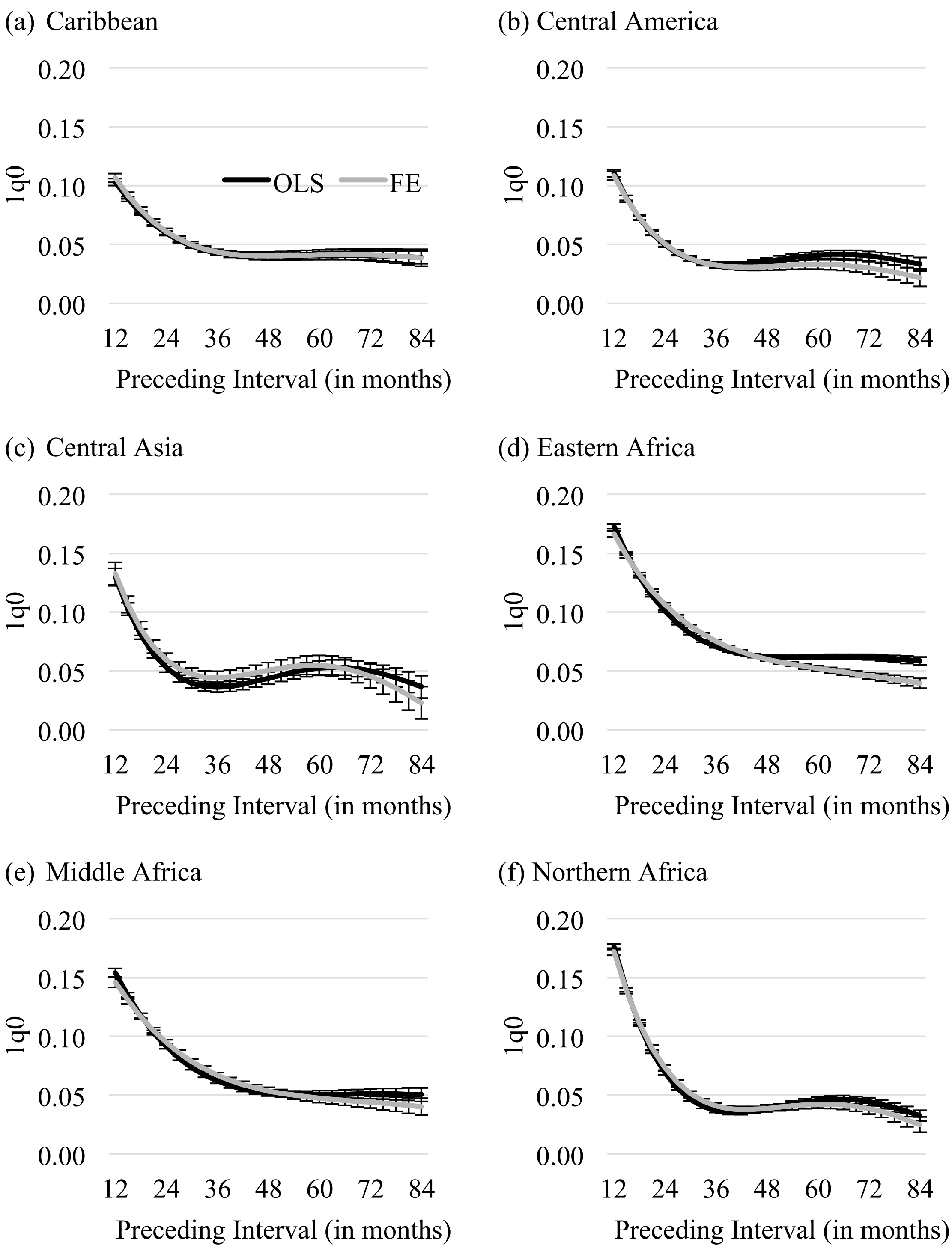
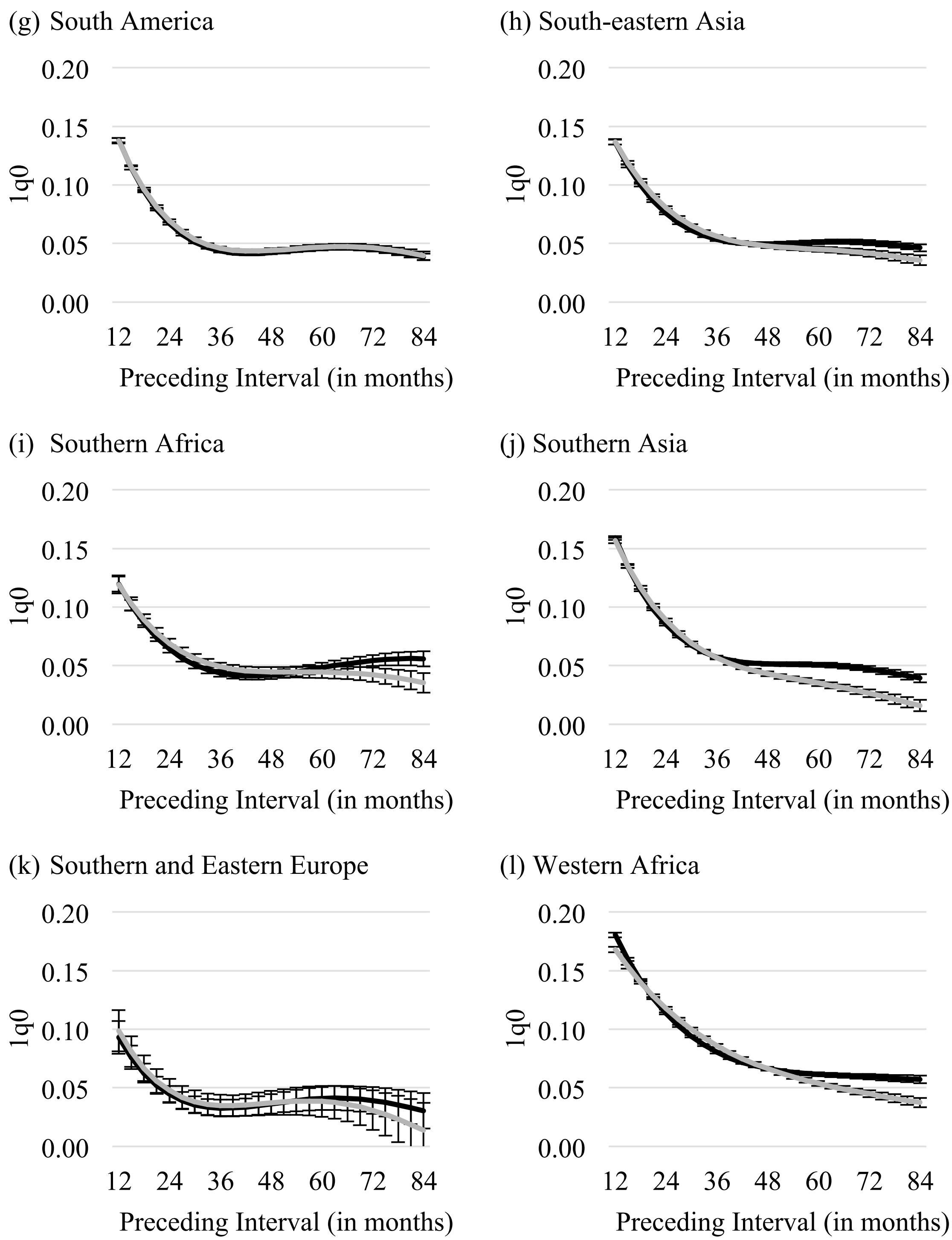


Figure 3. Probability of dying before age one at different preceding birth interval lengths predicted by OLS and FE models.

Note: 95% confidence intervals presented.



(a) Caribbean
(b) Central America
(c) Central Asia
(d) Eastern Africa
(e) Middle Africa
(f) Northern Africa
OLS
FE
1q0
0.00
0.05
0.10
0.15
0.20
12
24
36
48
60
72
84
Preceding Interval (in months)


(g) South America
(h) South-eastern Asia
(i) Southern Africa
(j) Southern Asia
(k) Southern and Eastern Europe
(l) Western Africa
1q0
0.20
0.15
0.10
0.05
0.00
12 24 36 48 60 72 84
Preceding Interval (in months)

(m) Western Asia

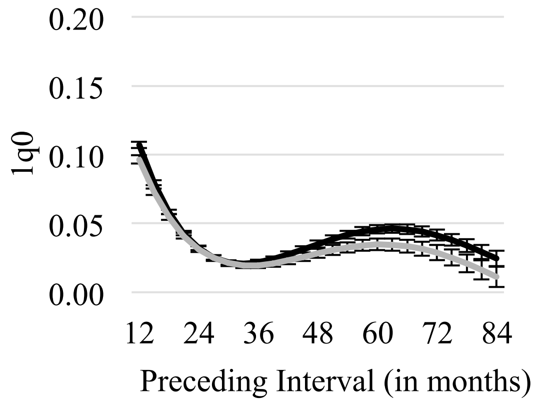


Figure 4. Predicted probabilities of dying before age one at different interval lengths and UN sub-region.

Note: Estimates are from models stratified by UN sub-region. A full list of countries included in regional groupings may be found in the appendix.

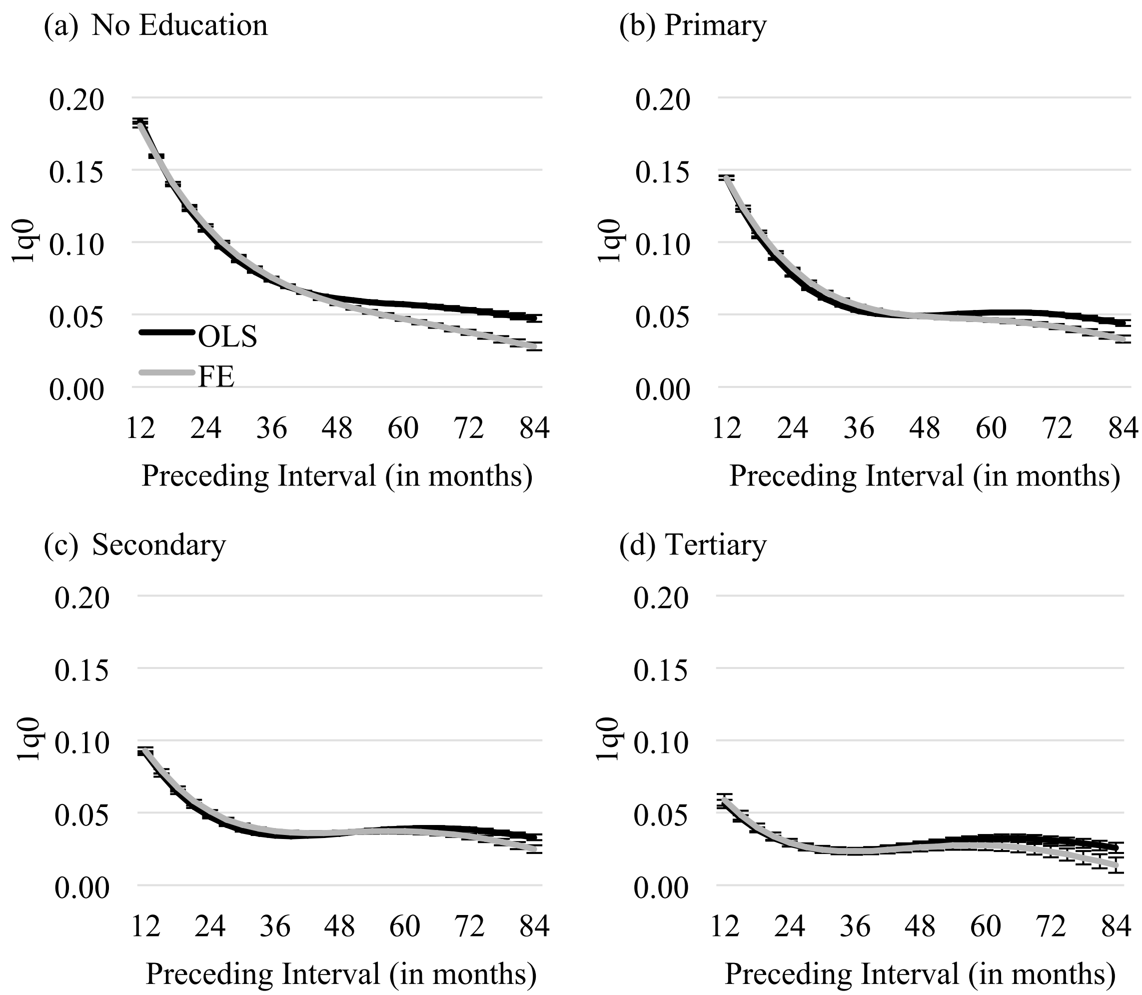
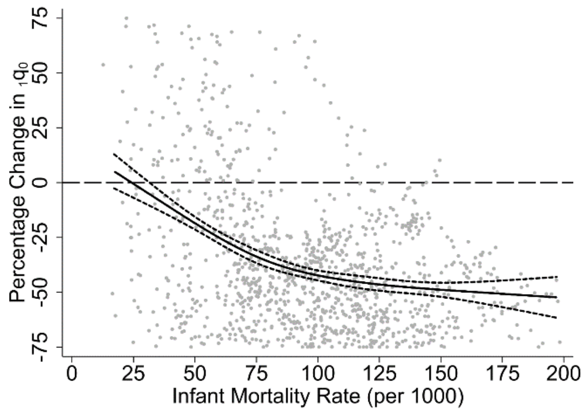


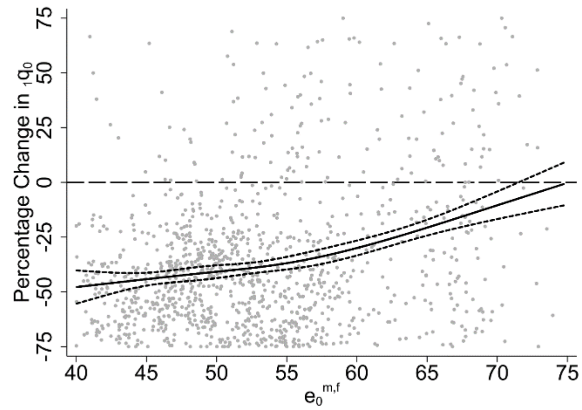
Figure 5. Predicted probabilities of dying before age one at different interval lengths and by mother's educational attainment.

Note: Estimates are from models stratified by a woman's highest level of education.

(a) Infant Mortality Rate



(b) Life Expectancy at Birth



(c) *ln* GNI per capita

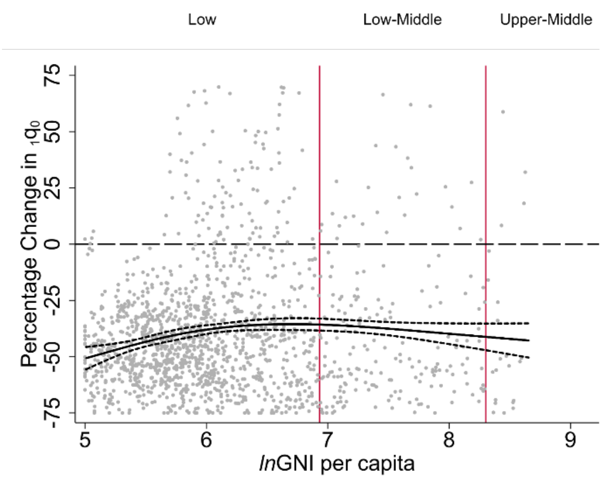


Figure 6. Marginal effect of increasing interval length from one to two years by the (a) infant mortality rate, (b) life expectancy at birth, and (c) *ln* GNI per capita.

Note: A cubic spline trend has been super imposed with 95% confidence intervals.

Appendix

Table 1A. Country-specific samples included in final analysis.

Country	Waves	Years	Women	Children	Deaths
Albania	1	2008-09	2,090	5,537	239
Armenia	2	2000; 2010	4,610	11,063	705
Azerbaijan	1	2006	2,267	5,895	502
Bangladesh	4	1993-94; 2004; 2007; 2011	32,372	117,981	11,012
Benin	4	2006; 2001; 1996; 2011-12	20,578	84,554	6,398
Bolivia	4	1989; 2003; 1994; 2008	24,423	102,322	10,389
Brazil	3	1986; 1991; 1996	7,810	30,344	3,083
Burkina Faso	4	1998-99; 2003; 1993; 2010	20,394	90,893	9,102
Burundi	2	1987; 2010	5,502	24,042	2,175
Cambodia	2	2000; 2005	18,965	73,996	6,951
Cameroon	4	2011; 1998; 2004; 1991	13,625	58,478	4,581
Central African Republic	1	1994-95	2,429	10,366	915
Chad	2	1996-97; 2004	6,678	31,008	3,626
Colombia	5	1995; 1990; 2000; 1986; 2010	34,775	112,410	4,074
Comoros	2	2012; 1996	2,710	11,795	748
Congo Brazzaville	2	2005; 2011-12	7,125	26,962	1,589
Cote D'Ivoire	2	1994; 2011-12	8,530	36,090	3,121
DR Congo	2	2007; 2013-14	12,539	54,897	4,507
Dominican Republic	6	1986; 2013; 2002; 1996; 1991; 2007	31,976	105,802	5,390
Ecuador	1	1987	1,776	7,348	694
Egypt	6	1992; 1995; 1988; 2008; 2014; 2000	56,254	208,862	16,910
El Salvador	1	1985	713	1,618	121
Ethiopia	2	2005; 2000; 2016; 2011	18,654	85,094	8,931
Gabon	2	2000; 2012	5,367	22,518	1,106
Gambia	1	2013	3,953	16,520	748
Ghana	5	2003; 1993; 2008; 1998; 1988	10,006	40,482	3,288
Guatemala	2	1995; 1987	8,148	36,547	3,195
Guinea	3	2012; 1999; 2005	11,191	47,818	5,425
Guyana	1	2009	1,753	6,083	235

Haiti	4	2005-06; 2000; 1994-95; 2012	13,217	55,962	4,660
Honduras	2	2005-06; 2011-12	15,179	58,138	1,989
India	3	2005-06; 1998-99; 1992-93	136,798	469,405	40,373
Indonesia	6	2007; 1994; 1987; 2002-03; 1991; 2012	84,009	296,467	22,846
Jordan	5	1990; 2002; 2007; 2012; 1997	25,272	114,643	3,532
Kazakhstan	1	1995	2,027	6,218	420
Kenya	5	1989; 1993; 2003; 1998; 2008-09	16,818	73,647	5,107
Kyrgyzstan	2	2012	4,271	13,380	735
Lesotho	2	2009; 2004	4,295	14,674	1,038
Liberia	3	1986; 2013; 2007	9,816	41,861	5,549
Madagascar	4	2003-04; 1997; 2008-09; 1992	15,931	68,670	5,435
Malawi	3	2010; 1992; 2000	23,952	98,928	10,011
Maldives	1	2009	3,006	11,855	588
Mali	5	1987; 2006; 2012-13; 1995-96; 2001	25,909	119,983	14,658
Mexico	1	1987	3,412	14,515	1,023
Moldova	1	2005	960	2,415	109
Morocco	3	2003-04; 1992; 1987	11,678	52,337	4,320
Mozambique	3	2003; 1997; 2011	14,806	59,647	6,824
Namibia	4	2006-07; 2000; 1992; 2013	9,627	34,763	1,931
Nicaragua	2	2001; 1998	10,487	44,498	2,627
Niger	4	2012; 1998; 2006; 1992	18,063	87,474	9,474
Nigeria	4	2003; 2008; 2013; 1990	39,157	174,953	17,302
Pakistan	3	2012-13; 2006-07; 1990-91	18,056	79,833	6,324
Paraguay	1	1990	2,123	9,496	408
Peru	6	1986; 1991-92; 2004-06; 2009; 1996; 2000	70,369	265,456	17,699
Phillipines	5	1993; 1998; 2003; 2008; 2013	25,432	97,364	4,331
Rwanda	3	1992; 2010; 2000	16,331	71,932	6,912

Sao Tome & Principe	1	2008-09	1,147	4,429	230
Senegal	4	2010-11; 1992-93; 2005; 1986	27,206	121,012	8,963
Sierra Leone	2	2008; 2013	10,168	40,206	5,269
South Africa	1	1998	3,307	11,018	717
Sri Lanka	1	1987	2,954	10,287	371
Sudan	1	1989-90	3,474	17,516	1,375
Swaziland	1	2006-07	1,610	6,304	396
Tajikistan	1	2012	3,478	11,475	641
Tanzania	4	2010; 1991-92; 2004-05; 1996	17,761	77,778	6,723
Thailand	1	1987	2,775	9,398	511
Timor Leste	1	2009-10	5,555	25,111	1,904
Togo	3	2013-14; 1998; 1988	9,088	38,040	2,965
Trinidad & Tobago	1	1987	1,213	4,423	179
Tunisia	1	1988	2,537	10,811	851
Turkey	3	1993; 1998; 2003	9,003	32,347	2,923
Uganda	5	1995; 2011; 2000-01; 1988-89; 2006	16,980	78,993	6,986
Ukraine	1	2007	413	996	47
Uzbekistan	1	1996	1,604	5,476	272
Yemen	1	1991-92	3,851	20,542	2,468
Zambia	5	1992; 2007; 1996; 2001-02; 2013-14	19,949	87,004	7,042
Zimbabwe	5	1994; 2010-11; 1999; 1988; 2005-06	11,856	45,123	2,409
Total	207		1,154,143	4,564,128	369,227

Table 2A. OLS and FE models of the effects of birth intervals on infant mortality.

	OLS		FE	
	B	*S.E.*	*B*	*S.E.*
Preceding Interval	-0.173	0.002	-0.153	0.002
Preceding $Interval^2$	0.048	0.001	0.040	0.001
Preceding $Interval^3$	-0.006	0.000	-0.005	0.000
Preceding $Interval^4$	0.000	0.000	0.000	0.000
Female	-0.009	0.000	-0.009	0.000
Birth year	-0.001	0.000	0.008	0.000
Birth $year^2$	0.000	0.000	0.000	0.000
Birth $year^3$	0.000	0.000	0.000	0.000
Previous Sibling Died	0.094	0.000	-0.072	0.000
Birth Order				
2	(ref)		(ref)	
3	-0.006	0.000	-0.033	0.000
4	-0.001	0.000	-0.063	0.001
5	0.003	0.000	-0.090	0.001
6	0.007	0.001	-0.115	0.001
7	0.012	0.001	-0.139	0.001
8+	0.020	0.001	-0.181	0.002
Constant	0.270	0.001		
Children	4,564,128		4,564,128	
Mothers	1,154,143		1,154,143	
F-statistic for model fit	11,867.1		5,171.8	
R^2	0.038		0.022	
F-statistic for FE			1.12	
Rho			0.322	

Table 3A. OLS and FE models of the effects of birth intervals on infant mortality using a categorical operationalization of interval length.

	OLS		FE	
	Coefficient	S.E.	Coefficient	S.E.
Preceding Interval				
<12	0.104	0.001	0.092	0.001
12-14	0.059	0.001	0.057	0.001
15-17	0.034	0.001	0.035	0.001
18-20	0.019	0.001	0.019	0.001
21-23	0.010	0.001	0.008	0.001
24-26	(ref)		(ref)	
27-29	-0.004	0.001	-0.004	0.001
30-32	-0.010	0.001	-0.011	0.001
33-35	-0.017	0.001	-0.021	0.001
36-38	-0.023	0.001	-0.028	0.001
39-41	-0.024	0.001	-0.029	0.001
42-44	-0.024	0.001	-0.028	0.001
45-47	-0.029	0.001	-0.035	0.001
48-50	-0.033	0.001	-0.039	0.001
51-53	-0.032	0.001	-0.038	0.001
54-56	-0.032	0.001	-0.037	0.001
57-59	-0.034	0.001	-0.043	0.001
60-62	-0.035	0.001	-0.045	0.001
63-65	-0.033	0.001	-0.044	0.002
66-68	-0.034	0.001	-0.043	0.002
69-71	-0.034	0.001	-0.046	0.002
72-74	-0.035	0.002	-0.048	0.002
75-77	-0.034	0.002	-0.046	0.002
78-80	-0.033	0.002	-0.048	0.002
81-83	-0.032	0.002	-0.048	0.002
84+	-0.032	0.001	-0.052	0.001
Female	-0.009	0.000	-0.009	0.000
Birth year	-0.001	0.000	0.008	0.000
Birth year^2	0.000	0.000	0.000	0.000
Birth year^3	0.000	0.000	0.000	0.000
Previous Sibling Died	0.094	0.000	-0.072	0.000
Birth Order				
2	(ref)		(ref)	
3	-0.006	0.000	-0.033	0.000
4	-0.002	0.000	-0.063	0.001
5	0.003	0.000	-0.090	0.001
6	0.007	0.001	-0.116	0.001
7	0.012	0.001	-0.140	0.001

8+	0.020	0.001	-0.181	0.002
Constant	0.070	0.000		
Children	4,564,128		4,564,128	
Mothers	1,154,143		1,154,143	
F-statistic for model fit	4982.69		2168.45	
R^2	0.038		0.022	
F-statistic for FE			1.12	
Rho			0.322	

Table 4A. OLS and FE models of the effects of birth intervals on infant mortality for children born within 10 years of the survey.

	OLS		FE	
	B	*S.E.*	*B*	*S.E.*
Preceding Interval	-0.167	0.002	-0.131	0.003
Preceding Interval^2	0.046	0.001	0.034	0.001
Preceding Interval^3	-0.005	0.000	-0.004	0.000
Preceding Interval^4	0.000	0.000	0.000	0.000
Female	-0.008	0.000	-0.008	0.000
Birth year	0.000	0.000	0.008	0.000
Birth year^2	0.000	0.000	0.000	0.000
Birth year^3	0.000	0.000	0.000	0.000
Previous Sibling Died	0.077	0.001	-0.169	0.001
Birth Order				
2	(ref)		(ref)	
3	-0.013	0.001	-0.046	0.001
4	-0.009	0.001	-0.086	0.001
5	-0.005	0.001	-0.122	0.002
6	0.000	0.001	-0.155	0.002
7	0.005	0.001	-0.187	0.002
8+	0.015	0.001	-0.230	0.003
Constant	0.273	0.002		
Children	2,329,949		2,329,949	
Mothers	922,402		922,402	
F-statistic for model fit	4,835.6		5,479.3	
R^2	0.030		0.055	
F-statistic for FE			1.12	
Rho			0.434	

Table 5A. OLS and FE models of the effects of birth intervals on infant mortality using a categorical operationalization of interval length for children born within 10 years of the survey.

	OLS		FE	
	Coefficient	S.E.	Coefficient	S.E.
Preceding Interval				
<12	0.101	0.001	0.078	0.001
12-14	0.058	0.001	0.048	0.001
15-17	0.034	0.001	0.031	0.001
18-20	0.020	0.001	0.017	0.001
21-23	0.010	0.001	0.007	0.001
24-26	(ref)		(ref)	
27-29	-0.006	0.001	-0.007	0.001
30-32	-0.012	0.001	-0.013	0.001
33-35	-0.019	0.001	-0.022	0.001
36-38	-0.022	0.001	-0.026	0.001
39-41	-0.023	0.001	-0.027	0.001
42-44	-0.024	0.001	-0.027	0.001
45-47	-0.028	0.001	-0.033	0.001
48-50	-0.030	0.001	-0.036	0.001
51-53	-0.030	0.001	-0.033	0.002
54-56	-0.030	0.001	-0.032	0.002
57-59	-0.030	0.001	-0.035	0.002
60-62	-0.032	0.001	-0.039	0.002
63-65	-0.031	0.002	-0.039	0.002
66-68	-0.030	0.002	-0.036	0.002
69-71	-0.031	0.002	-0.036	0.003
72-74	-0.032	0.002	-0.038	0.003
75-77	-0.031	0.002	-0.039	0.003
78-80	-0.029	0.002	-0.034	0.003
81-83	-0.028	0.002	-0.037	0.004
84+	-0.028	0.001	-0.036	0.002
Female	-0.008	0.000	-0.008	0.000
Birth year	0.000	0.000	0.008	0.000
Birth year^2	0.000	0.000	0.000	0.000
Birth year^3	0.000	0.000	0.000	0.000
Previous Sibling Died	0.077	0.001	-0.169	0.001
Birth Order				
2	(ref)		(ref)	
3	-0.013	0.001	-0.046	0.001
4	-0.010	0.001	-0.086	0.001
5	-0.005	0.001	-0.122	0.002
6	0.000	0.001	-0.155	0.002

7	0.005	0.001	-0.187	0.002
8+	0.014	0.001	-0.230	0.003
Constant	0.079	0.001		
Children	2,329,949		2,329,949	
Mothers	922,402		922,402	
F-statistic for model fit	2027.8		2285.44	
R^2	0.030		0.055	
F-statistic for FE			1.12	
Rho			0.434	

Table 6A. Countries included in UN sub-regions.

UN sub-region	Countries
Caribbean	Dominican Republic; Haiti; Trinidad & Tobago
Central America	El Salvador; Guatemala; Honduras; Mexico; Nicaragua
Central Asia	Kazakhstan; Kyrgyzstan; Tajikistan; Uzbekistan
Eastern Africa	Burundi; Comoros; Ethiopia; Kenya; Madagascar; Malawi; Mozambique; Rwanda; Tanzania; Uganda; Zambia; Zimbabwe
Middle Africa	Cameroon; Central African Republic; Chad; Congo Brazzaville; DR Congo; Gabon; Sao Tome & Principe
Northern Africa	Egypt; Morocco; Sudan; Tunisia
South America	Bolivia; Brazil; Colombia; Ecuador; Guyana; Paraguay; Peru
South-eastern Asia	Cambodia; Indonesia; Phillipines; Thailand; Timor Leste
Southern Africa	Lesotho; Namibia; South Africa; Swaziland
Southern and Eastern Europe	Bangladesh; India; Maldives; Pakistan; Sri Lanka
Southern Asia	Albania; Moldova; Ukraine
Western Africa	Benin; Burkina Faso; Cote DIvoire; Gambia; Ghana; Guinea; Liberia; Mali; Niger; Nigeria; Senegal; Sierra Leone; Togo
Western Asia	Armenia; Azerbaijan; Jordan; Turkey; Yemen

Note: Regional groupings may be found at https://unstats.un.org/unsd/methodology/m49/.

I. STRATIFIED ANALYSES

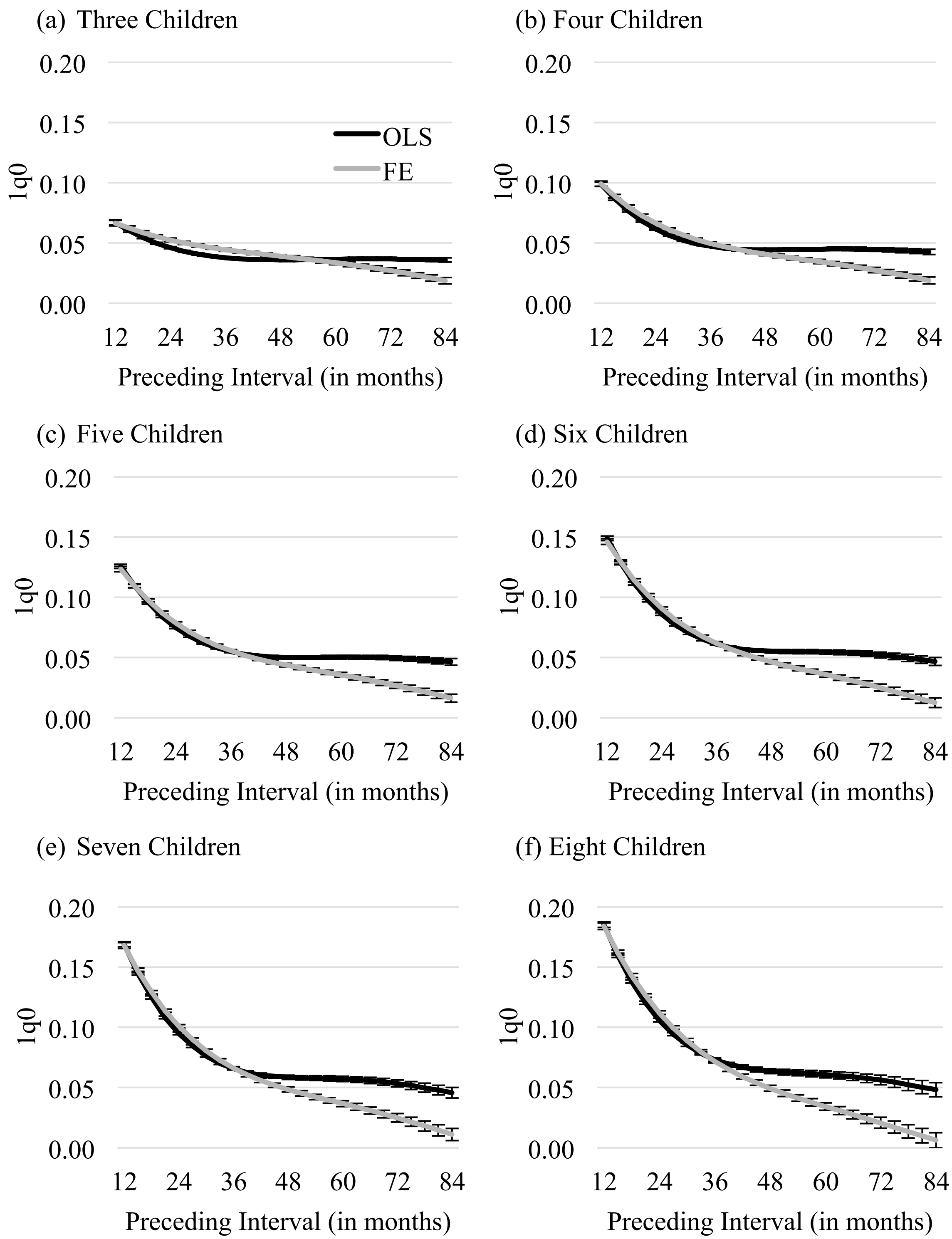


Figure 1A. Predicted probabilities of dying before age one at different interval lengths and by total children ever born at the time of interview.

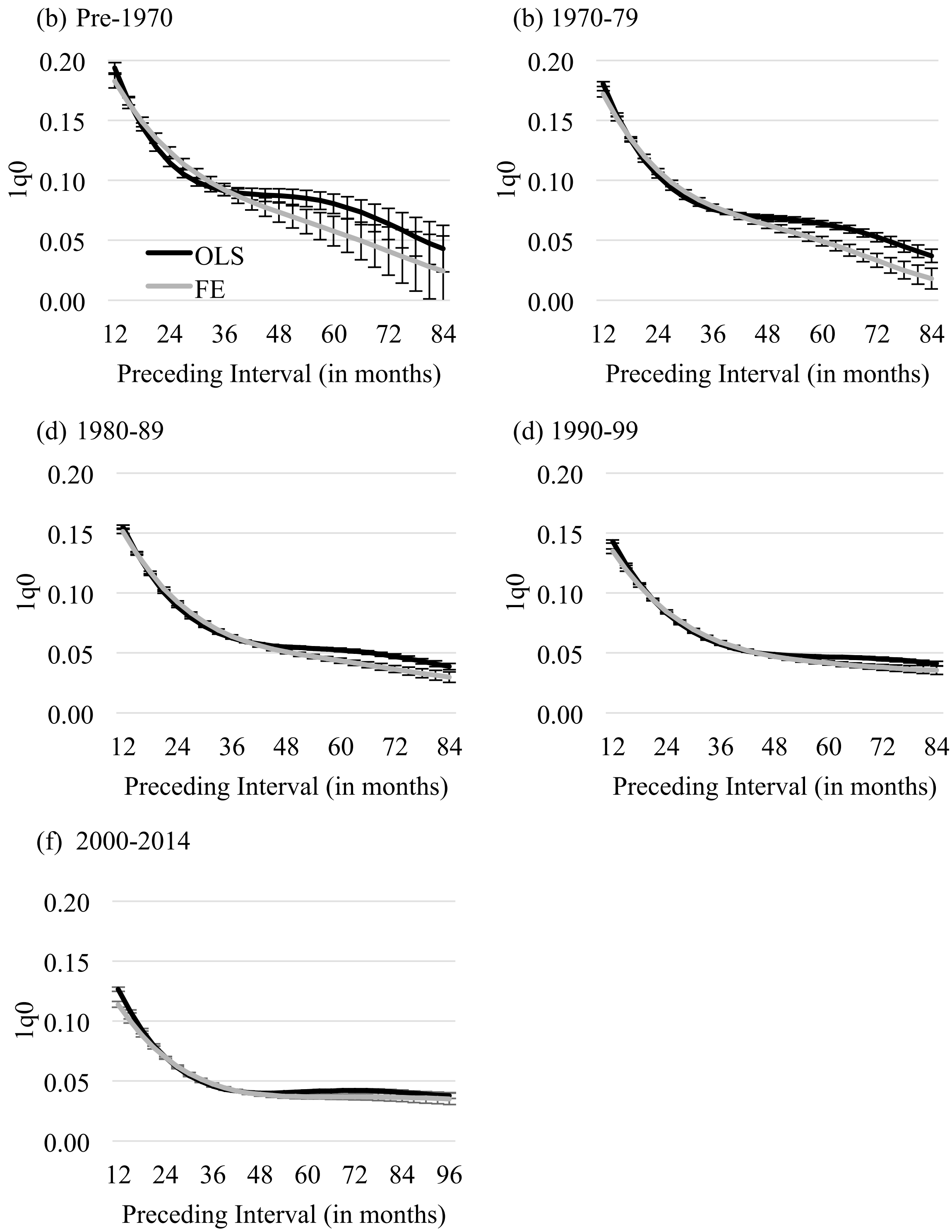


Figure 2A. Predicted probabilities of dying before age one at different interval lengths and by index child’s birth cohort.

II. Sub-Sample Analyses

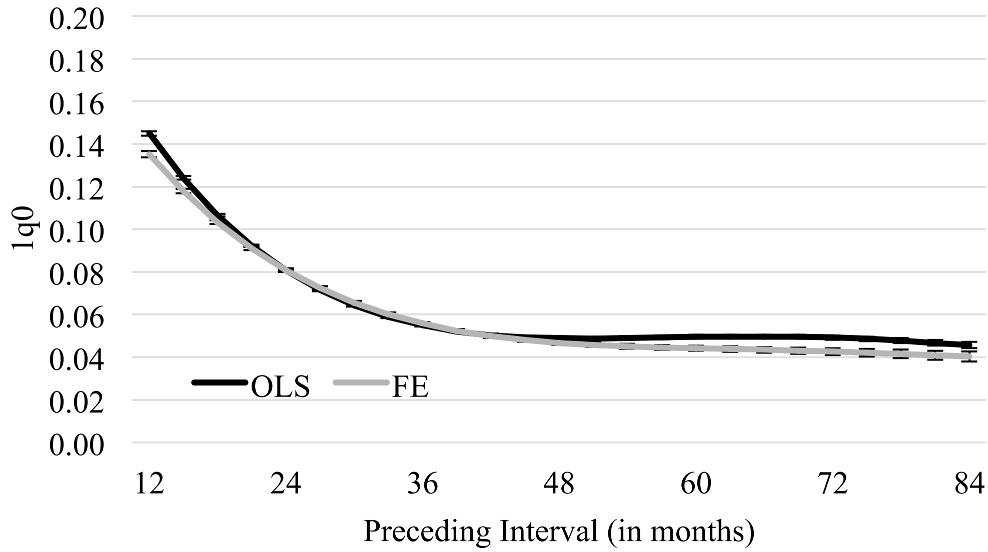


Figure 3A. Predicted probabilities of dying before age 1 at different interval lengths in OLS and FE models among children born within ten years of survey.

Note: A ten-year cutoff was chosen because it has been shown that the misreporting and displacement of births tends to be low within that time frame (see Schoumaker 2014).

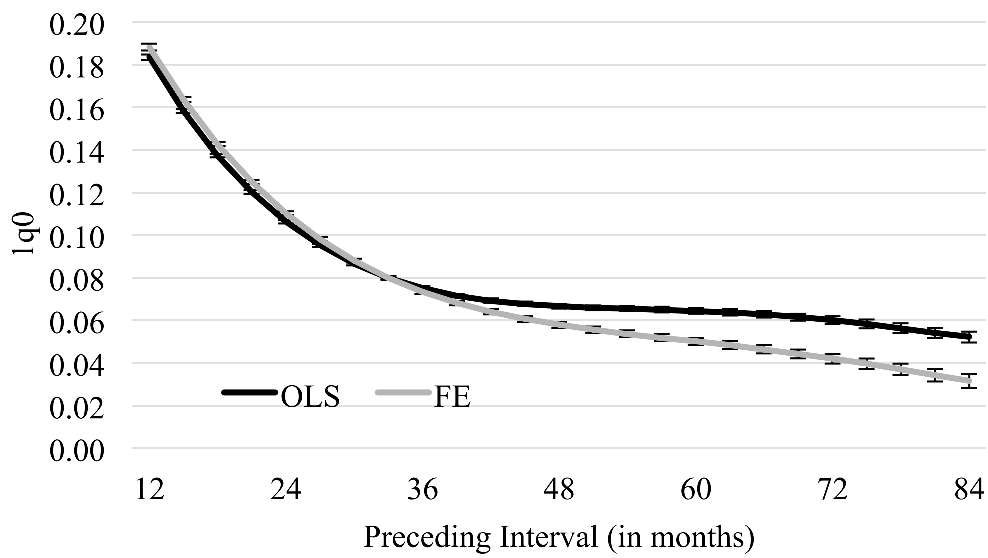


Figure 4A. Predicted probabilities of dying before age 1 at different interval lengths in OLS and FE models among children born at even parities.

Note: Predictions based on models including only children born at even parities in families of four or more. This is to be sure that the death of the index child cannot also be included as the death of the previous sibling variable.

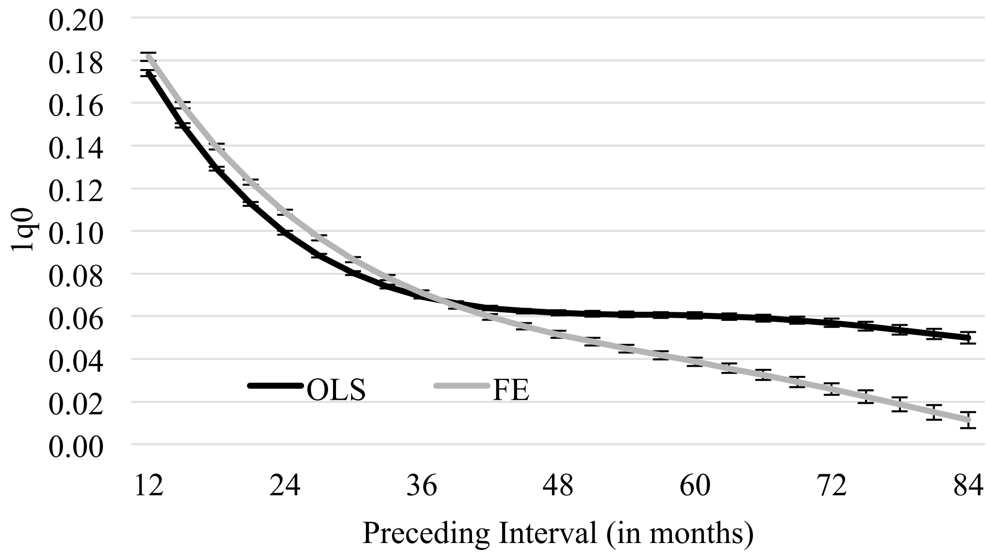
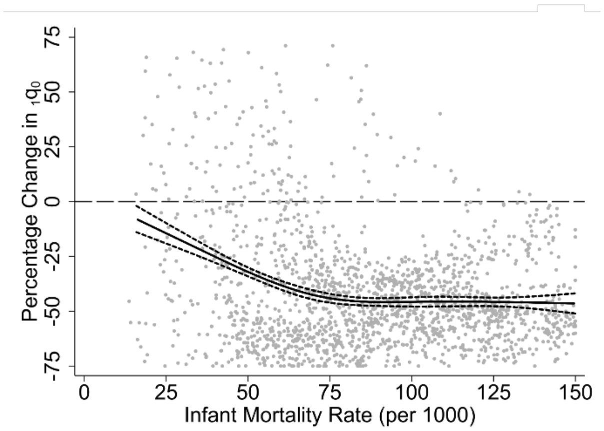


Figure 5A. Predicted probabilities of dying before age 1 at different interval lengths in OLS and FE models among children born at odd parities.

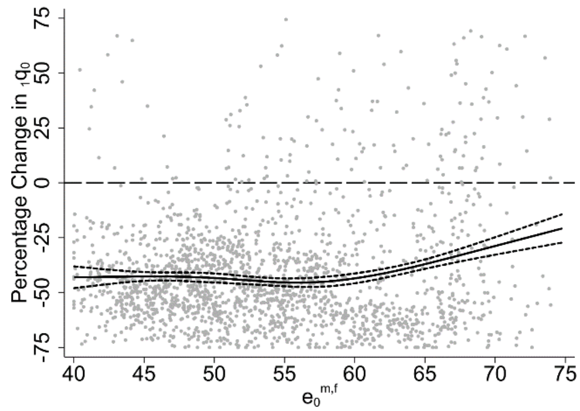
Note: Predictions based on models including only children born at odd parities in families of five or more. This is to be sure that the death of the index child cannot also be included as the death of the previous sibling variable.

III. Changing Reference Interval Length

(a) Infant Mortality Rate



(b) Life Expectancy



(c) GNI per Capita

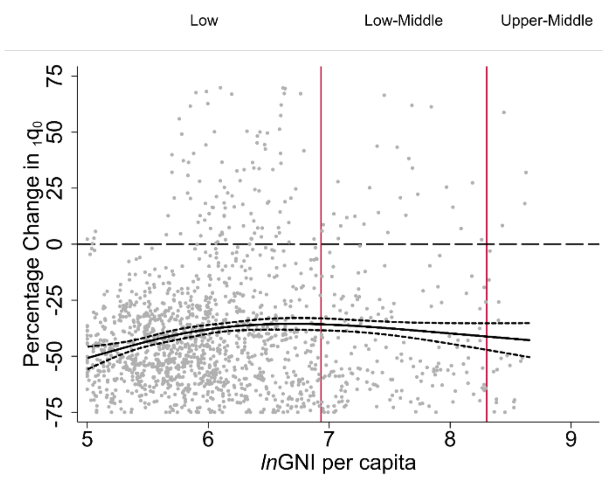
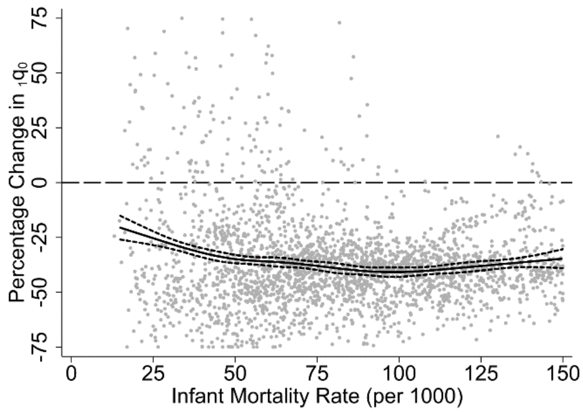


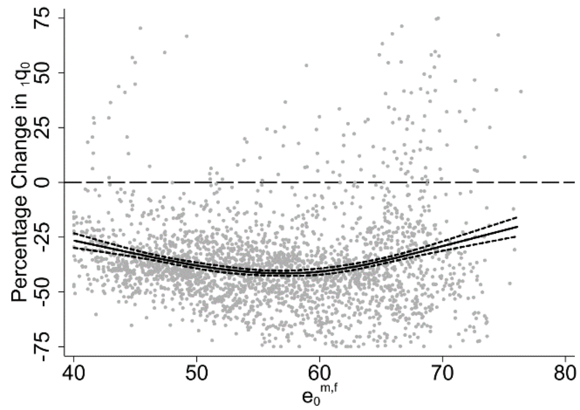
Figure 6A. Marginal effect of increasing interval length from 18 to 30 months by the (a) infant mortality rate, (b) life expectancy at birth, and (c) *ln* GNI per capita.

Note: A cubic spline trend has been super imposed with 95% confidence intervals.

(a) Infant Mortality Rate



(b) Life Expectancy



(c) GNI per Capita

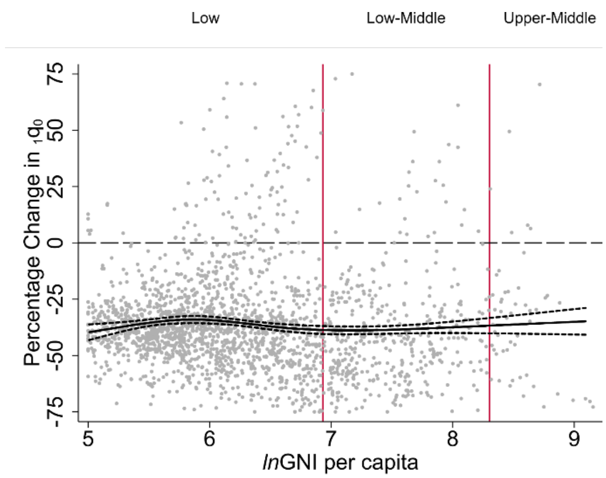


Figure 7A. Marginal effect of increasing interval length from 24 to 36 months by the (a) infant mortality rate, (b) life expectancy at birth, and (c) *ln* GNI per capita.

Note: A cubic spline trend has been super imposed with 95% confidence intervals.

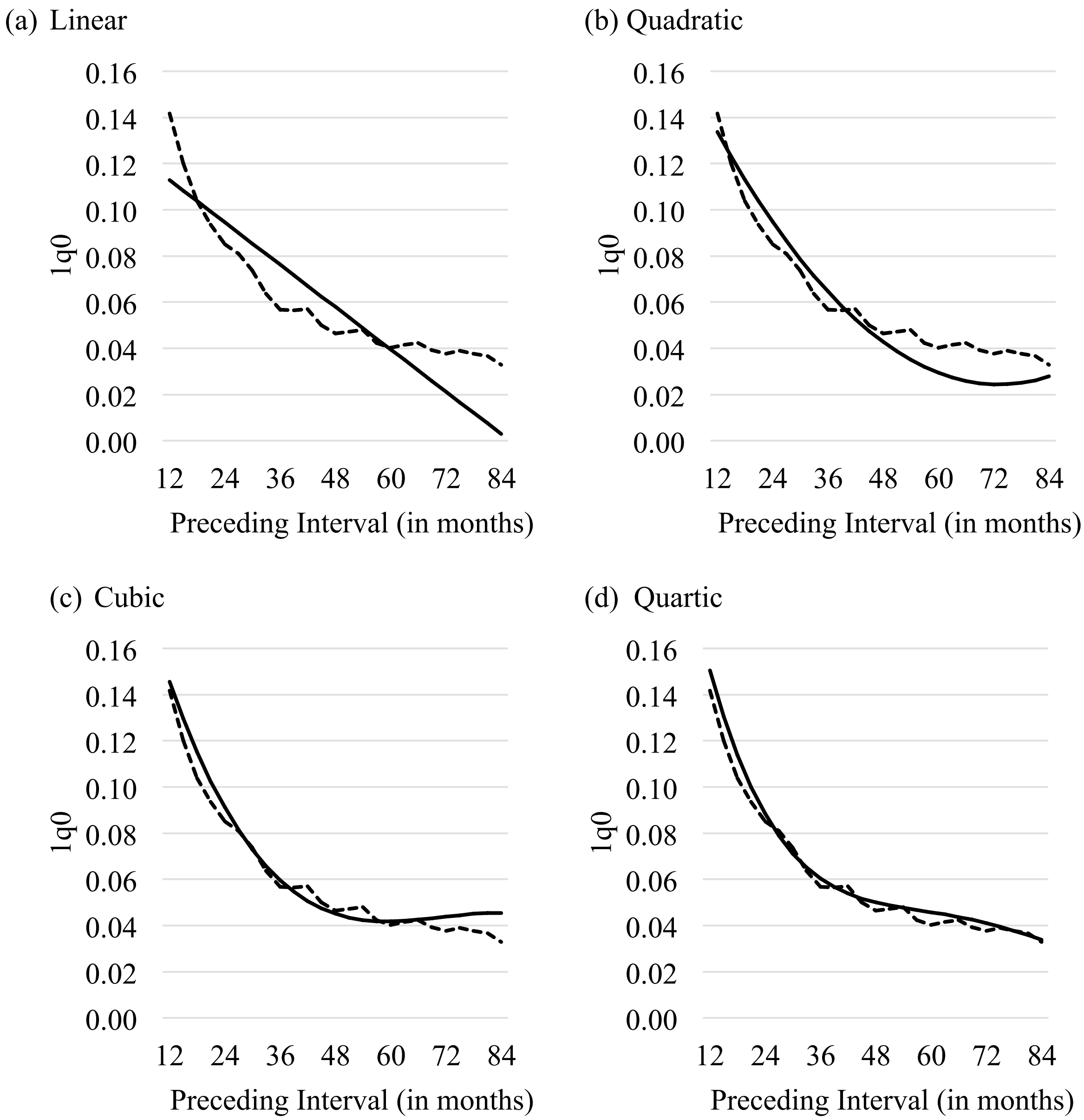


Figure 8A. Comparing predicted probabilities of infant mortality from models using categorical and polynomial operationalizations of the length of the previous interval.
Note: Dashed line refers to the categorical operationalization.