

REPORT

The enduring burden of short birth intervals on child survival: A 28-year trend analysis in Burkina Faso

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Abstract

Child survival has improved considerably over the past decades in sub-Saharan Africa, yet preventable risk factors such as closely spaced births continue to threaten progress, especially in fragile contexts. This study examines the association between interbirth intervals and under-five survival in Burkina Faso using data from five rounds of Demographic and Health Surveys conducted between 1993 and 2021. The study combines descriptive survival analysis (e.g., Kaplan–Meier curves, log-rank tests) with Cox proportional hazards models to estimate the adjusted association between interbirth intervals and under-five mortality, while controlling for maternal, child, household, and contextual characteristics. The results show that short interbirth intervals (<24 months) are consistently associated with excess mortality, following a markedly non-linear trajectory: the adjusted hazard ratio declines from 2.50 (95% confidence interval: 1.56–4.0) in 1993 to 1.49 (1.04–2.14) in 1998–1999 and 1.63 (1.27–2.1) in 2003, then rises to 1.97 (1.64–2.38) in 2010 and peaks at 3.20 (2.28–4.49) in 2021. In contrast, the protective advantage of long intervals (≥ 36 months), significant from 1998 to 2010, disappears in the most recent survey. These dynamics suggest that health system strengthening and expanded access to care have mitigated part of the vulnerability associated with closely spaced births, but that these gains remain fragile in the context of escalating security crisis and persistent socio-territorial inequalities. Over time, under-five survival appears increasingly shaped by socioeconomic (e.g., household wealth, maternal education), territorial (e.g., rural residence, conflict-affected regions), and biological factors (e.g., low birth weight, extreme maternal ages). In this context, the enduring burden of mortality linked to short interbirth intervals, particularly among poor rural households and adolescent mothers, calls for geographically and socially targeted family planning strategies, embedded within an integrated package of maternal and child health, nutrition, vaccination, and equitable access to quality healthcare services.

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1. Introduction

Under-five mortality remains a central indicator of human development and population well-being. It reflects not only the state of the health system, but also the social, economic, and environmental conditions in which families live (United Nations Inter-

agency Group for Child Mortality Estimation, 2026; World Health Organization, 2023). Since the early 2000s, significant progress has been made globally through the expansion of vaccination programs, improved health coverage, and the implementation of multisectoral policies aimed at enhancing child survival (Black *et al.*, 2010). According to the United Nations Inter-agency Group for Child Mortality Estimation (2026), the global under-five mortality rate declined from 76 deaths per 1,000 live births in 2000 to 38 in 2023, representing a 50% reduction over the past two decades.

However, these improvements mask persistent regional disparities. Sub-Saharan Africa remains the most affected region, accounting for 58% of under-five deaths despite representing only about 30% of global births (United Nations Inter-agency Group for Child Mortality Estimation, 2025). In 2024, the under-five mortality rate in the region still reached 71.6 deaths per 1,000 live births, nearly twice the global average (United Nations Inter-agency Group for Child Mortality Estimation, 2025). This excess mortality is driven by a combination of structural factors, including persistent poverty, weak health systems, armed conflicts, food insecurity, and inadequate maternal and child health services (Adedini *et al.*, 2024).

Burkina Faso illustrates this stark contrast. Under-five mortality declined sharply from 187 deaths per 1,000 live births in 1993 to 48 in 2021 (Institut National de la Statistique et de la Démographie & ICF, 2023). This achievement notwithstanding, the country remains well above the Sustainable Developmental Goals target of fewer than 25 deaths per 1,000 live births by 2030 (Organisation des Nations Unies, 2015). Efforts such as the removal of user fees for children and pregnant women, vaccination campaigns, and improvements in family planning (Ministère de la santé, 2011) have significantly enhanced child survival, but have not fully eliminated regional and social disparities. Despite these advances, no study has examined how the association between birth spacing and under-five survival has evolved over time in Burkina Faso. Existing work in the region relies predominantly on single cross-sectional surveys and does not address whether the excess mortality risk associated with short interbirth intervals has attenuated, persisted, or resurged as the country's health system developed between 1993 and 2021. This temporal evidence gap is the focus of the present study. In this context, birth spacing, or the interbirth interval, emerges as a key determinant of under-five survival. Numerous studies have shown that short intervals (<24 months) are associated with higher infant mortality risk, mainly due to competition for maternal resources, low birth weight, prematurity, and neonatal complications (Conde-

Agudelo *et al.*, 2006; DaVanzo *et al.*, 2007; Rutstein, 2005). Conversely, optimal spacing is associated with improved maternal and child health outcomes (Chekole *et al.*, 2025; Smith *et al.*, 2003). However, several studies also highlight that excessively long intervals may be associated with increased risk of perinatal complications (Conde-Agudelo, 2000).

It is important to note that the associations between birth spacing and child survival are neither uniform over time nor across contexts. They vary depending on health system conditions, socioeconomic environments, and social norms (Bocquier, 1991; Bongaarts & Sinding, 2011; Cleland *et al.*, 2012). In countries that have strengthened their health systems and improved access to care, the strength of the association between short birth intervals and child survival tends to decrease, although it does not disappear entirely (Coulibaly *et al.*, 2024). In Burkina Faso, the increase in the proportion of assisted deliveries—from 37% in 1993 to 96% in 2021 (Institut National de la Statistique et de la Démographie & ICF, 2023)—illustrates these major transformations that may help attenuate the excess risk associated with closely spaced births.

This study aims to analyze the evolution of the association between birth spacing and the survival of children under five in Burkina Faso between 1993 and 2021, using successive Demographic and Health Survey (DHS; *Enquête Démographique et de Santé*) data. Beyond examining changes in the effect of birth spacing on child survival, the study also assesses how maternal, child, household, and contextual characteristics help explain changes in the estimated association between birth spacing and under-five mortality when they are progressively introduced into the Cox proportional hazard models.

Building on this empirical and conceptual background, the present study pursues two main research questions:

- (i) RQ1: This study examines how the association between short interbirth intervals and under-five mortality has evolved between 1993 and 2021 in Burkina Faso.
- (ii) RQ2: This study explores to what extent maternal, child, household, and contextual characteristics mediate the association between birth spacing and under-five mortality, as reflected in the attenuation of the birth spacing effect when these covariates are progressively introduced into the Cox proportional hazard models.

In line with the existing literature and the specificities of the Burkinabè context, two working hypotheses are tested:

- (i) H1: The excess mortality risk associated with short interbirth intervals has persisted across all survey rounds, with a marked resurgence in 2021 in the

context of the ongoing security crisis.

- (ii) H2: This excess risk is only partially mediated by maternal, child, household, and contextual characteristics: the hazard ratio associated with short intervals decreases as these covariates are added stepwise to the models, but remains significantly above 1 in all survey rounds.

Together, these questions and hypotheses aim to document both the long-term dynamics of the birth spacing effect and the extent to which broader social and health system changes have reshaped its magnitude over time.

2. Methodology

2.1. Data sources

The analysis is based on the DHS conducted in Burkina Faso in 1993, 1998–1999, 2003, 2010, and 2021. These surveys, implemented by the National Institute of Statistics and Demography (INSD) in collaboration with ICF International, provide nationally representative data on fertility, mortality, maternal and child health, as well as utilization of health services (Institut National de la Statistique et de la Démographie & ICF, 2023; Institut National de la Statistique et de la Démographie & ORC Macro, 2004; Institut National de la Statistique et de la Démographie & Macro International Inc, 2000).

2.2. Study population

Our target population consists of all children aged 0 to 4 years, identified from the DHS conducted in Burkina Faso in 1993, 1998–1999, 2003, 2010, and 2021. The analysis is based on data derived from the birth histories of women of reproductive age. The initial sample sizes of children under five are 5,828 in 1993, 5,953 in 1998–1999, 10,645 in 2003, 15,045 in 2010, and 12,242 in 2021.

To address the specific objectives of the study, we excluded children from multiple births as well as single children. This selection helps limit biases related to specific obstetric complications and allows us to focus on children belonging to sibling groups, thereby enabling a more rigorous analysis of the association between birth spacing and child survival. After applying these criteria, the final sample sizes retained are 3,703 children in 1993, 3,577 in 1998–1999, 5,911 in 2003, 8,302 in 2010, and 5,092 in 2021. The marked increase in initial sample sizes across survey rounds reflects deliberate methodological decisions made by the DHS program and INSD in each successive wave.

The 1993 DHS used a stratified two-stage design based on the 1985 General Population and Housing Census (*Recensement Général de la Population et de l'Habitation*

[RGPH]), selecting primary sampling units proportionate to size within urban and rural strata. The 1998–1999 DHS maintained the same design with updated enumeration areas. The 2003 DHS, based on the 1996 RGPH, expanded coverage substantially, drawing 400 primary sampling units from over 350 departments and targeting 10,000 households, with 12,477 women interviewed (96% response rate) and a gross sample of 10,645 children under five. The 2010 DHS–Multiple Indicator Cluster Survey IV, based on the 2006 RGPH, extended to 574 clusters (176 urban, 398 rural), with 15,000 households selected and 17,087 women interviewed (98% response rate), explaining the peak sample of 15,045 children under five. The 2021 DHS updated enumeration zones from the 2019 RGPH and interviewed approximately 17,659 individuals, producing an initial sample of 12,242 children under five.

The reduction in the 2021 final analytical sample (5,092 after exclusions) relative to 2010 (8,302) reflects the combined effect of a smaller overall survey, itself partly attributable to incomplete cluster coverage in regions facing security challenges: 86 of the 600 sampled clusters (14.3%) could not be surveyed due to insecurity, with the Sahel and Est regions being the most affected (28 of 41 clusters [68.3%] and 14 of 41 [34.1%] unreachable, respectively), leading to the full exclusion of the provinces of Oudalan, Yagha, Tapoa, Komandjoari, and Kompienga (Institut National de la Statistique et de la Démographie & ICF, 2023), and a higher proportion of first-born or single children excluded under our selection criteria. All analyses apply DHS complex-survey weights to ensure national representativeness and comparability across rounds.

Among these children, 3,137 were survivors in 1993 (85.69%), 2,962 in 1998–1999 (82.8%), 5,062 in 2003 (85.6%), 7,461 in 2010 (89.9%), and 4,839 in 2021 (95%). These figures constitute the main analytical sample used in the analysis of the evolution of under-five survival in Burkina Faso.

2.3. Ethical considerations

This study is based on secondary data analysis and does not include any information that could identify participants. The DHS survey protocols received ethical approval and the necessary administrative authorizations. Data collection, processing, and analysis were conducted in strict accordance with ethical principles, particularly with respect to informed consent, confidentiality, and anonymity.

2.4. Measures

The dependent variable in this study is the survival time of children. This duration may be fully observed or subject to

censoring. When observed, it corresponds to the number of months between the child's birth and death. In the case of censoring, when the child is still alive at the time of the survey, the duration is measured by the child's age at the time of the interview. As in any event history analysis, accounting for censoring requires the creation of an indicator variable to distinguish censored from observed cases. This binary variable takes the value 1 when the child died before the survey (observed event) and 0 when the child was still alive at the time of the interview (right-censored).

The main independent variable is the interbirth interval (spacing between two successive births), defined in accordance with the World Health Organization (2007) and Rutstein (2005) recommendations. It is categorized into three groups: short interval (<24 months), medium interval (24–35 months, reference category), and long interval (≥ 36 months).

Other explanatory variables are included as control factors, in line with the conceptual framework of Mosley and Chen (1984). These include maternal characteristics (e.g., age at childbirth, education level, employment status, religion), child characteristics (e.g., sex, birth order, birth weight), and household and contextual characteristics (e.g., household wealth, place of residence [urban or rural]). These variables were selected based on their documented association with child mortality in the literature (Cleland *et al.*, 2012; Fotso *et al.*, 2013).

2.5. Analytical methods

The analysis included several complementary approaches. First, we conducted a descriptive analysis using non-parametric survival analysis with Kaplan–Meier curves and the log-rank test to compare children's survival probabilities by birth spacing. Following the descriptive analysis, a multivariable analysis was conducted using Cox proportional hazards regression models to estimate the net association between birth spacing and child survival while controlling for explanatory variables.

The advantage of the Cox model lies in the fact that it does not impose a specific functional form for $h_0(t)$, thus allowing for a flexible estimation of the relative effects of covariates without constraining the shape of the hazard function over time.

In this study, the model was used to assess the effect of interbirth spacing on the risk of under-five mortality, while controlling for a set of explanatory variables. This specification is particularly well-suited to DHS data, which provides duration data measured in months and includes a substantial proportion of censored observations (children still alive at the time of the survey).

To capture temporal changes, analyses were conducted separately for each survey and then compared. The evolution of the birth spacing effect was assessed descriptively by comparing the hazard ratios (HRs) across survey-specific Cox proportional hazards models estimated with the same set of covariates (Bocquier, 1991; Cleland *et al.*, 2012).

An important methodological consideration concerns the unequal intervals between the five surveys (five years: 1993–1998; five years: 1998–2003; seven years: 2003–2010; 11 years: 2010–2021). The analysis does not rely on a pooled Cox proportional hazards model combining all survey rounds, but on five independent Cox models estimated separately for each round. Each model was adjusted for the same set of covariates (M0 to M9), which ensures the comparability of estimates across rounds. The evolution of the birth spacing effect is therefore assessed descriptively through comparison of the fully adjusted HR (M9) across rounds, and is not formally tested through interaction terms in a single pooled model. This approach imposes no temporal structure on the data and makes no assumption about the linearity of change between periods. The unequal inter-survey intervals do not affect the internal validity of each model but call for caution in interpreting the pace and direction of change in the birth spacing effect. In particular, the 11-year gap between 2010 and 2021 may conceal non-monotonic trends within this decade that cannot be detected using the available data; sub-period analyses were not feasible given data availability but are recommended for future research using civil registration or health facility data.

2.6. Verification of model assumptions

Prior to fitting the Cox models, the proportional hazards assumption was assessed for each covariate using Schoenfeld residuals and the scaled Schoenfeld residual test. Variables showing evidence of non-proportionality were handled through stratification or by including time-by-covariate interaction terms.

Collinearity among explanatory variables was examined using variance inflation factors (VIF): all VIF values were well below the critical threshold of 5 in multivariable analyses, ranging from 1.27 to 1.35 across surveys, indicating very low intercorrelation between the explanatory variables and confirming that each variable contributes distinct, non-redundant information to the model. Model selection proceeded stepwise across 10 successive models (M0–M9):

- (i) M0 estimated unadjusted effects.
- (ii) M1 added place of residence.
- (iii) M2 added maternal religion.
- (iv) M3 added maternal employment status.

- (v) M4 added household wealth.
- (vi) M5 added maternal age at childbirth.
- (vii) M6 added maternal education.
- (viii) M7 added child's birth weight.
- (ix) M8 added child's sex.
- (x) M9 added birth order.

This stepwise approach, consistent with the Mosley and Chen (1984) proximate determinants framework, allows the incremental contribution of each block of covariates to be assessed and the robustness of the birth spacing effect to be tracked as controls accumulate. All analyses were performed using Stata 16, with survey weights applied to account for the complex stratified two-stage sampling design of the DHS.

3. Results

3.1. Child survival by birth spacing (1993–2021)

Figure 1 shows that birth spacing was strongly associated with child survival in the 1990s, with a clear disadvantage for closely spaced births. Survival differentials narrowed in the late 1990s and early 2000s, but increased again in the most recent period, indicating a non-linear rather than a

steady attenuation over time.

3.2. Multivariable regression results: Dynamics of net effects

After adjusting for maternal, child, and household characteristics (M1–M8), the association between short birth intervals and child mortality remains significant throughout the study period, with a markedly non-linear trajectory: the HR for intervals below 24 months stood at 2.50 (95% confidence interval [CI]: 1.56–4.00) in 1993, declined to 1.49 (95% CI: 1.04–2.14) in 1998–1999, and 1.63 (95% CI: 1.27–2.10) in 2003, rose to 1.97 (95% CI: 1.64–2.38) in 2010, and reached its highest value in 2021 at 3.20 (95% CI: 2.28–4.49).

The protective advantage of long intervals (≥ 36 months) was significant from 1998 to 2010 but disappeared in 2021 (HR = 0.86, 95% CI: 0.63–1.19). These results reveal an enduring and, in the most recent period, amplified burden of short birth spacing on child survival. Rather than converging toward zero, the spacing effect resurged in 2021, likely reflecting the deterioration of health service access in the context of the ongoing security crisis. The comparison of survey-specific adjusted models confirms

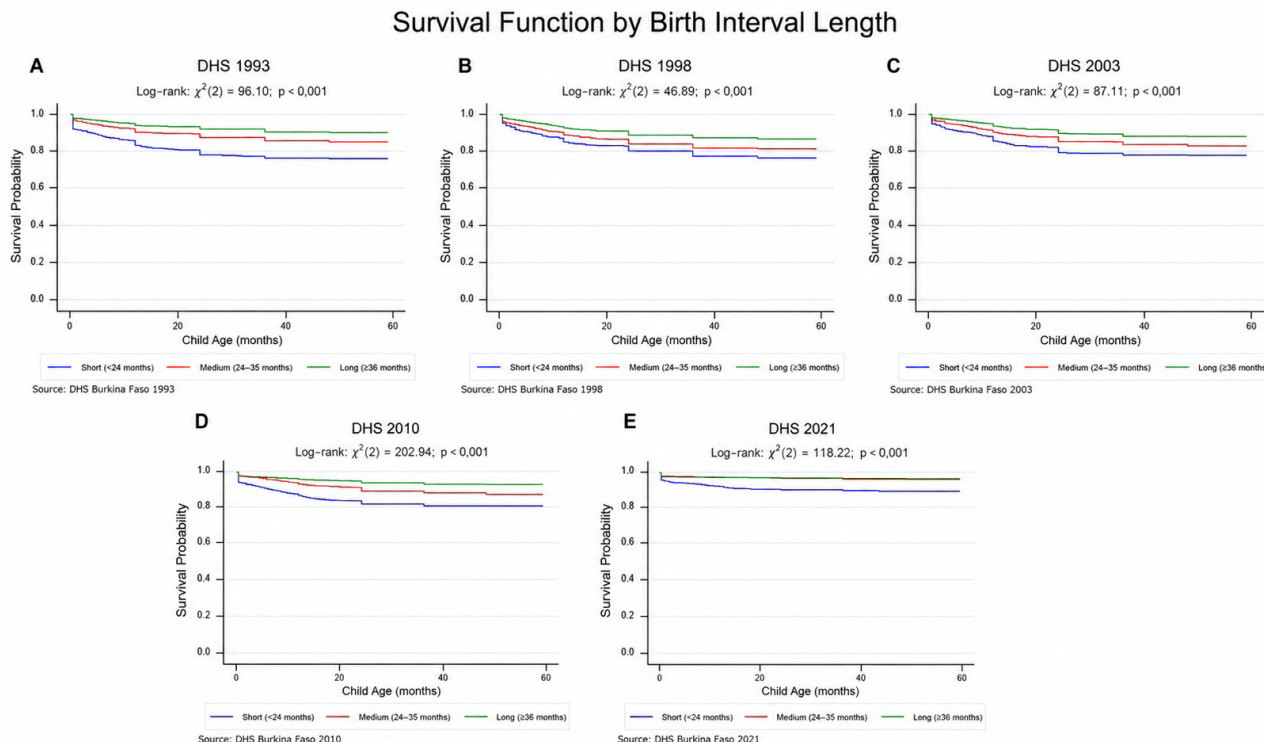


Figure 1. Survival rates of children under five by birth spacing, based on the Demographic and Health Survey (DHS) data for (A) 1993, (B) 1998–1999, (C) 2003, (D) 2010, and (E) 2021.

this non-linear pattern across the five survey rounds (Figure 2). Its trajectory remains, however, highly sensitive to contextual disruptions, as illustrated by the sharp resurgence observed in 2021.

3.3. Factors associated with under-five child survival

Beyond the association with birth spacing, several socioeconomic, demographic, and biological factors are significantly associated with under-five child survival in Burkina Faso. Their role has evolved over time, reflecting both structural progress and the persistence of contextual inequalities.

Place of residence is a key determinant whose effect has changed over time. In the fully adjusted model (M9), the urban protective effect was not statistically significant in any of the five survey rounds (1993: HR = 1.02, 95% CI: 0.59–1.75; 1998–1999: HR = 1.26, 95% CI: 0.76–2.11; 2003: HR = 0.77, 95% CI: 0.56–1.06; 2010: HR = 0.86, 95% CI: 0.64–1.16; 2021: HR = 0.93, 95% CI: 0.62–1.40), suggesting that once socioeconomic factors are controlled for, urban residence per se does not independently reduce child mortality risk (Table 1).

In partially adjusted models (M1), urban residence appeared protective in 1993 (HR = 0.74, 95% CI: 0.58–0.94), 2003 (HR = 0.70, 95% CI: 0.54–0.90), 2010 (HR = 0.72, 95% CI: 0.56–0.91), and 2021 (HR = 0.69, 95% CI: 0.50–0.95), but these effects attenuated and lost significance once all socioeconomic covariates were included, confirming that the urban advantage operates through education,

wealth, and access to care rather than through residence per se (Table A1). This trend reflects growing territorial disparities in access to healthcare services, vaccination, and improved living conditions.

The 2021 DHS reveals significant sub-national heterogeneity in under-five mortality rates: the Sud-Ouest region records the highest rate at 111 per 1,000 live births, followed by the Sahel at 98, both far above the national average of 48. Regions with the lowest rates include the Centre (39; which covers Ouagadougou), the Nord (40), and the Centre-Ouest (42). Intermediate levels are observed in the Cascades (77), Hauts-Bassins (61), Centre-Nord and Est (59 each). These regional disparities are compounded by the active security crisis, which since 2015 has disrupted health service delivery in the Sahel, Est, and Nord regions, displacing health workers and closing facilities.

Beyond insecurity, territorial inequalities in under-five mortality are structurally rooted in differences in physical access to health facilities, skilled birth attendance rates, access to safe water and nutrition, and vaccination coverage, all of which remain markedly lower in rural and conflict-affected zones than in urban centers. These conditions interact with birth spacing dynamics: in the most disadvantaged rural districts, the compensatory mechanisms that have reduced spacing-related risk in urban areas (e.g., medical supervision of closely-spaced pregnancies, postnatal follow-up, vaccination) remain insufficiently available, which is why the birth spacing effect is expected to persist longest among rural and vulnerable populations.

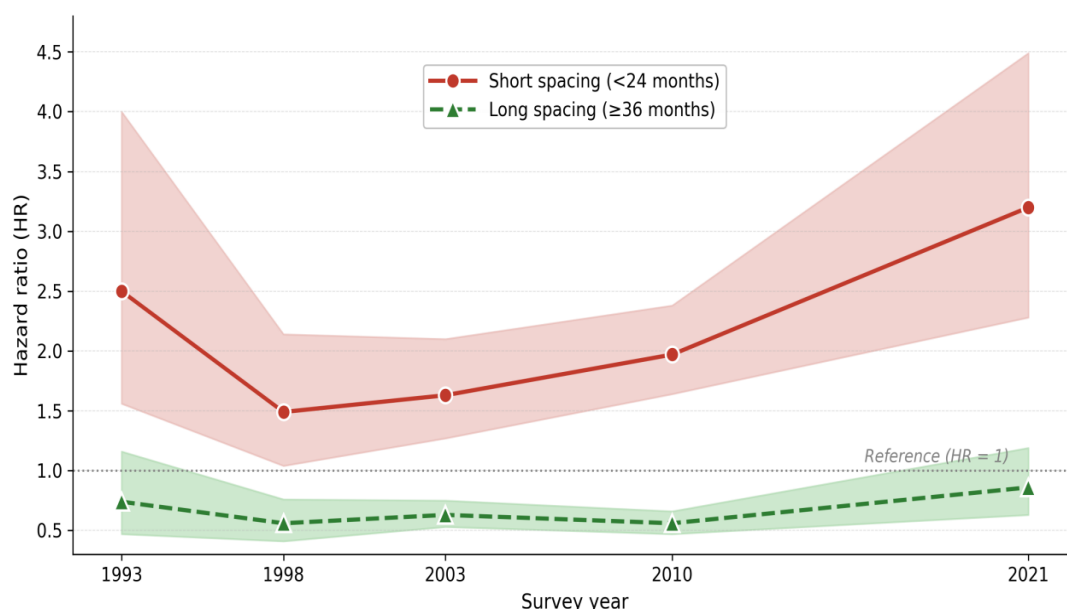


Figure 2. Temporal changes in the net effects of birth spacing on infant and child mortality. Shaded areas represent 95% confidence intervals. Hazard ratios (HRs) are relative to the 24–35 months reference category (M9; fully adjusted model).

Table 1. Adjusted hazard ratios (M9) and 95% confidence intervals (1993–2021)

Variable	Hazard ratio (95% confidence interval)				
	1993	1998–1999	2003	2010	2021
Interbirth interval					
<24 months	2.50*** (1.56–4.00)	1.49* (1.04–2.14)	1.63*** (1.27–2.10)	1.97*** (1.64–2.38)	3.20*** (2.28–4.49)
24–35 months	R	R	R	R	R
≥36 months	0.74 (0.47–1.16)	0.56*** (0.41–0.76)	0.63*** (0.53–0.75)	0.56*** (0.47–0.66)	0.86 (0.63–1.19)
Place of residence					
Urban	1.02 (0.59–1.75)	1.26 (0.76–2.11)	0.77 (0.56–1.06)	0.86 (0.64–1.16)	0.93 (0.62–1.40)
Rural	R	R	R	R	R
Religion					
Christian	0.89 (0.62–1.28)	0.96 (0.68–1.37)	1.46** (1.14–1.87)	0.97 (0.79–1.18)	0.83 (0.62–1.12)
Muslim	R	R	R	R	R
Other religion	1.88 (0.60–5.89)	1.31 (0.86–1.98)	1.37* (1.02–1.83)	1.34* (1.06–1.69)	1.33 (0.89–1.99)
Employment status					
Unemployed	0.99 (0.66–1.48)	0.95 (0.70–1.29)	0.90 (0.55–1.47)	1.27* (1.05–1.55)	0.94 (0.72–1.23)
Employed	R	R	R	R	R
Household wealth (reference category = Poor)					
Poor	R	R	R	R	R
Middle	1.17 (0.74–1.85)	0.99 (0.72–1.36)	0.97 (0.79–1.18)	0.90 (0.75–1.09)	0.95 (0.70–1.28)
Rich	0.77 (0.48–1.23)	1.14 (0.82–1.60)	1.00 (0.82–1.22)	0.79* (0.64–0.97)	0.77 (0.52–1.12)
Maternal age at childbirth					
<20 years	1.11 (0.46–2.70)	1.29 (0.58–2.86)	1.32 (0.88–1.98)	1.61* (1.08–2.41)	1.73* (0.90–3.33)
20–25 years	0.88 (0.53–1.46)	1.36* (0.95–1.96)	1.22 (0.97–1.53)	1.20* (0.97–1.49)	1.00 (0.72–1.38)
26–39 years	R	R	R	R	R
40 years and above	1.12 (0.54–2.30)	0.97 (0.51–1.82)	1.24 (0.93–1.66)	1.35* (0.98–1.86)	2.15*** (1.49–3.10)
Mother's education					
None	R	R	R	R	R
Primary	0.78 (0.44–1.37)	1.86* (1.08–3.18)	0.82 (0.58–1.14)	0.77 (0.55–1.07)	1.01 (0.67–1.50)
Secondary or higher	0.41 (0.12–1.45)	0.53 (0.16–1.74)	0.89 (0.47–1.66)	0.70 (0.39–1.27)	0.88 (0.54–1.43)
Child's birth weight					
<2,500 g (low)	3.45** (1.69–7.07)	1.07 (0.43–2.64)	0.92 (0.53–1.59)	1.68*** (1.30–2.16)	1.14 (0.64–2.04)
≥2,500 g	R	R	R	R	R
Child's sex					
Male	R	R	R	R	R
Female	0.88 (0.62–1.25)	1.13 (0.87–1.47)	1.02 (0.88–1.18)	0.85* (0.74–0.99)	0.81 (0.62–1.07)
Birth order					
Order 2	1.27 (0.67–2.39)	0.73 (0.45–1.18)	0.88 (0.65–1.19)	0.81 (0.61–1.06)	0.95 (0.67–1.35)
Order 3	1.41 (0.81–2.45)	0.98 (0.65–1.46)	0.94 (0.74–1.20)	0.96 (0.76–1.22)	0.52*** (0.36–0.74)
≥Order 4	R	R	R	R	R

Notes: Hazard ratio from the fully adjusted Cox proportional hazards model (M9), estimated separately for each survey round. “R” indicates reference category. Statistical significance: * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$, and **** $p < 0.001$.

Household wealth emerged as a major differentiating factor from 2010 onward. In 2021, in the fully adjusted model (M9), children from middle-income households had a 5% lower risk (HR = 0.95, 95% CI: 0.70–1.28), and those from wealthy households had a 23% lower risk (HR = 0.77, 95% CI: 0.52–1.12) compared to poor households, neither statistically significant (Table A1). The slightly stronger effects in partially adjusted models (M4: HR = 0.93, 95% CI: 0.69–1.25 for middle-income households; M5: HR = 0.75, 95% CI: 0.51–1.09 for wealthy households) reflect confounding by maternal education, which was not yet controlled at that stage: once education is included, the independent effect of wealth attenuates substantially, suggesting that the survival advantage of wealthier households operates largely through maternal education and access to care rather than wealth per se. These gaps, which were minimal or absent in the 1990s, indicate a rise in economic inequalities in child survival, though the fully adjusted estimates suggest that these inequalities are structurally mediated rather than direct.

Maternal education shows a consistent protective trend, although estimates in the fully adjusted model (M9) rarely reach conventional significance thresholds. In the fully adjusted model (M9), mothers with secondary education or higher reduced the risk of child death by 59% in 1993 (HR = 0.41, 95% CI: 0.12–1.45) (Table A1). The effect observed in 2021 could not be reliably estimated in M9 due to very small cell sizes for secondary-educated mothers in the analytical sample; the partially adjusted estimate (M6: HR = 0.87, 95% CI: 0.54–1.43) suggests a residual protective effect, but should be interpreted with caution. In contrast, primary education remains less decisive, with effects ranging from neutral to slightly protective depending on the period (e.g., in 2021, M9: HR = 1.01, 95% CI: 0.67–1.50), with no statistically significant association overall ($p > 0.05$).

From a biological perspective, birth weight emerges as a major predictor. In the fully adjusted model (M9), children with low birth weight show consistently elevated mortality risks across survey rounds: HR of 3.45 (95% CI: 1.69–7.07) in 1993, 0.92 (95% CI: 0.53–1.59) in 2003, 1.68 (95% CI: 1.30–2.16) in 2010, and 1.14 (95% CI: 0.64–2.04) in 2021 (Tables A1–A4). Although these estimates do not reach conventional significance thresholds in M9, likely due to reduced statistical power after full adjustment, the magnitude and consistency of the pattern confirm the extreme vulnerability of low-birthweight children.

It is worth noting that in the partially adjusted model M7 (birth weight entered without subsequent controls), HRs were 3.40 (95% CI: 1.67–6.95) in 1993, 0.92 (95% CI: 0.53–1.59) in 2003, 1.66 (95% CI: 1.29–2.14) in 2010,

and 1.15 (95% CI: 0.64–2.06) in 2021—the attenuation in M9 reflects partial mediation by birth order and child sex. Maternal age also accentuates disparities, in 2021, the M9 estimate for adolescent mothers confirms a substantial excess risk (HR = 1.73, 95% CI: 0.90–3.33), while older mothers (40–49 years) exhibited the strongest excess risk (HR = 2.15, 95% CI: 1.49–3.10, $p < 0.001$) (Table A5).

Overall, while closely spaced births remain a consistent source of vulnerability, under-five survival in Burkina Faso is increasingly shaped by socioeconomic, territorial, and biological conditions. These findings call for an integrated approach combining family planning, inequality reduction, and improvements in perinatal health.

4. Discussion

The findings of this study confirm well-established conclusions in the empirical literature: short interbirth intervals are significantly associated with higher infant mortality risk (Conde-Agudelo *et al.*, 2006; Rutstein, 2005). This relationship can be explained by several mechanisms, including maternal depletion, competition for resources, and higher perinatal risks (DaVanzo *et al.*, 2007). The descriptive and multivariable analyses conducted using Burkina Faso data confirm the robustness of these mechanisms during the early years of the study period (1993–2003).

One of the key contributions of this study is the precise quantification and documentation of the non-linear evolution of the impact of short birth intervals on child survival. In 1993, children born after intervals shorter than 24 months carried the highest excess mortality risk, a risk that declined substantially in the late 1990s and early 2000s before rising again in 2010 and resurging sharply in 2021, the most recent and most acute inflection in the series. Conversely, the survival advantage of long intervals, which was significant across most of the study period, disappeared entirely in the 2021 round. Rather than a convergence, what is observed is an asymmetric pattern: the risk associated with short intervals resurged in 2021, while the advantage of long intervals vanished simultaneously. This finding is consistent with recent studies showing a weakening of the relationship between birth spacing and mortality in contexts where access to healthcare has improved (Chekole *et al.*, 2025; Fotso *et al.*, 2013). This evolution can be partly explained by the expansion of contraceptive use, improvements in reproductive health policies, and increased female education (Institut National de la Statistique et de la Démographie & ICF, 2023). However, disparities persist. Rural women, poor households, and uneducated mothers continue to have shorter birth intervals compared to their

urban and educated counterparts. These results confirm the importance of socioeconomic determinants in shaping birth spacing dynamics, as highlighted by Cleland *et al.* (2012) and Bongaarts and Sinding (2011).

This evolution can also be attributed to several structural transformations: the sharp increase in skilled birth attendance (from 37% in 1993 to 96% in 2021; Institut National de la Statistique et de la Démographie & ICF, 2023); the expansion of immunization programs and malaria control interventions, which target major causes of child mortality (United Nations Children's Fund, 2023; World Health Organization, 2023), and the implementation of free healthcare policies for children under five and pregnant women (Ministry of Health, 2018). These structural changes suggest that stronger health systems mitigate the biological and social vulnerabilities associated with short birth intervals.

The study also shows that the advantage of long intervals (≥ 36 months), which became significant from 1998 to 2010, disappears entirely in 2021. This disappearance is consistent with a saturation of protective effects in contexts where health system coverage reduces perinatal risks across all spacing categories, but the persistence of elevated risk for short intervals simultaneously suggests that biological mechanisms of maternal depletion and resource competition remain operative. Similar trends have been observed in contexts with major health improvements, where better prevention of obstetric and neonatal complications reduces differences across spacing categories (Cleland *et al.*, 2012; Conde-Agudelo, 2000).

Despite the attenuation of the birth spacing effect, socioeconomic and territorial inequalities remain strong determinants of child mortality. Children born in rural areas, in poor households, or to uneducated mothers continue to face higher risks of death, confirming patterns observed in other countries in the region (Bocquier, 1991). This highlights that overall gains in survival are not equally distributed across social groups.

Beyond the direct association of birth spacing with child survival, several variables consistently emerge as important modifiers of this relationship. Place of residence remains a key explanatory factor, with a protective effect associated with urban settings, although this effect weakens in adjusted models. This finding is consistent with studies by United Nations Children's Fund (2019) and Coulibaly *et al.* (2024), which document persistent urban–rural differentials in under-five survival.

Similarly, maternal employment status does not appear to play a decisive role in moderating the association between birth spacing and child survival. Children of

employed mothers have mortality risks comparable to those of children of unemployed mothers, regardless of whether they are born after long or medium intervals. These findings suggest that, over time, differences between employed and unemployed mothers in terms of the association of optimal spacing with child survival tend to diminish. On one hand, unemployed mothers appear to have benefited from awareness campaigns, maternal and child health policies, and improved dissemination of good practices, enabling them to better care for their children. On the other hand, working mothers continue to face persistent constraints related to working conditions, domestic workload, and limited access to adequate childcare services, thereby limiting the expected positive effects of economic empowerment on child health. This observation is consistent with evidence from sub-Saharan Africa showing that when women's work is concentrated in precarious informal jobs and childcare options are scarce or of poor quality, the potential benefits of women's economic autonomy for child health are significantly reduced (Clark *et al.*, 2019).

Household wealth, introduced into the models from 2003 onward, shows a protective effect in crude models, particularly in 2010 and 2021. Children from wealthier households exhibit higher survival rates. However, this effect gradually diminishes in adjusted models, indicating that wealth is associated with child survival indirectly through intermediate determinants such as maternal education, access to healthcare, and the quality of the household environment. This finding reinforces the idea that wealth plays a moderating rather than a direct role in the association between birth spacing and child survival. A similar result from a study conducted in Niger (Hamadou Daouda, 2012) highlights this mediation, emphasizing that improvements in economic well-being alone are insufficient without parallel investments in health and education.

Maternal age at childbirth is a crucial determinant of under-five survival. The analyses show that children born to very young mothers (<20 years) face a higher risk of death, particularly between 1993 and 2003. This vulnerability may be attributed to adolescents' physiological immaturity, limited access to health information and reproductive services, and reduced decision-making autonomy. At the other end of the age spectrum, children born to older mothers (aged 40 and above) are also more exposed to mortality risks, regardless of the interbirth interval (Finlay *et al.*, 2011; Fofana, 2017). These findings confirm that extreme maternal ages constitute contexts of heightened vulnerability for child survival. Thus, maternal age has a dual effect: young age amplifies risks associated with short

intervals, while advanced age increases mortality risk regardless of spacing. This aligns with previous studies showing higher mortality risks among children born to mothers under 20 or over 35 compared to those born to mothers in intermediate age groups (Finlay *et al.*, 2011; Fofana, 2017).

Maternal education also plays a central role, as evidenced by multivariable analyses. It is associated with attenuation of the excess mortality risk linked to short birth intervals, through better access to healthcare and the adoption of appropriate health-seeking behaviors (Andriano & Monden, 2017; Cleland & Van Ginneken, 1988). However, this effect is constrained by household wealth, as highlighted by Fofana (2017) and others, indicating that education alone is insufficient when material conditions remain unfavorable.

Religion, as a cultural factor, also shows differentiated effects over time. In some surveys, children of Christian mothers exhibit lower mortality risks than those from other religious groups. Although this pattern varies across periods, it may reflect differences in access to reproductive health information, healthcare seeking behaviors, and attitudes toward contraception. However, these effects appear to be indirect, shaping the association between birth spacing and child survival through differentiated sociohealth practices (Costa *et al.*, 2020; Karlsson, 2019).

Finally, birth order—typically associated with higher mortality at higher parities—showed a more nuanced pattern. Children of intermediate birth orders (two or three) are sometimes more exposed, particularly in adjusted models for certain years such as 1993 and 2021. This may reflect an unfavorable combination of short birth intervals and limited maternal experience, or increased reproductive pressure in specific contexts, consistent with findings from other studies where higher-order births are generally more at risk (Ettarh & Kimani, 2012).

4.1. Operational implications and policy recommendations

These findings carry several operational implications. First, family planning interventions should be geographically and socially targeted: the greatest gains are achievable in poor rural households, districts with ongoing security crises, and communities with high concentrations of adolescent mothers, precisely those groups where risk convergence has been least pronounced. Second, the growing explanatory power of socioeconomic determinants calls for an integrated package that links birth spacing counseling with complementary interventions in nutrition, water and sanitation, vaccination, and access to skilled delivery care. Third, the marked rise of low birth weight as a predictor

in 2021 underscores the need for strengthened antenatal screening and management of high-risk pregnancies. Finally, the absence of comparable sub-national data and the study's cross-sectional design highlight the need for sustained, longitudinal surveillance systems and mixed-methods research to capture the micro-level pathways through which spacing and socioeconomic context jointly shape child survival.

4.2. Strengths and limitations of the study

Several limitations should be acknowledged. First, the data used rely on maternal reports, which may be subject to recall bias, particularly regarding the dates of children's births and deaths (Pullum, 2008). In addition, it was not possible to control for all biomedical factors that may influence child survival, such as maternal nutritional status, obstetric history, or neonatal conditions. Furthermore, although the semi-parametric Cox proportional hazards model is widely used, it relies on the proportional hazards assumption, which may be challenged in high-mortality settings (Allison, 2010).

Second, the study is based on cross-sectional data, which limits the ability to establish causal relationships between birth spacing and under-five mortality. It also does not distinguish between intended and unintended birth intervals, which may result from unplanned pregnancies or fertility issues, potentially affecting the interpretation of the findings. Moreover, some key contextual factors such as effective access to healthcare, social norms, or local health policies could not be included due to data limitations. Finally, methodological and contextual differences across successive DHS surveys may affect the temporal comparability of the results.

Third, the 2021 DHS was conducted in a particularly challenging security context that compromised the geographic completeness of the survey. Of the 600 clusters initially sampled, 86 (14.3%) could not be reached due to insecurity, with the Sahel and Est regions disproportionately affected: only 13 of 41 clusters (31.7%) were successfully surveyed in the Sahel and 27 of 41 (65.9%) in the Est, resulting in the complete exclusion of the provinces of Oudalan, Yagha, Tapoa, Komandjoari, and Kompienga from the analytical sample (Institut National de la Statistique et de la Démographie & ICF, 2023). Partial non-coverage also affected the Boucle du Mouhoun (7 out of 52 clusters unreached), the Cascades (8/36), the Centre-Nord (10/45), and the Nord (9/48). As a consequence, the 2021 estimates may underrepresent the most vulnerable and hard-to-reach populations, who are also likely to have the shortest birth intervals and the highest child mortality risks. This introduces a potential downward bias in

mortality estimates for 2021 and may affect the measured magnitude of the spacing effect in the most recent survey round. Future analyses should incorporate geospatial imputation methods or model-based small area estimation to account for systematically missing clusters in conflict-affected areas.

Despite these limitations, DHS data remain a robust and widely used source for studying child mortality in Africa (Rutstein, 2005), providing a solid empirical basis for analyzing trends and determinants of child survival.

5. Conclusion

This study analyzed the evolution of the association between birth spacing and under-five child survival in Burkina Faso using DHS data from 1993 to 2021. The results confirm that short interbirth intervals are consistently associated with excess child mortality across all five survey rounds, while medium intervals offer the best survival outcomes.

The findings do not support a simple convergence narrative. Rather, they reveal a non-linear trajectory in which the excess mortality risk associated with short intervals declined in the early 2000s before rising sharply in 2021, while the protective advantage of long intervals disappeared in the most recent survey. These patterns suggest that earlier gains in child survival remain highly sensitive to contextual disruption.

Socioeconomic and structural inequalities persist: children from poor households, rural areas, and mothers with no formal education remain disproportionately vulnerable. This confirms that birth spacing, while an important lever, cannot substitute for broader structural interventions targeting the social and territorial determinants of child mortality in Burkina Faso.

Health policies must therefore consolidate existing reproductive health gains while addressing these deeper inequalities through integrated, equity-oriented approaches.

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Conflict of interest

The author declares no competing interests.

Author contributions

This is a single-authored article.

Ethics approval and consent to participate

This study is based entirely on secondary analysis of de-identified, publicly available DHS data. No primary data collection involving human subjects was conducted. The DHS surveys were implemented with approval from relevant national ethics committees and Institutional Review Boards. As this study does not involve direct human subjects research, separate institutional ethics approval was not required. All analyses were conducted in compliance with the DHS data use agreement.

Consent for publication

Not applicable.

Availability of data

Data Availability Statement: The data used in this study are secondary data from the Demographic and Health Surveys (DHS) Programme in Burkina Faso (1993, 1998, 2003, 2010, 2021). These data are publicly available upon request from the DHS Programme website: <https://dhsprogram.com>. No proprietary data were used.

References

- Adedini, S. A., Anjorin, S. S., Mobolaji, J. W., Okolie, E. A., & Yaya, S. (2024). Assessing Africa's child survival gains and prospects for attaining SDG target on child mortality. *PLOS Global Public Health*, 4(7), e0003022.
<https://doi.org/10.1371/journal.pgph.0003022>
- Allison, P. D. (2010). *Survival analysis using SAS: A practical guide* (2nd ed.). SAS Institute. Accessed June 16, 2026. https://books.google.com/books/about/Survival_Analysis_Using_SAS.html?id=RmbZ2y1KLwUC
- Andriano, L., & Monden, C. W. S. (2019). The causal effect of maternal education on child mortality: Evidence from a quasi-experiment in Malawi and Uganda. *Demography*, 56(5), 1765–1790.
<https://doi.org/10.1007/s13524-019-00812-3>
- Black, R. E., Cousens, S., Johnson, H. L., Lawn, J. E., Rudan, I., Bassani, D. G., Jha, P., Campbell, H., Walker, C. F., Cibulskis, R., Eisele, T., Liu, L., & Mathers, C. (2010). Global, regional, and national causes of child mortality in 2008: A systematic analysis. *The Lancet*, 375(9730), 1969–1987.
[https://doi.org/10.1016/s0140-6736\(10\)60549-1](https://doi.org/10.1016/s0140-6736(10)60549-1)
- Bocquier, P. (1991). Les relations entre mortalité des enfants et espacement des naissances dans la banlieue de Dakar (Sénégal). *Population (French Edition)*, 46(4), 813.

- <https://doi.org/10.2307/1533297>
- Bongaarts, J., & Sinding, S. (2011). Population policy in transition in the developing world. *Science*, 333(6042), 574–576.
- <https://doi.org/10.1126/science.1207558>
- Chekole, F. A., Tesfu, A. A., Beyene, F. Y., & Balcha, W. F. (2025). The effect of antenatal care on low birth weight and neonatal mortality: Protocol for umbrella review of meta-analysis. *Journal of Health, Population and Nutrition*, 44(1).
- <https://doi.org/10.1186/s41043-025-00904-4>
- Clark, S., Kabiru, C. W., Laszlo, S., & Muthuri, S. (2019). The impact of childcare on poor urban women's economic empowerment in Africa. *Demography*, 56(4), 1247–1272.
- <https://doi.org/10.1007/s13524-019-00793-3>
- Cleland, J., Conde-Agudelo, A., Peterson, H., Ross, J., & Tsui, A. (2012). Contraception and health. *The Lancet*, 380(9837), 149–156.
- [https://doi.org/10.1016/S0140-6736\(12\)60609-6](https://doi.org/10.1016/S0140-6736(12)60609-6)
- Cleland, J. G., & Van Ginneken, J. K. (1988). Maternal education and child survival in developing countries: The search for pathways of influence. *Social Science & Medicine*, 27(12), 1357–1368.
- [https://doi.org/10.1016/0277-9536\(88\)90201-8](https://doi.org/10.1016/0277-9536(88)90201-8)
- Conde-Agudelo, A. (2000). Maternal morbidity and mortality associated with interpregnancy interval: Cross sectional study. *BMJ*, 321(7271), 1255–1259.
- <https://doi.org/10.1136/bmj.321.7271.1255>
- Conde-Agudelo, A., Rosas-Bermúdez, A., & Kafury-Goeta, A. C. (2006). Birth spacing and risk of adverse perinatal outcomes: A meta-analysis. *JAMA*, 295(15), 1809.
- <https://doi.org/10.1001/jama.295.15.1809>
- Costa, J. C., Weber, A. M., Darmstadt, G. L., Abdalla, S., & Vitoria, C. G. (2020). Religious affiliation and immunization coverage in 15 countries in Sub-Saharan Africa. *Vaccine*, 38(5), 1160–1169.
- <https://doi.org/10.1016/j.vaccine.2019.11.024>
- Coulibaly, A., Baguiya, A., Meda, I. B., Millogo, T., Koumbem, A. M. A., Garanet, F., & Kouanda, S. (2024). Prevalence and factors associated with short birth interval in the semi-rural community of Kaya, Burkina Faso: Results of a community-based survey. *Archives of Public Health*, 82(1).
- <https://doi.org/10.1186/s13690-024-01398-y>
- DaVanzo, J., Hale, L., Razzaque, A., & Rahman, M. (2007). Effects of interpregnancy interval and outcome of the preceding pregnancy on pregnancy outcomes in Matlab, Bangladesh. *BJOG: An International Journal of Obstetrics & Gynaecology*, 114(9), 1079–1087.
- <https://doi.org/10.1111/j.1471-0528.2007.01338.x>
- Ettarh, R. R., & Kimani, J. (2012). Determinants of under-five mortality in rural and urban Kenya. *Rural and Remote Health*, 12(1), 3–11.
- <https://doi.org/10.22605/RRH1812>
- Finlay, J. E., Özaltin, E., & Canning, D. (2011). The association of maternal age with infant mortality, child anthropometric failure, diarrhoea and anaemia for first births: Evidence from 55 low-and middle-income countries. *BMJ Open*, 1(2), e000226. Accessed June 16, 2026. <https://bmjopen.bmj.com/content/1/2/e000226.short>
- Fofana, A. (2017). *Déterminants de la mortalité néonatale au Burkina Faso* [Determinants of neonatal mortality in Burkina Faso]. Master's thesis, Université Saint Thomas d'Aquin. Accessed June 16, 2026. <https://www.ifris-bf.org/wp-content/uploads/2018/12/20-FOFANA-D%C3%A9terminants-de-la-mortalit%C3%A9-n%C3%A9onatale-au-Burkina.pdf>
- Fotso, J. C., Cleland, J., Mberu, B., Mutua, M., & Elungata, P. (2013). Birth spacing and child mortality: An analysis of prospective data from the Nairobi urban health and demographic surveillance system. *Journal of Biosocial Science*, 45(6), 779–798.
- <https://doi.org/10.1017/S0021932012000570>
- Hamadou Daouda, Y. (2012). Déterminants de la mortalité infantile et infanto-juvénile et la pauvreté au Niger [Determinants of Infant and Child Mortality and Poverty in Niger]. *Revue d'Économie Théorique et Appliquée*, 2(1), 23–47.
- <https://doi.org/10.62519/reta.v2n1a2>
- Institut National de la Statistique et de la Démographie. (1994). *Burkina Faso: Enquête démographique et de santé 1993* [Demographic and Health Survey]. Accessed June 16, 2026. <https://www.dhsprogram.com/pubs/pdf/FR53/FR53.pdf>
- Institut National de la Statistique et de la Démographie, & ICF. (2023). *Enquête démographique et de santé du Burkina Faso 2021* [Demographic and Health Survey]. Accessed June 16, 2026. https://www.insd.bf/sites/default/files/2023-09/EDS%20Burkina%20Faso%202021_VF_07-09-23.pdf
- Institut National de la Statistique et de la Démographie, & Macro International Inc. (2000). *Burkina Faso: Enquête démographique et de santé, 1998-1999* [Demographic and Health Survey]. Accessed June 16, 2026. <https://microdata.worldbank.org/index.php/catalog/1332/>
- Institut National de la Statistique et de la Démographie, & ORC Macro. (2004). *Burkina Faso: Enquête démographique et de santé 2003* [Demographic and Health Survey]. Accessed June 16, 2026. <https://demostaf.web.ined.fr/index.php/catalog/103>
- Karlsson, O. (2019). Religion and child health in West and Central Africa. *Population and Development Review*, 45(4), 707–738.
- <https://doi.org/10.1111/padr.12296>

- Ministère de la santé. (2011). *Plan National de Développement Sanitaire (PNDS) 2011-2020* [National Health Development Plan 2011-2020].
- Ministère de la santé. (2018). *Compte rendu de la quatrième rencontre bilan national de mise en œuvre des mesures de gratuité des soins* [Report of the fourth national review meeting on the implementation of free healthcare measures]. Accessed June 16, 2026. https://p4h.world/app/uploads/2023/02/Rapport_4C3A8me20rencontre20bilan20gratuitC3A9_VF.x23411.pdf
- Mosley, W. H., & Chen, L. C. (1984). An analytical framework for the study of child survival in developing countries. *Population and Development Review*, 10(Suppl), 25–45.
<https://doi.org/10.2307/2807954>
- Organisation des Nations Unies. (2015). *Transformer notre monde: Le Programme de développement durable à l'horizon 2030*. Accessed June 16, 2026. https://unctad.org/system/files/official-document/ares70d1_fr.pdf
- Pullum, T. W. (2008). *An assessment of the quality of data on health and nutrition in the DHS surveys, 1993-2003* (DHS Methodological Reports No. 6). Accessed June 16, 2026. <https://dhsprogram.com/publications/publication-mr6-methodological-reports.cfm>
- Rutstein, S. O. (2005). Effects of preceding birth intervals on neonatal, infant and under-five years mortality and nutritional status in developing countries: Evidence from the demographic and health surveys. *International Journal of Gynecology & Obstetrics*, 89(S1).
<https://doi.org/10.1016/j.ijgo.2004.11.012>
- Smith, G. C., Pell, J. P., & Dobbie, R. (2003). Interpregnancy interval and risk of preterm birth and neonatal death: Retrospective cohort study. *BMJ*, 327(7410), 313.
<https://doi.org/10.1136/bmj.327.7410.313>
- United Nations Children's Fund. (2019). *The state of the world's children 2019: Children, food and nutrition*. Accessed June 16, 2026. <https://www.UNICEF.org/reports/state-of-worlds-children-2019>
- United Nations Inter-agency Group for Child Mortality Estimation (UN IGME). (2026). *Levels and trends in child mortality: Report 2025*. United Nations Children's Fund (UNICEF). Accessed June 16, 2026. <https://data.UNICEF.org/resources/levels-and-trends-in-child-mortality-2025/>
- World Health Organization. (2007). *Report of a WHO technical consultation on birth spacing: Geneva, Switzerland, 13-15 June 2005*. World Health Organization. Accessed June 16, 2026. <https://iris.who.int/server/api/core/bitstreams/58173bae-0761-4d56-aebf-1e959f88da1d/content>
- World Health Organization. (2023). *World health statistics 2023: Monitoring health for the SDGs, sustainable development goals*. Accessed June 16, 2026. <https://www.who.int/publications/i/item/9789240074323>

Appendix

Table A1. Results of Cox proportional hazards regression models for 1993

Variable	Unadjusted model			Adjusted models							
	M0	M1	M2	M3	M4	M5	M6	M7	M8	M9	
Interbirth interval (months)											
<24	2.33*** [1.80–3.01]	2.31*** [1.78–2.99]	2.31*** [1.79–2.99]	2.33*** [1.80–3.02]	2.63*** [1.70–4.06]	2.55*** [1.62–4.01]	2.53*** [1.60–4.01]	2.56*** [1.61–4.01]	2.54*** [1.60–4.04]	2.50*** [1.56–4.00]	
24–35	R	R	R	R	R	R	R	R	R	R	
≥36	0.60*** [0.47–0.77]	0.60*** [0.47–0.77]	0.60*** [0.47–0.77]	0.61*** [0.47–0.77]	0.75ns [0.48–1.17]	0.76ns [0.49–1.18]	0.76ns [0.49–1.18]	0.76ns [0.49–1.18]	0.76ns [0.49–1.18]	0.74ns [0.47–1.16]	
Place of residence											
Rural	–	R	R	R	R	R	R	R	R	R	
Urban	–	0.74* [0.58–0.94]	0.73* [0.57–0.93]	0.73* [0.57–0.93]	1.02ns [0.60–1.71]	1.01ns [0.60–1.70]	1.12ns [0.66–1.89]	1.04ns [0.61–1.77]	1.03ns [0.60–1.76]	1.02ns [0.59–1.75]	
Religion											
Muslim	–	–	R	R	R	R	R	R	R	R	
Christian	–	–	0.94ns [0.78–1.13]	0.94ns [0.78–1.14]	0.89ns [0.62–1.28]	0.89ns [0.63–1.27]	0.91ns [0.64–1.31]	0.90ns [0.63–1.30]	0.90ns [0.63–1.29]	0.89ns [0.62–1.28]	
Other religions	–	–	1.15ns [0.36–3.61]	1.14ns [0.36–3.58]	2.16ns [0.73–6.38]	2.21ns [0.74–6.58]	2.20ns [0.74–6.60]	1.87ns [0.56–6.18]	1.89ns [0.58–6.13]	1.88ns [0.60–5.89]	
Mother's employment status											
Employed	–	–	–	R	R	R	R	R	R	R	
Unemployed	–	–	–	0.96ns [0.78–1.19]	0.98ns [0.66–1.45]	0.98ns [0.66–1.45]	0.98ns [0.66–1.46]	0.97ns [0.66–1.43]	0.98ns [0.66–1.44]	0.99ns [0.66–1.48]	
Household wealth											
Poor	–	–	–	–	R	R	R	R	R	R	
Upper-middle class	–	–	–	–	1.16ns [0.74–1.84]	1.16ns [0.73–1.84]	1.16ns [0.73–1.83]	1.16ns [0.74–1.83]	1.16ns [0.74–1.83]	1.17ns [0.74–1.85]	
Rich	–	–	–	–	0.77ns [0.49–1.23]	0.78ns [0.49–1.25]	0.80ns [0.50–1.27]	0.76ns [0.48–1.22]	0.76ns [0.48–1.22]	0.77ns [0.48–1.23]	
(Cont'd...)											

(Cont'd...)

Table A1. (Continued)

Variable	Unadjusted model		Adjusted models								
	M0	M1	M2	M3	M4	M5	M6	M7	M8	M9	
Age at childbirth (years)											
<20	-	-	-	-	-	1.36ns [0.64–2.91]	1.36ns [0.63–2.91]	1.36ns [0.63–2.93]	1.38ns [0.64–2.96]	1.11ns [0.46–2.70]	
20–25	-	-	-	-	-	1.07ns [0.70–1.63]	1.08ns [0.71–1.64]	1.06ns [0.69–1.61]	1.06ns [0.69–1.63]	0.88ns [0.53–1.46]	
26–39	-	-	-	-	-	R	R	R	R	R	
40–49	-	-	-	-	-	1.07ns [0.52–2.23]	1.07ns [0.51–2.21]	1.08ns [0.52–2.24]	1.07ns [0.52–2.21]	1.12ns [0.54–2.30]	
Mother's level of education											
None	-	-	-	-	-	-	R	R	R	R	
Elementary school	-	-	-	-	-	-	0.82ns [0.45–1.48]	0.78ns [0.44–1.37]	0.78ns [0.44–1.36]	0.78ns [0.44–1.37]	
Secondary or higher	-	-	-	-	-	-	0.41ns [0.12–1.41]	0.45ns [0.13–1.53]	0.45ns [0.13–1.52]	0.41ns [0.12–1.45]	
Child's birth weight											
<2,500 g	-	-	-	-	-	-	-	3.40** [1.67–6.95]	3.49** [1.71–7.15]	3.45** [1.69–7.07]	
≥2,500 g	-	-	-	-	-	-	-	R	R	R	
Child's gender											
Male	-	-	-	-	-	-	-	-	R	R	
Female	-	-	-	-	-	-	-	-	0.87ns [0.62–1.22]	0.88ns [0.62–1.25]	
Birth order of the child											
Order 2	-	-	-	-	-	-	-	-	-	1.27ns [0.67–2.39]	
Order 3	-	-	-	-	-	-	-	-	-	1.41ns [0.81–2.45]	
Order ≥4	-	-	-	-	-	-	-	-	-	R	

Notes: Values are hazard ratios with 95% confidence intervals in brackets. Statistical significance: ^a $p < 0.10$, * $p < 0.05$, ** $p < 0.01$, *** $p < 0.001$. ns = not significant. R = reference category.

Table A2. Results of Cox proportional hazards regression models for 1998–1999

Variable	Unadjusted model			Adjusted models								
	M0	M1	M2	M3	M4	M5	M6	M7	M8	M9		
Interbirth interval (months)												
<24	1.48** [1.19–1.85]	1.47*** [1.18–1.84]	1.47** [1.18–1.83]	1.46** [1.17–1.83]	1.46* [1.04–2.06]	1.47* [1.04–2.09]	1.51* [1.06–2.14]	1.50* [1.06–2.14]	1.52* [1.06–2.18]	1.49* [1.04–2.14]		
24–35	R	R	R	R	R	R	R	R	R	R		
≥36	0.60*** [0.49–0.74]	0.61*** [0.50–0.75]	0.62*** [0.50–0.75]	0.62*** [0.51–0.76]	0.54*** [0.40–0.73]	0.55*** [0.41–0.75]	0.55*** [0.41–0.75]	0.55*** [0.41–0.75]	0.55*** [0.41–0.75]	0.56*** [0.41–0.76]		
Place of residence												
Rural	–	R	R	R	R	R	R	R	R	R		
Urban	–	0.64** [0.48–0.86]	0.67** [0.49–0.90]	0.67** [0.49–0.90]	1.35ns [0.84–2.17]	1.33ns [0.83–2.14]	1.25ns [0.75–2.08]	1.25ns [0.75–2.08]	1.25ns [0.75–2.08]	1.26ns [0.76–2.11]		
Religion												
Muslim	–	–	R	R	R	R	R	R	R	R		
Christian	–	–	1.19ns [0.95–1.48]	1.21a [0.97–1.49]	0.99ns [0.70–1.40]	0.99ns [0.69–1.40]	0.97ns [0.68–1.38]	0.97ns [0.68–1.38]	0.97ns [0.69–1.38]	0.96ns [0.68–1.37]		
Other religions	–	–	1.39* [1.04–1.86]	1.41* [1.05–1.90]	1.29ns [0.85–1.96]	1.31ns [0.86–2.00]	1.31ns [0.86–1.99]	1.31ns [0.86–1.99]	1.31ns [0.86–1.99]	1.31ns [0.86–1.98]		
Mother's employment status												
Employed	–	–	–	R	R	R	R	R	R	R		
Unemployed	–	–	–	0.88ns [0.72–1.08]	0.95ns [0.69–1.30]	0.95ns [0.69–1.29]	0.94ns [0.69–1.28]	0.94ns [0.69–1.29]	0.94ns [0.69–1.28]	0.95ns [0.70–1.29]		
Household wealth												
Poor	–	–	–	–	R	R	R	R	R	R		
Upper-middle class	–	–	–	–	1.02ns [0.75–1.39]	1.02ns [0.75–1.39]	1.01ns [0.74–1.38]	1.01ns [0.74–1.38]	1.00ns [0.73–1.36]	0.99ns [0.72–1.36]		
Rich	–	–	–	–	1.13ns [0.81–1.56]	1.14ns [0.83–1.59]	1.14ns [0.82–1.59]	1.14ns [0.82–1.59]	1.14ns [0.82–1.59]	1.14ns [0.82–1.60]		
(Cont'd...)												

(Cont'd...)

Table A2. (Continued)

Variable	Unadjusted model		Adjusted models								
	M0	M1	M2	M3	M4	M5	M6	M7	M8	M9	
Age at childbirth (years)											
<20	-	-	-	-	-	1.03ns [0.54–1.96]	0.99ns [0.53–1.87]	0.99ns [0.53–1.87]	1.00ns [0.53–1.90]	1.29ns [0.58–2.86]	
20–25	-	-	-	-	-	1.24ns [0.91–1.70]	1.22ns [0.89–1.65]	1.22ns [0.89–1.65]	1.21ns [0.89–1.65]	1.36a [0.95–1.96]	
26–39	-	-	-	-	-	R	R	R	R	R	
40–49	-	-	-	-	-	0.99ns [0.52–1.87]	0.98ns [0.52–1.87]	0.98ns [0.52–1.87]	0.98ns [0.52–1.85]	0.97ns [0.51–1.82]	
Mother's level of education											
None/Uneducated	-	-	-	-	-	-	-	-	-	-	
Elementary school	-	-	-	-	-	-	1.88* [1.09–3.24]	1.87* [1.10–3.20]	1.88* [1.10–3.22]	1.86* [1.08–3.18]	
Secondary or higher	-	-	-	-	-	-	0.48ns [0.15–1.55]	0.48ns [0.15–1.55]	0.48ns [0.15–1.55]	0.53ns [0.16–1.74]	
Child's birth weight											
<2,500 g	-	-	-	-	-	-	-	1.05ns [0.43–2.55]	1.03ns [0.42–2.53]	1.07ns [0.43–2.64]	
≥2,500 g	-	-	-	-	-	-	-	R	R	R	
Child's gender											
Male	-	-	-	-	-	-	-	-	R	R	
Female	-	-	-	-	-	-	-	-	1.12ns [0.86–1.45]	1.13ns [0.87–1.47]	
Birth order of the child											
Order 2	-	-	-	-	-	-	-	-	-	0.73ns [0.45–1.18]	
Order 3	-	-	-	-	-	-	-	-	-	0.98ns [0.65–1.46]	
Order ≥4	-	-	-	-	-	-	-	-	-	R	

Notes: Values are hazard ratios with 95% confidence intervals in brackets. Statistical significance: ^a $p < 0.10$, * $p < 0.05$, ** $p < 0.01$, *** $p < 0.001$. ns = not significant. R = reference category.

Table A3. Results of Cox proportional hazards regression models for 2003

Variable	Unadjusted model		Adjusted models								
	M0		M1	M2	M3	M4	M5	M6	M7	M8	M9
Interbirth interval (months)											
<24	1.72*** [1.33–2.21]		1.70*** [1.32–2.20]	1.67*** [1.29–2.16]	1.67*** [1.29–2.16]	1.67*** [1.30–2.16]	1.64*** [1.28–2.14]	1.64*** [1.28–2.11]	1.64*** [1.28–2.10]	1.64*** [1.28–2.10]	1.63*** [1.27–2.10]
24–35	R	R	R	R	R	R	R	R	R	R	R
≥26	0.61*** [0.51–0.72]		0.62*** [0.52–0.73]	0.62*** [0.52–0.74]	0.62*** [0.52–0.74]	0.62*** [0.52–0.74]	0.63*** [0.53–0.74]	0.63*** [0.53–0.74]	0.63*** [0.53–0.74]	0.63*** [0.53–0.74]	0.63*** [0.53–0.75]
Place of residence											
Rural	–	R	R	R	R	R	R	R	R	R	R
Urban	–		0.70** [0.54–0.90]	0.70** [0.55–0.91]	0.72* [0.55–0.95]	0.72* [0.53–0.98]	0.73* [0.53–0.99]	0.76a [0.55–1.05]	0.77ns [0.56–1.05]	0.76ns [0.56–1.05]	0.77ns [0.56–1.06]
Religion											
Muslim	–	–	R	R	R	R	R	R	R	R	R
Christian	–	–		1.49** [1.16–1.91]	1.49** [1.16–1.91]	1.49** [1.16–1.91]	1.48** [1.16–1.89]	1.46** [1.14–1.87]	1.46** [1.14–1.87]	1.46** [1.14–1.87]	1.46** [1.14–1.87]
Other religions	–	–		1.40* [1.05–1.88]	1.40* [1.05–1.87]	1.39* [1.04–1.86]	1.39* [1.04–1.86]	1.37* [1.03–1.83]	1.37* [1.02–1.83]	1.37* [1.02–1.83]	1.37* [1.02–1.83]
Mother's employment status											
Employed	–	–	–	–	R	R	R	R	R	R	R
Unemployed	–	–	–	–	0.90ns [0.56–1.47]	0.90ns [0.56–1.47]	0.90ns [0.55–1.47]	0.90ns [0.55–1.46]	0.90ns [0.55–1.47]	0.90ns [0.56–1.47]	0.90ns [0.55–1.47]
Household wealth											
Poor	–	–	–	–	–	R	R	R	R	R	R
Upper-middle class	–	–	–	–	–	0.96ns [0.78–1.17]	0.96ns [0.79–1.18]	0.96ns [0.79–1.18]	0.96ns [0.79–1.18]	0.96ns [0.79–1.18]	0.97ns [0.79–1.18]
Rich	–	–	–	–	–	0.98ns [0.80–1.20]	0.99ns [0.81–1.21]	1.00ns [0.82–1.22]	1.00ns [0.82–1.22]	1.00ns [0.82–1.22]	1.00ns [0.82–1.22]

(Cont'd...)

Table A3. (Continued)

Variable	Unadjusted model		Adjusted models							
	M0	M1	M2	M3	M4	M5	M6	M7	M8	M9
Age at childbirth (years)										
<20	-	-	-	-	-	1.18ns [0.86–1.61]	1.18ns [0.87–1.62]	1.19ns [0.87–1.62]	1.19ns [0.87–1.62]	1.32ns [0.88–1.98]
20–25	-	-	-	-	-	1.13ns [0.96–1.33]	1.14ns [0.96–1.34]	1.14ns [0.96–1.34]	1.14ns [0.96–1.34]	1.22a [0.97–1.53]
26–39	-	-	-	-	-	R	R	R	R	R
40–49	-	-	-	-	-	1.27ns [0.95–1.69]	1.26ns [0.94–1.68]	1.26ns [0.94–1.68]	1.26ns [0.94–1.68]	1.24ns [0.93–1.66]
Mother's level of education										
None	-	-	-	-	-	-	R	R	R	R
Elementary school	-	-	-	-	-	-	0.81ns [0.58–1.13]	0.81ns [0.58–1.13]	0.81ns [0.58–1.13]	0.82ns [0.58–1.14]
Secondary or higher	-	-	-	-	-	-	0.86ns [0.46–1.59]	0.86ns [0.46–1.58]	0.86ns [0.46–1.58]	0.89ns [0.47–1.66]
Child's birth weight										
<2,500 g	-	-	-	-	-	-	-	0.92ns [0.53–1.59]	0.92ns [0.53–1.59]	0.92ns [0.53–1.59]
≥2,500 g	-	-	-	-	-	-	-	R	R	R
Child's gender										
Male	-	-	-	-	-	-	-	-	R	R
Female	-	-	-	-	-	-	-	-	1.02ns [0.89–1.18]	1.02ns [0.88–1.18]
Birth order of the child										
Order 2	-	-	-	-	-	-	-	-	-	0.88ns [0.65–1.19]
Order 3	-	-	-	-	-	-	-	-	-	0.94ns [0.74–1.20]
Order ≥4	-	-	-	-	-	-	-	-	-	R

Notes: Values are hazard ratios with 95% confidence intervals in brackets. Statistical significance: ^a $p < 0.10$, * $p < 0.05$, ** $p < 0.01$, *** $p < 0.001$. ns = not significant. R = reference category.

Table A4. Results of Cox proportional hazards regression models for 2010

Variable	Unadjusted model		Adjusted models								
	M0		M1	M2	M3	M4	M5	M6	M7	M8	M9
Interbirth interval (months)											
<24	2.07*** [1.72–2.49]		2.07*** [1.71–2.49]	2.04*** [1.69–2.46]	2.01*** [1.67–2.42]	2.00*** [1.66–2.41]	1.98*** [1.64–2.38]	1.97*** [1.64–2.38]	1.97*** [1.64–2.38]	1.98*** [1.65–2.39]	1.97*** [1.64–2.38]
24–35	R		R	R	R	R	R	R	R	R	R
≥36	0.53*** [0.45–0.62]		0.54*** [0.46–0.63]	0.54*** [0.46–0.64]	0.55*** [0.47–0.64]	0.55*** [0.47–0.65]	0.55*** [0.47–0.65]	0.55*** [0.47–0.66]	0.55*** [0.47–0.66]	0.56*** [0.47–0.66]	0.56*** [0.47–0.66]
Place of residence											
Rural	–		R	R	R	R	R	R	R	R	R
Urban	–		0.72** [0.56–0.91]	0.71** [0.55–0.91]	0.69** [0.54–0.90]	0.81ns [0.61–1.07]	0.81ns [0.61–1.07]	0.87ns [0.65–1.17]	0.85ns [0.63–1.14]	0.85ns [0.63–1.14]	0.86ns [0.64–1.16]
Religion											
Muslim	–		–	R	R	R	R	R	R	R	R
Christian	–		–	0.92ns [0.76–1.12]	0.95ns [0.79–1.16]	0.94ns [0.77–1.13]	0.94ns [0.77–1.14]	0.96ns [0.79–1.17]	0.97ns [0.79–1.18]	0.97ns [0.79–1.18]	0.97ns [0.79–1.18]
Other religions	–		–	1.32* [1.05–1.67]	1.40** [1.11–1.77]	1.33* [1.05–1.68]	1.34* [1.06–1.70]	1.34* [1.06–1.69]	1.35* [1.06–1.71]	1.35* [1.07–1.71]	1.34* [1.06–1.69]
Mother's employment status											
Employed	–		–	–	R	R	R	R	R	R	R
Unemployed	–		–	–	1.27* [1.04–1.55]	1.24* [1.02–1.51]	1.25* [1.03–1.52]	1.26* [1.03–1.53]	1.27* [1.05–1.55]	1.28* [1.05–1.55]	1.27* [1.05–1.55]
Household wealth											
Poor	–		–	–	–	R	R	R	R	R	R
Upper-middle class	–		–	–	–	0.89ns [0.74–1.08]	0.90ns [0.75–1.09]	0.91ns [0.75–1.09]	0.90ns [0.74–1.08]	0.90ns [0.74–1.08]	0.90ns [0.75–1.09]
Rich	–		–	–	–	0.76* [0.62–0.94]	0.77* [0.62–0.95]	0.79* [0.64–0.97]	0.78* [0.64–0.96]	0.78* [0.64–0.96]	0.79* [0.64–0.97]

(Cont'd...)

Table A4. (Continued)

Variable	Unadjusted model		Adjusted models								
	M0	M1	M2	M3	M4	M5	M6	M7	M8	M9	
Age at childbirth (years)											
<20	-	-	-	-	-	1.34a [0.99–1.80]	1.34a [0.99–1.80]	1.34a [1.00–1.81]	1.35a [1.00–1.82]	1.61* [1.08–2.41]	
20–25	-	-	-	-	-	1.10ns [0.93–1.29]	1.10ns [0.94–1.30]	1.10ns [0.94–1.30]	1.10ns [0.94–1.30]	1.20a [0.97–1.49]	
26–39	-	-	-	-	-	R	R	R	R	R	
40–49	-	-	-	-	-	1.39* [1.01–1.91]	1.37a [1.00–1.89]	1.37a [1.00–1.88]	1.37a [1.00–1.89]	1.35a [0.98–1.86]	
Mother's level of education											
None	-	-	-	-	-	-	R	R	R	R	
Elementary school	-	-	-	-	-	-	0.75ns [0.54–1.06]	0.76ns [0.54–1.06]	0.76ns [0.54–1.07]	0.77ns [0.55–1.07]	
Secondary or higher	-	-	-	-	-	-	0.68ns [0.38–1.22]	0.67ns [0.37–1.20]	0.66ns [0.37–1.19]	0.70ns [0.39–1.27]	
Child's birth weight											
<2,500 g	-	-	-	-	-	-	-	1.66*** [1.29–2.14]	1.66*** [1.29–2.14]	1.68*** [1.30–2.16]	
≥2,500 g	-	-	-	-	-	-	-	R	R	R	
Child's gender											
Male	-	-	-	-	-	-	-	-	R	R	
Female	-	-	-	-	-	-	-	-	0.85* [0.74–0.99]	0.85* [0.74–0.99]	
Birth order of the child											
Order 2	-	-	-	-	-	-	-	-	-	0.81ns [0.61–1.06]	
Order 3	-	-	-	-	-	-	-	-	-	0.96ns [0.76–1.22]	
Order ≥4	-	-	-	-	-	-	-	-	-	R	

Notes: Values are hazard ratios with 95% confidence intervals in brackets. Statistical significance: ^a $p < 0.10$, * $p < 0.05$, ** $p < 0.01$, *** $p < 0.001$. ns = not significant. R = reference category.

Table A5. Results of Cox proportional hazards regression models for 2021

Variable	Unadjusted model			Adjusted models						
	M0	M1	M2	M3	M4	M5	M6	M7	M8	M9
Interbirth interval (months)										
<24	3.35*** [2.40–4.67]	3.37*** [2.41–4.71]	3.33*** [2.38–4.66]	3.33*** [2.38–4.66]	3.32*** [2.37–4.64]	3.20*** [2.29–4.49]	3.19*** [2.28–4.47]	3.20*** [2.28–4.49]	3.20*** [2.28–4.49]	3.20*** [2.28–4.49]
24–35	R	R	R	R	R	R	R	R	R	R
≥36	0.84ns [0.62–1.15]	0.87ns [0.64–1.19]	0.88ns [0.64–1.19]	0.87ns [0.64–1.19]	0.89ns [0.65–1.21]	0.85ns [0.61–1.17]	0.85ns [0.61–1.17]	0.85ns [0.61–1.16]	0.85ns [0.62–1.17]	0.86ns [0.63–1.19]
Place of residence										
Rural	–	R	R	R	R	R	R	R	R	R
Urban	–	0.69* [0.50–0.95]	0.71* [0.52–0.98]	0.71* [0.52–0.98]	0.87ns [0.58–1.31]	0.89ns [0.60–1.33]	0.91ns [0.60–1.36]	0.91ns [0.61–1.37]	0.91ns [0.62–1.37]	0.93ns [0.62–1.40]
Religion										
Muslim	–	–	R	R	R	R	R	R	R	R
Christian	–	–	0.85ns [0.63–1.14]	0.84ns [0.63–1.13]	0.83ns [0.62–1.11]	0.82ns [0.62–1.10]	0.83ns [0.62–1.12]	0.83ns [0.62–1.12]	0.83ns [0.62–1.11]	0.83ns [0.62–1.12]
Other religions	–	–	1.44a [0.96–2.16]	1.43a [0.95–2.15]	1.39ns [0.93–2.07]	1.33ns [0.89–2.01]	1.33ns [0.89–2.01]	1.33ns [0.89–2.00]	1.33ns [0.89–1.99]	1.33ns [0.89–1.99]
Mother's employment status										
Employed	–	–	–	R	R	R	R	R	R	R
Unemployed	–	–	–	0.95ns [0.72–1.24]	0.93ns [0.72–1.22]	0.94ns [0.72–1.22]	0.94ns [0.72–1.22]	0.94ns [0.72–1.22]	0.94ns [0.72–1.23]	0.94ns [0.72–1.23]
Household wealth										
Poor	–	–	–	–	R	R	R	R	R	R
Upper-middle class	–	–	–	–	0.93ns [0.69–1.25]	0.93ns [0.69–1.26]	0.94ns [0.69–1.27]	0.94ns [0.69–1.27]	0.94ns [0.69–1.27]	0.95ns [0.70–1.28]
Rich	–	–	–	–	0.72a [0.49–1.06]	0.75ns [0.51–1.09]	0.76ns [0.52–1.11]	0.76ns [0.52–1.11]	0.76ns [0.52–1.11]	0.77ns [0.52–1.12]
(Cont'd...)										

(Contd...)

Table A5. (Continued)

Variable	Unadjusted model			Adjusted models						
	M0	M1	M2	M3	M4	M5	M6	M7	M8	M9
Age at childbirth (years)										
<20	-	-	-	-	-	1.65a [0.91–2.97]	1.66a [0.91–3.00]	1.66a [0.92–3.00]	1.66a [0.92–3.01]	1.73a [0.90–3.33]
20–25	-	-	-	-	-	0.88ns [0.65–1.20]	0.89ns [0.65–1.20]	0.89ns [0.65–1.20]	0.88ns [0.65–1.20]	1.00ns [0.72–1.38]
26–39	-	-	-	-	-	R	R	R	R	R
40–49	-	-	-	-	-	2.33*** [1.63–3.33]	2.32*** [1.62–3.32]	2.32*** [1.62–3.32]	2.31*** [1.62–3.31]	2.15*** [1.49–3.10]
Mother's level of education										
None	-	-	-	-	-	-	R	R	R	R
Elementary school	-	-	-	-	-	-	1.01ns [0.68–1.51]	1.01ns [0.68–1.51]	1.02ns [0.68–1.52]	1.01ns [0.67–1.50]
Secondary or higher	-	-	-	-	-	-	0.87ns [0.54–1.43]	0.87ns [0.54–1.42]	0.88ns [0.54–1.43]	0.88ns [0.54–1.43]
Child's birth weight										
<2,500 g	-	-	-	-	-	-	-	1.15ns [0.64–2.06]	1.15ns [0.64–2.06]	1.14ns [0.64–2.04]
≥2,500 g	-	-	-	-	-	-	-	R	R	R
Child's gender										
Male	-	-	-	-	-	-	-	-	R	R
Female	-	-	-	-	-	-	-	-	0.81ns [0.62–1.05]	0.81ns [0.62–1.07]
Birth order of the child										
Order 2	-	-	-	-	-	-	-	-	-	0.95ns [0.67–1.35]
Order 3	-	-	-	-	-	-	-	-	-	0.52*** [0.36–0.74]
Order ≥4	-	-	-	-	-	-	-	-	-	R

Notes: Values are hazard ratios with 95% confidence intervals in brackets. Statistical significance: ^a $p < 0.10$, * $p < 0.05$, ** $p < 0.01$, *** $p < 0.001$. ns = not significant. R = reference category.