Beyond the National Decline: Inequality, Union Formation, and Contraceptive Protection in Ethiopia's Adolescent Reproductive Transition, 2000–2024/25

Full manuscript

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Abstract

Background. Ethiopia has achieved a substantial decline in teenage childbearing, but national progress may conceal persistent social and geographic inequality and does not by itself explain how schooling, household resources, union formation, and contraceptive protection intersect across the adolescent life course. We examined Ethiopia's adolescent reproductive transition over a quarter century, moving beyond national and regional trends to integrate socioeconomic inequality, temporal sequencing of first union and pregnancy, contraceptive protection and unmet need, pre-pregnancy contraceptive exposure, and multivariable associations.

Methods. We analyzed nationally representative women's Individual Recode data from the 2000, 2005, 2011, 2016, and 2024–25 Ethiopia Demographic and Health Surveys (EDHS). The principal longitudinal outcome was begun childbearing among all interviewed women aged 15–19 years, defined consistently as at least one live birth or current pregnancy with a first child. Survey weights, primary sampling units, and strata were incorporated throughout. We estimated national and regional prevalence, interval-specific and long-run change, and inequalities by residence, education, and wealth. First-union timing was compared with estimated conception among adolescent mothers. Modern contraceptive use and unmet need were examined among currently married/in-union adolescents, while contraceptive calendars from 2005 onward were used to reconstruct method status immediately before the index pregnancy. For 2024–25, survey-weighted logistic regression estimated adjusted associations and prespecified education-by-union, wealth-by-union, education-by-region, and wealth-by-region interactions.

Results. Begun childbearing declined from 16.3% (95% CI 14.4–18.2) in 2000 to 9.9% (8.3–11.4) in 2024–25, a 39.3% relative reduction; the decline was episodic rather than continuous. Regional trajectories diverged sharply, with Afar remaining high while Amhara experienced the largest long-run reduction. Residence inequality narrowed, but substantial educational and wealth gradients persisted. Among adolescent mothers, first union preceded or coincided with estimated conception in 91.8% in 2000 and 84.5% in 2024–25, indicating that union remained the dominant context of first childbearing while its coupling with pregnancy weakened. Among currently married/in-union adolescents, modern contraceptive use rose from 3.0% to 32.7% and unmet need fell from 36.9% to 14.6%. Yet only about 3–4% of adolescents who had begun childbearing had a modern method recorded one month before estimated conception in calendar-bearing surveys. In 2024–25, any formal education was associated with lower adjusted odds before marital-status adjustment (AOR 0.44, 95% CI 0.26–0.75) but not afterward (AOR 0.89, 0.44–1.82); low wealth remained associated with higher adjusted odds (AOR 2.00, 1.03–3.88). Education interacted with current union status (p=.026): adjusted childbearing prevalence differed sharply by education outside current union but was nearly identical among adolescents currently in union.

Conclusions. Ethiopia's national decline in teenage childbearing represents substantial progress, but the transition remains unequal and strongly structured by age, household resources, geography, and especially union formation. Rising contraceptive coverage among married adolescents has coincided with falling unmet need, yet pregnancies that occurred were overwhelmingly concentrated in periods without recorded contraceptive protection. The findings support a transition-centered policy framework that integrates school continuation, delayed union, poverty-sensitive prevention, adolescent-responsive contraception, and services for both married and unmarried adolescents.

Keywords: adolescent childbearing; teenage pregnancy; adolescent reproductive transition; child marriage; union formation; contraception; unmet need; education; wealth inequality; Demographic and Health Survey; Ethiopia

1. Introduction

Adolescence is a period in which biological maturation intersects with transitions in schooling, family formation, social identity, and economic opportunity. Pregnancy and childbearing during this period are therefore not isolated reproductive events; they can be both consequences and accelerators of broader life-course transitions. Adolescent mothers face higher risks of eclampsia, puerperal endometritis, and systemic infection than women aged 20–24 years, while their infants face higher risks of low birth weight, preterm birth, and severe neonatal conditions (World Health Organization [WHO], 2024). Early pregnancy can also interrupt education, narrow employment opportunities, and reinforce socioeconomic disadvantage. These health and social consequences make prevention of early pregnancy, together with appropriate care for pregnant and parenting adolescents, a continuing public-health and development priority.

The global burden remains substantial despite long-run progress. WHO estimates that approximately 21 million girls aged 15–19 years in low- and middle-income countries experience pregnancy each year, resulting in about 12 million births; roughly half of these pregnancies are unintended (WHO, 2024). The global adolescent birth rate fell markedly between 2000 and 2023, but the pace of decline has been uneven, and sub-Saharan Africa continues to have the highest regional adolescent birth rate (WHO, 2024). The updated WHO guideline consequently frames adolescent pregnancy prevention as a multisectoral problem involving child marriage, schooling, sexual coercion, contraceptive access and continuation, and access to quality maternal health services rather than as a problem solvable by a single clinical intervention (WHO, 2025).

A life-course perspective is especially useful because the determinants of adolescent childbearing are temporally and socially interconnected. Continued schooling may delay union formation and first birth, while early marriage or cohabitation can increase exposure to pregnancy and alter fertility expectations, autonomy, and contraceptive negotiation. Poverty can constrain educational continuity and access to services while increasing vulnerability to early union formation. Contraceptive need and use also vary across reproductive states: adolescents who are never sexually active, sexually active but unmarried, currently married, pregnant, or postpartum do not constitute a single risk population. WHO's current guidance therefore places particular emphasis on preventing child marriage and improving adolescents' access to and sustained use of contraception (WHO, 2025).

These intersections are highly relevant in Ethiopia. The country has undergone major changes in fertility, schooling, urbanization, family planning, marriage patterns, and reproductive-health services since 2000. Yet the adolescent reproductive transition has been uneven. Earlier analysis of the 2000–2016 EDHS documented a decline in the proportion of women aged 15–19 who had begun childbearing from 16.3% to 12.5%, while identifying older adolescent age, early sexual initiation, marital status, and geography as important correlates (Kassa et al., 2019a). Other analyses of the 2016 EDHS similarly found strong associations with age, education, marital status, household or community socioeconomic position, and place of residence or region (Birhanu et al., 2019; Kefale et al., 2020). Spatial analyses have further demonstrated that adolescent pregnancy is geographically clustered rather than evenly distributed across Ethiopia (Tigabu et al., 2021; Tebeje et al., 2024).

The 2024–25 EDHS extends this evidence by nearly a decade and allows a more comprehensive assessment of what lies beneath the national trend. Using a harmonized historical definition, the proportion of women aged 15–19 who had begun childbearing declined from 16.3% in 2000 to 9.9% in 2024–25, but the decline occurred in two distinct episodes rather than as a smooth secular reduction. Regional trajectories also diverged markedly: Amhara saw the largest long-run reduction, whereas Afar showed little sustained improvement and remained the highest-prevalence region in 2024–25. The 2024–25 regional range—from about 2% in Addis Ababa to 23% in Afar—shows that national improvement and reproductive equity are not synonymous (Ethiopian Statistical Service [ESS] & ICF, 2025).

Geographic divergence, however, is only one dimension of the transition. National decline can coexist with persistent or widening inequality by education and household resources. Education is particularly complex for ages 15–19 because educational attainment is constrained by age and intertwined with school continuation, union formation, and pregnancy. Likewise, current marital status is not an ordinary baseline covariate: union may precede conception, occur after pregnancy begins, or dissolve before interview. Treating marriage simply as a contemporaneous predictor can therefore obscure the temporal sequence that gives the association meaning. UNICEF’s work in Ethiopia similarly emphasizes that child marriage is sustained by interacting gender norms, educational constraints, poverty, and shocks, rather than by a single determinant (UNICEF, 2022).

Contraception poses a related measurement problem. Current contraceptive use among all adolescents is not a valid proxy for contraceptive protection at the time a prior pregnancy began. Current use may be observed after pregnancy, birth, postpartum return to fecundity, or a change in marital status. Conversely, modern contraceptive prevalence among currently married adolescents and unmet need among the same population answer important but different programmatic questions. A pregnancy-centered analysis requires reconstructing contraceptive status before conception when contraceptive-calendar data permit. This distinction is essential if the aim is to understand whether adolescent pregnancies occurred primarily during periods of contraceptive use or during periods without recorded protection.

The present study therefore conceptualizes Ethiopia’s adolescent reproductive transition as a sequence of linked but analytically distinct domains: social and educational context; entry into union; contraceptive protection and unmet need; pregnancy and childbearing; and the interaction of these processes across geographic and socioeconomic settings. This approach deliberately moves beyond a conventional prevalence-and-predictors model. It combines repeated cross-sectional trend analysis with inequality measurement, event sequencing, contraceptive-calendar reconstruction, and survey-weighted multivariable modeling, while preserving temporal caution where the DHS design cannot establish causal order.

The study had six objectives: (1) quantify the nationally comparable trend in begun childbearing from 2000 through 2024–25 and determine whether change was continuous or episodic; (2) characterize long-run regional trajectories and contemporary geographic inequality; (3) examine how residence, education, and household wealth inequalities changed over time on both absolute and relative scales;

(4) distinguish the cross-sectional marital composition of adolescent childbearing from the temporal ordering of first union and estimated conception; (5) assess changes in modern contraceptive use and unmet need among married adolescents and reconstruct contraceptive exposure immediately before index pregnancies; and (6) estimate adjusted 2024–25 associations and test whether education and wealth patterning differs by current union status or region. By integrating these domains, the study asks not only whether teenage childbearing declined, but how the social and reproductive architecture surrounding adolescent childbearing changed—and which inequalities remain after a quarter century of national progress.

2. Methods

2.1 Study design and data sources

We conducted a repeated cross-sectional secondary analysis of nationally representative data from the 2000, 2005, 2011, 2016, and 2024–25 Ethiopia Demographic and Health Surveys (EDHS). The analysis used the women's Individual Recode (IR) files. The DHS Program employs standardized questionnaires and probability sampling procedures designed to produce nationally and subnationally representative demographic and health estimates (Central Statistical Authority [Ethiopia] & ORC Macro, 2001; Central Statistical Agency [Ethiopia] & ORC Macro, 2006; Central Statistical Agency [Ethiopia] & ICF International, 2012; Central Statistical Agency [Ethiopia] & ICF, 2016; ESS & ICF, 2025). The 2019 Mini EDHS was not included in the principal five-round historical series because the analytic program was designed around the full-survey sequence and the harmonized modules required for the extended analyses were not uniformly available in that round.

2.2 Study population and age framework

The principal analytic population comprised all interviewed women aged 15–19 years, regardless of current or previous marital status. The unweighted samples for the harmonized childbearing outcome were 3,584 in 2000, 3,252 in 2005, 3,835 in 2011, 3,498 in 2016, and 4,841 in 2024–25, for a combined unweighted sample of 19,010 adolescents. Corresponding design-weighted analytic denominators were approximately 3,710, 3,266, 4,009, 3,381, and 4,875, respectively. Because DHS weights are normalized within surveys, use these weighted denominators for design-weighted estimation; do not interpret them as population counts.

The analysis concerns older adolescents aged 15–19, not the full WHO adolescent age range of 10–19. The EDHS women's individual questionnaire does not provide an equivalent nationally representative reproductive history for girls aged 10–14 across all five surveys. Within the 15–19 range, we retained single-year age for the 2024–25 multivariable analysis because reproductive exposure and the opportunity to complete schooling change rapidly between ages 15 and 19.

2.3 Primary outcome: historically comparable begun childbearing

The principal longitudinal outcome was a binary indicator of whether an adolescent had begun childbearing. A woman aged 15–19 was classified as having begun childbearing if she had experienced at least one live birth or, having had no prior live birth, was currently pregnant with her first child. Adolescents meeting neither condition were classified as not having begun childbearing. This definition matches the conventional DHS indicator used in earlier Ethiopian surveys and supports direct comparison over time (Central Statistical Authority [Ethiopia] & ORC Macro, 2001; Kassa et al., 2019a).

A deliberate distinction was made for 2024–25. The 2024–25 EDHS also reports a broader indicator of ever having been pregnant that includes pregnancy loss. That measure is valuable for contemporary surveillance but is not definitionally interchangeable with the historical begun-childbearing series. We therefore reconstructed the historical indicator directly from individual records and used it for every temporal analysis. For descriptive decomposition, begun childbearing was separated into two mutually exclusive components: at least one previous live birth and current first pregnancy without a previous live birth.

2.4 Geographic harmonization

Two geographic frameworks were used. For longitudinal regional comparisons, an 11-region historical geography was retained: Tigray, Afar, Amhara, Oromia, Somali, Benishangul-Gumuz, historical Southern Nations, Nationalities and Peoples' Region (SNNPR), Gambella, Harari, Addis Ababa, and Dire Dawa. For 2024–25, Central Ethiopia, Sidama, South Ethiopia, and South West Ethiopia were pooled to reconstruct historical SNNPR. The reconstruction was audited against unweighted denominators, design-weighted denominators, and design-weighted outcome numerators.

A second 2024–25 analysis retained the contemporary regional geography, including the southern successor regions separately. Historical aggregation was used only to protect longitudinal comparability; contemporary disaggregation was used for current epidemiologic and policy interpretation. Newly constituted regions were not assigned retrospective trends that the earlier surveys cannot support.

2.5 Socioeconomic inequality measures

We examined inequalities by residence, education, and household wealth. Residence was classified as urban or rural. Education was harmonized as no formal education versus any formal education. This binary contrast was chosen because four-level attainment categories are strongly constrained by age among 15–19-year-olds and can partly encode opportunity to have reached a given schooling level rather than educational exposure itself. Household wealth was contrasted as low wealth (bottom two DHS wealth quintiles, Q1–Q2) versus high wealth (top two quintiles, Q4–Q5); the middle quintile was excluded to preserve a distinct lower-versus-upper socioeconomic comparison. A comparable wealth series was analyzed from 2005 onward.

For each survey, group-specific prevalence was estimated using the complex survey design. Absolute inequality was defined as the survey-weighted prevalence difference in percentage points: rural minus urban, no education minus any education, and low wealth minus high wealth. Absolute gaps and their standard errors were estimated with survey-weighted identity-link quasibinomial models saturated for the two-group contrast. Relative inequality was expressed as a prevalence ratio and estimated with survey-weighted log-link quasipoisson models. Reporting both scales was prespecified because a narrowing percentage-point gap can coexist with a widening prevalence ratio when prevalence declines at different proportional rates across groups.

2.6 Marriage, union formation, and event sequencing

Marriage and union formation were analyzed in three stages. First, begun childbearing was described by current marital status. Second, the current marital composition of adolescents who had begun childbearing was estimated. These two analyses are cross-sectional and were not interpreted as establishing whether union preceded pregnancy.

Third, temporal sequencing was reconstructed among adolescent mothers using DHS century-month codes (CMCs). First union/cohabitation timing was obtained from the DHS first-union date variable. For adolescents with a first live birth, estimated conception was approximated as first-birth CMC minus nine months. Union was classified as preceding or coinciding with estimated conception when first-union CMC was less than or equal to estimated conception CMC. Additional categories distinguished conception before first union followed by union before or by first birth, first birth before first union, and no union by interview. A ±1-month sensitivity analysis around estimated conception was used because month-based birth histories cannot identify the exact day of conception. Current first pregnancies were examined as supporting evidence using interview CMC and reported pregnancy duration, but sparse sequence categories were not used for strong trend inference.

A survey-weighted logistic calendar-time model, with time scaled in decades, assessed whether the probability that first union preceded or coincided with estimated conception changed over the five surveys. The event-sequencing analysis was given greater interpretive weight than current marital status because it directly addresses temporal order.

2.7 Contraceptive protection and unmet need

Contraceptive analysis used denominator-specific measures rather than a single adolescent contraceptive prevalence. Modern contraceptive prevalence (mCPR) was estimated among currently married/in-union women aged 15–19. A secondary descriptive series estimated mCPR among sexually active unmarried adolescents, defined as unmarried women aged 15–19 reporting sexual intercourse within the preceding 30 days. The latter estimates were retained primarily as descriptive evidence because several survey-specific denominators and numbers of modern-method users were small.

Unmet need was estimated among currently married/in-union adolescents. The preferred revised DHS unmet-need variable (v626a) was available from 2011 onward; the 2000 and 2005 files used the legacy v626 measure. The five-round series is therefore interpreted as evidence of broad historical change rather than as perfectly definition-invariant. Survey-weighted logistic models with calendar time scaled in decades estimated long-run trends in married-adolescent mCPR and unmet need.

2.8 Pre-pregnancy contraceptive-calendar reconstruction

To avoid treating contraceptive use measured after pregnancy as a predictor of that pregnancy, we reconstructed contraceptive exposure before the index pregnancy using DHS contraceptive calendars in the 2005, 2011, 2016, and 2024–25 surveys. The index pregnancy was defined as the pregnancy leading to the first live birth among adolescent mothers or the current first pregnancy among adolescents with no prior live birth. For a first live birth, estimated conception was first-birth CMC minus nine months; for a current first pregnancy, conception was approximated from interview CMC minus reported pregnancy duration.

Calendar orientation was empirically validated against known first-live-birth months before exposure was classified. The selected reverse mapping reproduced known birth months with hit rates of approximately 99.8–100% across the four calendar-bearing surveys. The primary exposure was contraceptive status one month before estimated conception; three months before conception was analyzed as a sensitivity window. Calendar cells were classified as modern method, traditional method, or no method using DHS calendar codes. Because the analysis begins with pregnancies that occurred, the resulting percentages are exposure distributions among index pregnancies, not contraceptive-failure rates. Estimating failure would require person-time at risk among method users and nonusers, which this analysis does not provide.

2.9 National and regional trend analysis

For each survey, national prevalence of begun childbearing and 95% confidence intervals were estimated. Absolute and relative changes were calculated for the full 2000–2024/25 period and for adjacent survey intervals (2000–2005, 2005–2011, 2011–2016, and 2016–2024/25). Because survey rounds are independent cross-sectional samples, standard errors for absolute differences were derived from survey-specific variance estimates; relative-change confidence intervals were obtained using the delta method.

For overall temporal inference, records from all five surveys were pooled and calendar time was modeled continuously in survey-design quasibinomial logistic regression, scaled so that the exponentiated coefficient represents the odds ratio per decade. Survey round was also modeled categorically, and a design-adjusted Wald test assessed heterogeneity without imposing linearity. Equivalent per-decade trend models were estimated separately for each historical region. Endpoint change and five-round linear trend were treated as distinct inferential questions.

For pooled analyses, survey identifiers were incorporated into PSU and stratum identifiers so that identically numbered sampling units from different EDHS rounds were not treated as the same clusters. Weights were normalized within survey round for pooled regression so that arbitrary between-survey differences in DHS weight scaling did not determine each round's contribution.

2.10 2024–25 multivariable prevalence models

The multivariable analysis used the 2024–25 EDHS and the same historically comparable begun-childbearing outcome. Survey-weighted logistic regression was selected as the principal model because adjusted odds ratios are familiar in the epidemiologic literature and permit direct comparison with prior Ethiopian studies (Birhanu et al., 2019; Kassa et al., 2019a; Kefale et al., 2020). All models incorporated DHS weights, PSUs, and strata.

Model 1 represented structural/background characteristics and included single-year age as a categorical factor (15 as reference), residence, contemporary region, binary education (no education versus any formal education), and binary wealth (low Q1–Q2 versus high Q4–Q5). Model 2 added current marital status (never married/in union, currently married/in union, formerly married/in union). The middle wealth quintile was intentionally excluded from the primary regression contrast. Consequently, the full 2024–25 outcome-validation sample of 4,841 adolescents was reduced to 4,031 adolescents with 415 childbearing events for the regression analysis.

Model 1 and Model 2 were interpreted sequentially. Changes in education and wealth coefficients after adding marital status were described as attenuation or amplification, not mediation, because current marital status is temporally ambiguous: union may precede conception for some adolescents but follow pregnancy initiation for others. Current contraceptive use and unmet need were excluded from the principal determinants model because they are measured at interview and can occur after the pregnancy or birth defining the outcome.

2.11 Prespecified interaction analyses and adjusted predictions

Four interaction blocks were prespecified: education × current union status, wealth × current union status, education × contemporary region, and wealth × contemporary region. For interaction models, union status was simplified to currently in union versus not currently in union; education remained no education versus any formal education; and wealth remained low versus high. Each interaction block was evaluated with an overall design-adjusted Wald test rather than relying on isolated coefficient p-values.

When an interaction was statistically supported and sufficiently stable, standardized adjusted predicted prevalences and 95% confidence intervals were calculated to aid interpretation. Cell-level unweighted denominators and event counts were audited before substantive interpretation. Region-by-education and region-by-wealth models were additionally screened for aliased coefficients, unusually large absolute coefficients, and large standard errors as indicators of sparse-cell instability or separation. Statistically significant global interaction tests were not treated as sufficient evidence for precise region-specific estimates when diagnostics indicated sparse cells.

2.12 Survey design, precision, and statistical inference

DHS women's sampling weights (v005) were divided by 1,000,000. Primary sampling units were identified using v021 and strata using v022, consistent with DHS guidance for complex-sample estimation (The DHS Program, 2018). Survey-design objects incorporated weights, clustering, and stratification, with appropriate handling of strata containing a single sampled PSU. Unweighted sample sizes and event counts were retained as precision diagnostics.

Regional and interaction estimates were interpreted through confidence intervals and event counts rather than rank ordering alone. For regional descriptive analyses, estimates based on fewer than 100 unweighted adolescents or fewer than 25 childbearing events were flagged for caution. For interaction-cell audits, cells with fewer than 25 observations or fewer than 10 events were considered very sparse, and cells with 25–49 observations or 10–24 events were treated cautiously. Statistical tests were two-sided with alpha=.05. Because the analyses were prespecified and address related but distinct scientific questions, p-values were not mechanically adjusted for multiple comparisons; emphasis was placed on effect size, confidence intervals, global interaction tests, and consistency across complementary analyses.

2.13 Validation and quality assurance

Validation preceded substantive interpretation. The historically comparable begun-childbearing outcome was independently reconstructed in each survey and compared with published EDHS values for 2000, 2005, 2011, and 2016. Reconstructed estimates differed from the published values by less than 0.05 percentage points in every historical round. The same definition yielded 9.87% (95% CI 8.31–11.44) in 2024–25. Regional aggregation was required to reproduce the national series, and the reconstructed historical SNNPR in 2024–25 was required to reproduce exactly the sum of the four contemporary successor regions.

The 2024–25 married-adolescent mCPR reconstruction (32.68%) and unmet-need reconstruction (14.57%) were checked against certified survey values of 32.7% and 14.6%, respectively. Contraceptive-calendar orientation was validated empirically against known first-birth months before preconception exposure estimates were accepted. Multivariable models underwent sample-retention, sparse-cell, coefficient-stability, and prediction-range audits. The restricted low-versus-high wealth regression sample was not expected to reproduce the full-sample national prevalence exactly because Q3 was intentionally excluded.

2.14 Statistical software

Data management and statistical analyses were conducted in R. The haven and dplyr packages were used for data import and management, and the survey package was used for design-weighted estimation, variance estimation, regression, and design-adjusted hypothesis testing. The survey package implements design-based methods for complex samples using Taylor linearization and related survey estimators (Lumley, 2004). Figures were produced in R with publication-oriented graphics. Estimates are generally reported to one decimal place for prevalence and to greater precision for inferential statistics where appropriate.

2.15 Ethical considerations

This study used de-identified secondary data from the Ethiopia Demographic and Health Surveys. The original surveys followed DHS ethical review, informed-consent, and confidentiality procedures. Access to the datasets was obtained under applicable DHS data-use requirements. The present analysis involved no direct contact with participants and made no attempt to identify individual respondents.

2.16 Interpretation framework

The study was designed to estimate population patterns and temporal associations, not causal effects. Repeated cross-sectional surveys can establish whether prevalence and inequality changed across survey rounds, but they do not follow the same adolescents over time. Education, wealth, residence, and region are interpreted as social and geographic correlates. Current marital status is treated as both a major life-course context and a temporally ambiguous state; event-sequencing analyses are used to clarify whether union preceded estimated conception among adolescent mothers. Current contraception is not interpreted as an antecedent of prior pregnancy. Pre-pregnancy calendar exposure is temporally better ordered but remains conditioned on having experienced an index pregnancy and therefore does not estimate contraceptive effectiveness or failure. This layered interpretation was prespecified to avoid assigning causal meaning to associations that the DHS design cannot identify.

2.17 Analytic domains and principal estimands

Domain	Primary population	Principal estimand
National trend	All women 15–19	Begun-childbearing prevalence; absolute/relative change; OR per decade
Regional trend	All women 15–19	Historical-region prevalence and OR per decade; contemporary 2024–25 prevalence
Socioeconomic inequality	All women 15–19 within contrast	Absolute prevalence gap and prevalence ratio
Union sequencing	Adolescent mothers / current first pregnancies	First union before/on estimated conception
Contraceptive protection	Currently married/in-union adolescents	mCPR and unmet need; OR per decade
Pre-pregnancy exposure	Adolescents with observable index pregnancy calendar	Modern/traditional/no method 1 month before conception
Multivariable model	2024–25 low/high wealth regression sample	AORs; design-adjusted Wald interaction tests; adjusted predicted prevalence

3. Results

3.1 Analytic sample and validation of the harmonized outcome

Across the five EDHS rounds, the harmonized childbearing analysis included 3,584 adolescents in 2000, 3,252 in 2005, 3,835 in 2011, 3,498 in 2016, and 4,841 in 2024–25 (combined unweighted n=19,010). The corresponding design-weighted analytic denominators were approximately 3,710, 3,266, 4,009, 3,381, and 4,875. Reconstruction of the historically comparable outcome—at least one live birth or current pregnancy with a first child—reproduced the published 2000, 2005, 2011, and 2016 EDHS estimates to within 0.05 percentage points. The same definition yielded 9.87% (95% CI 8.31–11.44) in 2024–25. The broader 2024–25 "ever pregnant" indicator was not substituted into the historical series because it additionally includes pregnancy loss.

3.2 National decline: substantial, but episodic

Nationally, begun childbearing was 16.3% (95% CI 14.4–18.2) in 2000, 16.6% (14.7–18.6) in 2005, 12.4% (10.5–14.2) in 2011, 12.5% (10.5–14.4) in 2016, and 9.9% (8.3–11.4) in 2024–25. The 2000–2024/25 decline was 6.39 percentage points (95% CI for the decline 3.93–8.85), equivalent to a 39.3% relative reduction (95% CI 27.3–51.2). The linear calendar-time model confirmed a downward long-run trajectory (OR per decade 0.780, 95% CI 0.717–0.848; $p<.001$), while the categorical survey-round test rejected constant prevalence across rounds ($F(4,2473)=9.58$; $p<.001$).

The decline was not continuous. There was no statistically detectable change from 2000 to 2005 (+0.37 percentage points, 95% CI −2.35 to 3.10) or from 2011 to 2016 (+0.11, −2.59 to 2.82). Significant declines occurred from 2005 to 2011 (−4.27 points, −6.97 to −1.56) and from 2016 to

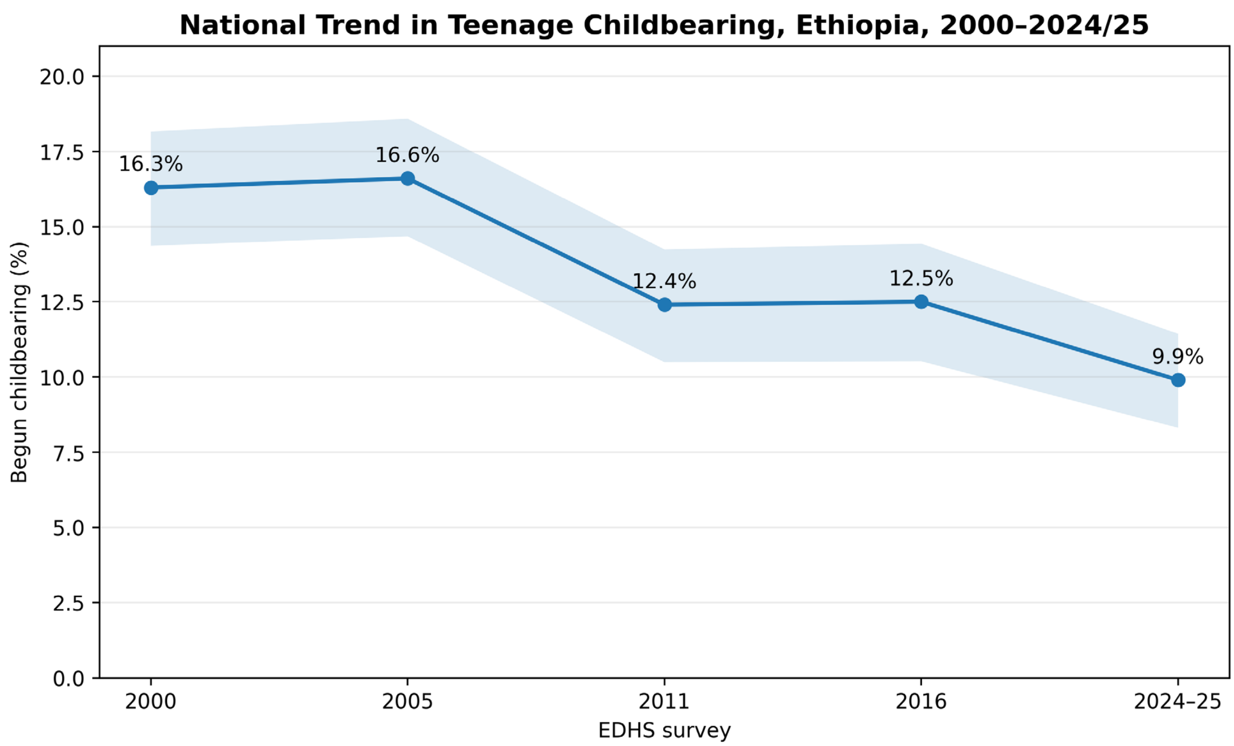
2024–25 (−2.60 points, −5.10 to −0.10). Thus, the quarter-century transition is better characterized as stair-step progress than as a smooth secular decline.

Table 1. National prevalence and interval-specific change in begun childbearing, Ethiopia, 2000–2024/25

EDHS	Begun childbearing, % (95% CI)	Change from prior round, pp
2000	16.3 (14.4–18.2)	—
2005	16.6 (14.7–18.6)	+0.37 (−2.35 to 3.10)
2011	12.4 (10.5–14.2)	−4.27 (−6.97 to −1.56)
2016	12.5 (10.5–14.4)	+0.11 (−2.59 to 2.82)
2024–25	9.9 (8.3–11.4)	−2.60 (−5.10 to −0.10)

Note. Begun childbearing = at least one live birth or current pregnancy with a first child. All percentages are survey-design weighted. Long-run OR per decade=0.780 (95% CI 0.717–0.848; p<.001).

Figure 1. National trend in begun childbearing among women aged 15–19, Ethiopia, 2000–2024/25.



Note. Points are survey-design-weighted prevalence estimates; error bars show 95% confidence intervals. The 2024–25 estimate uses the harmonized historical definition.

3.3 National progress concealed divergent regional transitions

Regional change was highly heterogeneous. Amhara showed the largest transformation, declining from 25.0% in 2000 to 6.3% in 2024–25, an 18.8-point decline (95% CI −24.8 to −12.7) and a 75.0% relative reduction. Tigray declined from 20.9% to 10.9% (−10.0 points, 95% CI −15.6 to −4.4). These were the only statistically significant endpoint declines. Five-round trend models additionally identified significant downward trajectories in Benishangul-Gumuz (OR/decade 0.706, p=.026) and Gambella (0.660, p=.007), alongside Amhara (0.485, p<.001) and Tigray (0.737, p=.001).

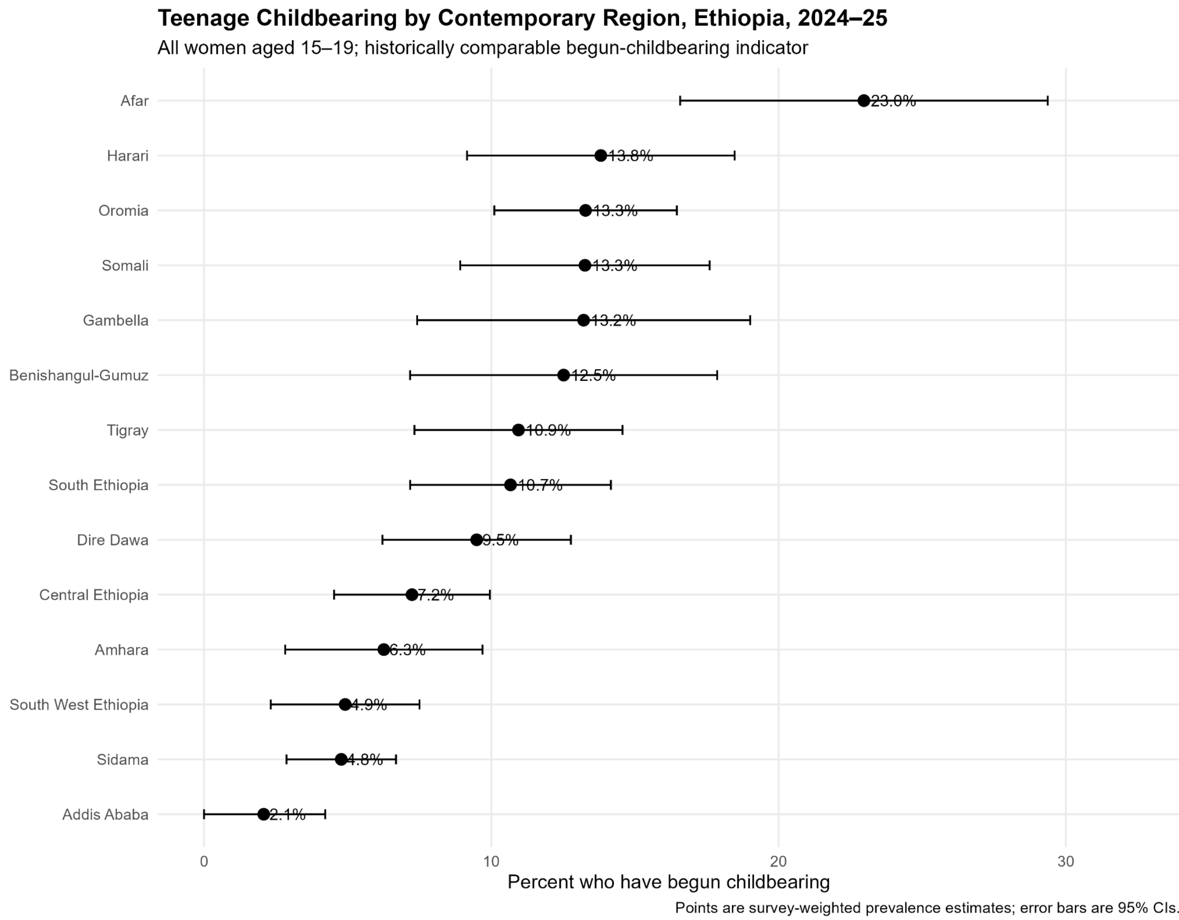
Afar stood apart from the national pattern: prevalence was 21.1% in 2000 and 23.0% in 2024–25, with no evidence of a downward five-round trend (OR/decade 1.073, p=.481). In the contemporary 2024–25 geography, Afar had the highest point estimate (23.0%, 95% CI 16.6–29.4) and Addis Ababa the lowest (2.1%, 0.0–4.2), an approximately eleven-fold point-estimate contrast. The reconstructed historical SNNPR prevalence of 7.49% concealed a more than twofold range across its successor regions, from 4.8% in Sidama to 10.7% in South Ethiopia.

Table 2. Long-run regional change and five-round trend inference

Historical region	2000 %	2024–25 %	Change, pp	OR/decade	Trend
Amhara	25.0	6.3	−18.8	0.485	Downward
Tigray	20.9	10.9	−10.0	0.737	Downward
Gambella	26.0	13.2	−12.8	0.660	Downward
Benishangul-Gumuz	22.2	12.5	−9.7	0.706	Downward
Oromia	15.8	13.3	−2.5	0.903	NS
Afar	21.1	23.0	+1.9	1.073	NS
Historical SNNPR	8.1	7.5	−0.6	0.939	NS
Addis Ababa	4.7	2.1	−2.6	0.707	NS

Note. Selected regions shown for parsimony. Endpoint contrasts and five-round linear trend tests answer different questions; "NS" indicates no statistically detectable linear trend at p<.05.

Figure 2. Begun childbearing by contemporary region, Ethiopia, 2024–25.



Note. Error bars show 95% confidence intervals. Rankings are descriptive and are not pairwise significance tests.

3.4 Social inequality changed on different scales

Residence showed the clearest convergence. In 2000, begun childbearing was 18.3% in rural areas and 9.1% in urban areas, a 9.1-point gap (95% CI 5.1–13.2) and rural/urban prevalence ratio (PR) of 2.00. By 2024–25, rural prevalence had fallen to 11.0% and urban prevalence was 7.5%; the gap narrowed to 3.5 points (95% CI 0.3–6.8; p=.034) and the PR to 1.47 (95% CI 0.99–2.18; p=.054).

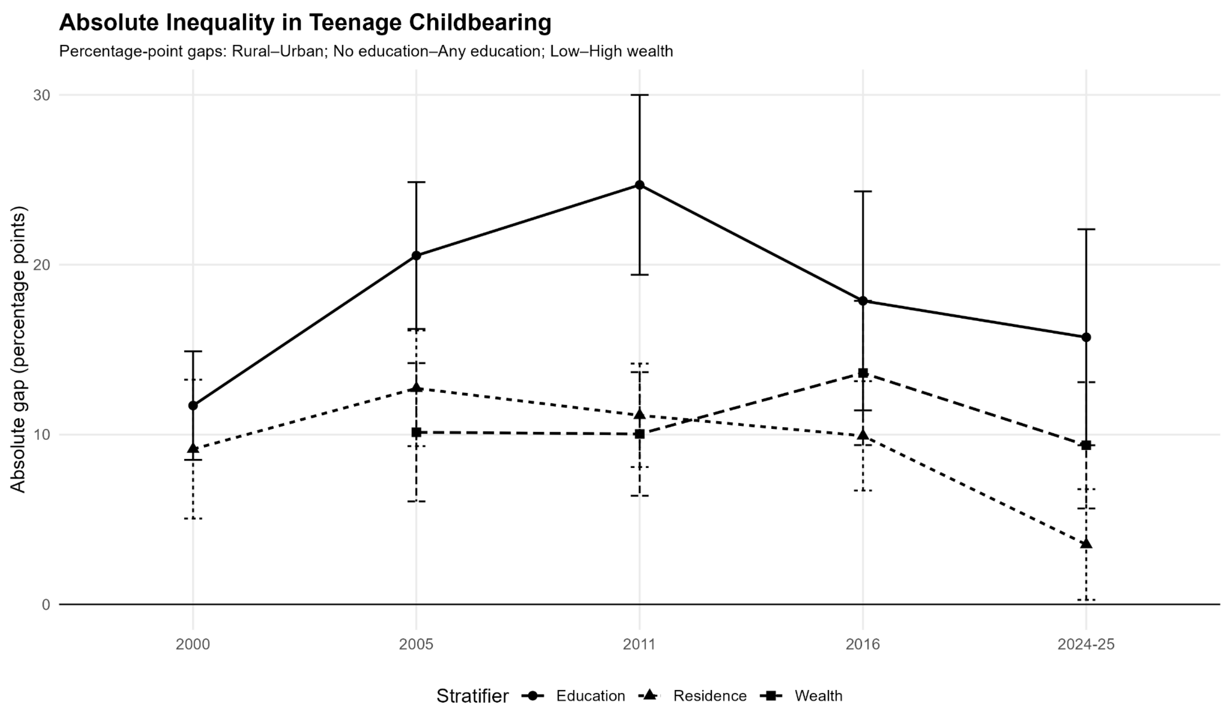
Educational inequality remained pronounced. In 2024–25, prevalence was 24.1% among adolescents with no formal education and 8.4% among those with any formal education, an absolute gap of 15.7 points (95% CI 9.4–22.1; p<.001) and PR of 2.88 (95% CI 2.10–3.93; p<.001). The education gap widened through 2011, reaching 24.7 points, before partially narrowing. Household wealth showed a different pattern: from 2005 to 2024–25, the low-versus-high wealth gap changed little in absolute terms (10.1 to 9.4 points), while the PR increased from 1.85 to 2.53. National progress therefore coexisted with residence convergence, persistent educational stratification, and mixed absolute-versus-relative wealth inequality.

Table 3. Endpoint socioeconomic inequalities in begun childbearing

Stratifier	Earlier contrast	2024–25 contrast	Absolute gap, pp	Prevalence ratio
Residence	2000: Urban 9.1%; Rural 18.3%	Urban 7.5%; Rural 11.0%	9.1 → 3.5	2.00 → 1.47
Education	2000: Any 9.1%; None 20.8%	Any 8.4%; None 24.1%	11.7 → 15.7	2.28 → 2.88
Wealth	2005: High 12.0%; Low 22.1%	High 6.1%; Low 15.5%	10.1 → 9.4	1.85 → 2.53

Note. Wealth compares Q1–Q2 with Q4–Q5 and excludes Q3. Endpoint changes in the inequality measures are descriptive; they are not formal tests of change in the gap.

Figure 3. Absolute inequality in begun childbearing by residence, education, and wealth.



Note. Gaps are rural minus urban, no education minus any education, and low wealth minus high wealth. Wealth begins in 2005.

3.5 Union formation remained the dominant reproductive context, but the coupling weakened

Current marital status showed an extreme concentration of childbearing in union: in 2024–25, 60.7% of currently married/in-union adolescents and 47.2% of formerly married/in-union adolescents had begun childbearing, compared with 0.6% of never-married adolescents. Conversely, 83.9% of adolescents who had begun childbearing were currently married/in union in 2024–25, down from 90.4% in 2000. These cross-sectional distributions, however, do not establish whether union preceded conception.

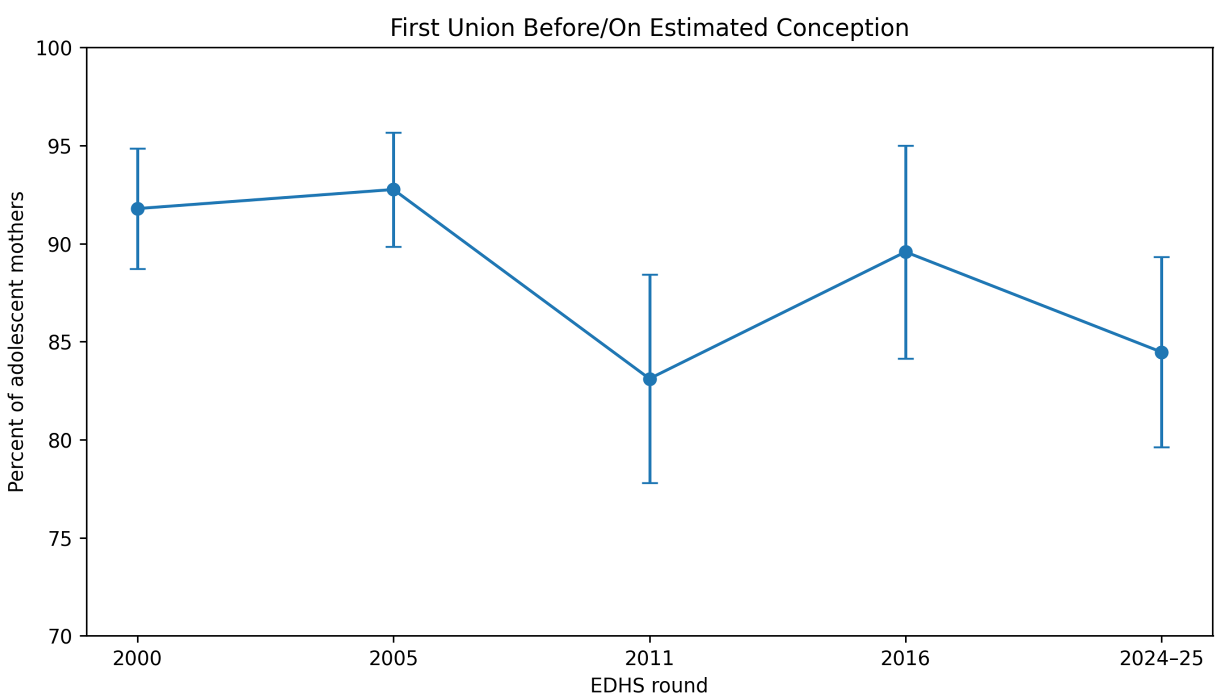
Event sequencing provided stronger temporal evidence. Among adolescent mothers, first union preceded or coincided with estimated conception in 91.8% in 2000, 92.8% in 2005, 83.1% in 2011, 89.6% in 2016, and 84.5% in 2024–25. The odds of union preceding estimated conception declined over calendar time (OR per decade 0.740, 95% CI 0.607–0.902; p=.003), with significant overall round heterogeneity (F(4,852)=4.34; p=.0018). In 2024–25, 6.8% conceived before first union but entered union by first birth, 2.9% had a first birth before first union, and 5.9% had never entered union by interview. A ±1-month sensitivity analysis around estimated conception preserved the substantive conclusion.

Table 4. Temporal ordering of first union and first pregnancy/birth among adolescent mothers (%)

EDHS	Union before/on conception	Pregnancy before union; union by birth	Birth before first union	Never in union
2000	91.8	6.6	0.6	1.0
2005	92.8	3.7	1.2	2.3
2011	83.1	11.4	3.4	2.1
2016	89.6	5.0	1.3	4.2
2024–25	84.5	6.8	2.9	5.9

Note. Estimated conception is first-birth CMC minus nine months. Categories are mutually exclusive and sum to approximately 100% within survey. OR per decade for union before/on conception=0.740 (95% CI 0.607–0.902; p=.003).

Figure 4. First union before or at estimated conception among adolescent mothers, Ethiopia, 2000–2024/25.



Note. Estimated conception is approximated from first-birth timing. Error bars show 95% confidence intervals.

3.6 Contraceptive protection improved among married adolescents, while pregnancies remained concentrated in unprotected periods

Modern contraceptive prevalence among currently married/in-union adolescents increased from 3.0% (95% CI 1.6–4.5) in 2000 to 32.7% (26.0–39.3) in 2024–25. The long-run increase was strong (OR per decade 2.46, 95% CI 2.10–2.89; $p<.001$). Over the same period, unmet need declined from 36.9% to 14.6% (OR per decade 0.60, 95% CI 0.51–0.69; $p<.001$). Because the 2000 and 2005 files use the legacy unmet-need variable and later rounds use the revised definition, the unmet-need series reflects broad historical change rather than a perfectly definition-invariant indicator.

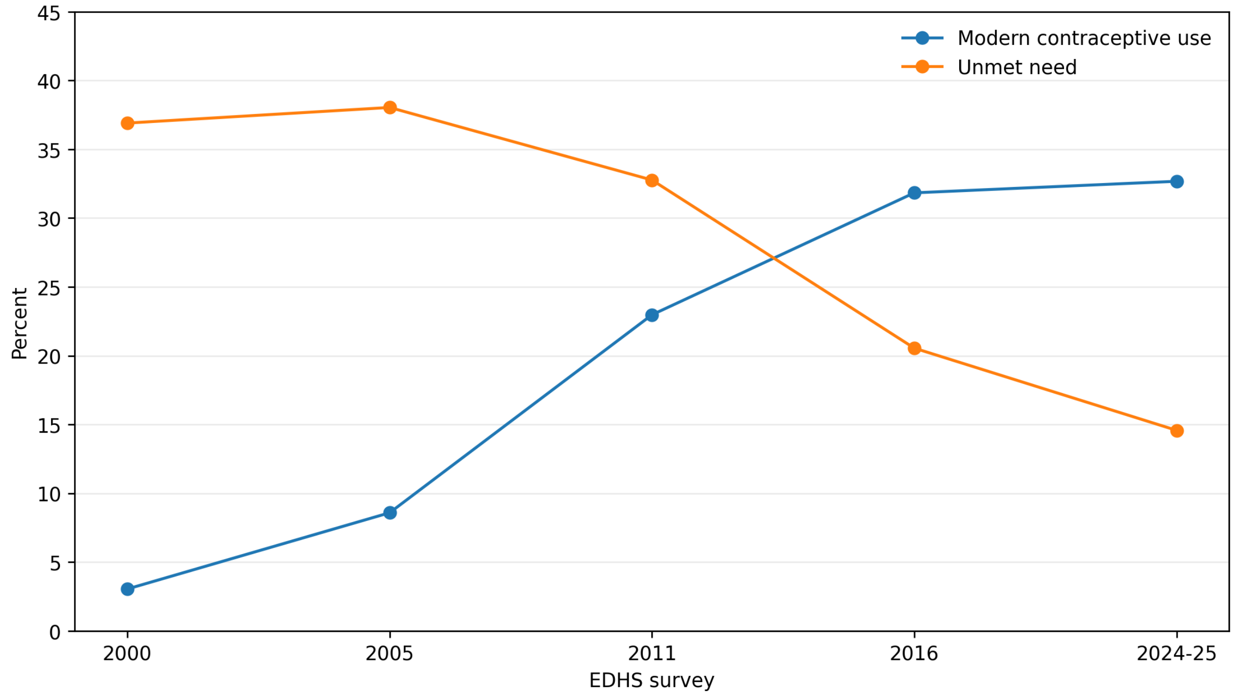
The pregnancy-centered calendar analysis produced a different, complementary result. Among adolescents who had begun childbearing and whose index pregnancy was observable in the calendar, modern-method exposure one month before estimated conception was 3.1% in 2005, 3.4% in 2011, 4.3% in 2016, and 3.0% in 2024–25. No-method exposure ranged from 95.3% to 97.0% across these surveys. A three-month sensitivity analysis produced the same substantive pattern. The one-month estimates should not be interpreted as contraceptive-failure rates: the analysis is conditioned on pregnancies that occurred and does not contain the person-time denominator needed to estimate failure.

Table 5. Contraceptive protection, unmet need, and pre-pregnancy exposure

EDHS	Married mCPR % (95% CI)	Married unmet need % (95% CI)	Modern method 1 month before conception %
2000	3.0 (1.6–4.5)	36.9 (32.3–41.5)	—
2005	8.6 (5.9–11.3)	38.0 (32.1–44.0)	3.1
2011	23.0 (18.1–27.8)	32.8 (27.0–38.6)	3.4
2016	31.8 (26.2–37.5)	20.5 (15.8–25.3)	4.3
2024–25	32.7 (26.0–39.3)	14.6 (9.4–19.7)	3.0

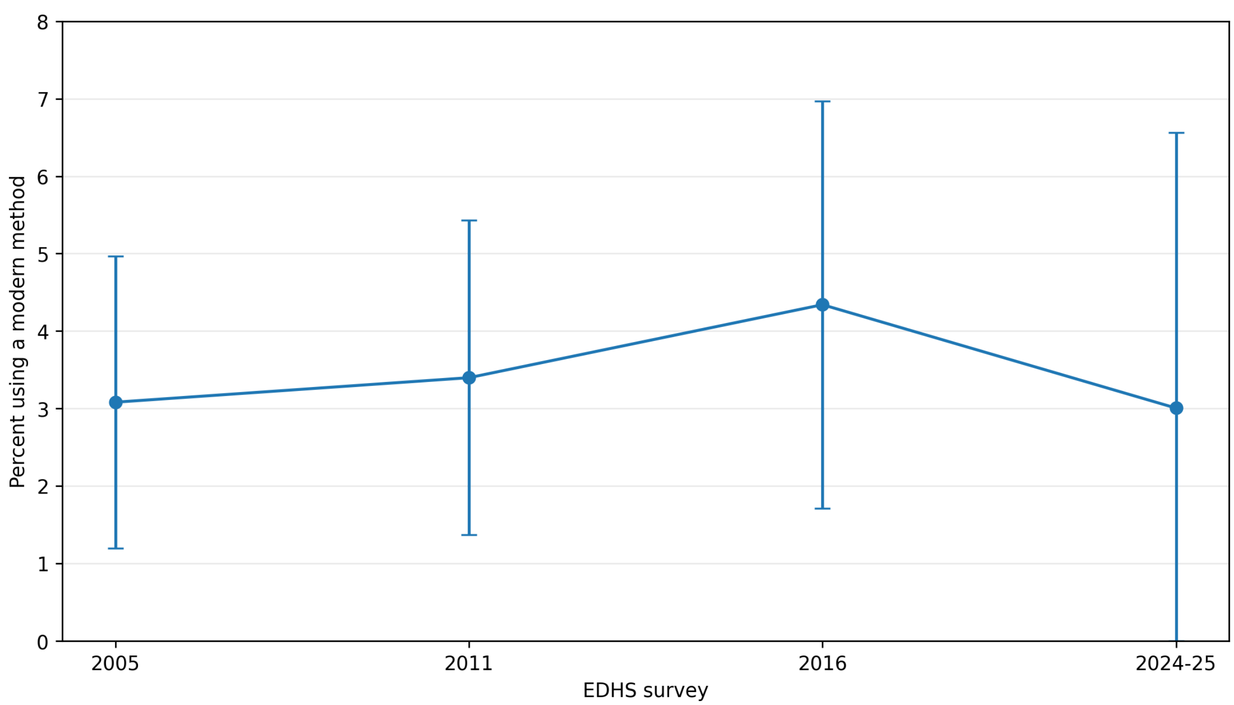
Note. mCPR and unmet need refer to currently married/in-union adolescents. Pre-pregnancy exposure refers only to adolescents with an observable index pregnancy in contraceptive-calendar surveys. It is not a contraceptive-failure rate.

Figure 5. Modern contraceptive use and unmet need among currently married/in-union adolescents, Ethiopia, 2000–2024/25.



Note. Survey-design-weighted estimates with 95% confidence intervals.

Figure 6. Modern contraceptive exposure one month before estimated conception among adolescents who had begun childbearing, 2005–2024/25.



Note. The denominator is adolescents with an observable index pregnancy in the contraceptive calendar. These estimates do not measure contraceptive failure.

3.7 Multivariable results: age, union formation, wealth, and place

The outcome was the historically comparable indicator of having begun childbearing: at least one live birth or current pregnancy with a first child among women aged 15–19. Survey-weighted logistic regression incorporated the DHS sampling weights, primary sampling units, and strata. Model 1 estimated structural and background associations using single-year age, urban/rural residence, contemporary region, education, and household wealth. Education contrasted no education with any formal education. Wealth contrasted the bottom two quintiles (Q1–Q2) with the top two quintiles (Q4–Q5); Q3 was intentionally excluded to create a clear low-versus-high contrast. Model 2 then added marital status (never, currently, or formerly married/in union).

The two-model sequence is substantively important. Model 1 asks how age, education, wealth, residence, and region pattern childbearing before current marital status is introduced. Model 2 asks how those adjusted associations change after accounting for current union status. Because current marital status can be measured after pregnancy initiation for some adolescents, this comparison is descriptive rather than a causal mediation analysis.

3.7.1 Analytic sample and outcome validation

The full 2024–25 outcome reconstruction included 4,841 adolescents and 488 childbearing events, with a design-weighted analytic denominator of 4,874.8 and a national prevalence of 9.87% (95% CI 8.31–11.44), reproducing the certified 9.9% estimate. The prespecified wealth contrast excluded Q3, leaving 4,031 adolescents and 415 events in the regression sample (83.3% of the full unweighted sample; design-weighted denominator 3,884.0). Prevalence in this intentionally restricted regression sample was 10.13% (95% CI 8.35–11.90). The 0.23-percentage-point difference from the full-sample estimate is therefore a denominator consequence, not an outcome-reconstruction discrepancy.

3.7.2 Primary adjusted associations

Model 1 showed pronounced age and socioeconomic patterning. Compared with age 15, adjusted odds rose steeply across later adolescence, becoming statistically clear by age 17 and especially at ages 18 and 19. Any formal education was associated with substantially lower adjusted odds of begun childbearing (AOR 0.44, 95% CI 0.26–0.75; $p=.003$), while low household wealth was associated with higher adjusted odds (AOR 2.91, 95% CI 1.43–5.90; $p=.003$). Rural residence was not independently associated with the outcome after the other background characteristics were held constant.

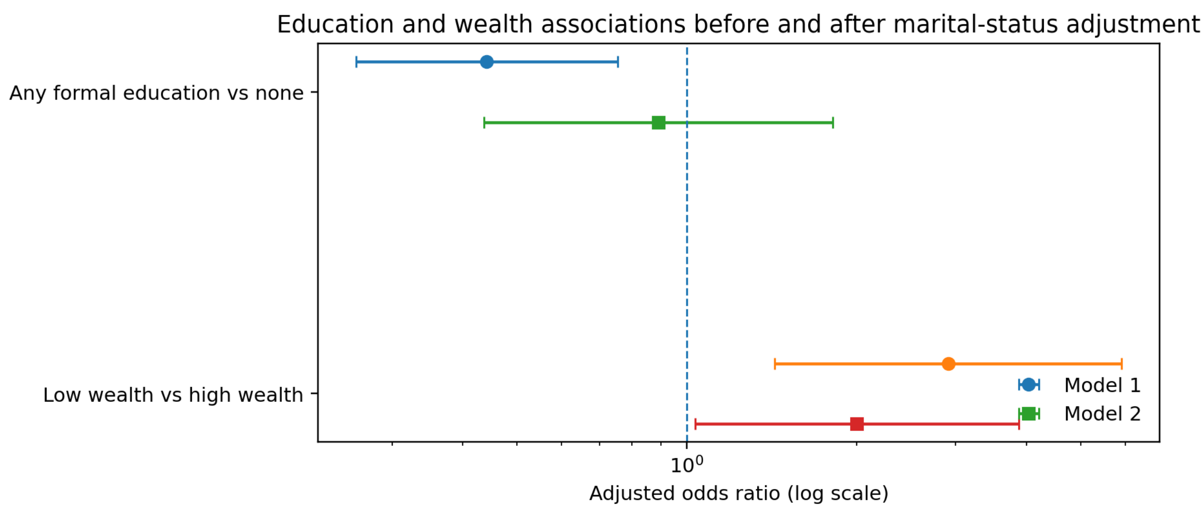
Introducing marital status in Model 2 materially changed the socioeconomic picture. The education association attenuated from AOR 0.44 to 0.89 (95% CI 0.44–1.82; $p=.750$), whereas low wealth remained associated with approximately twice the adjusted odds of childbearing (AOR 2.00, 95% CI 1.03–3.88; $p=.040$). Age remained strongly patterned: relative to age 15, the AOR was 12.01 (95% CI 4.54–31.73; $p<.001$) at age 18 and 17.51 (6.22–49.32; $p<.001$) at age 19. The estimates at ages 16 and 17 were no longer statistically significant in the fully adjusted model.

Table 6. Sequential multivariable associations with begun childbearing, Ethiopia DHS 2024–25

Contrast	Model 1 AOR (95% CI)	p	Model 2 AOR (95% CI)	p
Age 16 vs 15	2.26 (0.84–6.02)	.105	1.62 (0.43–6.03)	.474
Age 17 vs 15	5.26 (2.23–12.39)	<.001	2.61 (0.89–7.65)	.081
Age 18 vs 15	21.33 (9.05–50.27)	<.001	12.01 (4.54–31.73)	<.001
Age 19 vs 15	32.93 (14.65–73.99)	<.001	17.51 (6.22–49.32)	<.001
Rural vs urban	0.98 (0.46–2.09)	.959	0.52 (0.24–1.09)	.083
Any formal education vs none	0.44 (0.26–0.75)	.003	0.89 (0.44–1.82)	.750
Low wealth vs high wealth	2.91 (1.43–5.90)	.003	2.00 (1.03–3.88)	.040

Note. Model 1 adjusts for age, residence, contemporary region, education, and wealth. Model 2 additionally adjusts for marital status. Region coefficients are presented separately in Table M2. High wealth=Q4–Q5; low wealth=Q1–Q2; Q3 excluded. AORs are odds ratios, not risk ratios.

Figure 7. Education and wealth associations before and after marital-status adjustment.



Note. Points are adjusted odds ratios; horizontal bars are 95% confidence intervals. The reference line is AOR=1.0. The pronounced attenuation of education after marital-status adjustment should not be interpreted as proof of causal mediation.

3.7.3 Union formation: the dominant contemporaneous correlate

The largest coefficients in Model 2 were associated with marital status. Relative to never-married adolescents, the AOR was 484.39 (95% CI 191.71–1223.90; p<.001) for those currently married/in union and 340.41 (91.82–1262.05; p<.001) for those formerly married/in union. These values are exceptionally large because childbearing was extremely rare among never-married adolescents. They are odds ratios—not risk ratios—and should not be translated as "times more likely." Their substantive value is to show the extraordinary concentration of adolescent childbearing in the union/ever-union life-course state. The event-sequencing analysis elsewhere in the study is required to determine whether union preceded conception; current marital status alone cannot establish that temporal ordering.

3.7.4 Does place matter?

Answer: Yes—but not as a simple rural-versus-urban story.
After adjustment, rural residence itself was not statistically significant, yet several regional coefficients remained markedly different from Oromia. Place therefore matters at the regional/contextual level even when a coarse urban–rural indicator adds little independent information. At the same time, the model does not prove that region is causal; region is a contextual marker that can capture unmeasured social, cultural, service, marriage, educational, economic, and program environments.

The fully adjusted model directly answers the study's recurring question. Relative to Oromia, adjusted odds were significantly lower in Tigray (AOR 0.23, 95% CI 0.11–0.48), Amhara (0.21, 0.09–0.47), Sidama (0.20, 0.07–0.64), South West Ethiopia (0.14, 0.06–0.32), and Addis Ababa (0.10, 0.02–0.46). By contrast, Afar, Somali, Benishangul-Gumuz, Central Ethiopia, South Ethiopia, Gambella, Harari, and Dire Dawa did not differ statistically from Oromia at $p<.05$ after adjustment. Thus, compositional differences in age, education, wealth, residence, and marital status do not fully erase geographic heterogeneity.

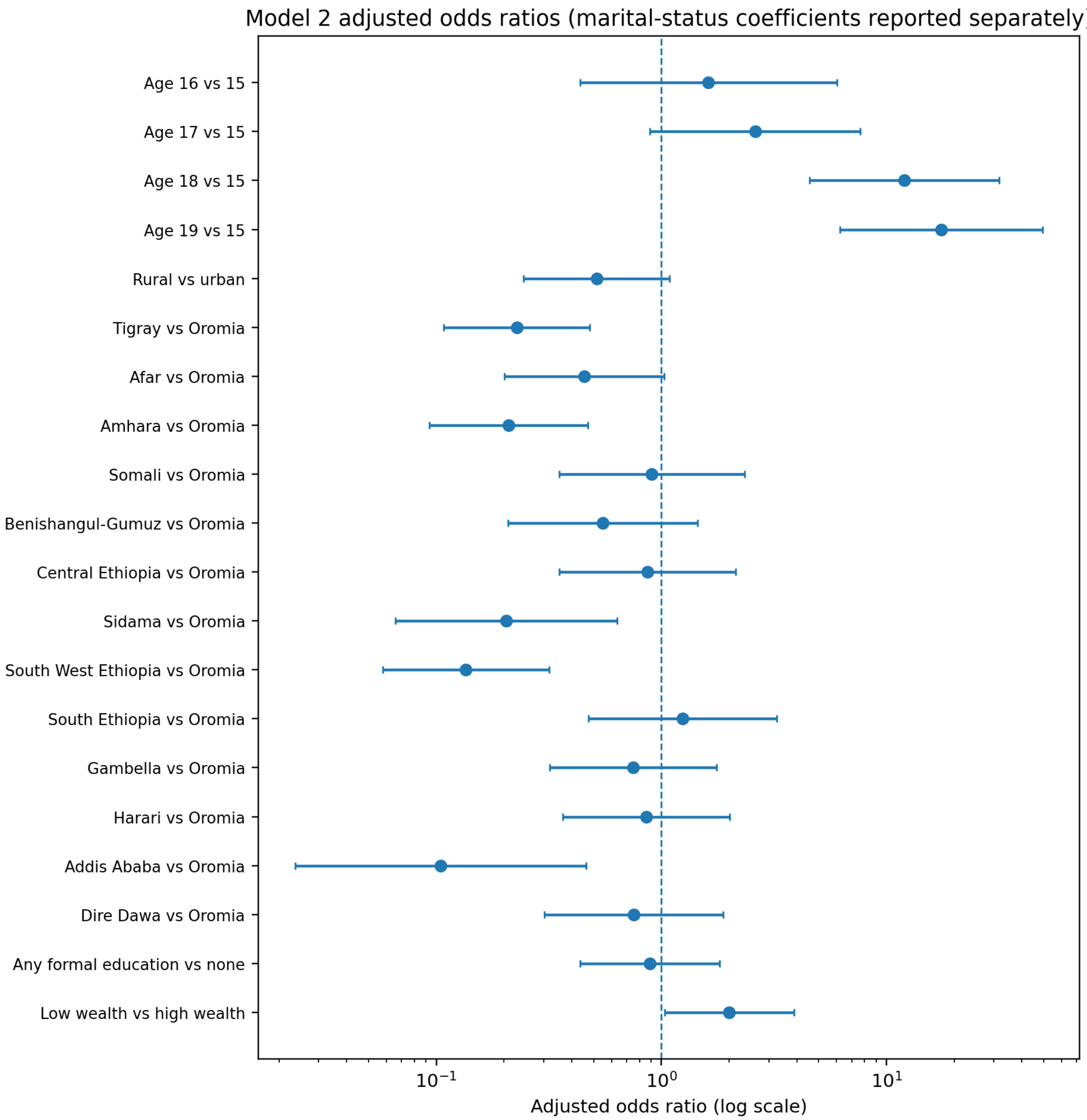
This is more informative than asking only whether rural adolescents differ from urban adolescents. Residence is a two-category spatial proxy; region represents a much richer bundle of context. The coexistence of a nonsignificant rural coefficient with multiple significant regional contrasts indicates that geographic patterning is not reducible to urbanicity. The result is also consistent with prior Ethiopian spatial analyses reporting nonrandom geographic clustering of adolescent pregnancy and regional or local variation in its correlates (Tigabu et al., 2021; Tebeje et al., 2024).

Table 7. Does place matter? Fully adjusted regional associations with begun childbearing

Region vs Oromia	AOR	95% CI	p	Interpretation
Tigray	0.23	0.11–0.48	<.001	Lower adjusted odds
Afar	0.45	0.20–1.03	.058	Not significant
Amhara	0.21	0.09–0.47	<.001	Lower adjusted odds
Somali	0.91	0.35–2.34	.841	Not significant
Benishangul-Gumuz	0.55	0.21–1.45	.226	Not significant
Central Ethiopia	0.87	0.35–2.14	.757	Not significant
Sidama	0.20	0.07–0.64	.006	Lower adjusted odds
South West Ethiopia	0.14	0.06–0.32	<.001	Lower adjusted odds
South Ethiopia	1.24	0.47–3.27	.657	Not significant
Gambella	0.75	0.32–1.76	.506	Not significant
Harari	0.86	0.36–2.01	.720	Not significant
Addis Ababa	0.10	0.02–0.46	.003	Lower adjusted odds
Dire Dawa	0.75	0.30–1.88	.541	Not significant

Note. Model 2 adjusts for single-year age, residence, education, low-versus-high wealth, and marital status. Oromia is the reference region. These are conditional associations; they do not identify a causal "regional effect," and individual pairwise coefficients should not be confused with a single omnibus test of region.

Figure 8. Fully adjusted Model 2 associations with begun childbearing, Ethiopia DHS 2024–25.



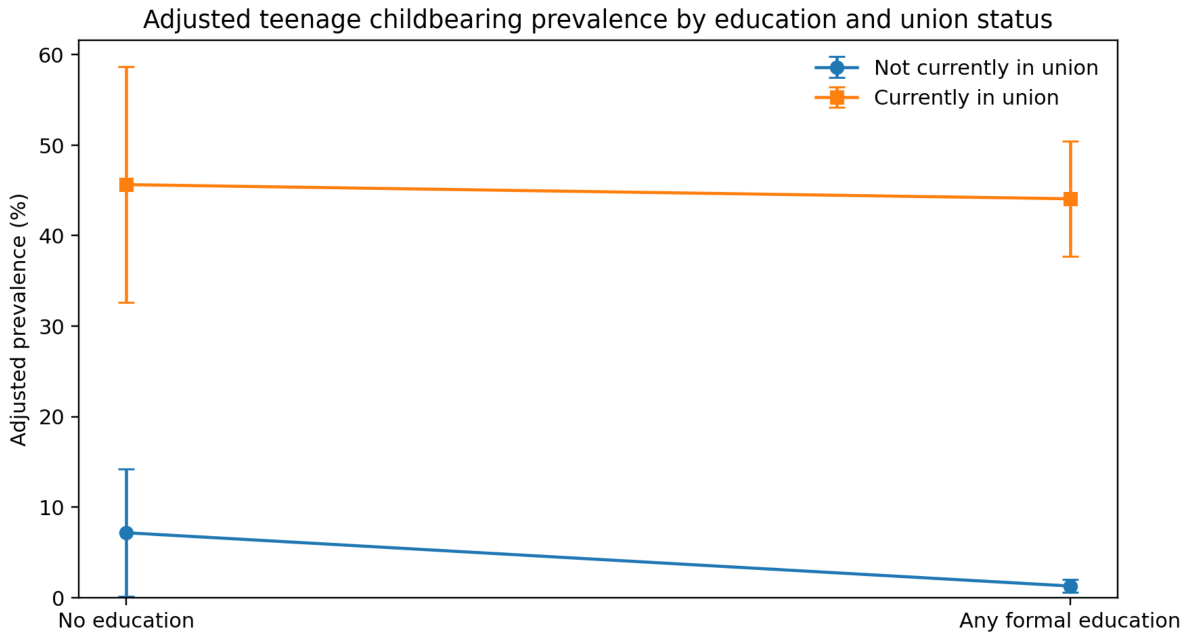
Note. Marital-status coefficients are omitted from the plotting scale because their extreme magnitude would compress the remaining estimates. The figure emphasizes age, residence, regional, education, and wealth associations.

3.7.5 Does the importance of education and wealth depend on union status?

The prespecified interaction analysis moves beyond average adjusted associations. Education varied significantly by current union status (design-adjusted Wald F(1,723)=4.95; p=.026). Standardized predicted prevalence among adolescents not currently in union was 7.15% (95% CI 0.10–14.21) for those with no education and 1.26% (0.55–1.98) for those with any formal education. Among adolescents currently in union, predicted prevalence was 45.64% (32.60–58.67) with no education and 44.07% (37.72–50.43) with any education. The education gradient is therefore pronounced outside current union but nearly absent among adolescents already in union.

This interaction helps explain why the education coefficient changes so sharply between Models 1 and 2. Education is not simply "protective" by a uniform amount across all adolescent social states. Its association with childbearing is strongly contingent on whether an adolescent is currently in union. This is compatible with a life-course interpretation in which schooling, delayed union, and delayed childbearing are interlocking transitions, but the cross-sectional model cannot determine the causal direction among them.

Figure 9. Adjusted prevalence of begun childbearing by education and current union status.



Note. Predicted probabilities are standardized over the analytic sample; error bars are 95% confidence intervals. The education-by-current-union interaction is statistically significant (p=.026).

The corresponding wealth-by-current-union interaction was not statistically significant ($F(1,723)=2.26$; $p=.134$). Predicted prevalence was 1.04% (95% CI 0.28–1.80) among high-wealth adolescents not currently in union and 3.67% (1.11–6.24) among low-wealth adolescents not currently in union; among those currently in union, it was 40.88% (32.50–49.27) and 49.33% (41.26–57.41), respectively. These differences are descriptively relevant but do not support a formal claim that the wealth association differs by current union status.

Table 8. Prespecified interaction tests

Interaction	Design-adjusted Wald F	df	p	Conclusion
Education × current union	4.95	1, 723	.026	Supported
Wealth × current union	2.26	1, 723	.134	Not supported
Education × region	27.78	13, 711	<.001	Global heterogeneity; unstable sparse cells
Wealth × region	1.05	12, 712	.397	Not supported

Note. Interaction tests are global design-adjusted Wald tests. The education-by-region model failed the prespecified coefficient-stability screen because numerous no-education regional cells were sparse or contained zero events; individual region-specific education predictions should therefore not be ranked or emphasized.

3.7.6 Does the education gradient itself vary by place?

Statistically, the global education-by-region interaction was highly significant ($F(13,711)=27.78$; $p<.001$), suggesting that the education-childbearing association is not homogeneous across all 14 regions. Substantively, however, this result cannot be converted into a reliable regional league table. The interaction model was the only prespecified model flagged for possible sparse-cell instability/separation: many region-by-no-education cells contained fewer than 25 adolescents or fewer than 10 childbearing events, and Sidama and Addis Ababa had zero events in the no-education cell. The resulting region-specific predicted probabilities can therefore become implausibly extreme.

Accordingly, the defensible answer to "Does place matter?" has two levels. First, yes: region remains associated with childbearing in the fully adjusted main-effects model. Second, there is evidence that the education gradient may itself vary geographically, but the available 2024–25 sample is not sufficiently dense to estimate every regional education contrast with acceptable stability. This is an important distinction between evidence of heterogeneity and precision about where, exactly, each heterogeneous effect lies.

3.7.7 Model diagnostics and inferential safeguards

Table 9. Coefficient-stability diagnostics for the multivariable and interaction models

Model	Aliased coefficients	Max |β|	Max SE	Assessment
Model 1	0	3.78	0.61	No instability flag
Model 2	0	6.84	0.76	No instability flag
Education × current union	0	3.79	0.81	No instability flag
Wealth × current union	0	4.92	0.78	No instability flag
Education × region	0	14.92	1.88	Review: sparse-cell instability/separation
Wealth × region	0	4.47	1.04	No instability flag

Note. The stability screen is diagnostic rather than a formal hypothesis test. It supports retaining the main models and education-by-union interaction while interpreting the education-by-region interaction cautiously.

4. Discussion

4.1 Principal findings

This study reframes Ethiopia's adolescent reproductive transition as more than a national prevalence decline. Over approximately 25 years, begun childbearing among women aged 15–19 fell by about 39%, but the improvement was episodic, geographically concentrated, and socially unequal. Residence disparities narrowed substantially, whereas educational inequality remained large and wealth inequality persisted—especially on the relative scale. First union continued to precede the great majority of adolescent first pregnancies, although that coupling weakened over time. At the same time, contraceptive protection among married adolescents improved markedly and unmet need declined, yet pregnancies that occurred remained overwhelmingly concentrated in months with no recorded contraceptive method. In 2024–25, age and union status dominated the adjusted pattern, low household wealth remained independently associated with childbearing, and the crude educational gradient largely disappeared after marital status was introduced. In the fully adjusted 2024–25 analysis, regional heterogeneity persisted even though rural residence was not independently significant, indicating that place mattered at a contextual regional level rather than through urbanicity alone.

4.2 The national decline is real, but it is not a sufficient account of the transition

The long-run decline extends the direction reported by Kassa et al. (2019a), who documented falling teenage childbearing through the 2016 EDHS. Adding 2024–25 shows renewed improvement after the 2011–2016 plateau, but also demonstrates that progress has not been continuous. This matters programmatically: a favorable endpoint comparison can conceal intervals of stagnation, and a single per-decade coefficient summarizes direction without describing when progress occurred. Ethiopia's experience therefore supports sustained implementation rather than assuming that adolescent childbearing will continue to fall automatically.

The outcome definition also matters. The historical DHS indicator used here captures entry into motherhood or a current first pregnancy; the broader 2024–25 ever-pregnant measure additionally captures pregnancy loss. Preserving the harmonized outcome protects the validity of temporal inference. At the same time, the broader contemporary indicator remains important for clinical and reproductive-rights surveillance because it captures pregnancy experiences omitted by the historical series. The two indicators should therefore be reported in parallel rather than treated as interchangeable.

4.3 Geographic divergence requires differentiated policy

The regional findings reinforce prior evidence that adolescent pregnancy and childbearing in Ethiopia are spatially patterned. Earlier 2016 EDHS analyses identified strong regional differences (Birhanu et al., 2019; Kefale et al., 2020), while spatial studies identified high-risk clustering in Afar, Somali, Oromia, Harari, and other settings (Tigabu et al., 2021; Tebeje et al., 2024). The present five-round series adds a longitudinal dimension: Amhara underwent the largest transformation, whereas Afar showed no evidence of long-run improvement and remained the highest-prevalence region in 2024–25. This divergence argues against a single national intervention intensity or delivery model.

The 2024–25 multivariable results sharpen that geographic conclusion and directly answer the recurring question, “Does place matter?” Yes—but not as a simple rural-versus-urban story. Rural residence was not independently associated with begun childbearing after adjustment, whereas Tigray, Amhara, Sidama, South West Ethiopia, and Addis Ababa retained significantly lower adjusted odds than Oromia after simultaneous control for single-year age, residence, education, wealth, and marital status. Thus, measured compositional differences did not erase regional heterogeneity. Region should not be interpreted as a causal exposure; rather, it functions here as a contextual marker for social norms, union and fertility regimes, schooling opportunity, economic conditions, service access, mobility, conflict exposure, and program implementation. The significant global education-by-region interaction further suggests that the education gradient may vary geographically, although sparse cells preclude stable region-by-region ranking of that interaction.

Regional prevalence alone cannot explain the mechanisms behind regional divergence. Differences in schooling, age at union, poverty, mobility, service availability, conflict exposure, social norms, and contraceptive access may operate in different combinations. UNICEF’s Ethiopia program similarly characterizes child marriage as arising from interrelated social norms, educational constraints, economic incentives, and shocks such as conflict and drought (UNICEF, 2022). The appropriate next step is therefore not to attribute high prevalence to a single regional “culture,” but to diagnose the local configuration of constraints and protective factors.

4.4 Education appears deeply intertwined with the transition into union

The descriptive education gradient was among the largest inequalities observed: in 2024–25, adolescents with no formal education had nearly three times the prevalence of those with any education. This is consistent with Ethiopian multilevel studies linking lower educational attainment to adolescent pregnancy (Birhanu et al., 2019; Kefale et al., 2020) and with WHO’s broader observation that adolescent pregnancy is concentrated among girls with less education and lower economic status (WHO, 2024). Yet the sequential 2024–25 models add an important qualification. The education AOR changed from 0.44 before marital-status adjustment to 0.89 afterward, and the education-by-union interaction showed a strong gradient outside current union but almost none within union.

These findings should not be reduced to a causal mediation claim. Education may delay union, union may interrupt schooling, pregnancy may precipitate school exit or union, and unmeasured family or community conditions may influence all three. Nevertheless, the empirical pattern is consistent with a transition-centered interpretation: schooling and delayed union are tightly linked protective domains before entry into union, while adolescents already in union require direct reproductive-health protection regardless of educational background. WHO’s 2025 guideline similarly emphasizes both ending child marriage and strengthening girls’ schooling, while treating access to contraception as a complementary intervention rather than an alternative to social protection (WHO, 2025; Plesons et al., 2025).

4.5 Poverty remains consequential even after accounting for current marital status

Wealth inequality was persistent in both the descriptive and adjusted analyses. The percentage-point gap between low- and high-wealth adolescents changed little from 2005 to 2024–25, while the prevalence ratio widened because childbearing declined proportionately faster among the better-off. In the fully adjusted 2024–25 model, low wealth remained associated with approximately twice the odds of begun childbearing. This is consistent with Ethiopian evidence identifying household or community poverty as an important correlate (Kefale et al., 2020; Tigabu et al., 2021) and with

WHO's conclusion that early pregnancy is concentrated among socially and economically disadvantaged adolescents (WHO, 2024).

The policy implication is that adolescent pregnancy prevention cannot rely solely on information provision. Financial barriers, school discontinuation, limited employment prospects, transport costs, and household dependence may all shape reproductive choices and access to services. The 2025 WHO guideline explicitly recommends reducing financial barriers to adolescent contraception and highlights broader social and economic interventions relevant to early marriage and pregnancy (WHO, 2025). Poverty-sensitive adolescent programming should therefore be considered part of reproductive-health strategy rather than external to it.

4.6 Union formation remains central, but services cannot be restricted to married adolescents

The event-sequencing analysis is one of the study's most important extensions beyond conventional cross-sectional modeling. Current marital status alone could easily be misread as evidence that pregnancy occurred within marriage. By reconstructing first-union and first-birth timing, we found that union preceded or coincided with estimated conception for the large majority of adolescent mothers—84% to 93% across surveys—but that the probability declined over time. The increasing minority of pregnancies beginning before union or occurring among adolescents who remained outside union means that a married-adolescent-only service model will miss a meaningful and potentially growing group.

This finding aligns with WHO's updated framework, which places prevention of child marriage alongside universal, adolescent-responsive access to contraception and sexual and reproductive health services (WHO, 2025). It also supports a dual strategy in Ethiopia: continue preventing early marriage and supporting girls to remain in school, while ensuring confidential contraception, pregnancy testing, counseling, and referral pathways are accessible regardless of marital status. The extremely large marital-status odds ratios in the 2024–25 model are best understood as markers of this life-course concentration, not as causal effect estimates.

4.7 Rising contraceptive coverage and low pre-pregnancy exposure are complementary findings

The contraceptive findings initially appear paradoxical: modern contraceptive prevalence among married adolescents rose more than tenfold while modern-method exposure immediately before pregnancies that occurred remained only about 3–4%. The two measures answer different questions. Current mCPR describes protection among a defined population at interview; the calendar analysis starts with pregnancies that occurred and asks what method was recorded before conception. It therefore identifies the exposure context of pregnancies, not contraceptive effectiveness.

The programmatic signal is nonetheless strong. Approximately 95–97% of observable index pregnancies occurred after a calendar month in which no method was recorded. This points primarily to absence, discontinuation, inconsistent use, or lack of timely uptake rather than to widespread contraceptive failure. WHO notes that adolescents can face stigma, cost, provider bias, restrictive consent rules, lack of knowledge, and high discontinuation related to side effects or changing circumstances (WHO, 2024). The 2025 guideline consequently emphasizes access, uptake, and continued use—not simply method availability—and recommends attention to financial barriers, gender and social norms, service quality, enabling policies, and adolescent engagement (WHO, 2025).

The small and unstable samples of sexually active unmarried adolescents in the EDHS are themselves informative for surveillance. Their contraceptive needs are programmatically important, but conventional household surveys may not provide sufficiently precise annual or round-specific estimates for this subgroup. Complementary service data, targeted surveys, or pooled multi-round analyses may be required to monitor protection outside marriage with adequate precision.

4.8 Implications for Ethiopia's adolescent and youth health strategy

The findings support a policy architecture organized around transitions rather than isolated indicators. Ethiopia's National Adolescents and Youth Health Strategy already adopts a multisectoral approach to adolescent and youth health (Ministry of Health [Ethiopia], 2021). The present evidence suggests five linked priorities: sustain school participation and prevent educational exclusion; delay first union and enforce protections against child marriage; ensure contraceptive access before and across transitions into sexual activity and union; target poverty-related barriers; and differentiate implementation geographically, particularly in regions where national progress has not translated into comparable local improvement.

This transition-centered framing is closely aligned with WHO's 2025 guideline, which integrates child-marriage prevention, contraception, protection from coercion, and quality care for pregnant adolescents (WHO, 2025). It also implies that monitoring should not rely on a single national prevalence. A compact adolescent reproductive-transition dashboard could track begun childbearing, age at first union, married-adolescent mCPR and unmet need, school participation, poverty gradients, and contemporary regional disparities. Such a dashboard would make it harder for national improvement to conceal stalled regions or disadvantaged social groups.

5. Limitations

Several limitations should guide interpretation. First, the EDHS rounds are repeated cross-sectional surveys rather than a cohort. They provide strong population-level estimates of change but cannot establish individual causal pathways. The 2024–25 multivariable models therefore estimate adjusted cross-sectional associations. Current marital status is particularly temporally ambiguous and was introduced sequentially for descriptive interpretation rather than causal mediation.

Second, the historically harmonized outcome deliberately excludes prior pregnancies ending in miscarriage, induced abortion, or stillbirth when the adolescent had no live birth and was not currently pregnant. It therefore measures historically comparable entry into childbearing rather than the full burden of adolescent pregnancy. The broader 2024–25 ever-pregnant measure should be used alongside, not substituted into, the historical series.

Third, the women's IR files provide comparable individual reproductive histories for ages 15–19 but not girls aged 10–14. The study therefore concerns older adolescents and cannot quantify pregnancy or childbearing among younger adolescents. Fourth, administrative boundaries changed substantially. Reconstructing historical SNNPR protects longitudinal comparability but suppresses contemporary heterogeneity; presenting both historical and contemporary geographies mitigates but does not eliminate this limitation.

Fifth, some regional estimates and interaction cells contain few events. The education-by-region interaction was statistically significant globally but showed sparse-cell instability and possible separation; region-specific education predictions were therefore not used for ranking or strong inference. Sixth, the binary education measure improves interpretability across ages 15–19 but cannot distinguish schooling level, current enrollment, quality, interruption, or completion. The low-versus-high wealth model likewise excludes Q3, sharpening the socioeconomic contrast at the cost of sample size.

Seventh, first conception was approximated from month-level first-birth or current-pregnancy timing. The ±1-month sensitivity analysis reduced concern about classification near the union-conception boundary but cannot recover exact dates. Eighth, contraceptive-calendar estimates are subject to recall and coding error. Although calendar orientation was strongly validated against known birth months, the pregnancy-centered analysis contains few modern-method events, especially in 2024–25. Finally, pre-pregnancy method distributions are conditioned on pregnancies that occurred and must not be interpreted as contraceptive-failure rates. Estimating failure requires person-time at risk among users and nonusers.

6. Conclusion

Ethiopia achieved a substantial reduction in begun childbearing among women aged 15–19 between 2000 and 2024–25, but the national decline is only part of the story. Progress occurred in episodes rather than continuously, differed sharply across regions, and left persistent educational and economic inequalities. Union formation remained the dominant context of adolescent first childbearing, although its temporal coupling with pregnancy weakened. Contraceptive protection among married adolescents improved dramatically and unmet need declined, yet pregnancies that occurred remained overwhelmingly concentrated in periods with no recorded contraceptive method.

The combined evidence supports a transition-centered approach to adolescent reproductive health. School continuation, delayed union, poverty-sensitive prevention, and contraceptive protection should be treated as mutually reinforcing, not competing, strategies. Services must remain accessible to adolescents already in union while also reaching those who are unmarried or formerly in union. National monitoring should preserve a historically comparable childbearing indicator, but policy should be guided simultaneously by contemporary regional and socioeconomic inequalities. Ethiopia's next gains are likely to depend less on a uniform national program than on sustained, geographically differentiated action across the transitions that move adolescents from schooling and social dependence into union, sexual exposure, pregnancy risk, and parenthood.

Declaration of Generative AI and AI-Assisted Technologies. During the preparation of this manuscript, the author used OpenAI to assist with analytical planning, manuscript organization, language refinement, methodological exposition, and presentation of the study findings.

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