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Intersectional inequalities in demand for family planning satisfied with modern contraceptive methods in Ethiopia: a cross-sectional MAIHDA analysis of the 2024–25 Ethiopia Demographic and Health Survey

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Research Article

Keywords: family planning, modern contraception, demand satisfied, intersectionality, MAIHDA, health inequality, Ethiopia, Demographic and Health Survey

Posted Date: November 12th, 2026

DOI: https://doi.org/10.21203/rs.3.rs-11109715/v1

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Additional Declarations: No competing interests reported.

Abstract

Background

National family-planning averages may conceal inequalities across combinations of reproductive age and socioeconomic position. We examined intersectional inequalities in demand for family planning satisfied with modern contraceptive methods (mDFPS) in Ethiopia and assessed how much between-stratum heterogeneity was accounted for by additive social gradients versus residual intersectional departures.

Methods

This cross-sectional study used data from the 2024–25 Ethiopia Demographic and Health Survey Individual Recode dataset. The analytic sample comprised 6,611 women aged 15–49 years with family-planning demand. Survey-weighted mDFPS was estimated by age, wealth, education and residence. The primary MAIHDA specification cross-classified four age groups, five wealth quintiles, three education categories and two residence categories, yielding 120 theoretical intersectional strata. Null and additive multilevel logistic models estimated between-stratum variance, the variance partition coefficient (VPC) and proportional change in variance (PCV). Sensitivity analyses used a more granular 210-strata specification and an unweighted primary model.

Results

National survey-weighted mDFPS was 67.3%. mDFPS was 61.9% (95% CI 53.4–70.4) at ages 15–19 years, 74.8% (70.2–79.4) at ages 20–24, 72.6% (69.6–75.6) at ages 25–34 and 56.9% (53.2–60.7) at ages 35–49. Coverage increased from 43.8% (38.5–49.0) in the poorest wealth quintile to 77.3% (73.1–81.6) in the richest, and from 53.7% (49.7–57.6) among women with no education to 78.3% (75.4–81.2) among those with secondary or higher education. In the primary MAIHDA model, between-stratum variance declined from 0.858 (VPC 20.7%) in the null model to 0.175 (VPC 5.1%) after inclusion of the additive social-position variables, corresponding to a PCV of 79.6%. Among strata with at least 20 women, the lowest model-predicted mDFPS was 19.6% (approximate 95% interval 9.6–36.1) among rural women aged 15–19 years in the poorest quintile with no education. The 210-strata sensitivity specification produced a similar reduction in between-stratum variance (PCV 77.0%).

Conclusions

mDFPS in Ethiopia is strongly patterned by age, household wealth and education. Most between-stratum heterogeneity was accounted for by additive socioeconomic and life-course gradients, while a smaller residual intersectional component remained. Equity-oriented monitoring and programming should

address these broad structural gradients while retaining intersectional disaggregation to identify combinations of women with particularly low coverage.

1. Introduction

Access to contraception is fundamental to reproductive autonomy and enables individuals and couples to determine whether and when to have children. It reduces unintended pregnancy and associated maternal and perinatal risks and contributes to gender equality and wider social and economic participation [1, 2]. Yet access remains uneven. Recent global estimates indicate that, despite sustained gains in modern contraceptive use, more than 250 million women of reproductive age who want to avoid pregnancy are not using a modern method, with the largest remaining gaps concentrated in low- and middle-income settings, particularly sub-Saharan Africa [3, 4].

Demand for family planning satisfied with modern methods (mDFPS), Sustainable Development Goal (SDG) indicator 3.7.1, is the proportion of women aged 15–49 years who have a demand for family planning and are using a modern contraceptive method [3, 5]. Unlike contraceptive prevalence alone, mDFPS conditions the denominator on women who wish to delay or avoid pregnancy and therefore provides a more direct measure of whether expressed family-planning demand is being met. It is increasingly used for monitoring progress toward universal access to sexual and reproductive health services and for assessing equity in service coverage [3–5].

Global progress in mDFPS has not eliminated socioeconomic inequality. A recent population-based analysis across 48 FP2030 focus countries documented persistent wealth-related inequalities in modern contraceptive use and mDFPS despite overall improvement [6]. Multi-country analyses in sub-Saharan Africa similarly show that demand satisfied by modern methods varies by age, educational attainment, household wealth, residence and service-access conditions [7]. Such inequalities are especially important in settings where national averages can conceal population groups with substantially lower coverage.

Ethiopia has made important gains in family planning, but recent evidence continues to show marked heterogeneity in contraceptive demand and access. Analyses of Ethiopian survey data have reported lower family-planning coverage or greater unmet need among adolescents and young women, women in poorer households, women with less education and populations in underserved geographic areas [8–11]. A 2024 analysis across four Ethiopian regions documented substantial socioeconomic inequality in contraceptive demand and unmet need [8], while recent studies among young women and national mDFPS have reinforced the importance of age, wealth and geographic context [9–11]. Geospatial work across Ethiopia, Kenya and Nigeria has further shown that national and regional averages can conceal pronounced subnational variation in mDFPS [12].

Most previous Ethiopian studies, however, have examined social characteristics separately or as additive covariates in conventional regression models. Such approaches estimate average associations but may obscure how social positions combine. Intersectionality emphasizes that health opportunities and

constraints are produced through intersecting social locations rather than through isolated dimensions considered independently [13]. Quantitative intersectional analysis therefore requires methods that can estimate inequalities across combinations of characteristics without relying on unstable estimates from small cross-classified subgroups.

Multilevel analysis of individual heterogeneity and discriminatory accuracy (MAIHDA) provides such a framework. Individuals are nested within intersectional strata defined by combinations of social characteristics, allowing total between-stratum heterogeneity to be quantified and then partitioned into variation accounted for by additive main effects and residual stratum-specific departures from additive expectations [14, 15]. Partial pooling stabilizes estimates for sparsely populated strata, making MAIHDA particularly useful when multiple social dimensions are cross-classified in survey data [14]. Importantly, large differences between intersectional strata do not necessarily imply large residual interactions: MAIHDA can distinguish strong structural additive gradients from smaller residual intersectional departures.

Using the EDHS 2024–25, this study examined intersectional inequalities in mDFPS among women aged 15–49 years with demand for family planning. We aimed to: (1) describe survey-weighted variation in mDFPS across age, household wealth, education and residence; (2) quantify total between-stratum heterogeneity using MAIHDA; (3) estimate the extent to which this heterogeneity was accounted for by additive social gradients; (4) identify adequately represented intersectional strata with particularly low predicted mDFPS; and (5) assess robustness to a more granular age specification and alternative treatment of survey weights.

2. Methods

2.1 Study design and data source

This was a cross-sectional secondary analysis of the Ethiopia Demographic and Health Survey (EDHS) 2024–25 Individual Recode dataset [16]. The EDHS 2024–25 used standardized DHS questionnaires and a nationally representative, stratified, multistage cluster sampling design. The analysis retained the sampling weight, primary sampling unit and survey-stratum variables required to account for the complex survey design in descriptive estimation. The Individual Recode file contained 21,395 women aged 15–49 years.

The DHS Program provided access to the de-identified dataset after authorization for research use.

2.2 Study population and analytic sample

The target population comprised women aged 15–49 years with demand for family planning. Total demand was defined according to the revised DHS unmet-need framework: women were considered to have demand if they had an unmet need for spacing or limiting births, or were currently using contraception for spacing or limiting [17].

Women were included in the analytic sample if they had a defined mDFPS outcome and non-missing information on age, household wealth, educational attainment, residence, primary sampling unit, survey stratum and sampling weight. A total of 6,611 women met these criteria and constituted the final analytic sample.

2.3 Outcome definition

The primary outcome was demand for family planning satisfied with modern contraceptive methods (mDFPS). Among women with demand for family planning, mDFPS was coded 1 if the woman was currently using a modern contraceptive method and 0 otherwise.

Modern contraceptive methods were classified according to DHS/United Nations measurement guidance and included female and male sterilization, intrauterine devices, implants, injectables, oral contraceptive pills, male and female condoms, emergency contraception, lactational amenorrhea, and other methods classified as modern in the DHS recode. Fertility-awareness methods were classified according to the modern-method coding used in the EDHS 2024–25 recode. The resulting indicator therefore represented the proportion of women with family-planning demand whose demand was satisfied by a modern method, consistent with SDG indicator 3.7.1 and international measurement guidance [1, 5].

2.4 Intersectional dimensions and strata

Four social dimensions were selected a priori based on their substantive relevance to family-planning access and routine availability in DHS data: age, household wealth, educational attainment and place of residence.

For the primary analysis, age was categorized as 15–19, 20–24, 25–34 and 35–49 years; household wealth as poorest, poorer, middle, richer and richest; educational attainment as no education, primary, and secondary or higher; and residence as rural or urban. Cross-classification of these dimensions generated 120 theoretically possible intersectional strata (4 × 5 × 3 × 2).

To examine whether the relatively broad older age categories masked age-related heterogeneity, a sensitivity specification used seven age categories (15–19, 20–24, 25–29, 30–34, 35–39, 40–44 and 45–49 years), generating 210 theoretically possible strata (7 × 5 × 3 × 2).

2.5 Survey-weighted descriptive analysis

National and subgroup mDFPS estimates were calculated using the EDHS 2024–25 complex survey design, incorporating the individual sampling weights, primary sampling units and sampling strata. Estimates were produced nationally and by age, household wealth quintile, educational attainment and residence.

Results are presented as survey-weighted percentages with design-based 95% confidence intervals (CIs). Unweighted analytic sample sizes are reported alongside weighted estimates to make subgroup

precision and intersectional-stratum sparsity transparent.

The seven-category age analysis was conducted using the same design-based survey estimation framework and was used to examine age-pattern sensitivity independently of the MAIHDA models.

2.6 MAIHDA modelling strategy

Binary logistic multilevel analysis of individual heterogeneity and discriminatory accuracy (MAIHDA) models were fitted with women at level 1 nested within intersectional strata at level 2, following established guidance for logistic MAIHDA [14, 15].

For the primary weighted MAIHDA analysis, the DHS individual sampling weights were normalized to have a mean of 1 within the analytic sample:

$w_i^* = w_i / \bar{w}$

where w_i is the original DHS sampling weight, $\bar{w}$ is its analytic-sample mean, and w_i^* is the normalized weight. The normalized weights were supplied to the multilevel logistic models as prior/pseudo-likelihood weights. This approach incorporates differential individual selection probabilities but does not reproduce the full stratification and clustering structure of the complex DHS design within lme4. Accordingly, the MAIHDA estimates are interpreted as survey-weight-adjusted model-based estimates, rather than fully design-based multilevel estimates.

Two models were estimated for each intersectional specification. The null model contained only a random intercept for intersectional stratum and quantified total between-stratum heterogeneity. The additive model additionally included fixed effects for age group, household wealth quintile, educational attainment and residence while retaining the stratum-level random intercept.

The additive model therefore decomposed overall between-stratum heterogeneity into variation associated with the additive contributions of the social dimensions defining the strata and residual stratum-specific departures from those additive expectations.

Fixed-effect associations are reported as adjusted odds ratios (aORs) with 95% Wald confidence intervals and Wald P values.

2.7 Measures of between-stratum heterogeneity

For logistic MAIHDA models, the variance partition coefficient (VPC) was calculated using the latent-variable method:

$VPC = \sigma_u^2 / (\sigma_u^2 + \pi^2/3)$

where σ_u^2 is the estimated between-stratum random-intercept variance and $\pi^2/3$ is the level-1 logistic variance.

The proportional change in variance (PCV) quantified the reduction in between-stratum variance between the null and additive models:

$$PCV = [(\sigma_u,null^2 - \sigma_u,additive^2) / \sigma_u,null^2] \times 100$$

The null-model VPC was interpreted as the proportion of latent-scale variation partitioned between intersectional strata before adjustment for their constituent characteristics. The residual VPC from the additive model quantified the remaining between-stratum heterogeneity after accounting for additive age, wealth, education and residence gradients.

Stratum-specific empirical-Bayes random-intercept estimates were extracted from the additive model to characterize residual departures from additive expectations. Positive residuals indicated higher and negative residuals lower mDFPS than expected from the additive fixed-effects structure. These residuals were interpreted descriptively and not as causal interactions.

2.8 Stratum-specific predicted probabilities and policy-facing rankings

For each observed primary intersectional stratum, model-predicted mDFPS was calculated from the combined fixed-effects linear predictor and empirical-Bayes stratum-specific random intercept and transformed from the log-odds to the probability scale.

The additive expectation for each stratum was calculated using the fixed-effects component alone. The residual difference was defined as:

Residual difference (percentage points) = [P(full stratum prediction) − P(additive expectation)] × 100

and expressed in percentage points.

Approximate model-based 95% intervals around the stratum-specific predictions were calculated by combining uncertainty from the fixed-effects component with the conditional uncertainty of the estimated stratum random effect on the linear-predictor scale, before transformation to the probability scale. These intervals are therefore approximate model-based intervals and should not be interpreted as fully design-based survey confidence intervals.

All observed intersectional strata were retained in model estimation regardless of cell size. For the ranked policy-facing display only, strata containing fewer than 20 women were omitted to reduce the risk of overinterpreting rankings driven by very small cells. No such exclusion was applied during MAIHDA estimation.

2.9 Sparse strata and sensitivity analyses

Stratum sparsity was characterized using the number of theoretical and observed strata, the number of empty strata, minimum, median and mean stratum sample size, and the proportions of observed strata

containing fewer than 5, 10 and 20 women.

Sensitivity analyses included:

1. a more granular 210-strata MAIHDA specification based on seven age categories; and
2. an unweighted version of the primary 120-strata additive MAIHDA model.

The 210-strata specification evaluated whether the primary variance decomposition was sensitive to finer age stratification. The unweighted analysis assessed sensitivity of empirical-Bayes stratum residuals to incorporation of normalized DHS sampling weights.

Concordance between weighted and unweighted stratum residuals was evaluated using the Spearman rank correlation coefficient (ρ). Model singularity was assessed for all fitted multilevel models. Akaike information criterion (AIC) values were reported descriptively and compared only between models fitted using the same weighting framework, outcome, analytic sample and likelihood specification; weighted and unweighted AIC values were not directly compared.

2.10 Statistical software

All analyses were conducted in R version 4.6.0. Complex-survey descriptive analyses were performed using the survey and srvyr frameworks, and multilevel logistic MAIHDA models were fitted using lme4. Additional data management, model extraction, tabulation and visualization were undertaken using supporting R packages.

Statistical results are reported as survey-weighted percentages with design-based 95% CIs, aORs with 95% CIs, between-stratum variances, VPCs, PCVs, empirical-Bayes stratum residuals and model-predicted probabilities.

Because the study was cross-sectional and descriptive, associations, residual stratum effects and predicted probabilities were not interpreted as causal effects.

OpenAI ChatGPT was used during manuscript preparation for language editing, organization and clarity. It was not used to generate or alter the study data or statistical results. The author reviewed and verified all scientific content, analyses, interpretations and references and remains fully responsible for the manuscript.

2.11 Ethical considerations

This study used de-identified secondary EDHS 2024–25 data obtained through the DHS Program following authorization for research use. The original survey followed established DHS ethical procedures, including review by relevant ethical bodies, informed consent from participants and safeguards for respondent confidentiality [16].

No direct participant contact, collection of new data or access to personally identifiable information occurred in this secondary analysis.

3. Results

3.1 Analytic sample and intersectional structure

The EDHS 2024–25 Individual Recode dataset contained 21,395 women aged 15–49 years. Of these, 6,611 women had demand for family planning and constituted the final analytic sample. The primary intersectional classification generated 120 theoretically possible strata, of which 112 were observed and eight were empty. In the more granular 210-strata sensitivity specification, 192 strata were observed and 18 were empty. All fitted MAIHDA models included the full analytic sample of 6,611 women, and none showed evidence of a singular fit (Supplementary Table S1).

3.2 Survey-weighted inequalities in mDFPS

National survey-weighted mDFPS was 67.3%. Substantial variation was evident across age, household wealth, educational attainment and residence (Table 1).

mDFPS showed a distinctly non-linear age pattern. Among adolescents aged 15–19 years, mDFPS was 61.9% (95% CI 53.4–70.4). It increased to 74.8% (70.2–79.4) among women aged 20–24 years and remained comparatively high at 72.6% (69.6–75.6) among those aged 25–34 years, before declining to 56.9% (53.2–60.7) among women aged 35–49 years (Fig. 1).

Clear socioeconomic gradients were also observed. Survey-weighted mDFPS increased progressively from 43.8% (38.5–49.0) among women in the poorest wealth quintile to 77.3% (73.1–81.6) among those in the richest quintile. Similarly, mDFPS increased from 53.7% (49.7–57.6) among women with no formal education to 78.3% (75.4–81.2) among those with secondary or higher education. Rural women had lower mDFPS than urban women: 62.8% (59.6–65.9) versus 75.1% (71.2–79.0), respectively.

The seven-category age analysis revealed substantial heterogeneity within the broad 35–49-year category used in the primary MAIHDA specification. mDFPS peaked at 74.8% among women aged 20–24 years and 74.6% among those aged 25–29 years, then declined progressively to 69.9% at ages 30–34, 62.6% at 35–39, 54.8% at 40–44 and 38.5% at 45–49 years (Supplementary Figure S4). Thus, the lower mDFPS observed in the broad 35–49-year group was driven in part by a pronounced decline across later reproductive age.

Table 1
Survey-weighted demand satisfied with modern contraceptive methods by population characteristics

Population characteristic	Unweighted n	mDFPS, % (95% CI)
Age group		
15–19	397	61.9 (53.4–70.4)
20–24	1115	74.8 (70.2–79.4)
25–34	2975	72.6 (69.6–75.6)
35–49	2124	56.9 (53.2–60.7)
Household wealth		
Poorest	1354	43.8 (38.5–49.0)
Poorer	1185	61.0 (56.0–66.0)
Middle	1133	67.2 (62.8–71.6)
Richer	1167	73.2 (68.4–78.0)
Richest	1772	77.3 (73.1–81.6)
Education		
No education	2144	53.7 (49.7–57.6)
Primary	2553	71.0 (67.3–74.8)
Secondary or higher	1914	78.3 (75.4–81.2)
Residence		
Rural	4159	62.8 (59.6–65.9)
Urban	2452	75.1 (71.2–79.0)

Note: Values are unweighted analytic sample sizes and survey-weighted mDFPS percentages with design-based 95% confidence intervals. Estimates account for DHS sampling weights, primary sampling units and survey strata. The analytic population comprises women aged 15–49 years with demand for family planning. mDFPS, demand for family planning satisfied with modern contraceptive methods.

3.3 Additive social gradients in the primary MAIHDA model

The primary additive 120-strata MAIHDA model showed strong independent gradients by age, household wealth and educational attainment (Table 2). Compared with adolescents aged 15–19 years, women aged 20–24 years had nearly twice the adjusted odds of mDFPS (aOR 1.97, 95% CI 1.30–2.99; P = 0.001), while women aged 25–34 years also had higher adjusted odds (aOR 1.75, 1.19–2.57; P = 0.005).

Women aged 35–49 years did not differ statistically from adolescents after adjustment (aOR 0.93, 0.62–1.38; P = 0.708).

A graded wealth association remained after mutual adjustment. Relative to women in the poorest quintile, the adjusted odds of mDFPS were higher among those in the poorer (aOR 1.96, 95% CI 1.28–3.01), middle (2.65, 1.75–4.01), richer (3.41, 2.29–5.07) and richest (3.76, 2.36–5.97) quintiles.

Educational attainment was also independently associated with mDFPS. Compared with women with no education, those with primary education had 1.57 times the adjusted odds of mDFPS (95% CI 1.17–2.11), while those with secondary or higher education had more than twice the adjusted odds (aOR 2.21, 1.57–3.11).

In contrast, the crude urban–rural difference was substantially attenuated after adjustment for age, wealth and education (urban vs rural aOR 0.98, 95% CI 0.73–1.32; P = 0.891). This attenuation indicates that the observed residence gradient was largely accounted for by differences in the other social-position characteristics included in the model.

Table 2
Fixed effects from the primary additive 120-strata MAIHDA model

Comparison	Adjusted OR (95% CI)	P value
Age 20–24 vs 15–19	1.97 (1.30–2.99)	0.001
Age 25–34 vs 15–19	1.75 (1.19–2.57)	0.005
Age 35–49 vs 15–19	0.93 (0.62–1.38)	0.708
Poorer vs poorest	1.96 (1.28–3.01)	0.002
Middle vs poorest	2.65 (1.75–4.01)	< 0.001
Richer vs poorest	3.41 (2.29–5.07)	< 0.001
Richest vs poorest	3.76 (2.36–5.97)	< 0.001
Primary vs no education	1.57 (1.17–2.11)	0.003
Secondary or higher vs no education	2.21 (1.57–3.11)	< 0.001
Urban vs rural	0.98 (0.73–1.32)	0.891

Note: Adjusted odds ratios are conditional on the other additive social-position variables and the stratum-level random intercept. Reference categories were age 15–19 years, poorest household wealth quintile, no education and rural residence. CIs are Wald 95% confidence intervals; P values are Wald tests.

3.4 Between-stratum heterogeneity and proportional change in variance

Substantial between-stratum heterogeneity was present before adjustment for the constituent social-position characteristics. In the primary 120-strata null model, the estimated between-stratum variance was 0.858, corresponding to a VPC of 20.7%. Thus, approximately one-fifth of latent-scale variation in mDFPS was partitioned between intersectional strata before adjustment (Table 3).

Following inclusion of additive age, household wealth, education and residence effects, the between-stratum variance declined to 0.175 and the residual VPC to 5.1%. This corresponded to a PCV of 79.6%, indicating that approximately four-fifths of the initial between-stratum heterogeneity was accounted for by the additive social dimensions defining the strata.

The 210-strata sensitivity specification produced a similar pattern. The VPC decreased from 23.6% in the null model to 6.6% after additive adjustment, with a corresponding PCV of 77.0%. The close correspondence between the two specifications indicates that the principal variance-decomposition finding was robust to finer age stratification.

Taken together, these findings indicate that substantial inequalities existed across intersectional social positions, but most initial between-stratum heterogeneity reflected strong additive age and socioeconomic gradients. A smaller component of residual intersectional heterogeneity remained after these gradients were accounted for.

Table 3
Between-stratum variance, variance partition coefficient and proportional change in variance in primary and sensitivity MAIHDA models

Model specification	Between-stratum variance	VPC (%)	PCV (%)
Primary 120-strata: null	0.858	20.7	-
Primary 120-strata: additive	0.175	5.1	79.6
Sensitivity 210-strata: null	1.017	23.6	-
Sensitivity 210-strata: additive	0.234	6.6	77.0

Note: VPC, variance partition coefficient; PCV, proportional change in between-stratum variance relative to the corresponding null model. VPCs are calculated on the latent logistic scale.

3.5 Partial pooling and sparse intersectional strata

Comparison of directly estimated survey-weighted stratum proportions with MAIHDA-predicted probabilities demonstrated the expected partial-pooling behaviour of the multilevel model (Fig. 2). Predictions for well-represented strata generally remained close to their observed survey-weighted estimates, whereas predictions for sparsely represented strata were more strongly displaced from extreme observed values toward estimates supported by the multilevel model.

This pattern was particularly evident for small strata with observed mDFPS estimates near 0% or 100%. The MAIHDA predictions for these strata were generally less extreme, illustrating how partial pooling

reduced instability associated with small intersectional cells while retaining all observed strata in estimation.

Sparsity was substantial in both intersectional specifications (Supplementary Figure S2). In the primary 120-strata specification, 22 of 112 observed strata (19.6%) contained fewer than five women, 37 (33.0%) contained fewer than 10, and 55 (49.1%) contained fewer than 20. Sparsity increased under the 210-strata specification: 110 of 192 observed strata (57.3%) contained fewer than 20 women, and the median stratum sample size declined from 22 to 13.

All observed strata were retained in the MAIHDA models regardless of sample size. The n ≥ 20 criterion was applied only to ranked policy-facing displays and did not determine inclusion in model estimation.

3.6 Residual intersectional heterogeneity

Residual between-stratum heterogeneity remained after adjustment for additive social gradients, although it was considerably smaller than in the null model. The primary additive model had a residual between-stratum variance of 0.175 and VPC of 5.1%.

Empirical-Bayes stratum-specific random-intercept residuals were generally concentrated around zero, and many conditional 95% intervals crossed zero. Figure 3 presents the largest positive and negative residual estimates as descriptive departures from additive expectations. The full residual distribution is provided in Supplementary Figure S1.

Positive residuals represented strata with modelled mDFPS higher than expected from the additive combination of age, household wealth, education and residence, whereas negative residuals represented lower-than-expected mDFPS. However, several of the most extreme estimates had wide conditional intervals, indicating substantial uncertainty.

Accordingly, the residual distribution supports evidence of modest remaining intersectional heterogeneity after additive adjustment, rather than widespread large residual interactions. Individual residual rankings should be regarded as descriptive signals for further investigation rather than as definitive or causal interactions.

3.7 Intersectional strata with the lowest predicted mDFPS

The ranked policy-facing analysis was restricted to intersectional strata containing at least 20 women, while all observed strata remained included in the fitted MAIHDA model (Fig. 4; Table 4).

Among displayed strata, the lowest model-predicted mDFPS was observed among rural adolescents aged 15–19 years in the poorest wealth quintile with no education: 19.6% (approximate model-based 95% interval 9.6–36.1). This estimate was substantially below the national survey-weighted mDFPS of 67.3%.

Other strata with particularly low predicted mDFPS included poorest, uneducated women aged 35–49 years living in urban settings (28.8%, 15.3–47.5); poorest, uneducated rural women aged 35–49 years (33.5%, 24.0–44.6); and poorest, uneducated rural women aged 20–24 years (34.4%, 21.7–49.9).

Additional low-ranked strata frequently combined poorer household wealth, no or primary education and rural residence. The consistency of these combinations is more informative than the exact ordering of individual strata: low predicted mDFPS was concentrated particularly where economic disadvantage, limited education and vulnerable life-course positions overlapped.

Notably, low predicted mDFPS did not necessarily imply a large negative residual intersectional effect. For example, several disadvantaged strata had low predicted probabilities primarily because of their additive age, wealth and educational profiles rather than unusually negative residual departures. The ranking therefore identifies population groups with low predicted coverage, whereas the residual-effect analysis addresses a different question—whether strata perform above or below their additive expectations.

Table 4
Intersectional strata with the lowest model-predicted mDFPS among strata containing at least 20 women

Intersectional stratum	n	Predicted mDFPS, % (approx. 95% interval)	Additive expectation (%)	Residual difference (percentage points)
15–19 | Poorest | No education | Rural	26	19.6 (9.6–36.1)	29.1	-9.4
35–49 | Poorest | No education | Urban	29	28.8 (15.3–47.5)	27.1	+ 1.7
35–49 | Poorest | No education | Rural	305	33.5 (24.0-44.6)	27.5	+ 6.0
20–24 | Poorest | No education | Rural	79	34.4 (21.7–49.9)	44.7	-10.3
15–19 | Poorest | Primary | Rural	48	38.6 (24.5–54.9)	39.2	-0.6
25–34 | Poorest | No education | Rural	299	40.5 (30.0-52.1)	41.7	-1.2
25–34 | Poorest | No education | Urban	23	42.2 (24.8–61.7)	41.2	+ 0.9
35–49 | Poorest | Primary | Rural	69	43.5 (28.4–59.9)	37.4	+ 6.1
35–49 | Richer | Primary | Rural	101	51.4 (39.7–62.9)	67.0	-15.6
35–49 | Poorer | Primary | Rural	95	51.7 (38.1–65.0)	53.9	-2.2
20–24 | Poorest | Primary | Rural	121	53.3 (39.6–66.6)	56.0	-2.6
35–49 | Middle | No education | Rural	252	53.9 (43.3–64.2)	50.2	+ 3.7
25–34 | Poorer | No education | Rural	164	55.4 (43.5–66.8)	58.4	-2.9
35–49 | Poorer | No education | Rural	288	56.3 (45.3–66.8)	42.6	+ 13.7
15–19 | Poorer | Primary | Rural	67	56.7 (41.7–70.6)	55.8	+ 0.9

Note: All observed strata were retained in model estimation. The n ≥ 20 criterion applies only to this ranked display. Predicted mDFPS values are accompanied by approximate model-based 95% intervals and are not fully design-based survey confidence intervals. The additive expectation is the predicted probability

obtained from the fixed-effects component of the additive MAIHDA model. Residual difference is the full stratum-specific predicted probability minus the additive expectation, expressed in percentage points.

3.8 Sensitivity analyses and model robustness

Sensitivity analyses generally supported the primary findings. The more granular 210-strata specification reproduced the marked reduction in between-stratum heterogeneity after additive adjustment: VPC declined from 23.6% in the null model to 6.6% in the additive model, corresponding to a PCV of 77.0%, compared with 79.6% in the primary 120-strata specification.

The more granular specification generated greater sparsity, however. As shown in Supplementary Figure S2 and detailed in Supplementary Table S2, 57.3% of observed 210-strata cells contained fewer than 20 women, compared with 49.1% in the primary specification, and median stratum size declined from 22 to 13. These findings support use of the 120-strata specification as the primary model while retaining the 210-strata specification as a robustness analysis.

All five fitted models were non-singular. Within comparable weighted specifications, additive models had lower AIC values than their corresponding null models: 6,860.6 versus 6,908.2 for the 120-strata models and 6,782.5 versus 6,868.4 for the 210-strata models (Supplementary Table S1). These within-framework comparisons were consistent with improved model fit following inclusion of age, wealth, education and residence. AIC values from weighted and unweighted models were not compared directly.

Empirical-Bayes stratum residuals from the weighted and unweighted primary additive models showed moderate rank concordance (Spearman $\rho = 0.60$; Supplementary Figure S3). Thus, evidence of residual stratum heterogeneity was retained across weighting approaches, but the magnitude and rank ordering of individual residuals showed some sensitivity to incorporation of DHS sampling weights.

The complete residual distribution (Supplementary Figure S1), together with the moderate weighted–unweighted rank concordance, reinforces the interpretation that residual effects represent modest model-based departures from additive expectations rather than stable causal interactions. Exact residual rankings should therefore not be overinterpreted.

The seven-category age sensitivity analysis also supported the primary conclusion of non-linear age patterning. mDFPS peaked among women aged 20–29 years and subsequently declined across older reproductive ages, reaching 38.5% among women aged 45–49 years (Supplementary Figure S4).

4. Discussion

4.1 Principal findings

This study examined intersectional inequalities in demand for family planning satisfied with modern contraceptive methods among women aged 15–49 years with family-planning demand in Ethiopia. Four

findings are particularly important. First, national survey-weighted mDFPS was 67.3%, but this average concealed substantial inequalities by reproductive age, household wealth, educational attainment and residence. Second, approximately one-fifth of latent-scale variation in mDFPS was partitioned between intersectional strata before adjustment (VPC 20.7%). Third, inclusion of additive age, wealth, education and residence effects reduced between-stratum variance by 79.6%, leaving a residual VPC of 5.1%. Thus, most initial between-stratum heterogeneity was accounted for by additive social gradients rather than large residual departures from additivity. Fourth, despite this reduction, several combinations of social disadvantage had very low predicted mDFPS, particularly those combining poverty, limited education and younger or older reproductive age.

These findings refine how intersectional inequality in contraceptive coverage should be interpreted. Large differences between socially defined population groups do not necessarily imply correspondingly large statistical interactions among their constituent characteristics. In the present analysis, inequalities across intersectional strata were substantial, but much of their structure reflected the accumulation of disadvantages associated with age, wealth and education. Residual stratum-specific departures from these additive expectations were comparatively modest. This distinction is consistent with the MAIHDA framework, which separates heterogeneity across intersectional strata from the additional heterogeneity remaining after additive social-position effects have been modelled [14, 15].

4.2 Age and life-course inequalities

The relationship between age and mDFPS was distinctly non-linear. Coverage was lower among adolescents aged 15–19 years, highest during the twenties and substantially lower in later reproductive age. The seven-category age analysis showed that the broad 35–49-year category concealed a pronounced decline, with mDFPS decreasing from 62.6% at ages 35–39 to 54.8% at ages 40–44 and 38.5% at ages 45–49.

The lower coverage observed among adolescents is consistent with Ethiopian evidence showing elevated unmet need and age-related differences in contraceptive use among adolescent girls and young women [9, 10]. Adolescents may encounter barriers associated with confidentiality, social stigma, provider attitudes, limited decision-making autonomy and uneven access to youth-responsive contraceptive services. These barriers are also consistent with broader evidence that contraceptive use varies substantially by age and life-course stage [1, 18, 19].

The decline in mDFPS at older reproductive ages is equally important. Family-planning equity discussions frequently emphasize adolescents and young women, whereas women approaching the end of reproductive age may receive less attention despite continuing pregnancy risk and contraceptive need. Global age-specific evidence indicates substantial variation in contraceptive use and method patterns across reproductive age [18, 19]. Lower mDFPS among older women could reflect differences in fertility intentions, perceived pregnancy risk, sexual activity, method discontinuation, changing method preferences or reduced engagement with family-planning services. These mechanisms cannot be

distinguished using the present cross-sectional data and should therefore be investigated rather than inferred from the observed association.

The finding nevertheless has an important measurement implication: grouping all women aged 35–49 years together may obscure substantial within-group heterogeneity. Future family-planning monitoring should therefore consider more granular age disaggregation where sample size permits.

4.3 Wealth and educational inequalities

Household wealth showed one of the clearest gradients in the analysis. Survey-weighted mDFPS increased from 43.8% among women in the poorest quintile to 77.3% among those in the richest, and a strong graded association remained after adjustment. Women in the richest quintile had almost four times the adjusted odds of mDFPS compared with those in the poorest quintile.

These findings are consistent with recent evidence documenting persistent socioeconomic inequalities in modern contraceptive use and demand satisfied across FP2030 focus countries [6]. Ethiopian studies have similarly identified socioeconomic inequality in contraceptive demand, unmet need and modern-method use [8, 11]. The persistence of the wealth gradient indicates that improvements in national contraceptive coverage do not necessarily translate into equitable satisfaction of family-planning demand.

Educational inequality was similarly pronounced. mDFPS increased from 53.7% among women with no education to 78.3% among those with secondary or higher education, and education remained strongly associated with mDFPS after adjustment. Previous Ethiopian and sub-Saharan African studies have likewise identified education as an important correlate of modern contraceptive use and demand satisfaction [7–11].

Education may operate through several pathways, including health literacy, contraceptive knowledge, economic opportunity, reproductive autonomy, communication with partners, and confidence in navigating health services. Wealth may influence transport costs, opportunity costs, access to information and ability to overcome service-access barriers. These pathways are likely interrelated and should not be interpreted as independent causal mechanisms from the present cross-sectional analysis.

The crude urban–rural difference was substantial, but residence was close to the null after adjustment for age, wealth and education. This attenuation suggests that differences in these social characteristics accounted for much of the observed residence gradient. It does not imply that geography or health-service accessibility is unimportant. Ethiopian and regional evidence demonstrates substantial geographic and subnational variation in contraceptive coverage and unmet need [8, 11, 12]. Rather, the finding indicates that residence alone may be an insufficient marker of family-planning disadvantage when socioeconomic composition is not considered.

4.4 Intersectional heterogeneity: additive structure versus residual departures

A central contribution of this study is the distinction between inequality across intersectional strata and residual intersectional heterogeneity after accounting for additive social gradients. In the null model, the VPC of 20.7% indicated meaningful clustering of latent-scale variation in mDFPS across combinations of age, wealth, education and residence. After these constituent dimensions were included as additive fixed effects, the between-stratum variance declined from 0.858 to 0.175 and the VPC to 5.1%, corresponding to a PCV of 79.6%.

Contemporary MAIHDA guidance emphasizes that this reduction can be used to distinguish heterogeneity associated with the additive distribution of social characteristics from residual stratum-specific departures beyond those additive expectations [14, 15]. The present findings therefore support a more nuanced interpretation of intersectional contraceptive inequality. Intersectional strata were highly informative for locating disadvantaged population combinations, but most initial between-stratum variation was accounted for by strong additive life-course and socioeconomic gradients rather than by large residual statistical interactions.

This distinction is substantively important. A population group may experience exceptionally low contraceptive coverage because it simultaneously occupies several disadvantaged social positions even when there is little additional departure from the additive combination of those characteristics. Intersectionality in population-health analysis should therefore not be reduced to the presence or absence of statistical interaction [13–15].

The empirical-Bayes residual distribution reinforced this interpretation. Most residual estimates were relatively close to zero, many conditional 95% intervals crossed zero, and weighted and unweighted residual rankings showed only moderate concordance (Spearman $\rho = 0.60$). Consequently, the most extreme residual estimates should not be interpreted as definitive evidence of synergistic or antagonistic interactions. Their primary value is diagnostic: they identify combinations whose observed modelled performance differs from additive expectations and may therefore merit further investigation.

4.5 Partial pooling and the value of MAIHDA

The substantial number of sparsely populated intersectional strata illustrates an important methodological advantage of MAIHDA. Almost half of the observed primary strata contained fewer than 20 women, and sparsity was even greater under the 210-strata sensitivity specification. Conventional subgroup analyses based on direct estimates would therefore generate unstable proportions for many combinations, including values close to 0% or 100%.

Through partial pooling, MAIHDA borrows information across intersectional strata and shrinks estimates from sparsely represented groups toward values supported by the broader multilevel structure [14, 15]. The observed-versus-predicted analysis demonstrated this clearly: predictions for large strata remained relatively close to their direct survey estimates, whereas estimates from small strata were more strongly moderated.

Partial pooling does not remove uncertainty or make very small strata equivalent to well-represented groups. Rather, it reduces the influence of sampling noise while allowing socially meaningful combinations to remain in the analysis. This is particularly valuable for equity research, where the populations of greatest substantive interest may also be those least well represented in conventional survey subgroup analyses.

4.6 Population groups with low predicted mDFPS

The predicted-probability analysis translated the multilevel findings into a policy-relevant scale. Among strata containing at least 20 women, rural adolescents aged 15–19 years in the poorest household wealth quintile with no education had the lowest model-predicted mDFPS, at 19.6%, compared with the national survey-weighted estimate of 67.3%.

Several other low-predicted strata combined poverty, no or primary education, rural residence and either adolescent or later reproductive age. The recurrence of these characteristics is more informative than the precise ordering of individual strata. It indicates that particularly low contraceptive coverage is concentrated where established dimensions of socioeconomic and life-course disadvantage accumulate.

Importantly, low predicted mDFPS and large negative residual intersectional departures are not equivalent. Some strata had low predicted coverage primarily because of their additive age, wealth and educational profiles, while their residual departure from the additive expectation was small or even positive. The predicted-probability analysis therefore identifies who has low coverage, whereas the residual analysis addresses whether a stratum performs better or worse than expected from its constituent characteristics.

The strata identified here should consequently not be interpreted as a definitive causal ranking of populations. Model-based intervals remained relatively wide for some combinations, and the ranking of residual effects showed some sensitivity to weighting. Rather, these estimates provide an equity-monitoring framework for identifying population combinations that warrant closer programmatic assessment.

4.7 Implications for policy and programming

The findings suggest that family-planning equity strategies in Ethiopia should address both broad structural gradients and the concentration of disadvantage within particular combinations of social positions.

First, routine family-planning monitoring should move beyond national averages and report mDFPS across reproductive age, wealth and educational groups. Where data permit, carefully selected intersectional disaggregation may reveal population combinations that remain obscured within single-axis indicators.

Second, the pronounced non-linear age pattern suggests a need for life-course-responsive family-planning services. Adolescents may require confidential, non-judgmental and youth-responsive services, while women in later reproductive age may require greater attention to continuing contraceptive need, changing fertility preferences, appropriate method choice and counselling about pregnancy risk [1, 18, 19].

Third, the strong wealth and education gradients suggest that improving contraceptive commodity availability alone is unlikely to eliminate inequality. Equity-oriented programming may also need to address indirect financial barriers, information gaps, health literacy, informed choice, service responsiveness and continuity of care.

Fourth, intersectional targeting should complement rather than replace population-wide strengthening of voluntary family-planning services. The findings favour a proportionate equity approach: universal access should remain the objective, with greater programmatic attention directed toward groups experiencing accumulated socioeconomic and life-course disadvantage.

Finally, the attenuation of the residence association after socioeconomic adjustment suggests that broad rural–urban targeting may be insufficient on its own. Programmes may obtain greater equity gains by identifying the socioeconomic and educational disadvantages that occur within geographic settings rather than treating residence as a stand-alone proxy for need.

4.8 Robustness and methodological implications

Sensitivity analyses supported the principal interpretation. In the more granular 210-strata specification, the null-model VPC was 23.6% and declined to 6.6% after additive adjustment, corresponding to a PCV of 77.0%. This was close to the 79.6% PCV observed in the primary specification, indicating that the conclusion that additive social gradients accounted for most initial between-stratum heterogeneity was not dependent on the primary age categorization.

Greater granularity, however, increased sparsity. The proportion of observed strata containing fewer than 20 women increased from 49.1% in the primary specification to 57.3% in the 210-strata model, and median stratum size declined from 22 to 13. The 120-strata model therefore offered a more favourable balance between substantive differentiation and statistical support.

All fitted models were non-singular, and additive models had lower AIC values than corresponding null models within the same weighting framework. These diagnostics support model stability, although AIC values across weighted and unweighted analyses are not directly comparable.

Weighted and unweighted empirical-Bayes residuals showed moderate rank concordance ($\rho = 0.60$). This finding suggests that the broad presence of residual heterogeneity was not entirely dependent on the weighting approach, but the exact magnitude and ordering of individual stratum residuals were less stable. Accordingly, policy interpretation should emphasize the robust population-level patterns—

particularly age, wealth and educational inequalities and the variance decomposition—rather than the precise ranking of individual empirical-Bayes residuals.

4.9 Strengths and limitations

This study has several strengths. It used recent nationally representative Ethiopia DHS data and examined mDFPS, an internationally recognized family-planning indicator that conditions contraceptive coverage on expressed demand [3, 5, 16]. Design-based survey methods were used for national and subgroup descriptive estimates, while MAIHDA provided a framework for examining heterogeneity across multiple cross-classified social positions. Partial pooling allowed sparse observed strata to remain in the analysis rather than being discarded. The primary 120-strata specification was evaluated against a more granular 210-strata model, and sensitivity to the treatment of individual sampling weights was explicitly assessed. The study also distinguished low predicted coverage from residual departures from additive expectations, reducing the risk of treating every intersectional inequality as statistical interaction.

Several limitations require consideration. First, the cross-sectional design precludes causal inference or determination of temporal relationships between social position and contraceptive demand satisfaction. Second, DHS measures are self-reported and may be affected by recall, reporting or social-desirability biases.

Third, the four dimensions used to construct the intersectional strata—age, wealth, education and residence—capture only part of the social structure potentially relevant to contraceptive access. Other dimensions, including region, marital status, parity, disability, ethnicity, religion, migration, displacement, conflict exposure and partner characteristics, may also shape family-planning opportunities but were not incorporated into the primary intersectional classification.

Fourth, normalized DHS individual sampling weights were incorporated into the lme4 models as prior/pseudo-likelihood weights. This approach accounts for differential individual selection probabilities but does not reproduce the complete multistage survey design within the multilevel likelihood. The MAIHDA estimates should therefore be interpreted as survey-weight-adjusted model-based estimates rather than fully design-based multilevel estimates. The unweighted sensitivity analysis was included to assess the robustness of stratum residuals to this modelling choice.

Fifth, many intersectional strata were sparsely represented. Partial pooling reduces instability but does not eliminate uncertainty. The $n \geq 20$ threshold used for the policy-facing ranking was therefore a reporting criterion rather than a formal inferential threshold and was not used to determine inclusion in model estimation.

Sixth, the approximate model-based intervals presented for stratum-specific predicted probabilities are not fully design-based survey confidence intervals. Similarly, empirical-Bayes random-intercept residuals are shrinkage estimates and should not be interpreted as direct causal interaction parameters.

Finally, mDFPS measures whether contraceptive demand is met with a modern method but does not directly assess method preference, voluntariness, contraceptive autonomy, quality of counselling, side-effect management, continuity, discontinuation or person-centred experience. High mDFPS should consequently not be interpreted as a complete measure of high-quality or rights-based family-planning care [1].

4.10 Implications for future research

Future research should investigate the mechanisms underlying low mDFPS at both the younger and older ends of reproductive age and among women experiencing combined economic and educational disadvantage. Mixed-methods and longitudinal studies could examine the roles of fertility intentions, contraceptive autonomy, service accessibility, provider behaviour, method availability, discontinuation and financial barriers.

Future quantitative analyses could extend the MAIHDA framework to additional dimensions where theory and sample size support meaningful cross-classification. Geographic or spatial analyses may also help distinguish socioeconomic composition from contextual service-access effects, given the substantial subnational variation documented in Ethiopia and neighbouring settings [12].

Repeated application of intersectional inequality analyses across successive Ethiopian surveys would be particularly valuable for assessing whether disparities are narrowing, widening or changing composition over time. Such analyses could complement conventional equity measures by distinguishing changes in broad structural gradients from changes in residual intersectional heterogeneity.

5. Conclusion

Demand for family planning satisfied with modern contraceptive methods remains substantially unequal among women with family-planning demand in Ethiopia. Although national mDFPS was 67.3%, this average concealed pronounced inequalities across age, household wealth and educational attainment and across their intersectional combinations. Particularly low predicted coverage was concentrated among women experiencing overlapping economic, educational and life-course disadvantage.

MAIHDA provided an important distinction between inequalities across intersectional social positions and residual departures beyond additive social gradients. Approximately one-fifth of latent-scale variation in mDFPS was initially partitioned between intersectional strata, but adjustment for additive age, wealth, education and residence effects reduced between-stratum variance by nearly 80%, leaving comparatively modest residual heterogeneity. Thus, the dominant pattern was not widespread large residual interaction, but the accumulation of established socioeconomic and life-course disadvantages within the same population groups.

Family-planning policy and monitoring in Ethiopia should therefore complement national coverage indicators and single-dimension disaggregation with more granular equity assessment. Universal strengthening of voluntary, high-quality family-planning services should remain the foundation, with

proportionately greater attention to population groups in which poverty, limited education and vulnerable reproductive-age positions converge. Intersectional monitoring, supported by methods that appropriately account for sparse subgroups and uncertainty, can help identify populations at risk of being concealed by national averages and support more equitable progress toward satisfying contraceptive demand.

AIC, Akaike information criterion; aOR, adjusted odds ratio; CI, confidence interval; DHS, Demographic and Health Survey; EDHS, Ethiopia Demographic and Health Survey; MAIHDA, multilevel analysis of individual heterogeneity and discriminatory accuracy; mDFPS, demand for family planning satisfied with modern contraceptive methods; PCV, proportional change in variance; SDG, Sustainable Development Goal; VPC, variance partition coefficient.

Declarations

Ethics approval and consent to participate

This secondary analysis used de-identified data from the DHS Program. The original survey obtained ethical approvals and informed consent under DHS procedures. No new participant contact or collection of identifiable information occurred.

Consent for publication

Not applicable.

Competing interests

The author declares no competing interests.

Funding

This study received no specific funding.

Author Contribution

GMA conceived the study, developed the analytical approach, conducted the statistical analysis, interpreted the findings, prepared the figures and tables, drafted the manuscript, critically revised it for important intellectual content, and approved the final manuscript.

Data Availability

The EDHS 2024–25 data used in this study are available from the DHS Program subject to registration and approval of a data-access request.

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Figures

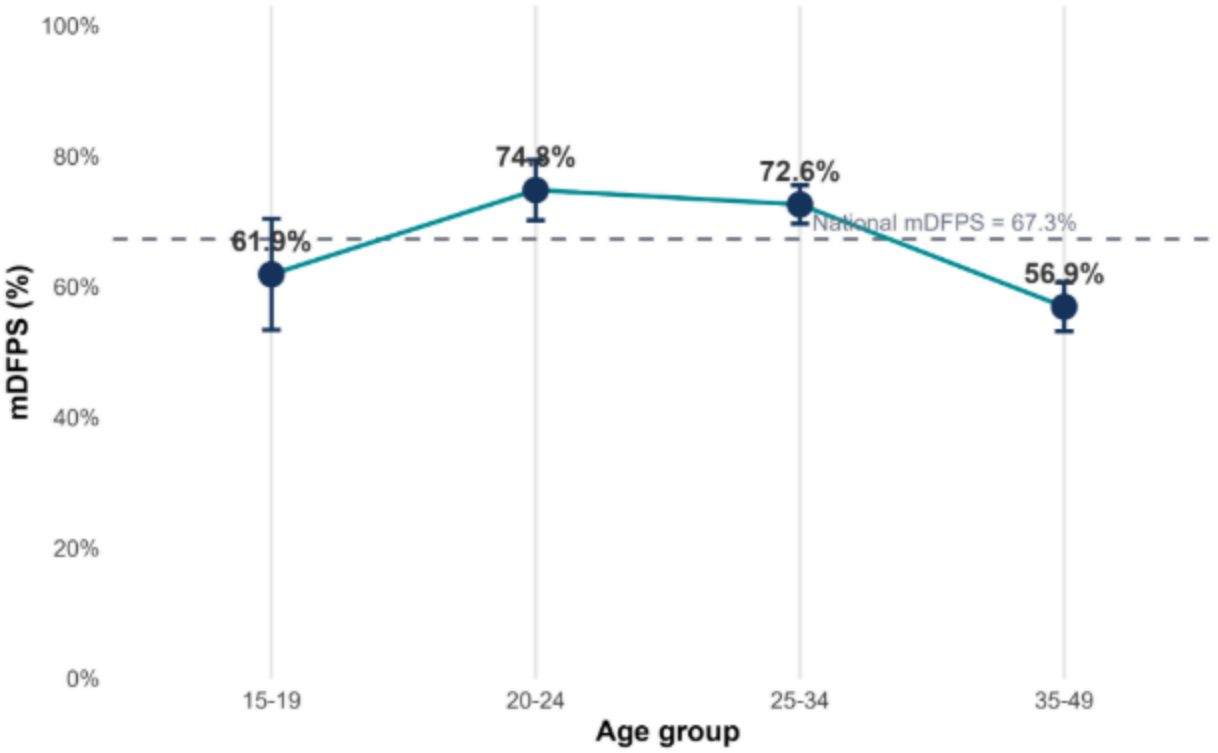


Figure 1

Survey-weighted mDFPS by primary age group, EDHS 2024–25. Points represent survey-weighted estimates and error bars represent design-based 95% confidence intervals among women aged 15–49 years with demand for family planning. The dashed horizontal line indicates the national survey-weighted mDFPS estimate of 67.3%.

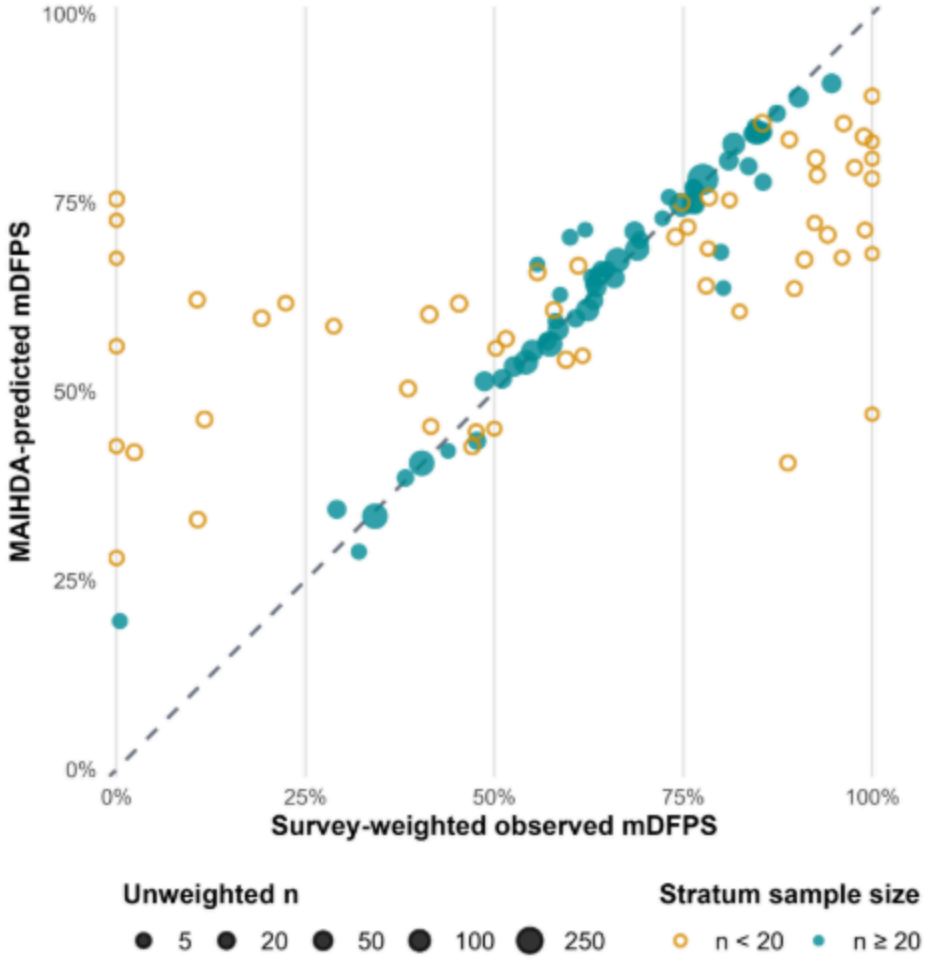


Figure 2

Survey-weighted observed versus MAIHDA-predicted mDFPS across primary intersectional strata. Each point represents an observed stratum in the primary 120-strata specification. The x-axis shows directly estimated survey-weighted mDFPS and the y-axis shows partially pooled probabilities predicted from the primary MAIHDA model. Point size represents unweighted stratum sample size; filled points indicate strata with at least 20 observations and open points indicate strata with fewer than 20 observations. The dashed diagonal denotes perfect agreement. Greater displacement from the diagonal among sparse strata illustrates model-based shrinkage through partial pooling.

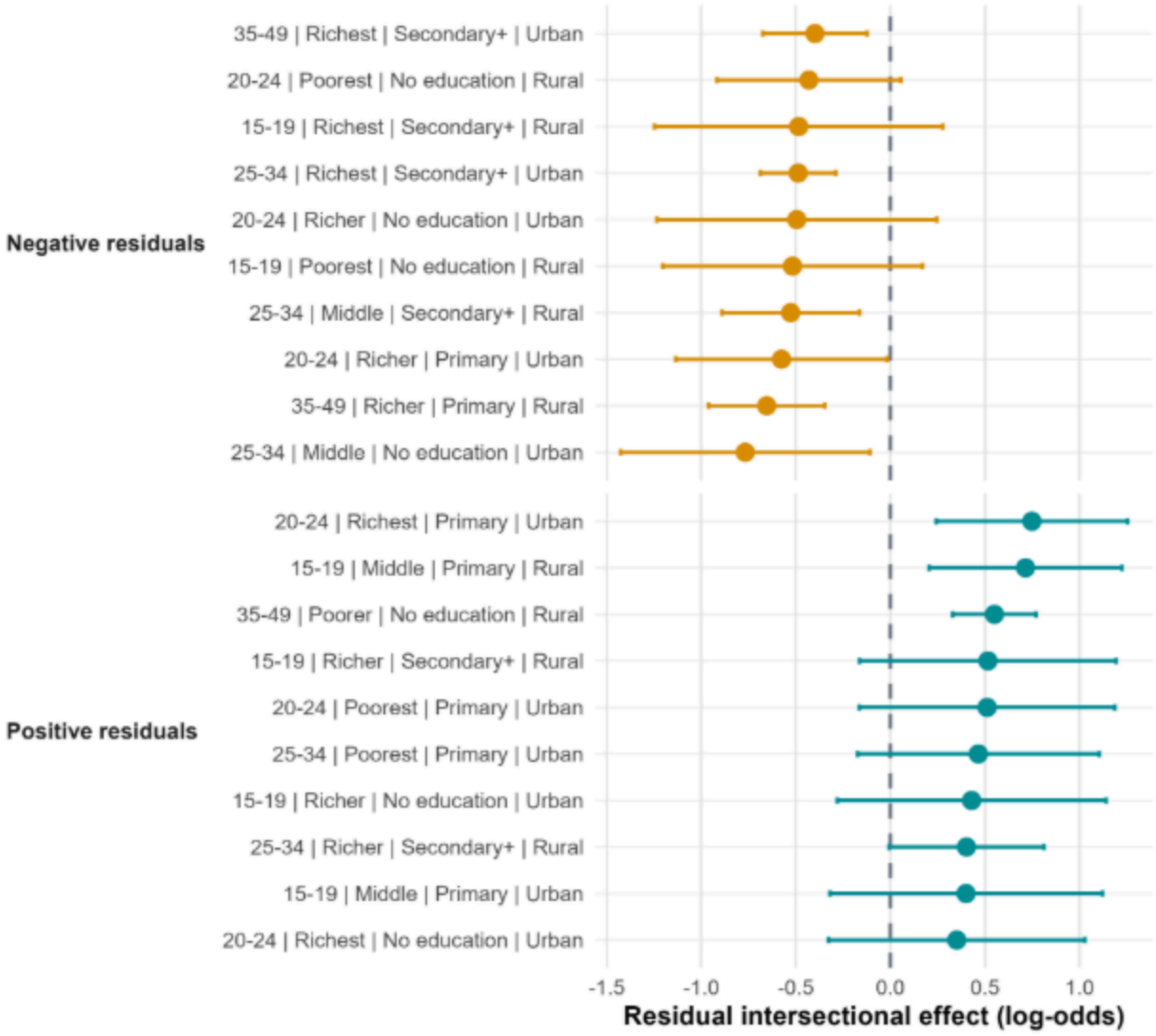


Figure 3

Largest positive and negative residual intersectional departures from the primary additive MAIHDA model. Points represent selected empirical-Bayes stratum-specific random-intercept estimates at the positive and negative extremes of the residual distribution; horizontal bars represent conditional 95% intervals. The dashed vertical line denotes zero residual effect. Positive values indicate higher and negative values lower mDFPS than expected from the additive combination of age, household wealth, education and residence. These residuals are model-based departures from additive expectations and should not be interpreted as causal interactions.

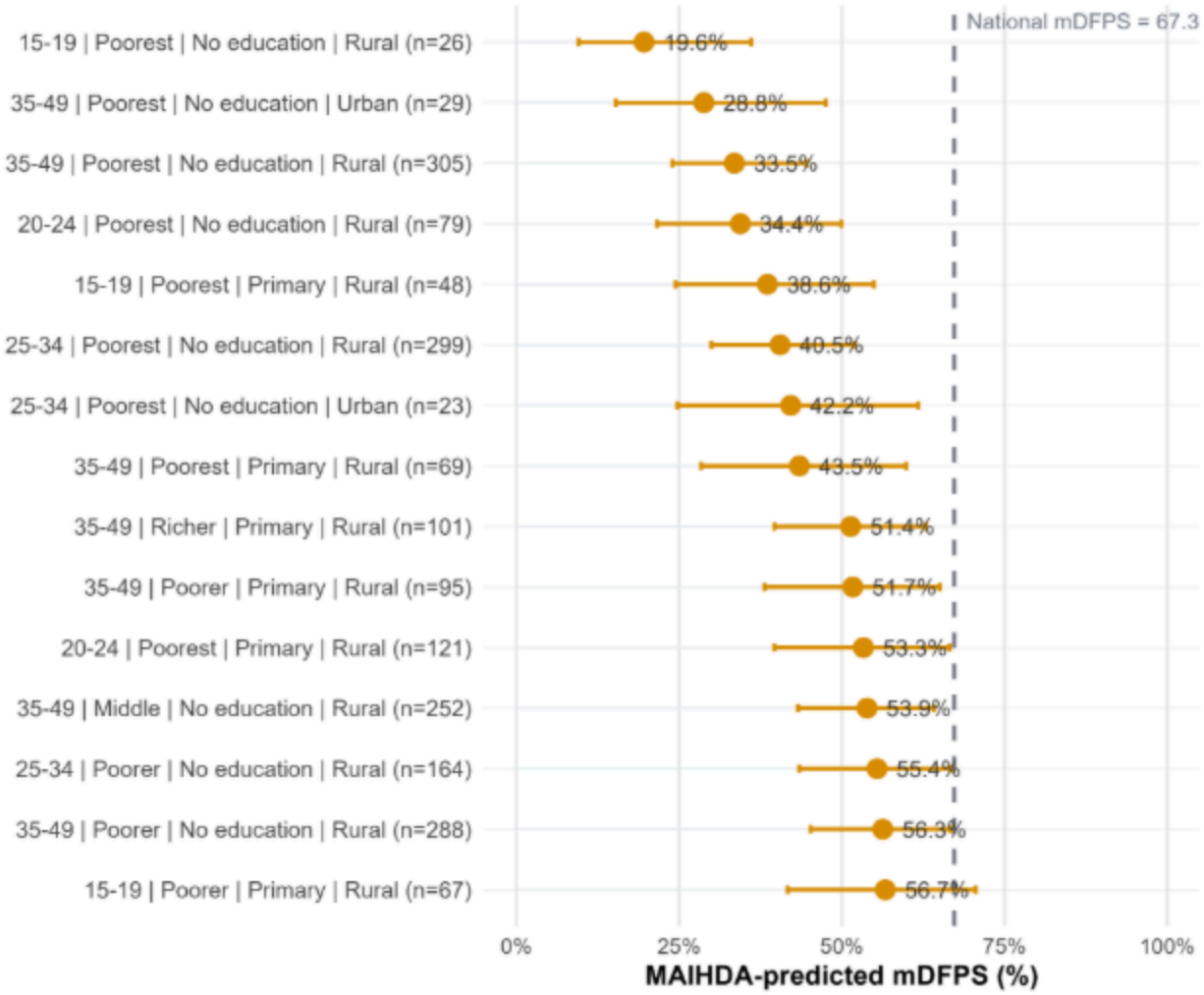


Figure 4

Intersectional strata with the lowest model-predicted mDFPS. Ranked predictions are shown for primary intersectional strata containing at least 20 women. Points represent MAIHDA-predicted probabilities and horizontal bars represent approximate model-based 95% intervals. The dashed vertical line indicates the national survey-weighted mDFPS estimate of 67.3%. All observed strata, including those containing fewer than 20 women, were retained in model estimation.

Supplementary Files

This is a list of supplementary files associated with this preprint. Click to download.

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