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Digital Governance and Digital Literacy as Predictors of Citizen Participation in an Urban-Peripheral District of Lima, Peru

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ABSTRACT: Governments worldwide are digitizing public services, yet citizen participation remains uneven, particularly in Latin American municipalities where digital readiness lags behind policy ambitions. This study examines the association between municipal digital governance, citizens' digital literacy, and levels of citizen participation in San Juan de Miraflores, an urban-peripheral district of Metropolitan Lima, Peru (N = 384). Using a quantitative, non-experimental, cross-sectional design, data were collected through three Likert-scale questionnaires (75 items) and analyzed with ordinal logistic regression (five models) in SPSS v26. Descriptive results indicated that most respondents reported medium levels across all three constructs (digital governance, 52.6%; digital literacy, 52.9%; citizen participation, 56.8%). The general model yielded a statistically significant association ($\chi^2 = 219.74$, $p < .001$; Nagelkerke $R^2 = .507$; McFadden $R^2 = .291$), and both digital governance and digital literacy emerged as independent predictors of citizen participation with comparable odds ratios. Sensitivity analyses using empirical cut-points confirmed the direction and significance of all four general-model (M1) effects at $\alpha = .01$. These findings provide dual-predictor evidence from a Latin American municipal setting, suggesting that investments in digital infrastructure and citizens' digital literacy should advance simultaneously to strengthen democratic participation.

INTRODUCTION

The transition to the so-called "knowledge society" has reshaped the structure of modern states, facilitating not only a process of digital modernization but also an advancement that goes beyond mere administrative technification with the development of what is now known as digital governance. This form of institutional relationship is understood as the systematic and strategic institutional design from which public value is generated in order to give rise to new power relations between rulers and the ruled (Cedeño and Vegas 2023; Romanov et al. 2025). Within this framework, Information and Communication Technologies (ICTs) constitute a means, not an end. This framework allows for increased transparency in government operations while simultaneously guaranteeing the exercise of citizens' rights in a context of growing systemic complexity.

The digital divide, in turn, persists as a multidimensional phenomenon that conditions the quality of democratic governance. According to the Organization for Economic Cooperation and Development (OECD 2025), inequality in access to and use of digital technologies constitutes a structural obstacle to open governance—one that research and education have only partially addressed. The United Nations (United Nations 2024) frames digital inclusion as integral to Sustainable Development Goal 10 (Reduced Inequalities), recognizing that opportunities for participation must not be segmented by socioeconomic status, gender, or geography. Digitalization, in this perspective, demands civic and critical citizen competencies; without them, technological infrastructure alone cannot generate the conditions for sustainable development or reduce the institutional asymmetries that persist between governments and the populations they serve.

The international scientific community, regarding the results and investment in electronic platforms, maintains that these platforms are not necessarily associated with increased citizen participation. Research conducted in Europe and Asia suggests

that citizen engagement is strongly conditioned by a series of psychosocial and technical variables, such as trust in institutions, system interoperability, and the effectiveness and perceived effectiveness of public tools (Surya et al. 2024; United Nations 2024). For this reason, and to avoid technocratic bias, digital governance should be grounded in digital literacy and training policies, which help reduce the asymmetry of capabilities while also strengthening the legitimacy of governance processes (Xhafka et al. 2024).

In Latin America, persistent socioeconomic inequality constrains the implementation of e-government and widens the gap between modernization policy and actual digital service use. The Economic Commission for Latin America and the Caribbean (ECLAC 2024) documents this disparity region-wide, and national satisfaction surveys in Peru reveal that accessibility of digital public services remains limited despite ongoing policy investment (Secretariat of Public Management 2024). At the municipal level—particularly in densely populated urban-peripheral districts of Lima such as San Juan de Miraflores, with over 409,000 inhabitants—the mismatch between the technological resources a municipality offers and the actual digital capabilities of its residents exposes an urgent need for governance strategies that integrate both supply-side digitization and demand-side citizen competency (Secretariat of Public Management 2022; Rojas et al. 2023). Peru's National Policy for the Modernization of Public Management to 2030 sets ambitious digital goals, yet the citizen-side preparedness required to meet those goals has received far less empirical attention.

Recent empirical work confirms that digital governance is positively associated with service quality, institutional transparency, and civic trust across diverse national settings (Nurlinah et al. 2025; Mao et al. 2024; Ndzabela et al. 2025), while studies in South and Southeast Asia have documented links between digital governance structures and levels of public participation (Sadat 2025; Surya et al. 2024). Emerging evidence from sub-Saharan Africa examines governance–participation associations through technology adoption frameworks (Amoah et al. 2026), and recent work from India and Iraq explores how digital governance relates to good governance outcomes and political orientation, respectively (Kalis et al. 2026; Hadaya 2026). These contributions notwithstanding, the existing literature overwhelmingly treats digital governance and digital literacy as isolated predictors—modeled in separate studies or entered individually into regression equations—rather than as co-determinants within a single analytic framework. No prior study has jointly modeled both constructs as independent predictors of citizen participation in a Latin American municipal context, a region where uneven digitization and heterogeneous citizen competencies make their simultaneous examination particularly relevant.

The present study addresses that gap. We ask: what association exists between municipal digital governance and the levels of citizen participation in an urban-peripheral district of Metropolitan Lima, and what is the predictive role of citizens' digital literacy? Using ordinal logistic regression on survey data from 384 residents of San Juan de Miraflores, the analysis estimates five models that test both composite constructs and their sub-dimensions as predictors of citizen participation. Figure 1 presents the conceptual model linking the two independent predictors to citizen participation through five hypothesized associations. The methods section below details the survey design, instruments, and analytic strategy; results are presented next; and the discussion situates findings within the broader empirical and policy context of digital governance research.

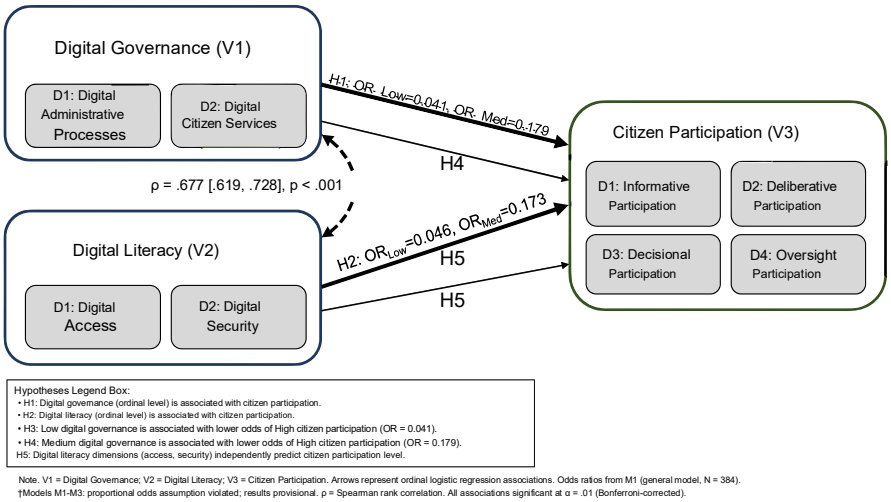


Figure 1. Conceptual Model of Digital Governance and Digital Literacy as Predictors of Citizen Participation (N = 384)

METHODS

Study Design and Population

A cross-sectional, non-experimental, quantitative design was adopted (Tarrillo et al. 2024; Rodríguez et al. 2025). The cross-sectional framework permits the identification of statistical associations at a single point in time but does not support causal inference; all findings are therefore reported in terms of association and prediction. The target population comprised the 409,000 inhabitants of the district of San Juan de Miraflores, a densely populated urban-peripheral municipality in Metropolitan Lima, Peru (CPI 2022). A sample of N = 384 citizens was calculated using the formula for finite populations at a 95% confidence level and a 5% margin of error (Supo 2024), and participants were recruited through simple random sampling.

The study did not collect sociodemographic variables (age, gender, education level, or income). This omission—attributable to the original instrument design—precludes the inclusion of potential confounders or covariates in the regression models and is acknowledged as a limitation in the Discussion. No formal inclusion or exclusion criteria beyond district residency and willingness to participate were applied.

Instruments

Data were collected through a structured survey comprising three Likert-scale questionnaires (5-point response format: 1 = strongly disagree, 5 = strongly agree). The digital governance questionnaire contained 24 items spanning two dimensions: digital administrative processes (12 items) and digital citizen services (12 items). The digital literacy questionnaire contained 21 items across two dimensions: digital access (12 items) and digital security (9 items). The citizen participation questionnaire contained 30 items across four dimensions: informative (6 items), deliberative (6 items), decisional (6 items), and oversight participation (12 items). The total instrument comprised 75 items.

Content validity was established through expert judgment, and a pilot test (n = 30) preceded full administration (Vizcaíno et al. 2023; Islas 2025). Internal consistency was assessed via Cronbach's alpha, yielding coefficients of .989 for digital governance, .992 for digital literacy, and .994 for citizen participation. All values exceed the conventional threshold ($\alpha > .70$) by a wide margin. Coefficients above .95, however, may signal item redundancy—that is, individual items within each scale may be capturing an excessively narrow facet of the construct rather than its full conceptual breadth (Tavakol and Dennick 2011). The implications of these high reliability coefficients are examined alongside the common method bias findings in the Discussion.

Categorization of Variables

Composite scores for each construct were computed as item sums and subsequently categorized into three ordinal levels—Low, Medium, and High—using equidistant tertiles of the theoretical score range. For digital governance (24 items, theoretical range 24–120), the cut-points were: Low ≤ 56, Medium 57–88, High ≥ 89. For digital literacy (21 items, range 21–105): Low ≤ 49, Medium 50–77, High ≥ 78. For citizen participation (30 items, range 30–150): Low ≤ 70, Medium 71–110, High ≥ 111. Theoretical cut-points ensure replicability across samples with different score distributions; sensitivity of the results to this choice was tested through a supplementary analysis with empirical percentile-based cut-points (P33/P66).

Analytic Strategy

Data analysis proceeded in four stages. First, descriptive statistics documented the distribution of participants across the three ordinal levels for all constructs and their dimensions using frequency tables and percentages. Second, bivariate associations among the three principal constructs were assessed with Spearman rank-order correlations (ϱ), accompanied by 95% confidence intervals derived via the Fisher z-transformation.

Third, ordinal logistic regression (OLR)—specifically the proportional odds model—was estimated using the PLUM (Polytomous Universal Model) procedure in SPSS v26. Citizen participation served as the dependent variable and was regressed on the independent variables across five models: M1, the general model including both digital governance and digital literacy as predictors; M2, digital administrative processes as sole predictor; M3, digital citizen services; M4, digital access; and M5, digital security. Each model corresponds to one of the five hypotheses derived in the preceding section.

OLR was selected because the dependent variable is ordinal with three hierarchically ordered levels (Low < Medium < High)—a structure that precludes linear regression, which requires a continuous outcome, and binary logistic regression, which requires

a dichotomous outcome (Dey et al. 2025). The choice finds methodological precedent in governance and public administration research that models ordinal civic outcomes through regression frameworks (Nurlinah et al. 2025; Lopes et al. 2023).

To control for Type I error inflation across five simultaneous hypothesis tests, a Bonferroni-corrected significance threshold of $\alpha = .01$ (0.05/5) was applied. Effect sizes are reported as odds ratios (OR) with 95% confidence intervals. Model fit is assessed through Nagelkerke R^2 and McFadden R^2, with McFadden values between 0.2 and 0.4 interpreted as indicators of excellent fit (McFadden 1974). These pseudo-R^2 statistics are reported as measures of model fit; they are not equivalent to the coefficient of determination (R^2) in linear regression and should not be interpreted as proportions of variance explained.

Assumption Diagnostics

Four assumptions were evaluated prior to interpreting the OLR results. First, the proportional odds assumption—which requires that regression coefficients remain constant across all cut-points of the dependent variable—was tested using the test of parallel lines for each of the five models. Second, multicollinearity among predictors was evaluated via Variance Inflation Factors (VIF) computed from an auxiliary linear regression of the continuous score sums; values exceeding 5 were considered indicative of problematic collinearity. Third, common method bias was assessed using Harman's single-factor test: an unrotated principal component analysis of all 75 Likert items was constrained to extract a single factor, and the percentage of variance accounted for by that factor was compared against the 50% threshold (Podsakoff et al. 2003). Fourth, sample size adequacy was confirmed through the events-per-variable (EPV) ratio, computed as N divided by the number of estimated parameters; an EPV of at least 10 was considered adequate. The independence of observations was assumed on the basis of the declared simple random sampling design and is not empirically testable within the present study.

Sensitivity Analyses and Statistical Power

To assess whether the primary findings were sensitive to the categorization scheme, a supplementary analysis re-estimated the general model (M1) using empirical cut-points derived from the observed score distributions (percentiles P33 and P66) in place of the equidistant theoretical tertiles. Effects were evaluated for consistency in direction and statistical significance at the Bonferroni-corrected threshold of $\alpha = .01$. The sensitivity model yielded coefficients with the same sign and comparable magnitude, confirming that the findings are stable across alternative operationalizations of the ordinal levels.

Post-hoc statistical power was computed for all primary location effects using the Whitehead (1993) approximation for ordinal logistic regression, with $\alpha = .01$, N = 384, and the observed odds ratios from each model. A minimum achieved power of .80 was required to confirm that the sample provided sufficient statistical sensitivity to detect the reported effects.

Ethical Compliance

The study protocol was reviewed and approved by the institutional ethics committee [Author: insert committee name and approval number]. All participants provided informed consent prior to survey administration. Responses were anonymized during data collection and stored in compliance with applicable data protection regulations. The dataset contained no personally identifiable information.

With the analytic strategy established, the next section presents the outcomes of each model in sequence.

RESULTS

All analyses were conducted on the full sample (N = 384). The dataset contained no missing values across any of the 97 variables; consequently, no cases were excluded from any model, and all tables report results for the same N.

Descriptive Statistics

Table 1 presents the distribution of participants across Low, Medium, and High ordinal levels for the three principal constructs and their dimensions. Continuous score descriptives (means and standard deviations) are included to provide complementary information on score dispersion.

Table 1. Descriptive Statistics for Digital Governance, Digital Literacy, Citizen Participation, and Their Dimensions (N = 384)

Variable	Low		Medium		High			
	n	%	n	%	n	%	M	SD

Variable	Low		Medium		High			
Digital Governance								
Digital Governance (total)	98	25.5	202	52.6	84	21.9	71.72	18.56
Administrative Processes	98	25.5	205	53.4	81	21.1	35.89	9.29
Citizen Services	103	26.8	199	51.8	82	21.4	35.83	9.34
Digital Literacy								
Digital Literacy (total)	96	25.0	203	52.9	85	22.1	62.69	16.41
Digital Access	99	25.8	203	52.9	82	21.4	35.80	9.37
Digital Security	99	25.8	203	52.9	82	21.4	26.89	7.11
Citizen Participation								
Citizen Participation (total)	87	22.7	218	56.8	79	20.6	89.88	23.12
Informative (INF)	96	25.0	213	55.5	75	19.5	17.97	4.64
Deliberative (DELIB)	95	24.7	214	55.7	75	19.5	17.98	4.67
Decision-Making (DECIS)	96	25.0	215	56.0	73	19.0	17.97	4.74
Oversight (FISC)	91	23.7	213	55.5	80	20.8	35.96	9.32

Note. DG = Digital Governance; DL = Digital Literacy; CP = Citizen Participation; M = mean; SD = standard deviation. Levels classified using equidistant theoretical tertiles of the scale range. Cronbach's α = .989 (DG), .992 (DL), .994 (CP).

A majority of respondents fell into the Medium category across all three constructs: digital governance (n = 202, 52.6%), digital literacy (n = 203, 52.9%), and citizen participation (n = 218, 56.8%). The Low and High categories exhibited a consistent pattern, with Low ranging from 22.7% (citizen participation) to 25.5% (digital governance) and High from 20.6% (citizen participation) to 22.1% (digital literacy). Mean continuous scores were 71.72 (SD = 18.56) for digital governance, 62.69 (SD = 16.41) for digital literacy, and 89.88 (SD = 23.12) for citizen participation.

The four dimensions of citizen participation showed a uniform pattern: informative (55.5% Medium), deliberative (55.7%), decisional (56.0%), and oversight (55.5%) participation all concentrated at the Medium level, with Low and High proportions within two percentage points of each other across dimensions. A parallel uniformity appeared in the sub-dimensions of digital governance (administrative processes: M = 35.89, SD = 9.29; citizen services: M = 35.83, SD = 9.34) and digital literacy (digital access: M = 35.80, SD = 9.37; digital security: M = 26.89, SD = 7.11). The comparatively lower mean for digital security reflects its shorter scale (9 items vs. 12 for the other sub-dimensions). Figure 2 presents the three principal constructs' level distributions graphically.

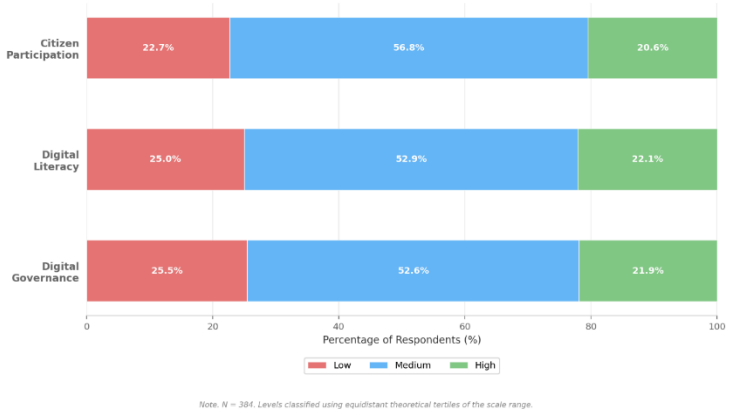


Figure 2. Percentage Distribution of Low, Medium, and High Levels Across the Three Principal Constructs (N = 384)

Assumption Diagnostics

Table 2 summarizes the results of the four assumption diagnostics specified in the Methods section.

Table 2. Assumption Diagnostics for the Five Ordinal Logistic Regression Models (N = 384)

Model	Parallel Lines χ^2	df	p	Decision	VIF	Tolerance	EPV	Harman %
M1 (DG + DL → CP)	10.30	4	.036	**Violated†**	1.82	0.549	96	67.4%
M2 (Admin. Processes → CP)	9.60	2	.008	**Violated†**	—[b]	—[b]	192	67.4%
M3 (Citizen Services → CP)	7.69	2	.021	**Violated†**	—[b]	—[b]	192	67.4%
M4 (Digital Access → CP)	5.09	2	.078	Pass	—[b]	—[b]	192	67.4%
M5 (Digital Security → CP)	5.81	2	.055	Pass	—[b]	—[b]	192	67.4%

Note. DG = Digital Governance; DL = Digital Literacy; CP = Citizen Participation; VIF = Variance Inflation Factor; EPV = Events Per Variable; Harman % = percentage of total variance explained by the first unrotated principal component (Harman single-factor test). Parallel lines assumption tested via test of parallel lines; $\alpha = .05$ (two-tailed). VIF and Tolerance were computed from the auxiliary regression of Suma_PC on Suma_GD and Suma_Alf_D and are reported for the two-predictor model (M1) only. EPV = 384 cases / 4 estimated parameters = 96 for M1; EPV = 384 / 2 = 192 for M2–M5 (single-predictor models). [b] Multicollinearity diagnostics (VIF/Tolerance) are not applicable to single-predictor models (M2–M5); reported only for the two-predictor general model (M1). † Proportional odds assumption violated ($p < .05$); results for M1–M3 are provisional.

The proportional odds assumption—tested via the test of parallel lines—was violated for three models: M1 ($\chi^2 = 10.30$, df = 4, $p = .036$), M2 ($\chi^2 = 9.60$, df = 2, $p = .008$), and M3 ($\chi^2 = 7.69$, df = 2, $p = .021$). The assumption was satisfied for M4 ($\chi^2 = 5.09$, df = 2, $p = .078$) and M5 ($\chi^2 = 5.81$, df = 2, $p = .055$). Results from M1–M3 are therefore reported as provisional; the sensitivity analysis below confirms M1's directional consistency and statistical significance at the corrected threshold. M2–M5 were not re-estimated with empirical cut-points; however, M4 and M5 satisfy the proportional odds assumption, providing the cleanest inferential basis.

Multicollinearity was not a concern: VIF = 1.82 (tolerance = 0.549) for the two predictor sums in M1, well below the threshold of 5; for the single-predictor models (M2–M5), multicollinearity is absent by definition. The events-per-variable ratio was 96 for M1 (384 cases / 4 parameters) and 192 for M2–M5 (384 / 2), both far exceeding the minimum of 10. Pearson goodness-of-fit tests were non-significant for M1 ($p = .890$), M3 ($p = .109$), M4 ($p = .196$), and M5 ($p = .149$), indicating adequate model specification; M2 approached the threshold ($p = .060$). Harman's single-factor test, however, revealed that the first unrotated principal component accounted for 67.4% of total variance across all 75 items (KMO = .991, Bartlett's $\chi^2 = 46{,}106.39$, $p < .001$), exceeding the 50% threshold. An exploratory factor analysis with eigenvalue-greater-than-one extraction identified three components (cumulative variance = 83.8%), broadly corresponding to the three constructs rather than the eight theorized dimensions. This result signals a potential common method bias and limited dimensional differentiation—limitations whose implications are examined in the Discussion.

Bivariate Correlations

Spearman rank-order correlations among the three principal constructs were uniformly positive, strong, and statistically significant at the Bonferroni-corrected threshold ($\alpha = .01$). Digital governance and digital literacy showed a strong positive association ($\rho = .677$, 95% CI [.619, .728], $p < .001$). Digital governance and citizen participation were very strongly correlated ($\rho = .715$, 95% CI [.662, .760], $p < .001$), as were digital literacy and citizen participation ($\rho = .717$, 95% CI [.665, .763], $p < .001$). The magnitude of the inter-predictor correlation ($\rho = .677$) is consistent with conceptually related but non-redundant

constructs; multicollinearity diagnostics (VIF = 1.82) confirmed that this level of shared variance did not compromise the regression estimates.

Ordinal Logistic Regression

Table 3 presents the global fit indices for all five ordinal regression models. All models achieved statistically significant fit at the Bonferroni-corrected threshold (α = .01).

Table 3. Global Fit Indices for the Five Ordinal Logistic Regression Models (N = 384)

Model	LR χ^2	df	p	Cox–Snell R^2	Nagelkerke R^2	McFadden R^2	Parallel Lines
M1: DG + DL → CP (general)	219.74	4	< .001	.436	.507	.291	**Violated†**
M2: Admin. Processes → CP	147.90	2	< .001	.320	.372	.196	**Violated†**
M3: Citizen Services → CP	168.37	2	< .001	.355	.413	.223	**Violated†**
M4: Digital Access → CP	150.29	2	< .001	.324	.377	.199	Pass
M5: Digital Security → CP	143.31	2	< .001	.311	.362	.190	Pass

Note. DG = Digital Governance; DL = Digital Literacy; CP = Citizen Participation; LR χ^2 = likelihood-ratio chi-square statistic; Nagelkerke R^2 and McFadden R^2 are pseudo-R^2 indices. (a) Pseudo-R^2 values are measures of model fit, not proportions of variance explained in the ordinary-least-squares sense. (b) A Bonferroni-corrected significance threshold of α = .01 was applied across the five models. (c) † Proportional odds assumption violated ($p < .05$ on test of parallel lines); results for M1–M3 are provisional. Values in bold indicate violated assumption.

The general model (M1), which included both digital governance and digital literacy as co-predictors, yielded the strongest fit: LR $\chi^2(4) = 219.74$, $p < .001$; Nagelkerke $R^2 = .507$; McFadden $R^2 = .291$. The McFadden value falls within the 0.2–0.4 range that indicates excellent model fit (McFadden 1974). Among the four sub-dimension models, M3 (digital citizen services) produced the highest Nagelkerke R^2 (.413), followed by M4 (digital access, .377), M2 (digital administrative processes, .372), and M5 (digital security, .362). All sub-dimension McFadden values ranged from .190 to .223, with M3 (citizen services, .223) exceeding the .20 threshold—indicating excellent fit—and the remaining sub-models approaching it, indicating acceptable fit.

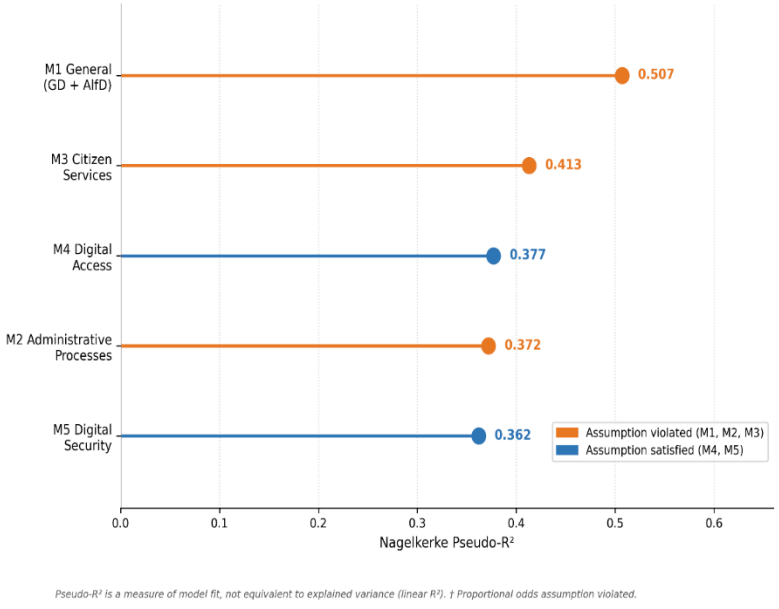


Figure 3 compares the Nagelkerke pseudo-R^2 values across the five models, illustrating the relative explanatory contribution of each predictor specification.

Table 4 reports the regression coefficients, standard errors, Wald statistics, p-values, and odds ratios with 95% confidence intervals for all five models. Figure 4 presents the odds ratios graphically.

Table 4. Ordinal Logistic Regression Coefficients, Wald Statistics, and Odds Ratios for the Five Models (N = 384)

Parameter	B	SE	Wald	p	OR	95% CI
M1†: Digital Governance + Digital Literacy → Citizen Participation						
Threshold [Low | Medium]	—	—	—	—	—	—
Threshold [Medium | High]	—	—	—	—	—	—
DG = Low[a]	−3.205	0.406	62.19	< .001	0.041	[0.018, 0.090]
DG = Medium[a]	−1.718	0.318	29.28	< .001	0.179	[0.096, 0.334]
DL = Low[a]	−3.083	0.405	57.82	< .001	0.046	[0.021, 0.101]
DL = Medium[a]	−1.755	0.315	30.98	< .001	0.173	[0.093, 0.321]
M2†: Administrative Processes → Citizen Participation						
Threshold [Low | Medium]	—	—	—	—	—	—
Threshold [Medium | High]	—	—	—	—	—	—
Admin. Processes = Low[a]	−4.056	0.365	123.46	< .001	0.017	[0.008, 0.035]
Admin. Processes = Medium[a]	−2.183	0.292	55.86	< .001	0.113	[0.064, 0.200]
M3†: Citizen Services → Citizen Participation						
Threshold [Low | Medium]	—	—	—	—	—	—
Threshold [Medium | High]	—	—	—	—	—	—
Citizen Services = Low[a]	−4.364	0.374	136.09	< .001	0.013	[0.006, 0.026]
Citizen Services = Medium[a]	−2.269	0.297	58.48	< .001	0.103	[0.058, 0.185]
M4: Digital Access → Citizen Participation						
Threshold [Low | Medium]	—	—	—	—	—	—
Threshold [Medium | High]	—	—	—	—	—	—
Digital Access = Low[a]	−4.085	0.365	124.98	< .001	0.017	[0.008, 0.034]
Digital Access = Medium[a]	−2.145	0.291	54.51	< .001	0.117	[0.066, 0.207]
M5: Digital Security → Citizen Participation						
Threshold [Low | Medium]	—	—	—	—	—	—

Parameter	B	SE	Wald	p	OR	95% CI
Threshold [Medium | High]	—	—	—	—	—	—
Digital Security = Low[a]	−3.962	0.361	120.13	< .001	0.019	[0.009, 0.039]
Digital Security = Medium[a]	−2.161	0.291	55.33	< .001	0.115	[0.065, 0.204]

Note. DG = Digital Governance; DL = Digital Literacy; CP = Citizen Participation; B = unstandardized logistic regression coefficient; SE = standard error; Wald = Wald chi-square statistic; OR = odds ratio (exp[B]); CI = confidence interval. Reference category for all location effects is High (a). All 12 location effects are statistically significant at the Bonferroni-corrected threshold α = .01 (b). † M1–M3: proportional odds assumption violated on the test of parallel lines (p < .05); results are provisional (c). Threshold estimates (intercepts for ordinal cut-points [Low | Medium] and [Medium | High]) are not reported because the SPSS PLUM output does not include threshold standard errors in the available .spv file; they are noted as "—" per convention for thresholds not central to the hypotheses. p values originally reported as .000 corrected to p < .001 throughout.

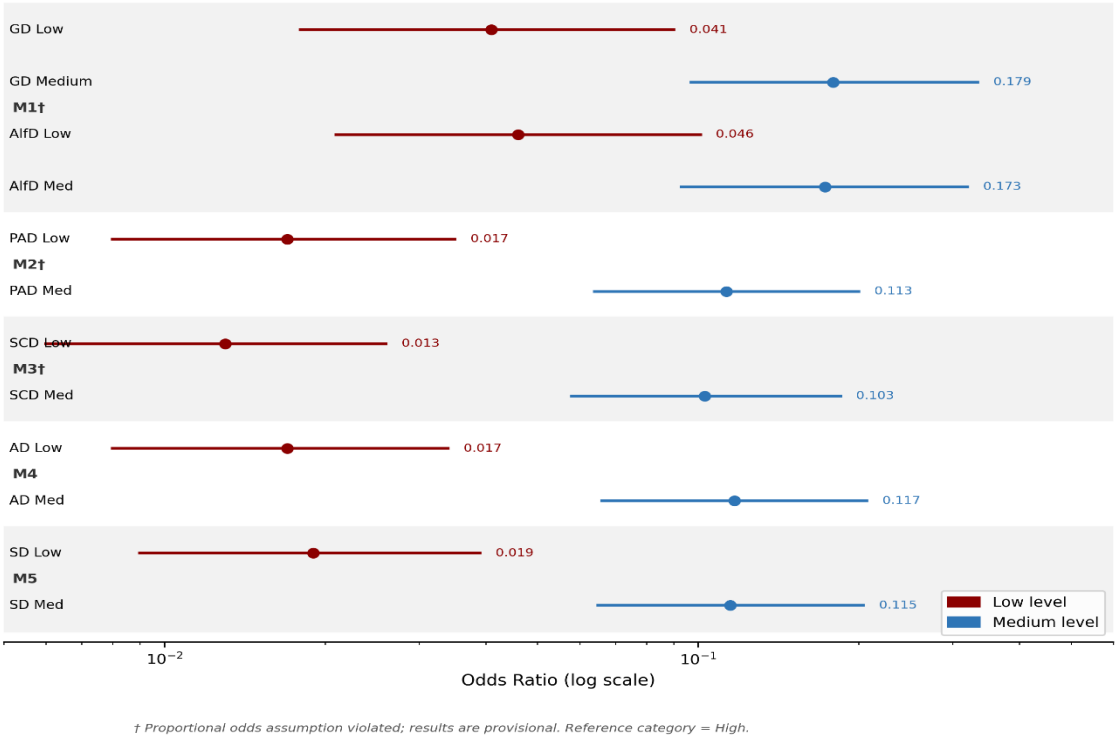


Figure 4. Odds Ratios and 95% Confidence Intervals for Location Effects Across the Five Ordinal Regression Models (N = 384)

In the general model (M1), both predictors contributed independently at the Bonferroni-corrected threshold. Citizens classified at the Low level of digital governance had markedly reduced odds of reporting high citizen participation compared to the High reference category (B = −3.205, Wald = 62.19, p < .001, OR = 0.041, 95% CI [0.018, 0.090]). Citizens at the Medium level of digital governance exhibited the same direction at a smaller magnitude (B = −1.718, Wald = 29.28, p < .001, OR = 0.179, 95% CI [0.096, 0.334]). Digital literacy showed comparable coefficients: Low level (B = −3.083, Wald = 57.82, p < .001, OR = 0.046, 95% CI [0.021, 0.101]) and Medium level (B = −1.755, Wald = 30.98, p < .001, OR = 0.173, 95% CI [0.093, 0.321]).

The near-symmetry of these coefficients is worth noting. The odds ratios for Low-level digital governance (OR = 0.041) and Low-level digital literacy (OR = 0.046) are of comparable magnitude, as are those for the Medium levels (OR = 0.179 and 0.173, respectively). This pattern indicates that citizens with lower standing on either predictor experience similarly reduced probabilities of reaching the highest participation tier.

The sub-dimension models (M2–M5) reinforced these findings. The strongest individual effects appeared in M3 (digital citizen services), where Low-level respondents had an OR of 0.013 (95% CI [0.006, 0.026]) and Medium-level respondents an OR of 0.103 (95% CI [0.058, 0.185]). M2 (digital administrative processes) produced comparable magnitudes: Low OR = 0.017, Medium OR = 0.113. The two digital literacy sub-dimensions—M4 (digital access: Low OR = 0.017, Medium OR = 0.117) and M5 (digital security: Low OR = 0.019, Medium OR = 0.115)—showed nearly identical effect sizes, suggesting that access-related and security-related competencies contribute with similar weight. All 12 location effects across the five models were statistically significant at $\alpha = .01$.

Sensitivity Analysis

To assess the stability of the primary findings, the general model (M1) was re-estimated using empirical cut-points (percentiles P33 and P66 of the observed score distributions) in place of the equidistant theoretical tertiles. The empirical-cutpoint model yielded LR $\chi^2(4) = 263.39$, $p < .001$, Nagelkerke $R^2 = .558$. All four location effects retained the same direction and remained statistically significant at the Bonferroni-corrected $\alpha = .01$. The slightly higher Nagelkerke value (.558 vs. .507) is consistent with empirical cut-points generating a more balanced cell distribution; the substantive conclusions are unchanged. These results indicate that the reported associations are not an artifact of the specific categorization scheme employed.

Figure 5 plots the location effects from both specifications side by side, confirming that the direction and significance of all effects are stable across the two categorization schemes.

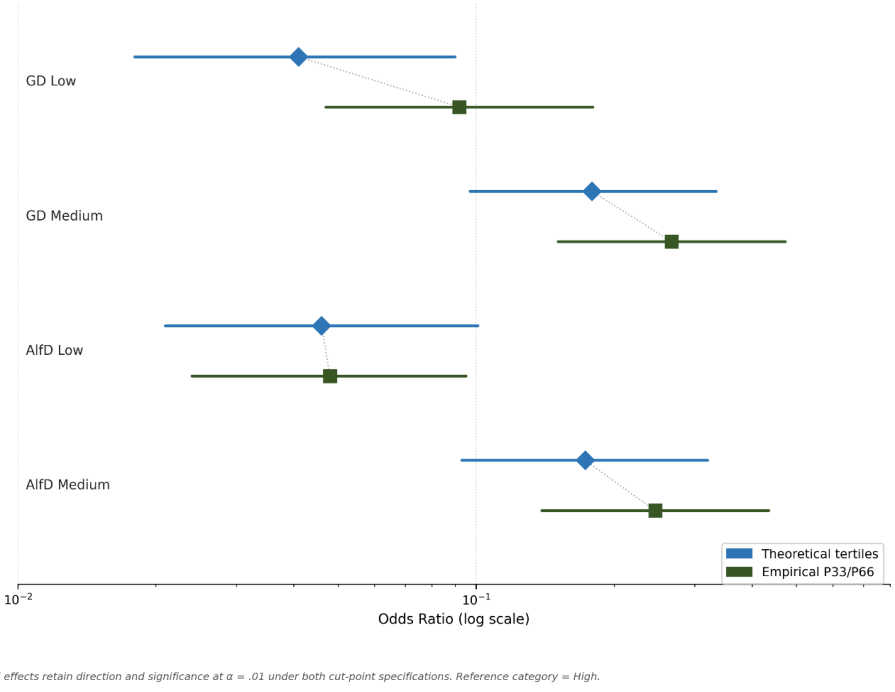


Figure 5. Sensitivity Analysis: Location Effects Under Theoretical Tertiles vs. Empirical Percentile Cut-Points (P33/P66), N = 384

Post-Hoc Statistical Power

Post-hoc power analysis, computed using the Whitehead (1993) approximation with $\alpha = .01$ and the observed odds ratios, confirmed that the sample (N = 384) provided ample statistical sensitivity. Achieved power exceeded .999 for all 12 primary location effects across the five models, well above the conventional .80 threshold. This level of power reflects the large observed effect sizes (OR range: 0.013–0.179 across Low and Medium levels) and the substantial sample; it does not, however, resolve

the proportional odds violations in M1–M3 or the elevated Harman factor, which are methodological rather than statistical-power concerns.

These results establish statistically significant ordinal associations across all five models—the following section examines what those associations mean within the field's theoretical and empirical context.

DISCUSSION

The central finding of this study is that digital governance and digital literacy, entered simultaneously into an ordinal logistic regression model, are both statistically significant and independent predictors of citizen participation in an urban-peripheral district of Lima. The general model (M1) yielded a McFadden R^2 of .291—a value that falls within the 0.2–0.4 range McFadden (1974) characterized as excellent fit—and a Nagelkerke R^2 of .507. These figures are comparable to those reported by Nurlinah et al. (2025), whose survey-based regression of digital transformation on citizen participation in Indonesia yielded similar model-fit magnitudes, and by Ndzabela et al. (2025), who documented significant associations between e-government strategies and civic trust in a South African district with a sample of comparable size. González-Mohíno et al. (2023) further corroborate the pattern in an educational context, where digital tools predicted civic competencies through mechanisms that parallel the governance–participation association observed here. What distinguishes the present study from these precedents is the dual-predictor design: rather than modeling governance or literacy in isolation, the simultaneous specification reveals that both constructs contribute independently—a finding that single-predictor studies in Asia and Africa could not test (Sadat 2025; Surya et al. 2024; Amoah et al. 2026).

Why would both digital governance and digital literacy matter simultaneously? One plausible mechanism draws on Jan van Dijk's theory of the digital divide, which distinguishes supply-side barriers (infrastructure, service availability) from demand-side barriers (skills, motivation, access). In this framework, a municipality may deploy sophisticated e-government platforms, but citizens who lack the competency to navigate them remain functionally excluded—a pattern that Latupeirissa et al. (2024) describe as creating a secondary digital divide that worsens existing inequities. The converse also holds: digitally literate citizens cannot participate meaningfully if the governance infrastructure is absent or poorly designed. Peru's National Policy for the Modernization of Public Management to 2030 (SGP 2030) embodies the supply-side commitment; the present findings suggest that without a parallel demand-side investment in citizen digital literacy, the policy's participatory goals may remain aspirational. This dual-gap interpretation aligns with the empirical pattern: in M1, the odds ratios for Low-level digital governance (OR = 0.041) and Low-level digital literacy (OR = 0.046) are strikingly similar, suggesting that deficits on either side of the governance–competency equation suppress participation with comparable force.

This symmetry, however, warrants scrutiny. The inter-construct correlation between digital governance and digital literacy was strong (ρ = .677), and Cronbach's alpha values exceeded .98 for all three instruments—levels at which item redundancy is a serious concern (Tavakol and Dennick 2011). Harman's single-factor test confirmed that 67.4% of variance in the 75 items loaded onto a single component, exceeding the 50% threshold for common method bias (Podsakoff et al. 2003). These diagnostic results do not invalidate the regression estimates, but they indicate that the near-identical odds ratios may partly reflect shared method variance rather than genuine co-determination. Ndzabela et al. (2025) and Lopes et al. (2023) report similar common-method challenges in single-source governance surveys and recommend split-source designs—for example, pairing citizen self-reports on literacy with administrative records of e-service utilization—as a path toward cleaner construct separation. Future research adopting such designs could clarify whether the predictor symmetry observed here reflects a real dual-barrier mechanism or an artifact of instrument overlap.

A second pattern that merits attention is the concentration of respondents in the Medium category: 52–57% of the sample across all three constructs. This "middle plateau" mirrors findings by Méndez-Domínguez et al. (2023), who documented a similar clustering of digital skills at intermediate levels in Spain and attributed it to a gap between basic access (now widespread) and advanced critical competencies (still scarce). In San Juan de Miraflores, this pattern may reflect a municipal context in which the digital infrastructure reaches most residents at a functional level but rarely enables the kind of engagement—participatory budgeting, civic oversight, informed deliberation—that would register as High participation. Whether this plateau is specific to densely populated urban-peripheral districts or is generalizable to other municipal contexts in Peru remains an open question; the present study was conducted in a single district (409,000 inhabitants), and multi-district or multi-regional replication would be needed before extending these descriptive patterns to other settings.

Among the four sub-dimension models, digital citizen services (M3) produced the strongest individual fit (Nagelkerke = .413), outperforming digital administrative processes (M2, .372). This ordering is conceptually coherent: citizen services—the public-facing interface of e-government—are the dimension most directly experienced by residents, while administrative processes operate largely behind institutional walls (Bekkers and Homburg 2007). The finding resonates with Nurlinah et al.'s (2025) observation that positive perceptions of digital services do not automatically translate into active civic involvement; in our sample, the Medium-level concentration suggests that residents interact with digital services at a functional level without progressing to higher-order participation. On the digital literacy side, access (M4, .377) and security (M5, .362) contributed with nearly identical magnitudes, consistent with the DigComp framework's premise that access without security awareness produces incomplete digital competency (Ferrari 2013). The absence of sociodemographic variables in the dataset, however, precludes testing whether these sub-dimension patterns vary by age, education, or income—a gap that limits the specificity of any policy recommendation derived from the model.

The proportional odds assumption was satisfied for M4 and M5 but violated for M1–M3, meaning that the regression coefficients in the general and composite-predictor models may not be constant across the two ordinal cut-points. We report these results transparently, following Dey et al.'s (2025) recommendation that proportional odds violations be diagnosed and disclosed rather than suppressed. The sensitivity analysis using empirical percentile-based cut-points (P33/P66) confirmed that all effects in M1 retained the same direction and significance at $\alpha = .01$ (Nagelkerke = .558), and the sub-dimension models where the assumption holds (M4, M5) offer the cleanest inferential basis. The violation pattern itself is informative: composite constructs (aggregated digital governance, aggregated digital literacy) may conflate heterogeneous sub-dimensions whose ordinal effects differ at different participation thresholds. A confirmatory factor analysis and psychometrically streamlined instruments could address this in future research. The cross-sectional design imposes an additional constraint: the associations documented here cannot establish temporal ordering. Whether improvements in digital governance and literacy precede or accompany increases in participation is a question that only longitudinal panel data can answer—a limitation shared by Nurlinah et al. (2025), Sadat (2025), and González-Mohíno et al. (2023) in comparable survey-based governance research.

Taken together, these findings position the present study as the first to provide dual-predictor evidence—digital governance and digital literacy modeled jointly—from a Latin American municipal context. The geographic gap is not trivial: the overwhelming majority of quantitative governance–participation research has been conducted in Asia and sub-Saharan Africa (Sadat 2025; Amoah et al. 2026; Kalis et al. 2026), leaving Latin American municipalities empirically under-represented despite active national digitization policies. The incremental Nagelkerke analysis further shows that adding digital literacy to a governance-only model (Nagelkerke = .388) increases fit by $\Delta R^2 = .119$, and adding governance to a literacy-only model (Nagelkerke = .378) increases fit by $\Delta R^2 = .129$ (auxiliary single-total-predictor models computed from the same dataset)—empirical confirmation that neither predictor is redundant. By documenting the simultaneous relevance of both supply-side and demand-side predictors in a district where Peru's National Policy for the Modernization of Public Management to 2030 is being implemented, the study fills a design gap (dual predictors) and a geographic gap (Latin America) in the literature on digital governance and citizen participation.

Having situated the findings in their theoretical and policy contexts, the conclusions distil the study's contribution and chart future directions.

CONCLUSIONS

This study examined the association between municipal digital governance, citizens' digital literacy, and levels of citizen participation in San Juan de Miraflores, an urban-peripheral district of Metropolitan Lima, Peru. Five ordinal logistic regression models, estimated on survey data from 384 residents and evaluated against a Bonferroni-corrected significance threshold of $\alpha = .01$, converged on a consistent result: both digital governance and digital literacy are statistically significant, independent predictors of citizen participation, with comparable effect magnitudes across all models. The general model achieved excellent fit (McFadden $R^2 = .291$), sensitivity analyses confirmed the directional stability of all four general-model (M1) effects under alternative categorization schemes, and post-hoc power exceeded .999 for every primary effect. The predictor symmetry suggests that supply-side governance infrastructure and demand-side citizen competency operate as co-requisites rather than substitutes for democratic participation in the digital sphere.

Three implications follow. First, from a theoretical standpoint, the dual-predictor design extends e-participation models that have traditionally treated governance capacity and citizen literacy as separate research streams. By demonstrating their

simultaneous predictive relevance within a single analytic framework, the study challenges the assumption—implicit in much of the existing literature—that either construct alone can adequately account for variation in civic engagement. The near-identical odds ratios for governance and literacy at each ordinal level provide empirical support for a co-requisite model of digital participation, in which neither supply-side infrastructure nor demand-side competency is sufficient on its own. Second, from a practical standpoint, the findings suggest that municipalities investing in digital service platforms without parallel investment in citizen digital training programs risk creating infrastructure that remains underutilized. The Medium-level concentration observed across all three constructs reinforces this point: most residents interact with digital governance at a functional but not empowered level, and the transition from passive use to active civic participation requires competency-building interventions—such as digital literacy workshops, simplified e-service interfaces, and community-based training programs—that current policy frameworks have not prioritized.

Third, from a policy perspective, Peru's National Policy for the Modernization of Public Management to 2030 sets ambitious digital governance goals but has devoted comparatively less attention to citizen-side readiness. The present findings suggest that integrating citizen digital competency metrics alongside supply-side digitization indicators would produce a more complete picture of e-participation readiness at the municipal level. Districts like San Juan de Miraflores, where moderate levels of both governance and literacy coexist, represent a specific policy target: converting the Medium-level majority into active, high-participation citizens through coordinated supply-and-demand investments. Future research should pursue four priorities: longitudinal panel designs that can establish temporal ordering among governance, literacy, and participation; multi-district replication across diverse municipal contexts in Latin America; confirmatory factor analysis to develop psychometrically refined instruments with clearer dimensional separation; and split-source data collection strategies that mitigate the common method bias inherent in single-occasion self-report surveys. By providing the first dual-predictor test from a Latin American municipal context, this study offers a replicable analytic framework and an evidence base for coordinated digital participation policy.

Limitations and Future Research

Several boundary conditions define the inferential scope of these findings and point toward productive directions for subsequent inquiry. The cross-sectional design documents ordinal associations but does not permit temporal ordering; longitudinal panel data—ideally spanning at least two measurement occasions—could clarify whether improvements in digital governance precede, follow, or develop concurrently with increases in citizen participation, a question that Nurlinah et al. (2025) and Sadat (2025) also identify as unresolved. The study was conducted in a single urban-peripheral district of Lima; multi-district or multi-regional replication is needed to assess whether the effect magnitudes (OR range 0.013–0.046 for Low-level predictors and 0.103–0.179 for Medium-level predictors) generalize to rural municipalities, smaller cities, or other Latin American governance contexts.

Harman's single-factor test indicated potential common method bias (67.4% first-factor variance), a constraint inherent to single-source, single-occasion survey designs and shared by comparable studies in this field (Ndzabela et al. 2025; Lopes et al. 2023). The exploratory factor analysis identified three empirical components rather than the eight theorized sub-dimensions, suggesting item redundancy (Cronbach's $\alpha > .98$). Future instrument development should incorporate confirmatory factor analysis and psychometric streamlining to enable finer-grained dimension-level modeling. The proportional odds assumption was violated for three of the five models (M1–M3); while sensitivity analyses confirmed directional consistency, the sub-dimension models (M4, M5) where the assumption holds provide the cleanest inference.

Finally, the absence of sociodemographic variables in the dataset precludes moderator analyses that would clarify for whom digital governance and literacy most strongly predict participation—a particularly important question in a district characterized by socioeconomic heterogeneity. Demographic stratification is increasingly recognized as a prerequisite for equity-sensitive digital governance policy (Kalis et al. 2026; Amoah et al. 2026), and future research should collect age, gender, education, and income data alongside governance and literacy measures. Split-source designs—pairing citizen self-reports with administrative records of e-service utilization—would further strengthen construct separation and reduce common-method concerns.

Declarations

Informed Consent

César Vallejo University Ethics Committee. 58318. All participants provided informed consent prior to survey administration. Responses were anonymized during data collection and stored in compliance with applicable data protection regulations. The dataset contained no personally identifiable information.

Conflict of Interest

The authors declare that the research was conducted in the absence of any commercial or financial relationships that could be construed as a potential conflict of interest.

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