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Perceived School Support and Self-Reported Digital Civic Responsibility Among Vocational Health Professions Students in China: A Cross-Sectional Study

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Abstract

Background

Institutional conditions associated with responsible digital participation remain underexamined in health professions education. This study estimated the covariation between perceived school support and self-reported digital civic responsibility among vocational health professions students in China.

Methods

A cross-sectional electronic survey was administered at one vocational health professions college in Jiangsu Province in 2026. The primary exposure was a five-item perceived school support score; the outcome was a ten-item digital civic responsibility score. Standardized ordinary least-squares regression with HC3 robust standard errors adjusted for prespecified background and social-media-use characteristics. Sex and academic program were excluded from the primary model because of extreme category imbalance (90.5% women; 52.9% nursing students). Ordinal measurement screening, Harman diagnostics, and sensitivity analyses assessed construct separation, same-source measurement, eligibility, use-hours coding, and plausible class clustering.

Results

The dataset included 840 respondents (mean age 19.41 years; 90.5% women). School support and responsibility were strongly correlated (Spearman's $\rho = 0.771$). Their adjusted association was large (standardized $\beta = 0.745$, 95% confidence interval 0.680–0.810; $p < 0.001$; $R^2=0.638$), remained similar after the eligibility restriction ($\beta = 0.750$) and categorical coding of use time ($\beta = 0.747$), and attenuated after literacy adjustment ($\beta = 0.314$). The heterotrait–monotrait ratio was 0.853, exceeding the stringent 0.85 threshold; discriminant validity was not established under that criterion (latent correlation 0.842). The first unrotated component explained 50.3% of variance across all 59 Likert items and 72.4% across the 15 focal items.

Conclusions

Students who rated school support more favorably also rated their digital civic responsibility more favorably. Same-source measurement, a dominant general response factor, and unestablished discriminant validity could substantially inflate this association; it should not be interpreted as an independent or causal effect. The result is hypothesis-generating and supports prospective, multi-source evaluation.

Background

Social media has become part of the ordinary learning environment of health professions students. It supports peer communication, access to educational resources, professional networking, and public engagement, but it also creates risks involving misinformation, privacy, blurred professional boundaries, hostile discourse, and unreflective amplification of algorithmically selected content [1–5]. These tensions are especially salient in health professions education because students' online conduct may affect patients, professional identity, public trust, and the credibility of health information before formal qualification.

Digital citizenship is a multidimensional concept rather than a synonym for safe device use. Ribble and colleagues' nine-element framework includes access, commerce, communication, literacy, etiquette, law, rights and responsibilities, health and wellness, and security [6,7]. A narrower youth-oriented formulation emphasizes respectful online conduct and online civic engagement [8], whereas Choi's concept analysis organizes the field around ethics, media and information literacy, participation and engagement, and critical resistance [9]. The present study uses the term digital civic responsibility for a deliberately narrower, study-specific construct: evidence-aware and respectful participation, attention to rights, privacy, and consequences, and corrective or community-oriented action. It does not claim to measure the full domain of digital citizenship.

Digital professionalism is related but not identical: it concerns how professional identity, confidentiality, boundaries, civility, accountability, and competence are enacted online [1–5]. A learner may participate respectfully in public digital life without yet demonstrating profession-specific boundary management, and vice versa. This conceptual distinction made it necessary to test whether the study's measures were empirically separable before interpreting associations involving either construct.

Institutional context could shape digital civic responsibility through several theoretical pathways. Social cognitive theory proposes that modeled conduct, reinforcement, and perceived environmental affordances influence self-regulation and behavior [12]. The theory of planned behavior likewise identifies subjective norms and perceived behavioral control as proximal determinants of intention and action [13]. Applied here, clear rules and teacher modeling may communicate norms; instruction and authentic cases may build digital and algorithm literacy; credible official platforms may supply informational resources; and accessible help routes may increase perceived capacity to respond responsibly. Recent educational research similarly links institutional or leadership support with digital citizenship through information and communication technology competence [14]. These mechanisms are hypotheses; a cross-sectional single-wave survey cannot establish their temporal order or mediation.

Chinese higher-education policy emphasizes coordinated educational work and integration of civic and ethical formation with disciplinary teaching [15,16]. A school can operationalize this mandate through rules, training, case-based guidance, credible communication channels, and accessible help. Yet evidence concerning how students' perceptions of this environment covary with digital civic responsibility is limited in vocational health professions education. More broadly, meta-analytic evidence

connects school belonging or perceived support with student outcomes, usually with small-to-moderate associations rather than effects approaching unity [17,18].

Substantial measurement caution is therefore required. Exposure and outcome were collected from the same respondent, at the same time, with the same response format, and both used favorable normative wording. Common-method variance, social desirability, acquiescence, and conceptual overlap can create or inflate observed associations [19,20]. High internal consistency also cannot establish validity and may indicate redundant items [21]. The primary aim was thus to estimate, not causally explain, adjusted covariation between perceived school support and self-reported digital civic responsibility. Digital professionalism was prespecified as a candidate secondary outcome, but its association would be reported only if measurement screening supported adequate separation. Digital and algorithm literacy were reserved for sensitivity analysis because they may be mechanisms, proxies for earlier opportunity, or overlapping self-perceptions.

Conceptual framework

Figure 1 summarizes the hypothesized educational pathways. Perceived institutional inputs were expected to relate to responsibility through information competence, perceived norms, and help-seeking efficacy, while background and use characteristics were treated as adjustment variables. Arrows represent theoretical propositions rather than causal paths tested by the present data.

Methods

Study design, setting, and reporting

This was a single-institution cross-sectional survey conducted at Jiangsu College of Nursing, a public vocational health professions college in Jiangsu Province, China. Data collection was conducted from June to July 2026, during the summer vacation period. Because students were not consistently present on campus during this period, face-to-face recruitment was not feasible, and recruitment was conducted exclusively online through institutionally organized class-group communication channels. The study is reported in accordance with the Strengthening the Reporting of Observational Studies in Epidemiology (STROBE) statement [22]; the completed checklist is provided as Additional file 2.

Participants and data collection

The intended population comprised enrolled students aged 18 years or older at Jiangsu College of Nursing, including students undertaking clinical internship. Institution-wide recruitment information was distributed online through class-group communication channels serving approximately 3,300 enrolled students. Available recruitment records suggested that more than 1,200 students accessed or otherwise engaged with the survey invitation. Because recruitment occurred during the summer vacation and relied exclusively on online class-group communication, some students may not have viewed or actively monitored the invitation. In addition, the numbers covered and reached were approximate, and unique

invitation views could not be independently verified. These counts were therefore used to describe the recruitment funnel rather than to calculate an exact response rate.

Participation was voluntary. Before viewing any questionnaire items, students were presented with an electronic participant-information and consent page describing the purpose of the study, the voluntary and anonymous nature of participation, the anticipated completion time, and the intended use of the data. Only students who actively selected the electronic consent option were permitted to enter the questionnaire; students who did not consent could not proceed. The questionnaire was designed to require approximately 20 minutes.

No a priori power calculation was used because recruitment was designed as an institution-wide survey rather than sampling to achieve a prespecified minimum effect size. The online invitation was distributed through class-group channels serving approximately 3,300 enrolled students. All students who met the stated eligibility criteria, provided electronic informed consent, and successfully submitted the questionnaire during the data-collection period were eligible for inclusion. The final analytic sample consisted of all 840 successfully submitted questionnaires.

Questionnaire and construct operationalization

The Chinese-language questionnaire was developed for this study with reference to scholarship on digital citizenship, digital professionalism, social-media behavior, and the institutional context of Chinese higher education [1–16]. It was not a direct translation or adaptation of a single established instrument. All focal construct items used five-point Likert response options from 1 (strongly disagree) to 5 (strongly agree), and composite scores were calculated as item means. Higher values indicate more of the named construct. The study-specific scales should not be interpreted as externally validated diagnostic measures. The full items are provided in Chinese and in an English translation prepared for reporting in Additional file 1; only the Chinese version was administered.

The primary exposure, perceived school support, comprised five items: clear rules concerning social-media conduct, online expression, and privacy; courses or training addressing misinformation, algorithms, and online risks; teacher or counsellor use of authentic cases to discuss digital ethics and professional boundaries; the usefulness and credibility of official school information platforms; and knowledge of where to seek help for rumors, privacy violations, online conflict, or ethical problems. The score showed high internal consistency in this sample (Cronbach's $\alpha=0.965$).

The primary outcome, digital civic responsibility, comprised ten items covering attention to public-health and community issues, consideration of different social groups, discrimination between evidence and emotional or attacking content, respectful expression, respect for rights and privacy, consideration of consequences before reposting or commenting, verification and correction of misinformation, shared responsibility for a truthful and civil online environment, digitally supported volunteering or health-promotion activity, and online collaboration to address community or public-health problems ($\alpha=0.965$).

The candidate secondary measure, digital professionalism, comprised ten items addressing privacy and confidentiality, appropriate handling of clinical content, verification of health information, disclosure of student status and limits, avoidance of advice beyond competence, differentiation of personal from institutional positions, boundaries in patient contact, civility, disclosure of commercial conflicts, and reporting harmful content (α=0.976). Because several of these items are conceptually close to institutional guidance about ethics and boundaries, measurement separation from perceived school support was evaluated before interpreting the secondary association. The planned outcome regression was not interpreted or reported when screening showed inadequate separation.

Additional multi-item covariates were purposeful social-media use (five items; α=0.838), social-media dependence (six items; α=0.911), digital literacy (12 items; α=0.965), and algorithm literacy (eight items; α=0.953). Dependence items reflected salience, difficulty reducing use, emotional reliance, interference with study or sleep, and related behavioral features, consistent with a components-based understanding of problematic use [23]. Daily social-media use was recorded categorically and midpoint-coded as 0.5, 1.5, 2.5, 3.5, 5, and 7 hours for <1, 1–<2, 2–<3, 3–<4, 4–<6, and ≥6 hours, respectively.

Covariate selection

The adjustment set was specified by theory, temporal plausibility, construct reliability, and avoidance of redundant variables rather than by stepwise significance testing. It included age, current intensive internship status, self-rated academic performance, an indicator for missing academic-performance data, daily social-media hours, purposeful use, and social-media dependence. Academic performance was midpoint-filled with a value of 3 for 58 respondents who selected “do not know/unwilling to answer,” and the missingness indicator was included alongside the filled value.

Sex and academic program were deliberately excluded from the primary model. Women constituted 90.5% of the sample, and more than half of respondents were nursing students, with several programs represented by small counts. Including these variables would have produced unstable or poorly generalizable category-specific adjustments. They are therefore fully described in Table 1 and treated as limitations of the sample rather than modeled as primary confounders. Study year, any previous internship, student leadership, and a low-reliability maladaptive/passive-use score were also omitted because of redundancy, limited incremental information, or inadequate internal consistency.

Measurement screening

Because the focal measures were ordinal and potentially overlapping, polychoric correlation matrices were used for an exploratory ordinal measurement screening. One-factor and correlated two-factor structures were compared with a diagonally weighted least-squares approximation. Standardized root mean square residual (SRMR), standardized loadings, latent-factor correlations, and the heterotrait–monotrait ratio (HTMT) were emphasized [24–27]. The approximate chi-square, comparative fit index, Tucker–Lewis index, and root mean square error of approximation were regarded as diagnostic rather

than definitive because a full mean- and variance-adjusted weighted least-squares implementation was not used. No post hoc correlated residuals or item deletions were introduced to improve fit.

Assessment of common-method variance

The questionnaire contained no prespecified marker variable, and the single-wave design prevented temporal or source separation. As a post hoc diagnostic, Harman's single-factor procedure was applied as unrotated principal-components analysis to (1) all 59 Likert items and (2) the 15 items forming perceived school support and digital civic responsibility. The percentage of variance captured by the first component and the number of components with eigenvalues greater than 1 were recorded. This procedure cannot quantify or statistically remove common-method bias and was not treated as proof of its absence; it was used only to identify whether a dominant general response factor was plausible [19,20].

Statistical analysis

Descriptive statistics were reported as mean and standard deviation for continuous scores and as count and percentage for categorical variables. Internal consistency was summarized with Cronbach's alpha. Spearman correlations described unadjusted monotonic associations. The primary analysis used ordinary least-squares regression with HC3 heteroskedasticity-consistent standard errors [28]. The outcome and model predictors were standardized, so regression coefficients are reported in standard-deviation units with 95% confidence intervals. Two-sided p values are provided as measures of compatibility rather than as a variable-selection rule.

The primary model regressed digital civic responsibility on perceived school support and the prespecified covariates. Incremental explanatory value was assessed by comparing the model R^2 with that of the covariate-only model. Model optimism was examined using shuffled 10-fold cross-validation with a fixed random seed. Variance inflation factors were inspected for multicollinearity. The eligibility-consistent analysis repeated the model in 822 respondents. A second sensitivity model added digital and algorithm literacy; because these individual capacities may be mediators, proxies for prior educational opportunity, or common causes, its school-support coefficient estimates a different, more conditional association and was not designated as the primary association estimate. A further sensitivity analysis represented the six daily-use categories with indicator variables, using 2–<3 hours as the reference, rather than imposing midpoint linearity.

Class identifiers, class sizes, and intraclass correlations were unavailable. To show the possible magnitude of underestimated uncertainty, two illustrative design-effect scenarios were calculated with $DEFF=1+(m-1)ICC$, using mean class size $m=40$ and $ICC=0.05$ and, more conservatively, $m=50$ and $ICC=0.10$. The primary HC3 standard error was multiplied by the square root of each DEFF. These are not cluster-robust estimates and do not recover the unknown sampling structure; they are transparent sensitivity bounds.

The candidate digital-professionalism regression was conditional on acceptable construct separation. Because the screening criterion was not met, its association with school support was not estimated for substantive interpretation. Analyses were performed in Python 3.12 using pandas, NumPy, SciPy, scikit-learn, and statsmodels-compatible matrix procedures. No item-level data were missing for the 59 Likert items used in scale scoring.

The electronic questionnaire required completion of all items before successful submission. Accordingly, no item-level data were missing from the 59 Likert-type items used for scale scoring among the 840 submitted questionnaires. The “do not know/unwilling to answer” option for self-rated academic performance was treated as an observed response category rather than as questionnaire non-completion; its analytic handling is described above. No imputation was required for the perceived-school-support exposure or the digital-civic-responsibility outcome.

Results

Participant recruitment and analytic sample

The online survey invitation was distributed through class-group communication channels serving approximately 3,300 students across the institution. More than 1,200 students were estimated to have accessed or engaged with the invitation, and 840 provided electronic informed consent, completed all required questionnaire items, and successfully submitted the survey. The approximate institution-wide recruitment yield was therefore 25.5%, and the completion proportion among students known to have encountered the invitation was no greater than approximately 70%. Because the coverage and access counts were approximate and could not be linked to verified unique individuals, these values describe participant flow and are not presented as formal response rates. No partially completed questionnaire was included in the analytic dataset.

The final sample included students from all principal stages of the three-year vocational education program: 464 first-year students, 272 second-year students, two third-year students who were not undertaking intensive internship, and 102 students undertaking concentrated internship. This distribution provided coverage across different stages of training, although it does not by itself establish statistical representativeness of the entire student population.

Participant characteristics

The analytic dataset comprised all 840 successfully submitted questionnaires. Respondents had a mean age of 19.41 years (standard deviation 0.91). Most were women (760/840, 90.5%), and 444 (52.9%) were nursing students. First-year students accounted for 464 (55.2%) of the sample, second-year students for 272 (32.4%), and 104 (12.4%) were in the third-year or concentrated-internship stage. More than four fifths reported at least two hours of daily social-media use, including 257 (30.6%) reporting four or more hours. Full participant characteristics are presented in Table 1.

Scale distributions and bivariate associations

Mean perceived school support was 3.93 (standard deviation 0.76), mean digital civic responsibility was 3.83 (0.68), and both scores had Cronbach's α=0.965. Perceived school support was strongly correlated with digital civic responsibility (Spearman's ρ=0.771, 95% confidence interval 0.742 to 0.797; p<0.001). The magnitude of alpha may reflect homogeneous or redundant wording as well as consistency and is not evidence of construct validity [21]. Descriptive values and correlations are shown in Table 2.

Measurement screening

For perceived school support and digital civic responsibility, a two-factor model showed substantially lower residual misfit than a one-factor model (SRMR 0.044 versus 0.089). The two-factor solution nevertheless had an approximate RMSEA of 0.117, latent correlation of 0.842, and HTMT of 0.853. The HTMT ratio exceeded the stringent 0.85 threshold recommended for conceptually distinct constructs [24]; discriminant validity was therefore not established under that criterion. The data could not exclude the possibility that the two scores largely reflect one broad favorable evaluative or normative construct. For school support and digital professionalism, the two-factor SRMR was 0.033, but the latent correlation (0.923) and HTMT (0.924) exceeded commonly used 0.90 warning thresholds. Splitting digital professionalism into privacy, professional-boundary, and civility/reporting subdomains did not resolve the overlap (maximum latent correlation approximately 0.98; maximum HTMT 0.997). These results supported retaining digital civic responsibility as the primary outcome, with strong qualification, but did not support interpreting digital professionalism as an independent secondary outcome. Its planned regression was therefore not performed or reported; the measurement evidence supporting this decision is presented in Table 5.

Common-method diagnostic

Across all 59 Likert items, the first unrotated principal component had an eigenvalue of 29.69 and explained 50.3% of total variance; seven components had eigenvalues greater than 1. Across the 15 focal school-support and civic-responsibility items, the first component explained 72.4% and two components had eigenvalues greater than 1. The presence of multiple components is consistent with some multidimensionality, but the dominant first component—especially among the focal items—indicates a substantial general response factor. The increase from 50.3% across all items to 72.4% across the focal items suggests that the two focal scales were particularly vulnerable to a shared response style, plausibly because both used favorable normative wording and the same response format. Because Harman's procedure is only a weak diagnostic and no marker variable was available, the amount of common-method variance could not be estimated or removed.

Primary association with digital civic responsibility

In the covariate-only model, R^2 was 0.184. Adding perceived school support increased R^2 by 0.454 to 0.638 (adjusted R^2=0.634). Higher perceived school support was associated with higher digital civic responsibility (standardized β=0.745, 95% confidence interval 0.680 to 0.810; p<0.001). Purposeful use was positively associated with the outcome (β=0.085, 95% confidence interval 0.032 to 0.138; p=0.002), whereas social-media dependence was negatively associated (β=−0.063, 95% confidence interval −0.119 to −0.008; p=0.026). Better self-rated academic performance also had a small positive association (β=0.058, 95% confidence interval 0.018 to 0.097; p=0.004). Other covariates were not clearly associated after adjustment (Table 3). The maximum variance inflation factor was 1.23, and the 10-fold cross-validated R^2 was 0.625, suggesting limited optimism under random partitions within this sample.

Sensitivity analyses

The eligibility-consistent analysis produced a nearly identical school-support estimate (n=822; β=0.750, 95% confidence interval 0.685 to 0.816; p<0.001; R^2=0.644). When digital and algorithm literacy were added, the school-support coefficient attenuated to 0.314 (95% confidence interval 0.243 to 0.386; p<0.001), while digital literacy (β=0.298, 95% confidence interval 0.220 to 0.376) and algorithm literacy (β=0.352, 95% confidence interval 0.272 to 0.432) were independently associated with digital civic responsibility. The expanded model explained 80.7% of outcome variance and had a maximum variance inflation factor of 3.21. The attenuation is consistent with shared explanatory pathways or overlapping self-reported competencies, but cross-sectional data cannot distinguish mediation from confounding or common-method effects.

Replacing midpoint-coded hours with indicators for the six reported use categories produced essentially the same school-support coefficient (β=0.747, 95% confidence interval 0.683 to 0.811; p<0.001; R^2=0.643). In the illustrative clustering scenarios, a mean class size of 40 and ICC of 0.05 yielded DEFF=2.95 and an inflated standard error of 0.057 (95% confidence interval 0.634 to 0.856); a mean class size of 50 and ICC of 0.10 yielded DEFF=5.90 and an inflated standard error of 0.080 (95% confidence interval 0.588 to 0.902). These intervals remained above zero, but they are scenario calculations rather than valid cluster-adjusted inference. Sensitivity results are summarized in Table 4.

Discussion

Principal findings

Students who gave more favorable ratings of school support also reported substantially greater digital civic responsibility after adjustment for selected background and social-media-use characteristics. The numerical estimate was stable under the eligibility restriction and categorical coding of use time and was smaller after digital and algorithm literacy were included. However, the substantive interpretation is constrained more by measurement than by conventional sampling error: discriminant validity was not

established under the stringent HTMT criterion, and a dominant first component means that the model may chiefly capture a shared favorable evaluative orientation rather than an association between clearly distinct constructs.

Accordingly, $\beta = 0.745$ is not an estimated intervention effect, nor does $R^2=0.638$ mean that school support causally explains 63.8% of responsibility. Both values are properties of a same-wave, same-source scoring system. The data cannot exclude that perceived school support and digital civic responsibility are manifestations of one broader normative self-report construct. Digital professionalism showed still poorer separation and was therefore removed from effect reporting rather than presented as a second confirmatory finding.

Interpretation in relation to prior work

Prior reviews and qualitative studies describe digital professionalism as an educational challenge involving privacy, boundaries, identity, accountability, and the tension between personal and professional online lives [1–5]. The present results extend that discussion toward digital civic responsibility: verifying health-related claims, participating respectfully, considering groups affected by online discourse, and contributing to public or community activity. This is compatible with frameworks spanning online rights and responsibilities, literacy, etiquette, ethical participation, and civic engagement rather than treating competence as technical proficiency alone [6–11]. The present outcome is narrower than those frameworks and should be described as self-reported digital civic responsibility, not as a comprehensive measure of digital citizenship.

Theoretically, institutional rules and teacher examples can communicate norms and model conduct, while instruction, credible information, and help routes can strengthen capability and perceived control [12–14]. The attenuation after literacy adjustment is compatible with such a pathway, but it is equally compatible with confounding or overlapping self-evaluation. A longitudinal mediation design would require the institutional exposure to precede literacy and responsibility measurements and should include independent or behavioral outcomes.

The observed coefficient is also much larger than the small-to-moderate correlations commonly summarized between school belonging or perceived support and broad student outcomes, often approximately 0.2–0.4 [17, 18]. Those prior correlations are not directly comparable with this adjusted standardized coefficient because constructs and models differ. Nevertheless, the contrast makes substantive caution more—not less—important and is consistent with inflation from item similarity, shared response format, social desirability, or a general positive-response factor.

The school-support items still identify potentially actionable institutional inputs: clear rules, case-based teaching, credible platforms, and help-seeking routes. These align with social cognitive and planned-behavior mechanisms and with Chinese policy approaches integrating civic and ethical formation throughout higher education [12–16]. The present composite did not determine which component

mattered, whether exposure preceded responsibility, or whether students' favorable orientation influenced ratings of both.

Implications for health professions education

For educators, the findings support evaluating school support as a coherent educational environment rather than limiting digital conduct to disciplinary prohibition. Programs could explicitly connect social-media rules to professional and civic rationales, use cases involving misinformation and algorithmic amplification, maintain credible student-facing channels, and publicize confidential routes for advice. These practices are plausible targets for curriculum development, but their effectiveness should be tested with prospective or quasi-experimental designs and outcomes that extend beyond favorable self-report.

Measurement development is equally important. Future instruments should separate institutional inputs from student knowledge, attitudes, intentions, and observed practices; include negatively and behaviorally anchored items where appropriate; and be evaluated across institutions and study programs. Cognitive interviewing, independent item development samples, confirmatory ordinal models, measurement invariance, and associations with behavioral criteria would strengthen validity. Digital professionalism and digital civic responsibility should not be combined merely because they correlate strongly; each has a distinct educational purpose.

Strengths and limitations

Strengths include a large analytic sample, institution-wide distribution of the recruitment invitation, and inclusion of students from all principal stages of the three-year vocational education program, including students undertaking concentrated internship. Although the final sample represented approximately one quarter of the institution's enrolled student population, its distribution across training stages provided broader within-institution coverage than a survey restricted to one year group or classroom. Additional strengths include transparent prioritization of one exposure and one primary outcome, theory-guided rather than significance-driven covariate selection, HC3 robust inference, cross-validation, an eligibility-consistent sensitivity analysis, and explicit measurement screening.

Several limitations constrain interpretation. First, the cross-sectional design does not establish temporal order and cannot exclude reverse causation: students who already behave responsibly may perceive institutional support more favorably. Second, all focal variables came from the same self-report questionnaire, creating risks of social desirability, acquiescence, and common-method variance [19, 20]. The Harman diagnostic showed a dominant component, but no marker variable or independent source was available, so common-method bias could not be quantified or removed. Third, the study-specific scales require external validation. Ordinal screening showed imperfect global fit, and discriminant validity for the primary construct pair was not established under the stringent HTMT criterion; separation between school support and digital professionalism was inadequate. Very high alpha coefficients may reflect item redundancy and construct narrowing rather than superior validity [21].

Fourth, the study came from one vocational health professions college. The sample was predominantly female and concentrated in nursing, while nine respondents selected other/prefer not to say for sex; these categories were described but not modeled. Nursing students may receive discipline-specific ethical or professional instruction that changes the level or range of both scores, limiting transportability to differently composed institutions. The response-rate denominator was also unavailable. Sex and academic program were therefore not included in the primary model, but their omission does not remove the corresponding limitation in transportability or residual confounding. Fifth, class identifiers were unavailable, so potential clustering by class could not be modeled. The reported HC3 standard errors and p values may therefore be anticonservative; the design-effect scenarios illustrate possible inflation but cannot replace cluster-robust or multilevel analysis. Sixth, coding ≥ 6 hours as 7 hours assumes a finite midpoint for an open-ended category, although categorical coding gave a similar focal estimate. Midpoint filling of missing academic-performance responses also simplified an ordered category. Finally, cross-validation under random partitions of one institution is internal validation only; it neither validates performance across institutions nor addresses transportability.

Conclusions

Students who perceived more school support also reported greater digital civic responsibility in this single-institution sample. The large coefficient cannot be interpreted as an independent institutional effect because same-source measurement, a dominant general component, and discriminant validity not established under the stringent HTMT criterion may have substantially inflated it. Rules, training, teacher guidance, credible platforms, and help-seeking channels are reasonable targets for prospective study, but multi-source measurement and independent construct validation are required before causal or programmatic claims.

Abbreviations

CI
confidence interval
CMV
common-method variance
DEFF
design effect
HC3
heteroskedasticity-consistent covariance estimator type 3
HTMT
heterotrait–monotrait ratio
ICC
intraclass correlation coefficient
OLS
ordinary least squares

PCA
principal-components analysis
RMSEA
root mean square error of approximation
SRMR
standardized root mean square residual
STROBE
Strengthening the Reporting of Observational Studies in Epidemiology
VIF
variance inflation factor.

Declarations

Ethics approval and consent to participate

Ethical approval was granted by the Ethics Committee of Jiangsu College of Nursing (approval number JSCN-ME-2026063019). Electronic informed consent was obtained from every participant before questionnaire access. Participation was voluntary and responses were collected anonymously. All procedures were conducted in accordance with applicable institutional requirements and the principles of the Declaration of Helsinki.

Consent for publication

Not applicable. The manuscript contains no identifiable individual data.

Availability of data and materials

The de-identified dataset and analytic code are available from the corresponding author on reasonable request, subject to institutional and ethics requirements.

Competing interests

The authors declare no competing interests.

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Authors' contributions

Y.L. contributed to conceptualization, methodology, investigation, formal analysis, interpretation, and writing the original draft. B.X. contributed to investigation, data curation, interpretation, and reviewing and editing the manuscript. Q.L. contributed to conceptualization, supervision, project administration, validation, and reviewing and editing the manuscript. All authors read and approved the final manuscript.

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Tables

Table 1. Participant characteristics (N=840)

Characteristic	Value
Age, mean (SD), years	19.41 (0.91)
Sex, n (%)	
Female	760 (90.5)
Male	71 (8.5)
Other/prefer not to say	9 (1.1)
Academic program, n (%)	
Nursing	444 (52.9)
Medical laboratory technology	124 (14.8)
Other medical/health program	102 (12.1)
Rehabilitation therapy	67 (8.0)
Preventive medicine/health management	35 (4.2)
Medical imaging technology	32 (3.8)
Pharmacy/traditional Chinese medicine	28 (3.3)
Other	8 (1.0)
Study stage, n (%)	
First year	464 (55.2)
Second year	272 (32.4)
Third year, not in intensive internship	2 (0.2)
Concentrated internship	102 (12.1)
Internship experience, n (%)	
None	231 (27.5)
Short-term placement	494 (58.8)
Currently in internship	109 (13.0)
Completed internship	6 (0.7)
Student leadership experience, n (%)	
Never	407 (48.5)
Current	277 (33.0)
Previous	156 (18.6)

Characteristic	Value
Home location, n (%)	
City	117 (13.9)
County	298 (35.5)
Township	220 (26.2)
Rural area	205 (24.4)
Daily social-media use, n (%)	
<1 hour	23 (2.7)
1–<2 hours	110 (13.1)
2–<3 hours	267 (31.8)
3–<4 hours	183 (21.8)
4–<6 hours	158 (18.8)
≥6 hours	99 (11.8)

Percentages may not sum to 100 because of rounding. Sex and academic program were described but not included in the primary regression model because of severe category imbalance.

Table 2. Descriptive statistics, reliability, and Spearman correlations for focal measures

Measure	Items	Mean (SD)	α	1	2	3	4	5	6
1. School support	5	3.93 (0.76)	0.965	1.000					
2. Digital civic responsibility	10	3.83 (0.68)	0.965	0.771	1.000				
3. Purposeful use	5	3.00 (0.62)	0.838	0.252	0.261	1.000			
4. Social-media dependence	6	2.48 (0.68)	0.911	−0.311	−0.302	0.121	1.000		
5. Digital literacy	12	3.74 (0.68)	0.965	0.662	0.792	0.286	−0.226	1.000	
6. Algorithm literacy	8	3.69 (0.65)	0.953	0.685	0.789	0.196	−0.224	0.765	1.000

α=Cronbach's alpha. All displayed Spearman correlations had p<0.001. Scores range from 1 to 5; higher scores indicate more of the named construct.

Table 3. Primary HC3-robust regression for digital civic responsibility (N=840)

Predictor	Standardized β	HC3 SE	95% CI	p value
Age	0.047	0.027	−0.006 to 0.099	0.083
Current intensive internship	−0.042	0.023	−0.086 to 0.003	0.066
Self-rated academic performance	0.058	0.020	0.018 to 0.097	0.004
Academic-performance missingness	0.043	0.023	−0.001 to 0.088	0.058
Daily social-media hours	0.026	0.021	−0.015 to 0.066	0.215
Purposeful social-media use	0.085	0.027	0.032 to 0.138	0.002
Social-media dependence	−0.063	0.028	−0.119 to −0.008	0.026
Perceived school support	0.745	0.033	0.680 to 0.810	<0.001

Outcome and predictors were standardized. R^2=0.638; adjusted R^2=0.634; covariate-only R^2=0.184; ΔR^2 after adding perceived school support=0.454; 10-fold cross-validated R^2=0.625; maximum variance inflation factor=1.23. CI=confidence interval; HC3 SE=heteroskedasticity-consistent standard error type 3.

Table 4. Sensitivity analyses for the school-support coefficient

Analysis	N	School-support β (95% CI)	R^2	Interpretation
Primary model	840	0.745 (0.680 to 0.810)	0.638	Primary adjusted covariation
Eligibility-consistent	822	0.750 (0.685 to 0.816)	0.644	Materially unchanged
Added digital and algorithm literacy	840	0.314 (0.243 to 0.386)	0.807	Different, conditional association
Daily-use categories	840	0.747 (0.683 to 0.811)	0.643	No midpoint-linearity assumption
Illustrative clustering: m=40, ICC=0.05	840	0.745 (0.634 to 0.856)	—	DEFF=2.95; SE multiplied by 1.72
Illustrative clustering: m=50, ICC=0.10	840	0.745 (0.588 to 0.902)	—	DEFF=5.90; SE multiplied by 2.43

Outcome=digital civic responsibility. Each fitted model adjusted for age, current intensive internship, self-rated academic performance, its missingness indicator, daily social-media use, purposeful use, and social-media dependence. The literacy-adjusted model additionally included digital and algorithm literacy. The last two rows are design-effect scenarios, not fitted cluster-robust models; actual class identifiers and ICCs were unavailable. β values are standardized. CI=confidence interval; DEFF=design effect; ICC=intraclass correlation coefficient; m=assumed mean class size; SE=standard error.

Table 5. Measurement-separation and common-method diagnostics

Diagnostic	Model/sample	Key result	Interpretation
Ordinal factor screening	Support + civic responsibility, one factor	SRMR=0.089; RMSEA†=0.226	Poor residual fit
Ordinal factor screening	Support + civic responsibility, two factors	SRMR=0.044; RMSEA†=0.117; latent r=0.842; HTMT=0.853	Better residual fit, but HTMT >0.85; discriminant validity not established
Ordinal factor screening	Support + digital professionalism, two factors	SRMR=0.033; latent r=0.923; HTMT=0.924	Insufficient separation; outcome regression not reported
Ordinal factor screening	Support + professionalism subdomains	Maximum latent r=0.980; maximum HTMT=0.997	Subdomain splitting did not resolve overlap
Harman unrotated PCA	All 59 Likert items	First component=50.3%; eigenvalue=29.69; 7 eigenvalues >1	Dominant general component; CMV remains plausible
Harman unrotated PCA	15 focal support/responsibility items	First component=72.4%; 2 eigenvalues >1	Strong general response factor among focal items

†Approximate diagnostics from polychoric correlations and diagonally weighted least-squares screening; they are not full mean- and variance-adjusted WLS estimates. Harman PCA is a post hoc diagnostic, not a statistical control for common-method variance. CMV=common-method variance; HTMT=heterotrait–monotrait ratio; PCA=principal-components analysis; RMSEA=root mean square error of approximation; SRMR=standardized root mean square residual.

Figures

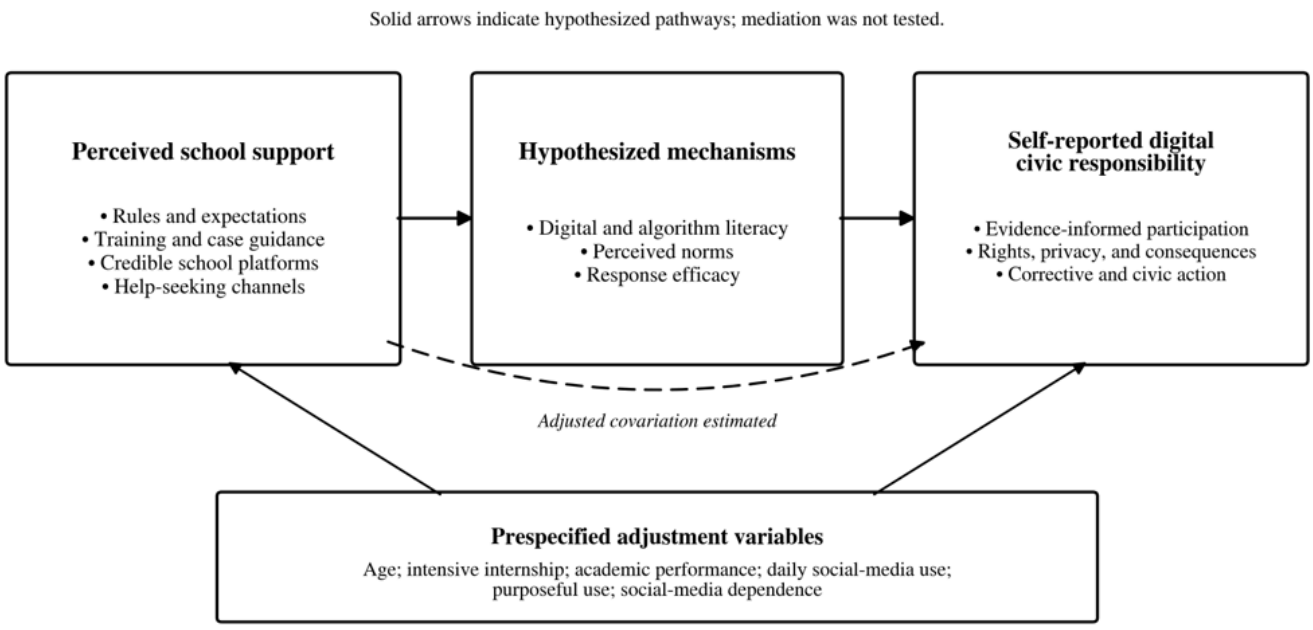


Figure 1

Hypothesized pathways linking school support and digital civic responsibility. The cross-sectional model estimated total adjusted covariation and did not test the depicted mediating mechanisms.

Supplementary Files

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

- Additionalfile1BilingualQuestionnaire.docx
- Additionalfile2STROBEChecklist.docx