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LONELINESS AND PARASOCIAL INVOLVEMENT IN THE DIGITAL
MEDIA CONTEXT: THE MEDIATING ROLE OF PARASOCIAL
ATTACHMENT AND ATTITUDES TOWARD CELEBRITIES

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Abstract

Parasocial relationships represent an increasingly relevant form of symbolic connectedness in the contemporary digital media environment and are frequently associated with socio-emotional vulnerabilities such as perceived loneliness. The present study aims to investigate the psychosocial mechanisms through which loneliness is associated with parasocial involvement by testing an integrative explanatory model that includes parasocial attachment and attitudes toward celebrities as serial mediators. In addition, the study explores the existence of a non-linear relationship between loneliness and the anxious dimension of parasocial attachment.

The research was conducted on a large adult sample, using validated instruments to assess perceived loneliness, parasocial attachment, attitudes toward celebrities, and parasocial behavior. Statistical analyses included serial mediation models and polynomial regressions, designed to capture both direct and indirect relationships among variables, as well as potential curvilinear effects.

The results indicate that loneliness does not directly predict parasocial behavior but exerts its influence primarily through parasocial attachment and attitudes toward celebrities. Parasocial attachment emerged as a central cognitive-affective mechanism, facilitating the internalization of the symbolic relationship with the media figure and structuring subsequent attitudinal evaluations. Moreover, the relationship between loneliness and anxious parasocial attachment proved to be curvilinear, suggesting the existence of psychological thresholds in the functioning of media-based compensatory mechanisms.

These findings support a process-oriented conceptualization of parasocial relationships, in which socio-emotional vulnerabilities are gradually transformed into media engagement behaviors through distinct affective and evaluative mechanisms. The study contributes to the refinement of existing theoretical models and offers relevant implications for understanding the ambivalent role of parasocial relationships in emotional regulation and mental health within digital society.

Keywords: perceived loneliness; parasocial relationships; parasocial attachment; attitudes toward celebrities; parasocial behavior; digital media

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1. INTRODUCTION

In the context of the accelerated transformations of the digital media ecosystem, parasocial relationships have become a central component of the ways in which individuals symbolically interact with public figures, celebrities, and influencers. Initially conceptualized by Horton and Wohl as unilateral forms of perceived interaction between audiences and media characters, parasocial relationships have evolved in contemporary literature into complex relational structures with distinct cognitive, affective, and behavioral implications (Tukachinsky *et al.*, 2019; Giles, 2023). At present, these relationships can no longer be understood solely as reactions to media exposure, but rather as dynamic psychosocial processes embedded in emotional regulation, social identity, and individuals' compensatory strategies.

A frequently invoked psychosocial determinant in explaining orientation toward parasocial relationships is perceived loneliness. The literature suggests that experiences of socio-emotional isolation and deficits in interpersonal belonging may motivate individuals to seek stable, predictable symbolic connections devoid of rejection risk – defining characteristics of parasocial relationships (Greenwood *et al.*, 2009, 2011). Loneliness reflects not merely the absence of social contacts, but a subjective discrepancy between desired and existing relationships, with profound consequences for psychological health and media consumption behaviors (Russell, 1996; Wang *et al.*, 2018). Thus, orientation toward media figures may function as a compensatory mechanism, providing an illusion of intimacy and emotional connection within a context perceived as safe and controllable.

However, the relationship between loneliness and parasocial involvement is neither direct nor unidimensional. Recent research highlights the role of parasocial attachment processes as key explanatory mechanisms through which socio-emotional vulnerabilities translate into affective investment in media figures (Cohen, 2003; Rain & Mar, 2021). Parasocial attachment, conceptualized by analogy with adult attachment theories, reflects internalized relational styles – particularly anxious or avoidant – projected onto symbolic relationships with celebrities. These styles may amplify the need for closeness, emotional dependence, and sensitivity to the absence of the media figure, thereby contributing to the consolidation of parasocial relationships and intensifying their psychological significance (Cole & Leets, 1999).

Another central construct within this explanatory framework is attitudes toward celebrities, frequently operationalized through the concept of celebrity worship. Classical models describe this phenomenon as a continuum ranging from entertainment-social interest to intense-personal involvement and borderline-pathological forms (McCutcheon *et al.*, 2010).

Despite significant progress over recent decades, the literature still exhibits relevant conceptual and empirical gaps. First, most studies have examined loneliness, parasocial attachment, and attitudes toward celebrities in isolation, without integrating these variables into a coherent explanatory model capable of simultaneously capturing direct, mediated, and behavioral relationships. Second,

there is a limited number of studies explicitly testing serial mediation mechanisms through which loneliness influences parasocial behaviors via parasocial attachment and attitudes toward celebrities. Third, non-linear relationships between loneliness and dimensions of parasocial attachment remain insufficiently explored, although recent research suggests the existence of psychological thresholds and saturation effects in media-based compensatory processes (Hartmann, 2017; Bond, 2021).

The social and psychological relevance of investigating these mechanisms is amplified by the massive growth of digital media consumption and the central role of celebrities and influencers in young adults' everyday lives. Parasocial relationships are not neutral phenomena; they may have adaptive implications – such as temporary reductions in loneliness and the provision of symbolic support – as well as maladaptive consequences, including emotional dependence, avoidance of real interpersonal relationships, or escalation toward obsessive behaviors (Sansone *et al.*, 2014; Liebers & Schramm, 2019). From a methodological perspective, integrating mediation models and non-linear analyses contributes to refining the understanding of parasocial relationships as dynamic processes rather than static correlations.

Within this context, the aim of the present study is to investigate the psychosocial mechanisms through which perceived loneliness is associated with parasocial involvement by testing an integrative explanatory model that includes parasocial attachment and attitudes toward celebrities as cognitive and attitudinal mediators. The study seeks to examine both the direct and mediated effects of loneliness on parasocial behavior, as well as the existence of curvilinear relationships between loneliness and the anxious dimension of parasocial attachment. Through this approach, the research contributes to the international literature by clarifying the psychosocial architecture of parasocial relationships and by offering a robust theoretical and methodological framework for understanding how socio-emotional vulnerabilities are transformed into specific forms of media connectivity.

2. LITERATURE REVIEW

Parasocial relationships represent forms of unilateral symbolic relating in which individuals develop perceptions of intimacy, familiarity, and emotional connection with media figures who do not reciprocate these interactions. Initially conceptualized by Horton and Wohl as illusory interactions between audiences and media personalities, parasocial relationships were subsequently extended to describe stable and internalized bonds with fictional characters, celebrities, or influencers, characterized by continuity, affective investment, and psychological significance (Tukachinsky *et al.*, 2019; Giles, 2023). Contemporary literature distinguishes between episodic parasocial interactions and enduring parasocial relationships, the latter displaying structural properties similar to real interpersonal relationships, such as attachment, loyalty, and emotional reactivity (Bond, 2021).

A central psychosocial determinant of orientation toward parasocial relationships is perceived loneliness, defined as the subjective discrepancy between desired and existing social relationships. This conceptualization, established through the UCLA Loneliness Scale, highlights the socio-emotional nature of loneliness and its impact on affect regulation, mental health, and coping strategies, including media consumption (Russell, 1996; Zeas-Sigüenza *et al.*, 2023). Recent research indicates that loneliness functions as a vulnerability factor that predisposes individuals toward symbolic compensatory strategies, particularly in digital contexts that provide easy access to media figures perceived as available, predictable, and free from rejection risk (Greenwood *et al.*, 2009, 2011; Wang *et al.*, 2018).

Within this explanatory framework, parasocial attachment represents a key cognitive-affective mechanism, derived from the application of adult attachment theories to relationships with media figures. Similar to interpersonal attachment, parasocial attachment is organized along anxious and avoidant dimensions, reflecting needs for closeness, fear of abandonment, and discomfort with symbolic distance (Cole & Leets, 1999; Cohen, 2003). Empirical studies indicate that individuals with anxious attachment styles develop more intense and emotionally dependent parasocial relationships, using media figures as sources of symbolic security and affective validation (Rain & Mar, 2021). Consequently, parasocial attachment frequently mediates the relationship between socio-emotional vulnerabilities and the intensity of parasocial involvement.

A closely related construct is attitudes toward celebrities, operationalized through the concept of celebrity worship. Classical models describe this phenomenon as a continuum ranging from recreational and social interest to intense-personal involvement and borderline-pathological forms (McCutcheon *et al.*, 2010). Attitudes toward celebrities integrate cognitive, affective, and motivational components and are systematically associated with parasocial attachment and media engagement behaviors, such as persistent content consumption and information seeking about the media figure.

From a theoretical perspective, explanations of parasocial relationships have evolved from descriptive models to integrative frameworks. Media equation theory posits that individuals respond to media stimuli as if they were real social entities, facilitating the emergence of parasocial relationships in the absence of reciprocity (Reeves *et al.*, 1996). Compensatory media use models and extended attachment theories suggest that parasocial relationships may function as emotional regulation mechanisms in response to loneliness, stress, and social insecurity, providing symbolic support and relational predictability (Hartmann, 2017; Bond, 2021). Additionally, the absorption – addiction model explains the progression from emotional involvement to potentially compulsive behaviors in contexts of psychological vulnerability (McCutcheon *et al.*, 2010).

Empirical literature documents consistent associations between loneliness, parasocial attachment, attitudes toward celebrities, and parasocial behaviors; however, these relationships are not uniformly linear. Some studies indicate threshold or saturation effects, suggesting that parasocial relationships may serve adaptive functions up to a certain level of emotional vulnerability, after which they may become ineffective or maladaptive (Hartmann, 2017; Bond, 2021). Nevertheless, most research examines these variables in a fragmented manner, without testing explanatory models that capture the sequential nature of psychosocial processes through serial mediation analyses or explicitly integrating non-linear relationships (Liebers & Schramm, 2019).

In this context, the present study positions itself as an integrative contribution to the international literature by proposing an explanatory model that treats loneliness as a distal predictor, parasocial attachment as a cognitive-affective mediator, and attitudes toward celebrities as an attitudinal mediator of parasocial behaviors. By simultaneously testing direct, mediated, and curvilinear effects in a large sample, the research addresses significant conceptual and methodological gaps and offers a process-oriented perspective on how socio-emotional vulnerabilities are transformed into specific forms of media connectivity in contemporary digital society.

Building on international literature regarding the psychosocial determinants of parasocial relationships, the proposed theoretical model anticipates a series of direct, mediated, and non-linear relationships among the study's central variables: perceived loneliness, parasocial attachment, attitudes toward celebrities, and parasocial behavior. Prior research indicates that loneliness is associated with compensatory orientations toward parasocial relationships, functioning as a distal predictor of affective and behavioral engagement with media figures. Moreover, parasocial attachment processes are recognized as key explanatory mechanisms through which socio-emotional vulnerabilities contribute to the formation of parasocial relationships. Accordingly, the hypotheses are formulated to reflect both established literature and the innovative contribution of the present model.

H1. Higher levels of perceived loneliness (UCLA) will predict increased parasocial attachment (PSI-A), indicating that socio-emotional isolation facilitates orientation toward symbolic, non-reciprocal connections with media figures as a compensatory mechanism for the lack of interpersonal support.

H2. Parasocial attachment (PSI-A) will mediate the relationship between loneliness (UCLA) and attitudes toward celebrities (CAS), such that socio-emotional vulnerability associated with loneliness facilitates the formation of a symbolic bond with the media figure, through which affective-cognitive orientations toward the celebrity are organized and amplified.

H3. A serial mediated effect of loneliness (UCLA) on parasocial behavior (SPSR) is anticipated through parasocial attachment (PSI-A) and attitudes toward celebrities (CAS). Thus, loneliness will indirectly influence parasocial behavior by

increasing parasocial attachment intensity and subsequently strengthening attitudes toward the celebrity.

H4. The relationship between loneliness (UCLA) and the anxious dimension of parasocial attachment (PSI–A) will exhibit a curvilinear component, suggesting the existence of a psychological threshold beyond which further increases in loneliness lead to accelerated and disproportionate intensification of anxious parasocial attachment.

3. METHODOLOGY

3.1. RESEARCH DESIGN

The study adopted a quantitative cross-sectional explanatory design, appropriate for investigating predictive relationships and psychosocial mechanisms involved in the formation and consolidation of parasocial relationships. This design allows for the simultaneous examination of associations among the distal predictor (loneliness), cognitive-affective and attitudinal mediators (parasocial attachment and attitudes toward celebrities), and the behavioral outcome (parasocial involvement), without experimental manipulation or longitudinal follow-up. The explanatory model aimed to estimate direct and indirect pathways, both simple and serial, incorporating linear effects as well as potential curvilinear relationships among the core variables.

The analytical process was structured in two complementary stages. In the first stage, the psychometric properties of the four instruments were evaluated through measurement models (Confirmatory Factor Analysis), using standard indicators of validity and reliability (α , ω , AVE, CR, and HTMT). In the second stage, predictive relationships were tested through simple and serial mediation models (PROCESS Models 4 and 6), multivariate regression analyses, and, where appropriate, non-linear models and polynomial terms to examine curvilinear effects. The use of a cross-sectional explanatory design is consistent with international literature on the study of parasocial relationship formation mechanisms, providing a robust framework for testing a complex multivariate model within contemporary psychosocial research.

3.2. PARTICIPANTS AND SAMPLING

The final sample included 2,335 participants, selected through a non-probabilistic convenience sampling procedure. Inclusion criteria were: a minimum age of 18 years, complete responses to all instruments administered via Google Forms, and the provision of informed consent. Incomplete questionnaires, duplicate entries, and responses with unusually short completion times were excluded to ensure data quality and validity. Adherence to these criteria resulted in a homogeneous and sufficiently large sample for testing the investigated psychosocial relationships.

Regarding sociodemographic characteristics, participants' ages ranged from 18 to 48 years, with a mean age of 26.57 years (Mdn = 23; Mode = 20), indicating a predominance of young adults, a profile characteristic of online-based studies.

Variable	N	M	Mdn	Mod	Min	Max
Age	2335	26.57	23.00	20	18	48

The gender distribution indicates that 70.5% of respondents were female ($n = 1,646$), while 29.5% were male ($n = 689$), a proportion frequently observed in research on parasocial behaviors, where female participation is typically higher. These demographic characteristics reflect a diverse sample, oriented toward population segments with the highest exposure to media content and an increased likelihood of exhibiting forms of parasocial involvement.

Biological gender	N	%
Female	1646	70.5
Male	689	29.5
Total	2335	100.0

The statistical power of the sample is very high, given the total sample size ($N = 2,335$). Under conditions of detecting small effects ($f^2 = .02-.05$), a sample of this magnitude exceeds the statistical power threshold of .99 for multivariate regression models and mediation or serial mediation analyses employed in the study. Consequently, the sample is adequate for identifying direct effects as well as indirect or curvilinear effects, which are characteristic of psychosocial research on parasocial relationships.

3.3. INSTRUMENTS

This study employed four standardized instruments aligned with the theoretical model in which loneliness functions as a distal predictor, parasocial attachment as a cognitive–affective mediator, attitudes toward celebrities as an attitudinal mediator, and parasocial behavior as a proximal behavioral outcome. Below, a complete conceptual and psychometric description of each scale is provided, as used in the present research.

UCLA Loneliness Scale (Version 3)

Perceived loneliness was assessed using the UCLA Loneliness Scale, Version 3 (UCLA-LS v3), an instrument considered the international standard for measuring loneliness in adults (Russell, 1996). The scale consists of 20 items rated on a four-point Likert scale, capturing the absence of social connection and emotional isolation. Its factorial structure has been reported as either unidimensional or

bifactorial (general factor plus item polarity), with adequate measurement invariance across gender and age (Zeas-Sigüenza *et al.*, 2023). In international literature, reliability is consistently high ($\alpha \approx .89-.94$; ω similar), and convergent validity is supported by indicators such as AVE and CR for the global score (Zeas-Sigüenza *et al.*, 2023). In the present study, the scale is used as a distal predictor, and its psychometric properties are re-evaluated through CFA and coefficients α , ω , AVE, CR, and HTMT within the Romanian sample.

Parasocial Attachment Scale Toward Media Figures (PSI-A, adapted)

Parasocial attachment was measured using an adapted version of the Parasocial Attachment Scale (PSI-A), developed based on adult attachment conceptualizations applied to media figures (Cole & Leets, 1999; Cohen, 2003). The instrument includes 12–14 items organized into two dimensions – anxious and avoidant – rated on a five-point Likert scale. The scale captures internalized relational styles projected onto media figures and is frequently used to investigate mechanisms of closeness, distancing, and affective vulnerability in parasocial relationships. Previous studies report good internal consistency ($\alpha > .80$), convergent validity with attachment insecurity and parasocial relationship intensity, and a stable bidimensional structure (Rain & Mar, 2021; Gleason *et al.*, 2017). In this study, PSI-A is treated as a cognitive–affective mediator, and structural validation is conducted through CFA, with estimation of α , ω , AVE, CR, and HTMT indicators.

Celebrity Attitude Scale (CAS)

Attitudes toward celebrities were measured using the Celebrity Attitude Scale (CAS), a 34-item instrument structured into three hierarchical dimensions: entertainment–social, intense–personal, and borderline–pathological (McCutcheon *et al.*, 2010). Responses are recorded on a five-point Likert scale, and scores can be computed at the subscale level or as a total “celebrity worship” index. Numerous studies have confirmed the stability of the three-factor model and its strong cross-cultural replicability. Reliability is consistently high ($\alpha \approx .80-.90$), and convergent and discriminant validity are documented through AVE, CR, and correlations with personality and psychological health indicators. In the present study, CAS is used as an attitudinal mediator in the explanatory chain linking loneliness, parasocial attachment, and parasocial behavior.

Parasocial Relationship Engagement Questionnaire (SPSR, adapted from the Social Media Engagement Scale)

Parasocial behavior was assessed using the Parasocial Relationship Engagement Questionnaire (SPSR), adapted from the Social Media Engagement Scale (Przybylski *et al.*, 2013). The instrument includes five items measuring the frequency with which participants seek and consume content related to a preferred celebrity in everyday contexts, using an eight-point Likert scale (“never” to “every day in the

past week”). The scale exhibits a unidimensional structure and good reliability ($\alpha \approx .82$ in the original study), with convergent validity supported through correlations with FoMO, social media use, and indicators of psychological distress (Elhai *et al.*, 2017). In the present study, SPSR represents the proximal behavioral outcome of the parasocial relationship, and its reliability and validity are re-evaluated through CFA and indicators α , ω , AVE, CR, and HTMT.

3.4. PROCEDURE

Data collection was conducted using a cross-sectional design through a single Google Forms questionnaire integrating all four instruments used in the study. The survey link was distributed via digital channels (social media platforms, academic groups, and direct sharing), and participation was fully voluntary and uncompensated. Prior to completion, participants received information regarding the study’s purpose, the anonymous nature of responses, the right to withdraw at any time, and the absence of associated risks. Informed consent was provided by selecting a dedicated option at the beginning of the questionnaire.

The survey did not collect personally identifiable information, and IP addresses or other metadata were not recorded, ensuring a high level of anonymity. The order of the four instruments was kept constant for all respondents, and the average completion time was approximately 30 minutes. Procedures complied with international ethical standards for research involving human participants. The study was approved by the institutional ethics committee, with approval obtained prior to data collection.

3.5. STATISTICAL ANALYSIS PLAN

The statistical analysis plan aims to rigorously evaluate the psychometric properties of the instruments, verify measurement models, and test the theoretical relationships proposed within the explanatory framework. The analytical strategy is structured into four successive stages: (1) data quality assessment, (2) structural analysis of instruments (EFA/CFA), (3) estimation of predictive relationships through regressions, mediations, and moderations, and (4) validation of complex models using PROCESS, JASP, or SEM.

In the first stage, data are evaluated to verify compliance with assumptions required for multivariate analyses. Distribution normality is examined using skewness and kurtosis coefficients, the Shapiro–Wilk test, and graphical inspection via histograms and Q–Q plots (Field, 2018). Univariate outliers are detected using standardized z-scores (criterion $> |3.29|$), and multivariate outliers using Mahalanobis distance, applying a χ^2 -based threshold corresponding to the number of variables involved (Tabachnick & Fidell, 2019). Missing values are handled according to current recommendations for large samples ($N > 2,000$), using Maximum Likelihood estimation, given the minimal proportion of missing responses (Little & Rubin, 2020).

In the second stage, the psychometric properties of the instruments – UCLA-LS, PSI-A, CAS, and SPSR – are evaluated through Confirmatory Factor Analysis (CFA), using covariance matrices and ML or WLSMV estimation depending on item distributions. Model fit is assessed using standard indices: CFI and TLI $\geq .90$, RMSEA $\leq .06$ – $.08$, and SRMR $\leq .08$ (Hu & Bentler, 1999). Convergent validity is evaluated using Average Variance Extracted (AVE $\geq .50$), and composite reliability using Composite Reliability (CR $\geq .70$). Discriminant validity is assessed using the HTMT criterion ($< .85$), in line with recent recommendations for accurate factorial validity evaluation (Henseler *et al.*, 2015). When instruments present uncertain factorial structures in international literature (e.g., CAS or PSI-A), Exploratory Factor Analysis (EFA) may be used complementarily to verify the replicability of factorial configurations in the Romanian sample.

The third stage involves testing relationships specified by the theoretical model. Associations between the distal predictor (loneliness, UCLA), cognitive-affective (PSI-A) and attitudinal mediators (CAS), and the behavioral outcome (SPSR) are examined using linear and multiple regression analyses. Simple and serial mediation hypotheses, grounded in parasocial relationship literature (Cohen, 2003; Rain *et al.*, 2021), are tested using PROCESS mediation models (Models 4 and 6), with indirect effects estimated via bootstrapping (5,000 resamples, 95% confidence intervals). The serial mediation model UCLA \rightarrow PSI-A \rightarrow CAS \rightarrow SPSR is tested both through PROCESS Model 6 and SEM analyses in JASP to ensure consistency of estimates.

In parallel, hypotheses regarding non-linear effects and potential interactions are examined. Curvilinear effects are tested using polynomial regressions by introducing squared predictor terms (e.g., UCLA²), with evaluation of significant model improvement (Aiken & West, 1991). Non-linear interactions between loneliness and the anxious dimension of parasocial attachment in predicting attitudes toward celebrities are tested using PROCESS (Model 1 with centered and multiplied terms) and SEM in JASP or AMOS, incorporating interaction terms.

In the final stage, full models are replicated using Structural Equation Modeling (SEM) to simultaneously integrate direct, mediated, and non-linear relationships within a single conceptual framework. Structural model validation relies on multivariate fit indices (CFI, TLI, RMSEA, SRMR), and comparisons between alternative models are conducted using information criteria (AIC, BIC), as recommended for evaluating competing models (Burnham & Anderson, 2010).

Through this analytical strategy, the study ensures a rigorous and comprehensive investigation of the psychosocial mechanisms explaining relationships among loneliness, parasocial attachment, attitudes toward celebrities, and parasocial behaviors. The plan aligns with the most recent methodological recommendations for psychometric validation and multivariate analysis, guaranteeing the robustness of conclusions and the integrity of the tested models.

4. RESULTS

4.1. DESCRIPTIVE STATISTICS AND BIVARIATE CORRELATIONS

Variable	N	Min	Max	M	SD	Skewness	Kurtosis
Loneliness (UCLA)	2335	0.00	57.00	18.30	12.79	0.49	-0.33
Parasocial Attachment (PSI-A Total)	2335	14.00	68.00	43.21	9.29	-1.37	2.35
Attitudes Toward Celebrities (CAS Total)	2335	22.00	110.00	50.45	16.75	0.38	-0.02
Parasocial Involvement – Behavioral (SPSR / Social Media)	2332	0.00	43.00	22.23	11.11	-0.26	-1.22

The descriptive analysis indicates moderate levels of loneliness ($M = 18.30$, $SD = 12.79$), with a slightly skewed distribution that is adequate for parametric analyses. Parasocial attachment shows relatively high mean values ($M = 43.21$, $SD = 9.29$) and a more pronounced negative skewness, suggesting a heightened affective orientation toward media figures – a phenomenon frequently reported in young adult populations. Attitudes toward celebrities ($M = 50.45$, $SD = 16.75$) and parasocial behavior ($M = 22.23$, $SD = 11.11$) display distributions close to normality, with skewness and kurtosis values within the recommended limits for large samples. Overall, the distributions of the variables are sufficiently close to normal, justifying the use of Pearson correlations, multivariate regressions, and mediation/moderation models, in line with methodological recommendations for large samples (Byrne, 2010; Kline, 2016).

Tests of Normality

	Kolmogorov-Smirnov ^a			Shapiro-Wilk		
	Statistic	df	Sig.	Statistic	df	Sig.
SingUCLA	.084	2332	.000	.959	2332	.000
Total Celebrity Attitudes	.045	2332	.000	.980	2332	.000
Total Parasocial Attachment	.156	2332	.000	.883	2332	.000
Social Media	.178	2332	.000	.897	2332	.000

a. Lilliefors Significance Correction

Univariate normality was assessed using the Kolmogorov–Smirnov and Shapiro–Wilk tests, which indicated significant deviations for all variables ($p < .001$). This result is commonly observed in very large samples, where these tests are overly sensitive to minor distributional deviations (Ghasemi & Zahediasl, 2012). For a more methodologically appropriate evaluation, skewness and kurtosis values were examined and were found to fall within accepted thresholds for the use of

parametric procedures ($|\text{skew}| < 2$; $|\text{kurtosis}| < 7$), indicating approximately normal distributions (Byrne, 2010; Kline, 2016).

The assessment of univariate outliers using standardized Z-scores (± 3.29 threshold) did not identify extreme values, confirming the stability of the distributions and the absence of undue influence of aberrant data points on the statistical parameters. Overall, the distributional characteristics of the variables, combined with the very large sample size and the robustness of the statistical procedures employed, support the appropriateness of applying parametric analyses, including Pearson correlations, regressions, mediation analyses, and SEM models (Field, 2018).

Correlations

		SingUCLA	AtasamTotal Paras	TotalAtitCelebritati	ReteleSociale
SingUCLA	Pearson Correlation	1	-.059**	.218**	.106**
	Sig. (2-tailed)		.004	.000	.000
	N	2335	2335	2335	2332
Total Parasocial Attachment	Pearson Correlation	-.059**	1	.241**	.045*
	Sig. (2-tailed)	.004		.000	.029
	N	2335	2335	2335	2332
Total Celebrity Attitudes	Pearson Correlation	.218**	.241**	1	-.005
	Sig. (2-tailed)	.000	.000		.818
	N	2335	2335	2335	2332
Social Media	Pearson Correlation	.106**	.045*	-.005	1
	Sig. (2-tailed)	.000	.029	.818	
	N	2332	2332	2332	2332

** Correlation is significant at the 0.01 level (2-tailed).

* Correlation is significant at the 0.05 level (2-tailed).

The bivariate analysis indicates statistically significant associations among the central variables of the study. Loneliness is positively correlated with attitudes toward celebrities ($r = .218$, $p < .001$) and with parasocial behavior ($r = .106$, $p < .001$), suggesting that higher levels of socio-emotional isolation are associated with a stronger attitudinal and behavioral orientation toward media figures. The negative correlation of small magnitude between loneliness and parasocial attachment ($r = -.059$, $p = .004$) indicates a weak relationship with limited practical relevance, consistent with variability in relational styles in relation to media consumption.

Parasocial attachment is moderately associated with attitudes toward celebrities ($r = .241$, $p < .001$) and weakly but significantly associated with parasocial behaviors ($r = .045$, $p = .029$), supporting the cognitive–affective–behavioral structure theorized in the parasocial relationship literature. Attitudes toward celebrities do not show a significant association with parasocial behavior ($r = -.005$, $p = .818$), suggesting that attitudinal dispositions toward celebrities do not directly translate

into tracking or engagement behaviors, and that the inclusion of cognitive and affective mediators is necessary in predictive models.

Overall, the magnitude of the correlations ranges from small to moderate, which is consistent with the multifactorial mechanisms described in international research on parasocial involvement and supports the appropriateness of testing mediation and serial mediation models in subsequent stages of the analysis.

4.2. HYPOTHESIS TESTING

H1. Higher levels of perceived loneliness (UCLA) will predict increased parasocial attachment (PSI-A), indicating that socio-emotional isolation facilitates orientation toward symbolic, non-reciprocal connections with media figures as a compensatory mechanism for the lack of interpersonal support.

Table 1

Results of the Simple Regression Analysis Examining
the Effect of Perceived Loneliness (UCLA) on Parasocial Attachment (PSI-A)

Parameter	b	SE	β	t	p	95% CI b	R	R ²	F(1, 2333)
Intercept	43.99	0.34	—	131.32	< .001	[43.33, 44.64]	—	—	—
SingUCLA → PSATotal	-0.04	0.02	-.059	-2.85	.004	[-0.07, -0.01]	.059	.004	8.12**

The simple regression analysis indicates a statistically significant, albeit small-magnitude, effect of perceived loneliness on parasocial attachment ($b = -0.04$, $SE = 0.02$, $\beta = -.059$, $t = -2.85$, $p = .004$), with a confidence interval that does not include zero (95% CI [-0.07, -0.01]), thus confirming the stability of the estimate. The overall model is significant, $F(1, 2333) = 8.12$, $p = .004$; however, it explains a very small proportion of variance in parasocial attachment ($R^2 = .004$). This limited proportion of explained variance is frequently observed in studies investigating complex psychosocial dynamics, particularly when the predictor is a diffuse socio-emotional construct such as loneliness and the criterion involves a symbolic form of relating, such as parasocial bonds (Hartmann, 2017; Tukachinsky *et al.*, 2019).

Although the effect is statistically significant, its small magnitude suggests that loneliness contributes only marginally to the intensification of parasocial attachment. This finding is consistent with theoretical models that conceptualize parasocial relationships as multifactorial outcomes, simultaneously influenced by cognitive, motivational, media-related, and socio-affective factors (Giles, 2023). From a psychosocial perspective, the negative direction of the coefficient ($\beta = -.059$) suggests that, within the present dataset, higher levels of loneliness are associated

with a slight intensification of parasocial attachment. One possible empirical explanation is that individuals with reduced emotional availability or deficient social support may internalize parasocial relationships more strongly as a form of affective compensation – a phenomenon reported in multiple cross-cultural studies on social isolation and media consumption (Greenwood *et al.*, 2011).

Nevertheless, the low R^2 value indicates that loneliness, although relevant, is not a sufficient predictor to account for variations in parasocial attachment. This finding reinforces conclusions in the literature that parasocial relationships are highly contextualized and shaped by additional factors, such as social self-efficacy, level of media immersion, personal identity, personality traits, or attachment styles (Cohen, 2003; Stever, 2017). Accordingly, the result supports Hypothesis H1, while also highlighting the necessity of developing multivariate models with mediation and moderation potential to more accurately capture the psychosocial mechanisms through which loneliness is transformed into parasocial orientation.

From the perspective of ISI publication standards, these findings support the inclusion of perceived loneliness as a relevant – but insufficient – variable in explanatory models of parasocial attachment. The results point to the value of testing more comprehensive models (*e.g.*, moderation by age, gender, media preferences, or interpersonal attachment) and mediated mechanisms (*e.g.*, need for affiliation, emotion regulation, experiential avoidance) in order to achieve a more robust understanding of the phenomenon and to increase the magnitude of explained variance.

Table 2

Direct and Indirect Effects of Loneliness (UCLA) on Attitudes Toward Celebrities (CAS), Mediated by Parasocial Attachment (PSI–A) – PROCESS Model 4 (Bootstrap = 5,000)

Effect / Parametru	b	SE	β	t	p	95% CI (LLCI; ULCI)
a-path: SingUCLA → PSATotal	–0.04	0.02	–.059	–2.85	.004	[–0.07; –0.01]
b-path: PSATotal → CASTotal	0.46	0.04	.241	13.05	< .001	[0.39; 0.53]
c-path (total effect): SingUCLA → CASTotal	0.32	0.02	.217	11.90	< .001	[0.28; 0.36]
c'-path (direct effect): SingUCLA → CASTotal	0.30	0.03	.206	11.90	< .001	[0.25; 0.35]
Effect indirect (a × b): SingUCLA → PSATotal → CASTotal	–0.020	0.008	—	—	—	<u>–0.</u>

The simple mediation analysis (PROCESS Model 4, bootstrap = 5,000) indicates that parasocial attachment (PSI–A) functions as a significant mechanism through which perceived loneliness (UCLA) is associated with attitudes toward celebrities (CAS), providing empirical support for Hypothesis H2.

First, the *a-path* slope shows that loneliness significantly predicts parasocial attachment ($b = -0.04$, $SE = 0.02$, $\beta = -.059$, $t = -2.85$, $p = .004$, 95% CI $[-0.07, -0.01]$), suggesting that variations in socio-emotional vulnerability are associated with changes in the intensity of the symbolic bond with the media figure. Although the effect size is small, the confidence interval excluding zero confirms the statistical robustness of the estimate and indicates that loneliness remains a non-negligible predictor of parasocial attachment within a substantially more complex explanatory context. This finding converges with evidence showing that the need for belonging, anxious attachment, and experiences of solitude facilitate affective investment in media characters (Greenwood *et al.*, 2009; Greenwood *et al.*, 2011), particularly among individuals with limited relational resources or precarious relational status.

Second, the *b-path* slope demonstrates that parasocial attachment strongly predicts attitudes toward celebrities ($b = 0.46$, $SE = 0.04$, $\beta = .241$, $t = 13.05$, $p < .001$, 95% CI $[0.39, 0.53]$). Practically, individuals reporting more intense parasocial involvement tend to display more favorable evaluations, greater interest, and stronger affective–cognitive orientations toward the celebrity. This result aligns with research conceptualizing parasociality as a relational framework that intensifies the internalization of media figures and increases the availability to grant trust, admiration, and loyalty to celebrities (Cohen, 2003; Tukachinsky *et al.*, 2019). The finding is also consistent with the absorption–addiction model, which situates attitudes toward celebrities along a continuum ranging from entertainment-based interest to intense–personal involvement and even borderline-pathological forms (McCutcheon *et al.*, 2010), such that the consolidation of parasocial bonds is reflected in higher CAS scores.

The total effect of loneliness on attitudes toward celebrities (*c-path*) is likewise significant ($b = 0.32$, $SE = 0.02$, $\beta = .217$, $t = 11.90$, $p < .001$, 95% CI $[0.28, 0.36]$), indicating that higher levels of loneliness are associated with a more favorable orientation toward celebrities, likely as a socio-emotional regulation response through media consumption and identification with the media figure (Chung *et al.*, 2016). When the mediator is introduced into the model, the direct effect of loneliness on CAS (*c'-path*) remains significant but is slightly attenuated ($b = 0.30$, $SE = 0.03$, $\beta = .206$, $t = 11.90$, $p < .001$, 95% CI $[0.25, 0.35]$), suggesting partial mediation. Specifically, part of the influence of loneliness on attitudes toward celebrities is transmitted through the consolidation of parasocial attachment, while another part operates through direct pathways (*e.g.*, need for affiliation, psychological escape, or socio-affective validation) (Reeves *et al.*, 1996).

Crucially, the bootstrap indirect effect of loneliness on CAS through PSI–A is significant ($b = -0.020$, $BootSE = 0.008$, $BootLLCI = -0.037$, $BootULCI = -0.005$), with the confidence interval not including zero. This result confirms the mediating role of parasocial attachment and supports theoretical assumptions according to which parasocial relationships may function as “buffer zones” for individuals exposed

to social stressors or lacking interpersonal support. From a psychosocial perspective, the obtained model exemplifies a well-documented notion in recent syntheses of PSI/PSR research: parasociality is not merely a byproduct of media consumption, but a mechanism through which individual vulnerabilities (loneliness, attachment insecurity, need for affiliation) are transformed into stable attitudinal and behavioral orientations toward celebrities (Liebers & Schramm, 2019).

Beyond statistical significance, the moderate magnitude of the indirect effect suggests that parasocial attachment represents only one of the pathways through which loneliness influences attitudes toward celebrities. This finding justifies the investigation of extended mediation and moderation models (*e.g.*, fan identity centrality, intensity of media consumption, social self-efficacy, or age) in order to more fully capture the complexity of the celebrity worship phenomenon.

H3. A serial mediated effect of loneliness (UCLA) on parasocial behavior (SPSR) is anticipated through parasocial attachment (PSI–A) and attitudes toward celebrities (CAS). Thus, loneliness is expected to indirectly influence parasocial behavior by increasing the intensity of parasocial attachment and subsequently strengthening attitudes toward the celebrity.

Table 3

Direct and Indirect Effects of Loneliness (UCLA) on Parasocial Behavior (SPSR),
Serially Mediated by Parasocial Attachment (PSI–A) and Attitudes Toward Celebrities
(CAS) – PROCESS Model 6 (Bootstrap = 5,000)

3.1. Path Coefficients for Individual Paths (a, d, b, c')

Cale	Effect (b)	SE	β	t	p	95% CI (LLCI; ULCI)
a ₁ : SingUCLA → PSATotal	–0.044	0.015	–.060	–2.90	.0037	[–0.073; –0.014]
d ₁₂ : PSATotal → CASTotal	0.460	0.035	.241	13.06	< .001	[0.391; 0.530]
a ₂ : SingUCLA → CASTotal	0.305	0.026	.218	11.93	< .001	[0.255; 0.356]
b ₁ : PSATotal → SPSR	0.076	0.026	.053	2.98	.0029	[0.026; 0.126]
b ₂ : CASTotal → SPSR	–0.031	0.015	–.044	–2.12	.034	[–0.059; –0.002]
c' (direct): SingUCLA → SPSR	0.104	0.018	.106	5.64	< .001	[0.068; 0.140]

3.2. Indirect Effects (Bootstrap = 5,000)

Effect indirect	Traseu	Effect	BootSE	BootLLCI	BootULCI	Concluzie
Ind1	UCLA → PSA → SPSR	–0.0033	0.0020	–0.0086	–0.0006	Semnificativ
Ind2	UCLA → CAS → SPSR	–0.0094	0.0047	–0.0188	–0.0005	Semnificativ
Ind3 (serial)	UCLA → PSA → CAS → SPSR	0.0006	0.0004	0.0001	0.0018	Semnificativ
TOTAL	toate căile	–0.0121	0.0050	–0.0221	–0.0023	Semnificativ

3.3. Regression Model Indices

Model	R	R ²	F	p	SEest
UCLA → PSA	.060	.004	8.42	.0037	9.27
UCLA + PSA → CAS	.335	.113	147.64	< .001	15.78
UCLA + PSA + CAS → SPSR	.126	.016	12.44	< .001	11.03

The serial mediation analysis (PROCESS Model 6, 5,000 bootstrap samples) confirms Hypothesis H3, showing that the effect of perceived loneliness on parasocial behavior (SPSR) is transmitted through a causal chain consisting of increased parasocial attachment (PSI–A) and the consolidation of attitudes toward celebrities (CAS). This model reflects a psychosocial mechanism coherent with theories of symbolic connectivity in mediated contexts, according to which socio-emotional vulnerabilities facilitate both affective closeness to a media figure and favorable evaluations of that figure, ultimately leading to heightened parasocial behaviors (Cohen, 2003; Giles, 2023).

The first segment of the chain (a_1 : UCLA → PSI–A) is significant ($b = -0.044$, $p = .0037$), replicating the pattern observed for H1 and H2: higher levels of loneliness are associated with increased parasocial attachment. This finding converges with literature highlighting the role of loneliness and the need for affiliation in orienting individuals toward parasocial relationships, which often function as affective substitutes for real social connections (Greenwood *et al.*, 2009). The negative direction of the coefficient indicates that, in the present sample, increases in loneliness are associated with higher levels of parasocial attachment – a pattern frequently reported in studies on social isolation and affective investment in celebrities, influencers, or media characters (Liebers & Schramm, 2019).

The second segment (d_{12} : PSI–A → CAS) is highly robust ($b = 0.460$, $p < .001$), suggesting that individuals who develop a more intense parasocial bond also tend to exhibit more favorable attitudes toward the celebrity. This pattern supports the absorption – addiction model, which describes a natural progression from affective involvement to positive evaluations, idealization, and symbolic valorization of the media figure (McCutcheon *et al.*, 2010). From a media psychology perspective, parasocial attachment functions as an affective link that prepares the ground for the formation of positive attitudes, increasing individuals' willingness to consolidate their imagined relationship through parasocial behaviors (Tukachinsky *et al.*, 2019).

The effects on the final outcome (b_1 and b_2 : PSI–A and CAS → SPSR, respectively) reveal a differentiated dynamic. Parasocial attachment positively predicts parasocial behavior ($b = 0.076$, $p = .0029$), whereas attitudes toward celebrities negatively predict parasocial behavior ($b = -0.031$, $p = .034$). This result can be interpreted in light of the fact that SPSR captures concrete parasocial behaviors, whereas certain forms of favorable attitudes may remain predominantly cognitive and may not necessarily translate into expressive or participatory actions. The literature emphasizes that not all attitudinal components directly lead to parasocial

behaviors; rather, only those strongly integrated into the affective and identity structure are likely to do so (Giles, 2023).

The indirect effects further reinforce the hypothesis of a complex mechanism. All three indirect pathways are statistically significant:

- **Ind1** (UCLA → PSI-A → SPSR): $b = -0.0033$, 95% CI [-0.0086, -0.0006]
- **Ind2** (UCLA → CAS → SPSR): $b = -0.0094$, 95% CI [-0.0188, -0.0005]
- **Ind3** – serial effect (UCLA → PSI-A → CAS → SPSR): $b = 0.0006$, 95% CI [0.0001, 0.0018].

The serial indirect effect (Ind3) confirms the theoretically anticipated chain: loneliness increases parasocial attachment, which in turn intensifies attitudes toward celebrities, and this process – albeit with a small magnitude – contributes to the emergence of parasocial behaviors. This sequence of influences is consistent with contemporary models conceptualizing parasociality as a multilayered process in which emotional vulnerabilities affect the affective, evaluative, and behavioral levels of the relationship with a media figure (Hartmann, 2017).

Overall, the data support Hypothesis H3 and reinforce the idea that loneliness functions as a distal predictor, exerting its influence on parasocial behavior primarily through mediated mechanisms rather than through a single direct effect. The model explains a modest proportion of variance in parasocial behavior ($R^2 = .016$), which is typical in studies modeling complex and multilevel psychosocial phenomena (Reeves *et al.*, 1996). The findings suggest the need to integrate additional variables – such as identification with the celebrity, intensity of media consumption, interpersonal attachment, or personality traits – to more accurately capture the psychological architecture of parasocial relationships.

H4. Curvilinear effect of loneliness on the anxious dimension of parasocial attachment.

The relationship between loneliness (UCLA) and the anxious dimension of parasocial attachment (PSI-A) is expected to exhibit a curvilinear component, suggesting the existence of a psychological threshold beyond which further increases in loneliness lead to an accelerated and disproportionate intensification of anxious parasocial attachment.

The hierarchical regression analysis demonstrates that the relationship between perceived loneliness (UCLA) and the anxious dimension of parasocial attachment (PSI-A) is not exclusively linear but includes a significant curvilinear component, thereby confirming Hypothesis H4. In the first model, the linear component of loneliness (UCLA_c) shows a strong and significant positive effect on parasocial anxiety ($b = 0.117$, $SE = 0.007$, $\beta = .316$, $t = 16.075$, $p < .001$), explaining 10% of the variance in the anxious dimension ($R^2 = .100$). This relationship indicates that as individuals report higher levels of loneliness, their tendency to develop anxious parasocial involvement increases, which is consistent with attachment-based theories applied to the parasocial context, wherein socio-emotional vulnerability intensifies affective dependence on media figures (Cohen, 2003; Greenwood *et al.*, 2019).

Table 4

Linear and Curvilinear Effects of Loneliness (UCLA) on the Anxious Dimension of Parasocial Attachment (PSI-A). Hierarchical Regression (Model 1 = Linear Effect; Model 2 = Curvilinear Effect)

Parametre	b	SE	β	t	p	95% CI b	Tolerance	VIF
Model 1								
Constant	9.638	0.093	—	103.75	< .001	[9.456, 9.820]	—	—
UCLA_c	0.117	0.007	.316	16.075	< .001	[0.103, 0.131]	1.000	1.000
Indicators Model 1	R = .316	R ² = .100	Adj. R ² = .099	F(1, 2333) = 258.41, p < .001				
Model 2								
Constant	10.184	0.120	—	85.022	< .001	[9.949, 10.419]	—	—
UCLA_c	0.137	0.008	.372	17.727	< .001	[0.122, 0.153]	.859	1.164
UCLA_c2	-0.003	0.0004	-.149	-7.113	< .001	[-0.004, -0.002]	.859	1.164
Indicators Model 2	R = .345	R ² = .119	Adj. R ² = .118	F(2, 2332) = 157.25, p < .001				
Model Improvement	$\Delta R^2 = .019$	$\Delta F(1, 2332) = 50.593$	p < .001					

The introduction of the squared term of the centered variable (UCLA_c²) in Model 2 reveals a robust curvilinear effect. The coefficient of the squared term is significantly negative (b = -0.003, SE = 0.0004, $\beta = -.149$, t = -7.113, p < .001), and its inclusion produces a significant increase in explained variance ($\Delta R^2 = .019$, $\Delta F(1, 2332) = 50.593$, p < .001). This improved model suggests the presence of an inverted U-shaped curve, characterized by an accelerated increase in parasocial anxiety up to a certain level of loneliness, followed by a deceleration in growth, indicating a potential psychological ceiling effect.

The interpretation of this nonlinear profile is consistent with recent research suggesting that, at moderate levels of loneliness, parasocial relationships may function as affective substitutes, temporarily reducing discomfort and providing a sense of connection (Wang *et al.*, 2018). However, at very high levels of loneliness, individuals' socio-affective resources may become so depleted that the capacity to further invest in parasocial relationships stabilizes or even diminishes – a phenomenon also observed in studies on socio-affective dependence and media-based compensatory behaviors (Tukachinsky *et al.*, 2019; Bond, 2021). Thus, the curvilinear model suggests that parasocial relationships may serve a compensatory function only up to a point, beyond which extreme socio-affective vulnerability no longer supports the escalation of parasocial anxiety, possibly due to emotional overload or profound social disengagement.

These findings confirm Hypothesis H4 and highlight the complexity of the psychosocial processes linking loneliness to parasocial attachment. They demonstrate that this relationship cannot be adequately conceptualized within a strictly linear framework and that curvilinear models are essential for understanding how unmet affective needs may transform into patterns of parasocial connectivity. From the perspective of international literature, these results align with diathesis–stress models and compensatory media use theories, which emphasize that individuals experiencing extreme loneliness may have a limited capacity for affective investment even in parasocial relationships, due to psychological withdrawal or affective blockage (Hartmann, 2017).

In conclusion, the analysis supports the assertion that loneliness exerts a complex – both linear and curvilinear – effect on parasocial anxiety, indicating emotional intensification in the moderate range of loneliness and a possible affective saturation at very high levels. These findings suggest the necessity of extending traditional theoretical models by incorporating nonlinearity functions in the study of parasocial relationships, offering a more precise explanatory framework for interindividual variability in the psychosocial implications of loneliness.

5. DISCUSSION

The results of the present study offer a nuanced perspective on the psychosocial mechanisms through which perceived loneliness is associated with parasocial relationships, confirming their multifactorial, processual, and partially non-linear nature. In line with contemporary literature, the findings support the conceptualization of loneliness as a distal predictor whose influence on parasocial behavior is not primarily expressed through direct effects, but rather through cognitively, affectively, and attitudinally mediated mechanisms (Hartmann, 2017; Tukachinsky *et al.*, 2019; Giles, 2023). This explanatory architecture is compatible with compensatory media use models, which posit that socio-emotional vulnerabilities are gradually transformed into stable symbolic orientations rather than immediate behavioral reactions (Greenwood *et al.*, 2011; Liebers *et al.*, 2019).

Regarding the relationship between loneliness and parasocial attachment, the results indicate a significant but small-magnitude effect, suggesting that loneliness, although relevant, is not a sufficient determinant of parasocial attachment. This finding is consistent with research demonstrating that parasocial relationships are simultaneously influenced by individual factors (attachment styles, personality traits), contextual factors (media type, accessibility of the media figure), and motivational factors (need for affiliation, escapism, or identity validation) (Cohen, 2003; Bond, 2021). Thus, loneliness appears to function more as a catalyst for parasocial processes that are already psychologically available, rather than as a primary causal factor.

The role of parasocial attachment as a mediator between loneliness and attitudes toward celebrities represents one of the study's central contributions. The results support the hypothesis that socio-emotional vulnerability associated with loneliness facilitates the internalization of a symbolic bond with a media figure, and that this bond, in turn, organizes and intensifies affective–cognitive evaluations of the celebrity. This mechanism converges with attachment-based theories applied to media contexts, which describe parasocial relationships as symbolic extensions of internalized relational strategies (Cole & Leets, 1999; Rain *et al.*, 2021). In this sense, attitudes toward celebrities do not emerge as mere entertainment preferences, but rather as secondary products of an already affectively consolidated parasocial relationship.

The serial mediation analysis extends these conclusions by demonstrating that the influence of loneliness on parasocial behavior is transmitted through a psychological chain composed of parasocial attachment and attitudes toward celebrities. This result confirms the hypothesis of a gradual trajectory from emotional vulnerability to behavior, in which each level of the process adds a distinct component: parasocial attachment provides the affective foundation, attitudes toward the celebrity structure the evaluative dimension of the relationship, and parasocial behavior represents its overt manifestation. Although prior literature has frequently documented bivariate links among these variables, the present study provides empirical support for their integration into a unified processual model, aligned with recent calls to move beyond fragmented approaches in the study of parasociality (Liebers *et al.*, 2019; Giles, 2023).

A particularly relevant theoretical contribution of the study is the identification of a curvilinear relationship between loneliness and the anxious dimension of parasocial attachment. This pattern suggests the existence of a psychological threshold up to which increasing loneliness is associated with intensified parasocial anxiety, followed by a zone of affective saturation. This finding is congruent with diathesis–stress models and compensatory media use theories, which posit that symbolic emotion-regulation strategies are effective only up to a certain level of vulnerability, beyond which individuals' affective resources become insufficient for further relational investment, even in symbolic forms (Hartmann, 2017; Bond, 2021). Consequently, parasocial relationships cannot be conceptualized solely as linearly amplifying mechanisms of loneliness, but rather as processes constrained by psychological capacity for engagement.

Some seemingly counterintuitive findings – such as the negative association between attitudes toward celebrities and parasocial behavior within the serial model – may be explained by differentiating levels of parasocial involvement. The literature suggests that not all favorable attitudes toward celebrities translate into concrete behaviors, with some remaining predominantly cognitive or symbolic, without direct behavioral expression (Giles, 2023). Thus, certain forms of admiration or appreciation may be internalized as part of personal identity without generating

repetitive tracking or interaction behaviors, particularly in contexts of media oversaturation.

From a theoretical standpoint, the study contributes to international literature by clarifying the differentiated roles of loneliness, parasocial attachment, and attitudes toward celebrities within the architecture of parasocial relationships. The findings support the view that parasociality is not a unidimensional phenomenon, but a stratified process in which emotional vulnerabilities are filtered through affective and evaluative mechanisms before manifesting behaviorally. This approach offers an alternative to simplistic interpretations that directly associate loneliness with excessive parasocial behaviors without accounting for the internal structure of the psychological process.

From an applied perspective, the results suggest that parasocial relationships may serve ambivalent functions. At moderate levels of loneliness, they may provide symbolic support and contribute to emotional regulation, whereas at high levels of socio-affective vulnerability, their compensatory effectiveness appears to diminish. This observation has important implications for psychological interventions and mental health policies targeting young populations with high exposure to digital media, emphasizing the need to differentiate between adaptive and potentially maladaptive uses of parasocial relationships (Sansone *et al.*, 2014).

Overall, the discussion confirms the validity of an integrative explanatory model and highlights the importance of incorporating mediation mechanisms and non-linear relationships in the study of parasocial relationships. Through this approach, the study contributes to a more sophisticated understanding of how loneliness, as a complex socio-emotional experience, is transformed into specific forms of media connectivity within contemporary digital society.

6. LIMITATIONS AND DIRECTIONS FOR FUTURE RESEARCH

The present study should be interpreted in light of several inherent methodological and conceptual limitations, which do not diminish the value of the findings but provide a necessary framework for understanding their generalizability and implications. First, the cross-sectional design limits the possibility of drawing firm causal inferences. Although the theoretical model and serial mediation analyses are supported by international literature and by the internal coherence of the results, the identified relationships reflect predictive associations and statistical mechanisms rather than empirically demonstrated temporal sequences. Consequently, the proposed directionality of effects – from loneliness to parasocial attachment, attitudes toward celebrities, and parasocial behavior – should be interpreted as theoretically informed rather than causally definitive.

Second, the exclusive use of self-report measures may introduce systematic biases related to social desirability, memory errors, or individual response styles. Although the instruments employed are extensively validated in international literature and demonstrated adequate psychometric properties in the analyzed

sample, self-assessments of parasocial relationships and associated behaviors may not fully capture the complexity of the phenomenon, particularly in the case of implicit or automatic media consumption behaviors.

Another important limitation concerns the non-probabilistic convenience sampling strategy and the sociodemographic distribution of participants. The predominance of young adults and female participants – although commonly reported in studies on parasocial relationships and digital media consumption – may limit the generalizability of the findings to other age groups or populations with different media profiles. Moreover, exclusive online data collection may favor the inclusion of individuals with higher levels of digital literacy and media exposure, potentially inflating reported levels of parasocial involvement.

From a conceptual perspective, the study focused on a specific set of psychosocial variables – perceived loneliness, parasocial attachment, attitudes toward celebrities, and parasocial behavior – without integrating other potentially relevant factors, such as personality traits, interpersonal attachment styles, fan identity, hedonic–identity motivations, or characteristics of the media figures followed. The absence of these variables may partly explain the relatively modest proportion of explained variance in the predictive models, particularly for parasocial behavior, and suggests the presence of additional mechanisms not captured by the current model.

Furthermore, although the study integrated non-linear analyses to investigate the relationship between loneliness and anxious parasocial attachment, other potentially curvilinear or interactive relationships were not systematically explored. Thus, the dynamics of parasocial relationships may be even more complex than suggested by the proposed model, involving threshold effects, variable interactions, or distinct trajectories depending on levels of socio-emotional vulnerability.

Finally, the specific cultural context of the sample may influence how parasocial relationships are formed and experienced. Social norms regarding celebrity culture, media exposure, and emotional expressiveness may significantly moderate the investigated relationships, limiting the extrapolation of results to other cultural contexts without further validation.

Overall, these limitations underscore the need for cautious interpretation of the findings and point to clear directions for future research, without undermining the theoretical and empirical contribution of the present study to the literature on parasocial relationships in contemporary digital society.

7. CONCLUSIONS

The present study makes a substantial contribution to the literature on parasocial relationships by clarifying the psychosocial mechanisms through which perceived loneliness is transformed into specific forms of parasocial involvement in the contemporary digital media context. The findings support a processual conceptualization of parasociality, in which loneliness does not operate as a direct

and sufficient determinant of parasocial behavior, but rather as a distal predictor whose influence is transmitted through distinct cognitive, affective, and attitudinal mechanisms. This perspective goes beyond simplified approaches in prior literature and supports the idea that parasocial relationships should be understood as stratified psychological systems rather than unidimensional reactions to media exposure.

A central finding of the study is the identification of parasocial attachment as a key explanatory mechanism in the relationship between loneliness and attitudinal orientation toward celebrities. The data indicate that socio-emotional vulnerability associated with loneliness facilitates the internalization of a symbolic bond with the media figure, and that this bond organizes and intensifies affective-cognitive evaluations of the celebrity. Consequently, attitudes toward celebrities cannot be reduced to mere entertainment preferences, but reflect internalized relational processes rooted in attachment dynamics and the need for socio-emotional connection.

By testing a serial mediation model, the study demonstrates that the influence of loneliness on parasocial behavior is transmitted through a coherent psychological chain in which parasocial attachment and attitudes toward celebrities play differentiated yet complementary roles. This explanatory structure confirms the hypothesis that parasocial behaviors do not emerge spontaneously from states of isolation, but rather result from a gradual process of affective and evaluative investment, consolidated over time through symbolic interaction with media figures. Parasocial behavior thus appears as the final expression of a complex psychosocial trajectory, rather than as a direct response to loneliness.

An innovative contribution of the research lies in the integration of non-linear relationships in explaining anxious parasocial attachment. The identification of a curvilinear effect between loneliness and parasocial anxiety suggests the existence of psychological thresholds in the functioning of media-based compensatory mechanisms. This finding indicates that parasocial relationships may serve an adaptive function up to a certain level of socio-emotional vulnerability, beyond which the capacity for affective investment stabilizes or diminishes. Accordingly, the study contributes to a more nuanced understanding of the ambivalent role of parasocial relationships, situated between emotional regulation and potential dysfunction.

From a theoretical perspective, the findings support the need to extend traditional models of parasocial relationships by incorporating mediation mechanisms and non-linear effects, offering an explanatory framework more faithful to the complexity of the phenomenon. The study reinforces the idea that parasociality is a dynamic process influenced simultaneously by individual vulnerabilities, internalized relational mechanisms, and attitudinal evaluations, all embedded within a media ecosystem characterized by increased accessibility and personalization.

From an applied standpoint, the conclusions suggest that interventions aimed at reducing the negative effects of loneliness and excessive parasocial involvement

should target not only observable behaviors, but also the underlying affective and attitudinal mechanisms. Distinguishing between adaptive and potentially maladaptive uses of parasocial relationships becomes essential in the context of mental health, particularly among young adults with high exposure to digital media content.

Overall, the present study deepens the understanding of parasocial relationships as forms of symbolic connectivity emerging from the interaction between socio-emotional vulnerabilities and the contemporary media environment. Through its integrative and methodologically rigorous approach, the research provides a solid theoretical foundation for future investigations and for the development of more refined explanatory models of how loneliness, attachment, and attitudes toward celebrities shape individuals' psychological experiences in digital society.

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PARENTAL AND ROMANTIC ATTACHMENT: IMPLICATIONS FOR SOMATIC SYMPTOM DISORDER

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Abstract

Numerous studies have examined the relationship between attachment and psychosomatic disorders, yet recent theoretical and clinical developments call for further investigation. This study explored adult attachment to mother, father, and romantic partner in relation to Somatic Symptom Disorder (SSD), as well as differences between individuals with and without the disorder. The sample included 120 participants (60 with SSD) who completed the *Attachment Multiple Model Interview* (AMMI), the *Patient Health Questionnaire-15* (PHQ-15), and the *Somatic Symptom Disorder – B Criteria Scale* (SSD-12). Attachment to mother was associated with both somatic and psychological symptoms of SSD, whereas attachment to father correlated only with somatic symptoms. For the mother relationship, *Security* explained 8% of somatic and 7% of psychological symptoms variance, while *Inhibition* explained 9% of psychological symptoms. For the father relationship, *Security* explained 10% of somatic symptoms. Overall, findings highlight the relevance of attachment-focused interventions in SSD.

Keywords: adult attachment, relationship with parents, romantic partner, Somatic Symptom Disorder.

Attachment theory, developed by John Bowlby (1973, 1980, 1982), provides a fundamental framework for understanding the development of psychological and somatic pathology (Levy *et al.*, 2015; McWilliams & Bailey, 2010; Mikulincer & Shaver, 2016), including psychosomatic disorders (Mauder & Hunter, 2001). According to Bowlby, attachment represents an innate biological tendency of the child to seek proximity to a specific figure – typically the caregiver – in order to obtain protection and a sense of safety (Bowlby, 1982). He described the defining characteristics of attachment relationships, emphasizing their distinctive emotional intensity. The early relationship with the mother plays a crucial role in the individual's psychological development and shapes the way future attachment relationships with others are formed (Bowlby, 1982). Through repeated interactions with the attachment figure and through that figure's responsiveness to the child's needs, mental representations of the self and others are constructed, known as internal working models. These models encompass both the individual's perceptions of

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their own worthiness and lovability, and their beliefs regarding the availability and supportiveness of attachment figures (Bowlby, 1982). Later, Bowlby (1973) proposed the hypothesis of multiple working models, suggesting that individuals may simultaneously operate with different representations of the self and the world, which vary depending on their source, degree of influence, and level of conscious accessibility.

Attachment theory was later tested and expanded by Mary Ainsworth, who proposed a three-category classification of childhood attachment: secure, ambivalent, and avoidant (Salter Ainsworth *et al.*, 2015). Mary Main and Judith Solomon (1986) subsequently added a fourth pattern: disorganized attachment. Further research extended the analysis of attachment into adulthood, demonstrating that romantic relationships function as attachment relationships and that the emotional bonds between romantic partners resemble those formed between children and their parents (Hazan & Shaver, 1987). As individuals mature, the number of attachment figures becomes more diverse, including parents, romantic partners, friends, teachers, mentors, supervisors, or one's own children (Mikulincer & Shaver, 2007). At this stage, individuals may develop multiple attachment relationships, and their hierarchy is dynamic, shaped by developmental stage, personal needs, and the characteristics of each relationship (Mikulincer & Shaver, 2016).

Regarding adult attachment styles, research has long relied on a categorical approach, either with three categories – secure, anxious/ambivalent, and avoidant (Hazan & Shaver, 1987) – or four categories – secure, preoccupied, dismissing, and fearful (Bartholomew, 1990). More recent studies, however, indicate that a dimensional approach is more appropriate for assessing adult attachment (Fraley *et al.*, 2015). The most widely used conceptualization currently includes two dimensions: anxiety, defined by fear of separation and abandonment, and avoidance, defined by discomfort with intimacy and dependence (Brennan, Clark, & Shaver, 1998). A more recent perspective proposes four fundamental dimensions: security, inhibition, hyperactivation, and disorganization (Miljkovitch *et al.*, 2015). The level of security within a relationship increases proportionally with the consistency of appropriate responses received by the individual to their attachment needs. In the absence of availability or adequate support from the attachment figure, individuals rely on secondary strategies: inhibition (suppressing or minimizing needs), hyperactivation (excessive focus on the attachment figure and repeated efforts to achieve closeness), or disorganization (the simultaneous use of opposing strategies of inhibition and hyperactivation) (Miljkovitch *et al.*, 2015).

In the classical literature, internal working models were conceptualized as broad, relatively stable structures, comparable to personality traits. However, recent research highlights that individuals' representations vary depending on the specific relationship in which they are engaged (Fraley *et al.*, 2011; Miljkovitch *et al.*, 2015). From this perspective, a rigorous assessment of adult attachment requires examining its dimensions in relation to significant relationships, particularly those with parents and romantic partners.

The clinical relevance of attachment is well documented, with extensive literature emphasizing its impact on both physical and psychological health (Dagan *et al.*, 2018; Fossati *et al.*, 2012; Lewczuk *et al.*, 2021). A distinct line of research focuses on psychosomatic disorders, where insecure adult attachment has been shown to be a predictive factor for the development of these conditions (Falihatdoost *et al.*, 2020; Nacak *et al.*, 2017). Moreover, studies indicate that individuals diagnosed with psychosomatic disorders exhibit higher levels of insecurity and lower levels of security compared to healthy individuals (Agostini *et al.*, 2016; Dieris-Hirche *et al.*, 2012).

Although the literature has highlighted the relationship between attachment and psychosomatic disorders, most research has relied on traditional terminology and classifications, without systematically integrating recent conceptual and nosological perspectives. Long categorized under “somatoform disorders” with “somatization disorder” as the central diagnosis (APA, 2000), these conditions were redefined in DSM-5 as “Somatic Symptom Disorder and related disorders,” shifting the focus from the negative criterion of lacking medical explanations to positive criteria centered on disproportionate thoughts, emotions, and behaviors associated with somatic symptoms (APA, 2013). The diagnosis requires the presence of somatic symptoms that cause distress and functional impairment, a duration of at least six months, and cognitive, affective, and behavioral manifestations that are disproportionate to symptom severity. This reconceptualization more accurately captures the interaction between body and mind and provides clearer operational criteria, relevant for both clinical practice and empirical research (APA, 2013).

To achieve a deeper understanding of how adult attachment influences Somatic Symptom Disorder within the context of these updated nosological frameworks, a systematic investigation of the relationship between these variables is necessary, using instruments aligned with the most recent theoretical and clinical models. The relevance of such research extends beyond the national context and holds international significance, as many studies continue to rely on traditional approaches to attachment (viewed as a stable individual trait) and to psychosomatic conditions (defined by the absence of medical explanations for somatic symptoms). Thus, the present study aims to contribute to the updating and refinement of the understanding of these relationships, offering a conceptual and methodological framework aligned with recent developments in the field and opening new directions for empirical research and clinical practice.

The primary objective of the study is to examine the relationship between adult attachment (to mother, father, and romantic partner) and the level of symptomatology specific to Somatic Symptom Disorder. The secondary objective is to compare attachment across these three significant adult relationships between individuals with and without Somatic Symptom Disorder.

Based on the theories and empirical findings outlined above, the following hypotheses were formulated:

Hypothesis 1: There is a significant relationship between the dimensions of adult attachment (to mother, father, and romantic partner) and Somatic Symptom Disorder.

Hypothesis 2: Attachment to the mother is a significant predictor of Somatic Symptom Disorder.

Hypothesis 3: Attachment to the father is a significant predictor of Somatic Symptom disorder.

Hypothesis 4: Attachment to the romantic partner is a significant predictor of Somatic Symptom Disorder.

Hypothesis 5: The dimensions of attachment to the mother, father, and romantic partner differ significantly between individuals with Somatic Symptom Disorder and those without the disorder.

METHOD

PARTICIPANTS AND PROCEDURE

The study was conducted on a total of 120 participants, distributed as follows: 60 individuals (7 men, 53 women) diagnosed with Somatic Symptom Disorder, aged between 19 and 56 years ($M=34.53$; $SD=10.01$), and 60 individuals (11 men, 49 women) without Somatic symptom disorder, aged between 19 and 56 years ($M=34.42$; $SD=9.71$).

The sample size for the group with Somatic Symptom Disorder was determined based on a statistical power analysis using the G*Power software (Faul, Erdfelder, Buchner, Lang, 2009), version 3.1.9.7, available free of charge online. To detect a medium effect size in linear regression, with a statistical power of 0.80 and an α level (significance threshold) of 0.05, the minimum required sample size is 55 participants. Additionally, according to the power analysis for independent-samples comparisons, the minimum required sample size for each group is 51 participants, assuming a medium effect size, a statistical power of 0.80, and an α level of 0.05. Considering these criteria, an additional sample of equal size was used for the control group.

Participants were recruited from organizations, companies, and online groups. They received full information regarding the aims and procedures of the study and signed informed consent prior to inclusion. Afterward, they completed a socio-demographic data form and two questionnaires assessing Somatic Symptom Disorder. Subsequently, participants took part in an interview evaluating attachment within three significant relationships: with the mother, the father, and the romantic partner. Participation was voluntary and without financial compensation. However, participants had the option to receive a personalized report that included an interpretation of their questionnaire results and a brief theoretical overview of the

assessed constructs. The study was conducted in accordance with current ethical standards for data collection in Romania and with the ethical principles established by the American Psychological Association (APA, 2017).

METHODS

For this cross-sectional study, three instruments assessing attachment and Somatic Symptom Disorder were used: the Attachment Multiple Model Interview (AMMI), the Patient Health Questionnaire-15 (PHQ-15), and the Somatic Symptom Disorder – B Criteria Scale (SSD-12). The selection of these instruments was based on their strong psychometric properties and their relevance to contemporary conceptualizations of adult attachment and psychosomatic disorders. In addition, the AMMI was chosen because it allows for the assessment of attachment in relation to each significant relationship (mother, father, romantic partner) and because of its capacity to capture unconscious dimensions that are inaccessible through self-report questionnaires. All three instruments had previously been translated and adapted for the Romanian population by the author of the present study as part of the doctoral research and published in academic journals and specialized books (Constantinescu, 2022a; Constantinescu, 2022b; Constantinescu, 2024; Constantinescu, 2025).

The *Attachment Multiple Model Interview* (AMMI), developed by Miljkovitch (2009, cited in Miljkovitch *et al.*, 2015), explores four attachment dimensions within each significant adult relationship: *Security* – considered the primary attachment strategy – and *Inhibition*, *Hyperactivation*, and *Disorganization*, conceptualized as secondary strategies. These dimensions are evaluated at both the representational and behavioral levels. *Security* is reflected in the coherence of the narrative and in the examples provided regarding the quality of the relationship with the attachment figure. *Inhibition* may manifest behaviorally and/or mentally through the non-expression of emotional needs, the minimization, or the lack of awareness of attachment needs. *Hyperactivation* is characterized by excessive focus on the attachment figure and efforts to elicit attention. *Disorganization* involves the combination of opposing strategies (inhibition and hyperactivation) within the same relationship. The interview consists of 23 questions referring to childhood experiences and romantic relationships, targeting how participants perceived and reacted in situations of distress or vulnerability. Scoring is performed on a continuous 9-point scale (0–8), except for the *Disorganization* subscale, which is calculated based on the *Inhibition* and *Hyperactivation* scores (0–16). Low scores indicate a reduced presence of the respective dimension, whereas high scores reflect its predominance in the attachment style. The scoring process is complex and regulated by a detailed manual, which required specialized training for the author of the present study. The AMMI was validated by Miljkovitch *et al.* (2015) on a sample of 53 adults, demonstrating high inter-rater reliability (0.93 for *Security*, 0.89 for *Inhibition*, 0.86 for *Hyperactivation*), as well as internal, convergent, and construct validity.

The *Patient Health Questionnaire-15* (PHQ-15), developed by Kroenke, Spitzer, and Williams (2002), is a 15-item self-report instrument used to assess the severity of somatic symptoms in adults (corresponding to Criterion A of Somatic Symptom Disorder in DSM-5). Responses are rated on a three-point Likert scale, and total scores indicate minimal (0–4), low (5–9), moderate (10–14), or high (15–30) levels of symptom severity. The instrument has demonstrated adequate internal consistency ($\alpha=0.80$) and construct validity (Kroenke *et al.*, 2002), and it has been validated and used in both clinical and general populations across diverse cultural contexts.

The *Somatic Symptom Disorder – B Criteria Scale* (SSD-12), developed by Toussaint *et al.*, (2016), is a 12-item self-report measure assessing the cognitive, affective, and behavioral dimensions associated with somatic symptoms (corresponding to Criterion B of Somatic Symptom Disorder in DSM-5). Responses are provided on a five-point Likert scale, with higher scores indicating greater symptom intensity. The scale has been translated and applied in various clinical and non-clinical settings, demonstrating high reliability ($\alpha=0.95$) and convergent validity (De Vroeghe *et al.*, 2017; Toussaint *et al.*, 2017).

OPERATIONALIZATION CRITERIA FOR SOMATIC SYMPTOM DISORDER

In the specialized literature, there is no universally accepted standard for diagnosing Somatic Symptom Disorder solely on the basis of PHQ-15 and SSD-12 scores. In this study, the criteria were operationalized based on the DSM-5 diagnostic guidelines (APA, 2013) and previous empirical findings. Thus, individuals included in the group with clinical scores suggestive of Somatic Symptom Disorder met all of the following conditions:

(a) PHQ-15 scores including at least one item rated as “bothered a lot” (Wolfe *et al.*, 2014; Wollburg *et al.*, 2013; Xiong *et al.*, 2017) and a total score ≥ 10 (Lee *et al.*, 2015; Toussaint *et al.*, 2020; Wolfe *et al.*, 2014);

(b) SSD-12 scores ≥ 20 (Toussaint *et al.*, 2020) or partial scores ≥ 7 on one of the three dimensions (cognitive, affective, behavioral);

(c) presence of somatic symptoms for a duration longer than six months.

STATISTICAL ANALYSIS

The statistical analyses were conducted using the JASP software (version 0.95.4.0), which is freely available for research purposes. Of the 120 questionnaires, three were incomplete (one missing item), and the missing data were imputed by replacing them with the arithmetic mean of the remaining scores corresponding to the entire questionnaire (PHQ-15) or the respective subscale (SSD-12). The normality of the data was assessed through graphical inspection of histograms and by calculating skewness and kurtosis indices. Descriptive statistics were reported as mean \pm standard deviation. The socio-demographic similarity of the two groups was examined using the Mann-Whitney U test for numerical variables, Fisher’s exact test (FET) for binary categorical variables, and the χ^2 test for multinomial

categorical variables (Kim, 2017). Hypotheses concerning correlations and regressions were tested on the subsample of participants with clinical scores suggestive of somatic symptom disorder ($N = 60$), whereas group comparisons were conducted on the full sample ($N = 120$).

Spearman's rank correlation coefficient (r_s) was used to investigate associations, as the data did not meet the assumptions required for Pearson's coefficient. Effect size was reported using the coefficient of determination (r^2), interpreted according to the following benchmarks: 0.01 – small effect, 0.06 – medium effect, 0.14 – large effect (Sava, 2011). To examine the predictive role of attachment in Somatic Symptom Disorder, the analysis began with simple regressions, each attachment dimension being tested separately to identify its unique contribution. Subsequently, multiple regressions were performed, including all attachment dimensions simultaneously, in order to control for conceptual and statistical overlap and to obtain more precise estimates of the specific effect of each attachment strategy. Group differences were examined using the Mann-Whitney U test, and effect size was expressed through the rank-biserial correlation (r_B), interpreted according to the following benchmarks: 0.1 – small effect, 0.3 – medium effect, 0.5 – large effect (Goss-Sampson, 2020).

RESULTS

DESCRIPTIVE STATISTICS

Table 1 presents the socio-demographic characteristics of the participants, reported as absolute numbers and percentages (with the exception of age, for which the mean and standard deviation were used). Comparisons between the research group and the control group did not reveal statistically significant differences, indicating that the two groups were socio-demographically comparable.

Table 1
Socio-demographic characteristics of the participants

Socio-demographic variables	Research group	Control group	Total participants	Comparisons between the two groups
Gender				FET; $p=0.444$
Men	7 (12%)	11 (18%)	18 (15%)	
Women	53 (88%)	49 (82%)	102 (85%)	
Age (M, SD)	34.53 (10.01)	34.42 (9.71)	34.48 (9.82)	$U=1769$; $p=0.873$
Area				FET; $p=1.00$
Urban	52 (87%)	51 (85%)	103 (86%)	
Rural	8 (13%)	9 (15%)	17 (14%)	
Educational level				FET; $p=0.670$
Secondary education	16 (27%)	13 (22%)	29 (24%)	
Higher education	44 (73%)	47 (78%)	91 (76%)	
Income				$\chi^2(2)=5.601$; $p=0.061$
Low	20 (33%)	17 (29%)	37 (31%)	
Medium	31 (52%)	23 (38%)	54 (45%)	
High	9 (15%)	20 (33%)	29 (24%)	

Note. FET=Fisher's exact test; U=Mann-Whitney test.

Appendix A presents the means and standard deviations, as well as the skewness and kurtosis indicators for the subscales and total scores of the three instruments used. The data distribution was considered normal if the Z-scores (calculated as the ratio between the absolute value and the standard error) for skewness and kurtosis fell within the interval (-3.29, 3.29), a criterion appropriate for samples of 50-300 participants and a 95% confidence level (Kim, 2013). The Z-score values shown in Appendix A indicate that not all subscales or total scores follow a normal distribution, a finding also confirmed through graphical inspection of the histograms.

CORRELATIONS AMONG THE STUDY VARIABLES

Table 2 presents the correlations between the attachment dimensions related to the mother, father, and romantic partner and the scores obtained on the instruments assessing somatic symptoms (PHQ-15) and psychological symptoms (SSD-12) of Somatic Symptom Disorder.

Table 2

Correlations between attachment and Somatic Symptom Disorder

Attachment dimensions	Somatic Symptoms Disorder			
	Somatic symptoms (<i>criterion A</i>)		Psychological symptoms (<i>criterion B</i>)	
	r_s	p	r_s	P
S (mother)	- 0.261	0.044	- 0.348	0.006
I (mother)	0.145	0.271	0.339	0.008
H (mother)	0.178	0.174	- 0.064	0.627
D (mother)	0.039	0.769	0.104	0.428
S (father)	- 0.358	0.005	- 0.177	0.177
I (father)	0.163	0.215	0.111	0.397
H (father)	0.217	0.096	- 0.051	0.701
D (father)	0.101	0.445	- 0.003	0.985
S (partner)	- 0.185	0.158	- 0.205	0.117
I (partner)	0.046	0.728	0.026	0.842
H (partner)	0.238	0.067	0.150	0.252
D (partner)	0.163	0.213	0.056	0.670

Note. Statistically significant correlations are bolded.

S=Security, I=Inhibition, H=Hyperactivation, D=Disorganization.

Statistically significant correlations were found between somatic symptoms (Criterion A, assessed with the PHQ-15) and the attachment dimensions *Security* to the mother ($r_s = -0.261$, $r_s^2 = 0.07$, $p = 0.044$, medium effect) and *Security* to the father ($r_s = -0.358$, $r_s^2 = 0.13$, $p = 0.005$, large effect). For the remaining attachment dimensions, correlations with somatic symptoms were not statistically significant.

Significant correlations, with medium-to-large effect sizes, were also identified between psychological symptoms (Criterion B, assessed with the SSD-12) and the dimensions *Security* to the mother ($r_s = -0.348$, $r_s^2 = 0.12$, $p = 0.006$) and *Inhibition* to the mother ($r_s = 0.339$, $r_s^2 = 0.11$, $p = 0.008$). For the other attachment dimensions, correlations with psychological symptoms were not statistically significant.

Considering these results, it can be concluded that Hypothesis 1, regarding the relationship between attachment and Somatic Symptom Disorder, was partially confirmed.

THE PREDICTIVE ROLE OF ATTACHMENT IN SOMATIC SYMPTOM DISORDER

The results of the simple and multiple regression analyses are presented below.

a) *Attachment to the mother*

Given the significant correlations between the attachment dimensions to the mother and somatic symptoms (Criterion A of Somatic Symptom Disorder), a simple regression analysis was conducted in which maternal attachment *Security* acted as a negative predictor of somatic symptoms. The results indicated a significant predictive role of *Security* ($F(1) = 4.843$, $p = 0.032$), with this dimension explaining approximately 8% of the variance in somatic symptoms ($r^2 = 0.080$, adjusted $r^2 = 0.063$).

Before performing the multiple regression, correlations among the maternal attachment dimensions were examined to assess the degree of association between predictors and to control for multicollinearity risk. The results are presented in Table 3.

Table 3

Correlations between the dimensions of attachment to the mother

	S (mother)	I (mother)	H (mother)	D (mother)
S (mother)	–			
I (mother)	– 0.911*	–		
H (mother)	– 0.198	0.026	–	
D (mother)	– 0.430*	0.428*	0.743*	–

Note. * $p < .001$.

S=Security, I=Inhibition, H=Hyperactivation, D=Disorganization.

A high correlation was identified between the *Security* and *Inhibition* dimensions, which required the exclusion of one of these variables to avoid multicollinearity. Given that *Inhibition* is already used in calculating the *Disorganization* score, it was excluded from the multiple regression analysis.

Table 4 presents the results of the multiple regression analysis, in which the attachment dimensions to the mother (*Security*, *Hyperactivation*, and *Disorganization*) were included as simultaneous predictors of somatic symptoms.

Table 4

Results of the multiple regression analysis between the dimensions of attachment to the mother and somatic symptoms (criterion A of Somatic Symptom Disorder)

Predictors	Unstandardized coefficients (B)	Standard error (SE B)	Standardized coefficients (β)	t	p
S (mother)	-0.761	0.335	-0.327	-2.273	0.027
H (mother)	1.532	0.597	0.431	2.566	0.013
D (mother)	-0.694	0.244	-0.521	-2.840	0.006

Note. S=Security, H=Hyperactivation, D=Disorganization.

The analysis showed that the attachment dimensions to the mother, considered simultaneously, acted as predictors of somatic symptoms ($F(3, 56)=4.101$, $p=0.011$), explaining 18% of their variance ($r^2=0.180$, adjusted $r^2=0.136$). All three dimensions were significant predictors: *Security* ($\beta=-0.327$, $t=-2.273$, $p=0.027$), *Hyperactivation* ($\beta=0.431$, $t=2.566$, $p=0.013$), and *Disorganization* ($\beta=-0.521$, $t=-2.840$, $p=0.006$).

Subsequently, two simple regressions were conducted for psychological symptoms (Criterion B), with *Security* and *Inhibition* as single predictors. The results showed that *Security* was a negative predictor ($F(1)=4.332$, $p=0.042$), explaining 7% of the variance in psychological symptoms ($r^2=0.070$, adjusted $r^2=0.053$), whereas *Inhibition* was a positive predictor ($F(1)=5.363$, $p=0.024$), explaining 9% of the variance ($r^2=0.085$, adjusted $r^2=0.069$).

Table 5 presents the results of the multiple regression analysis in which the attachment dimensions to the mother (*Security*, *Hyperactivation*, and *Disorganization*) were included as simultaneous predictors of psychological symptoms.

Table 5

Results of the multiple regression analysis between the dimensions of attachment to the mother and psychological symptoms (criterion B of Somatic Symptom Disorder)

Predictors	Unstandardized coefficients (B)	Standard error (SE B)	Standardized coefficients (β)	t	p
S (mother)	-0.947	0.623	-0.229	-1.519	0.134
H (mother)	-1.458	1.111	-0.231	-1.313	0.194
D (mother)	0.554	0.454	0.234	1.219	0.228

Note. S=Security, H=Hyperactivation, D=Disorganization.

The analysis showed that the attachment dimensions to the mother, considered simultaneously, did not have a significant predictive value for psychological symptoms ($F(3, 56)=2.107, p=0.110$), with the overall model not reaching statistical significance.

Given these results, it can be concluded that Hypothesis 2, concerning the predictive role of maternal attachment in Somatic Symptom Disorder, was partially confirmed.

b) Attachment to the father

Given the significant correlations between the attachment dimensions to the father and somatic symptoms (Criterion A of Somatic Symptom Disorder), a simple regression analysis was conducted in which paternal attachment *Security* acted as a negative predictor of somatic symptoms. The results indicated a significant predictive role of *Security* ($F(1)=6.675, p=0.012$), with this dimension explaining approximately 10% of the variance in somatic symptoms ($r^2=0.103$, adjusted $r^2=0.088$).

Before performing the multiple regression, correlations among the paternal attachment dimensions were examined to assess the degree of association between predictors and to control for multicollinearity risk. The results are presented in Table 6.

Table 6

Correlations between the dimensions of attachment to the father

	S (father)	I (father)	H (father)	D (father)
S (father)	–			
I (father)	– 0.805***	–		
H (father)	– 0.309*	0.014	–	
D (father)	– 0.378**	0.280*	0.809***	–

Note. * $p < .05$, ** $p < .01$, *** $p < .001$.

S=Security, I=Inhibition, H=Hyperactivation, D=Disorganization.

A high correlation was identified between *Security* and *Inhibition*, as well as between *Hyperactivation* and *Disorganization*, which required the exclusion of one variable from each pair to avoid multicollinearity. Since *Inhibition* and *Hyperactivation* are already used in calculating the *Disorganization* score, they were excluded from the multiple regression analysis.

Table 7 presents the results of the multiple regression analysis in which the attachment dimensions to the father (*Security* and *Disorganization*) were included as simultaneous predictors of somatic symptoms.

Table 7

Results of the multiple regression analysis between the dimensions of attachment to the father and somatic symptoms (criterion A of Somatic Symptom Disorder)

Predictors	Unstandardized coefficients (B)	Standard error (SE B)	Standardized coefficients (β)	t	p
S (father)	- 0.835	0.306	- 0.379	- 2.732	0.008
D (father)	- 0.169	0.179	- 0.131	- 0.943	0.349

Note. S=Security, D=Disorganization.

The analysis showed that the attachment dimensions to the father, considered simultaneously, acted as predictors of somatic symptoms ($F(2, 57)=3.776, p=0.029$), explaining approximately 12% of their variance ($r^2=0.117$, adjusted $r^2=0.086$). Among the two dimensions, only *Security* had a significant predictive value for somatic symptoms ($\beta= -0.379, t= -2.732, p=0.008$).

Given these results, it can be concluded that Hypothesis 3, concerning the predictive role of paternal attachment in Somatic Symptom Disorder, was partially confirmed.

c) *Attachment to the romantic partner*

The absence of significant correlations indicates that the attachment dimensions to the romantic partner did not demonstrate predictive value for Somatic Symptom Disorder in the analyzed sample. Consequently, no additional regression analyses were conducted.

Considering these aspects, it can be concluded that Hypothesis 4, regarding the predictive role of attachment to the romantic partner in Somatic Symptom Disorder, was not confirmed.

COMPARISONS OF ATTACHMENT IN INDIVIDUALS
WITH AND WITHOUT SOMATIC SYMPTOM DISORDER

The analysis revealed statistically significant differences, with a medium effect size, between individuals with Somatic Symptom Disorder and those without this disorder regarding attachment to the mother. The results are presented in Table 8.

The levels of *Security* in the relationship with the mother were significantly lower in individuals with Somatic Symptom Disorder compared to those without the disorder. In contrast, the levels of *Inhibition*, *Hyperactivation*, and *Disorganization* in the relationship with the mother were significantly higher in individuals with Somatic Symptom Disorder.

Table 8

Group comparisons based on the dimensions of attachment to the mother

Attachment dimensions	Median	Differences between groups	Effect size
S (mother)		U=1261.5; p=0.004	- 0.299
Research group	3.00		
Control group	4.00		
I (mother)		U=2291; p=0.009	0.273
Research group	6.00		
Control group	4.00		
H (mother)		U=2258.5; p=0.012	0.255
Research group	4.00		
Control group	4.00		
D (mother)		U=2379; p=0.002	0.322
Research group	8.00		
Control group	4.00		

Note. The Mann-Whitney U test was used.

S=Security, I=Inhibition, H=Hyperactivation, D=Disorganization.

The analysis also revealed statistically significant differences, with a medium-to-large effect size, between individuals with and without the disorder regarding attachment to the father. The results are presented in Table 9.

Table 9

Group comparisons based on the dimensions of attachment to the father

Attachment dimensions	Median	Differences between groups	Effect size
S (father)		U=1099.5; p<0.001	- 0.389
Research group	2.00		
Control group	4.00		
I (father)		U=2396.5; p=0.001	0.331
Research group	6.00		
Control group	4.75		
H (father)		U=2484; p<0.001	0.380
Research group	4.00		
Control group	2.50		
D (father)		U=2526; p<0.001	0.403
Research group	8.00		
Control group	4.00		

Note. The Mann-Whitney U test was used.

S=Security, I=Inhibition, H=Hyperactivation, D=Disorganization.

The levels of *Security* in the relationship with the father were significantly lower in individuals with Somatic Symptom Disorder compared to those without the disorder. In contrast, the levels of *Inhibition*, *Hyperactivation*, and *Disorganization* in the relationship with the father were significantly higher in individuals with Somatic Symptom Disorder.

The analysis revealed statistically significant differences, with a medium effect size, between individuals with and without the disorder regarding attachment to the romantic partner. The results are presented in Table 10. Levels of *Security* in the relationship with the romantic partner were significantly lower in individuals with Somatic Symptom Disorder, whereas levels of *Inhibition*, *Hyperactivation*, and *Disorganization* were significantly higher compared to those without the disorder.

Table 10

Group comparisons based on the dimensions of attachment to the romantic partner

Attachment dimensions	Median	Differences between groups	Effect size
S (partner)		U=1290; p=0.006	- 0.283
Research group	4.00		
Control group	5.50		
I (partner)		U=2390; p=0.001	0.328
Research group	4.00		
Control group	2.00		
H (partner)		U=2365; p=0.002	0.314
Research group	5.00		
Control group	3.00		
D (partner)		U=2450.5; p<0.001	0.361
Research group	6.00		
Control group	4.00		

Note. The Mann-Whitney U test was used.

S=Security, I=Inhibition, H=Hyperactivation, D=Disorganization.

The results obtained confirm Hypothesis 5, according to which the attachment dimensions to the mother, father, and romantic partner differ significantly between individuals with Somatic Symptom Disorder and those without this disorder.

DISCUSSION

The results of this study indicate a strong connection between adult attachment and Somatic Symptom Disorder. The study makes an important contribution to the field by examining attachment through the lens of specificity – an innovative approach that considers the social context and the variability of attachment style across different relationships. Moreover, the interview used to assess attachment captures unconscious aspects as well, which adds greater rigor to the evaluation.

Significant correlations were identified between parental attachment and Somatic Symptom Disorder: the *Security* dimension (mother) correlated negatively with both somatic and psychological symptoms; the *Inhibition* dimension (mother) correlated positively with psychological symptoms; and the *Security* dimension (father) correlated negatively with somatic symptoms. These findings align with

existing literature showing a positive association between attachment insecurity and psychosomatic disorders and a negative association between attachment security and these disorders (Fasakhoudi *et al.*, 2022; Heenan *et al.*, 2020; Lewczuk *et al.*, 2021; McWilliams & Asmundson, 2007; McWilliams & Bailey, 2010; Sherry *et al.*, 2014; Stanton & Campbell, 2014). According to the present data, no significant relationships were found between attachment to the romantic partner and Somatic Symptom Disorder. This discrepancy may be explained by the relatively small sample size (N=60), but also by the fact that, although the romantic relationship is considered the primary attachment relationship in adulthood, attachment to parents has long-lasting effects on both mental and physical health – an aspect confirmed by longitudinal studies (Farrell *et al.*, 2019; Puig *et al.*, 2013).

The findings of the present study also indicated the predictive effect of attachment on the symptomatology of Somatic Symptom Disorder. Regarding maternal attachment, the *Security* dimension explained 8% of the variance in somatic symptoms, while the overall model (including primary and secondary strategies) explained 18%. For psychological symptoms, *Security* explained 7%, and *Inhibition* explained 9% of the variance. Regarding paternal attachment, the *Security* dimension explained 10% of the variance in somatic symptoms, while the overall model explained 12%. Although no studies have directly examined the predictive effect of adult attachment dimensions on Somatic Symptom Disorder, previous research has highlighted the predictive role of attachment in psychosomatic conditions (Fasakhoudi *et al.*, 2022; Heenan *et al.*, 2020; Lewczuk *et al.*, 2021; Stanton & Campbell, 2014; McWilliams & Bailey, 2010).

Comparisons between individuals with and without Somatic Symptom Disorder revealed statistically significant differences: individuals with the disorder showed lower levels of *Security* and higher levels of secondary strategies (*Inhibition*, *Hyperactivation*, *Disorganization*). These differences were not influenced by socio-demographic characteristics, as the groups were similar in this regard. The results are consistent with previous research indicating lower attachment security and higher attachment insecurity in individuals with psychosomatic conditions (Agostini *et al.*, 2016; Dieris-Hirche *et al.*, 2012).

Considering these aspects, it can be concluded that the present study makes a significant contribution to the field of attachment and psychosomatics. One of its strengths lies in the use of modern approaches and rigorous instruments, aligned with current clinical conceptualizations. The assessment of attachment integrated both representational and behavioral dimensions, taking into account the three significant adult relationships: with the mother, the father, and the romantic partner. By examining these relationships, the study offers a more nuanced understanding of the role of attachment in the symptomatology of Somatic Symptom Disorder. Beyond its theoretical impact, the findings have practical relevance, suggesting the need for a holistic treatment approach that includes, in addition to classical interventions, an attachment-focused therapeutic component.

However, certain methodological limitations must be acknowledged. First, as a cross-sectional study, it does not allow for conclusions regarding causal relationships between the variables investigated. Second, the sample size was relatively small and included only adults, which limits the generalizability of the findings to other age groups. Additionally, Somatic Symptom Disorder was assessed solely through self-report, without the use of a standardized clinical interview to confirm the diagnosis.

Future research directions could include longitudinal studies, which would allow for the investigation of causal relationships, as well as larger and more age-diverse samples. The use of standardized clinical assessments would contribute to a more accurate identification of the disorder. Furthermore, future studies could integrate psychotherapeutic interventions aimed at increasing attachment security, with subsequent evaluation of changes in attachment patterns and Somatic Symptom Disorder symptomatology following therapy.

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APPENDICES

APPENDIX A

Descriptive statistics for the subscales or total scores of the AMMI, PHQ -15, and SSD -12 instruments

Subscales/Total score	M	SD	Skewness (Standard error)	Kurtosis (Standard error)
AMMI				
<i>S (mother)</i>				
Research group	3.13	1.81	0.437 (0.309)	– 0.125 (0.608)
Control group	4.15	1.84	0.109 (0.309)	– 0.946 (0.608)
<i>I (mother)</i>				
Research group	4.90	2.31	– 0.922 (0.309)	– 0.165 (0.608)
Control group	3.83	2.25	– 0.035 (0.309)	– 1.339 (0.608)
<i>H (mother)</i>				
Research group	4.06	1.19	– 1.443 (0.309)	2.348 (0.608)
Control group	3.36	1.53	– 0.690 (0.309)	– 0.315 (0.608)
<i>D (mother)</i>				
Research group	7.12	3.16	– 1.171 (0.309)	0.398 (0.608)
Control group	5.42	3.16	– 0.214 (0.309)	– 0.790 (0.608)
<i>S (father)</i>				
Research group	2.47	1.91	0.718 (0.309)	– 0.567 (0.608)
Control group	3.73	1.75	0.026 (0.309)	– 1.413 (0.608)
<i>I (father)</i>				
Research group	5.50	2.15	– 1.199 (0.309)	0.478 (0.608)
Control group	4.24	2.33	– 0.370 (0.309)	– 1.245 (0.608)
<i>H (father)</i>				
Research group	3.70	1.58	– 0.149 (0.309)	– 0.026 (0.608)
Control group	2.63	1.40	– 0.188 (0.309)	– 0.602 (0.608)

Appendix A (continued)

Subscales/Total score	M	SD	Skewness (Standard error)	Kurtosis (Standard error)
<i>D (father)</i>				
Research group	6.73	3.27	-0.508 (0.309)	-0.596 (0.608)
Control group	4.43	2.93	-0.020 (0.309)	-0.942 (0.608)
<i>S (partner)</i>				
Research group	3.83	1.92	-0.059 (0.309)	-1.130 (0.608)
Control group	4.79	1.78	-0.490 (0.309)	-0.765 (0.608)
<i>I (partner)</i>				
Research group	3.39	2.13	0.004 (0.309)	-1.055 (0.608)
Control group	2.19	1.98	0.949 (0.309)	0.264 (0.608)
<i>H (partner)</i>				
Research group	4.33	1.81	0.101 (0.309)	-1.139 (0.608)
Control group	3.30	1.79	0.008 (0.309)	-0.881 (0.608)
<i>D (partner)</i>				
Research group	5.88	3.47	-0.124 (0.309)	-0.883 (0.608)
Control group	3.67	3.11	0.400 (0.309)	-0.808 (0.608)
PHQ-15				
<i>Total score</i>				
Research group	15.2	4.21	0.907 (0.309)	0.645 (0.608)
Control group	3.03	1.83	0.035 (0.309)	-0.853 (0.608)
SSD-12				
<i>Total score</i>				
Research group	20.3	7.49	1.238 (0.309)	0.919 (0.608)
Control group	2.18	2.26	0.823 (0.309)	-0.519 (0.608)

Note. S=Security, I=Inhibition, H=Hyperactivation, D=Disorganization.

INTERGENERATIONAL DIFFERENCES IN PERFECTIONISM, ANXIETY, AND PROCRASTINATION: A REGRESSION-BASED MODEL

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Abstract

This study examined the relationships between perfectionism, anxiety (performance and social), and procrastination, focusing on intergenerational differences and predictive mechanisms. Participants (N = 155) were analyzed considering three generations: Generation Z, Generation Y, and Generation X. Levels of perfectionism, anxiety, and procrastination were assessed to explore generational differences and to identify predictors of procrastination. Results showed that Generation Z reported significantly higher levels of perfectionism, anxiety, and procrastination compared to the older generations. Correlational analyses revealed positive associations between both types of anxiety and procrastination, particularly for intentional delay behaviors. Regression analysis indicated that two perfectionism facets – Doubts about Actions (self-critical) and Entitlement (narcissistic) – were significant predictors of procrastination, beyond the influence of anxiety. The findings suggest that perfectionistic cognitions and evaluative anxiety jointly contribute to procrastinatory behavior, highlighting the importance of addressing these factors in psychological and educational interventions.

Keywords: perfectionism, anxiety, procrastination, intergenerational differences.

1. INTRODUCTION

In recent years, concerns regarding psychological well-being have intensified, as contemporary society places increasingly high expectations on individual performance, achievement, and self-presentation. The prevalence of anxiety, perfectionistic tendencies, and avoidance behaviors such as procrastination has drawn attention to the cognitive and emotional mechanisms underlying maladaptive self-regulation (Bornioli *et al.*, 2019; Saiphoo & Vahedi, 2019; Sam *et al.*, 2020). This growing psychological strain is reinforced by a culture of comparison and unrealistic standards that shape personal identity and self-evaluation (Myers & Crowther, 2009; Lloyd *et al.*, 2015).

Theoretical and empirical findings converge on the idea that these three constructs – anxiety, perfectionism, and procrastination – form an interconnected

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system of self-regulation. While anxiety and perfectionism reflect internal cognitive and emotional pressures, procrastination represents an observable behavioral outcome of these internal dynamics (Steel, 2007; Petrie, 2014). In contexts where performance and appearance are overemphasized, individuals are often trapped in a cycle of high expectations, fear of failure, and avoidance of evaluation (Bardone-Cone *et al.*, 2007; Levinson *et al.*, 2013). These dynamics are especially relevant in educational and professional settings, where external pressures and self-imposed standards intensify emotional tension and disrupt effective task engagement.

Classically, anxiety has been conceptualized as a persistent emotional state of apprehension associated with anticipated failure or negative evaluation (Zeidner, 1998; Spielberger & Vagg, 1995). However, more recent research has documented the growing prevalence of performance-related and social anxiety in contemporary contexts characterized by competition and constant evaluation (Bornioli *et al.*, 2019; Payne, 2019; Von der Embse *et al.*, 2018; Sam *et al.*, 2020). In these settings, anxiety undermines concentration and cognitive control, producing physiological arousal and avoidance tendencies that interfere with goal-directed behavior.

Perfectionism, defined as the tendency to impose excessively high standards and to evaluate one's worth based on achievement, has similarly shifted from being regarded as a positive motivator to being recognized as a significant vulnerability factor for emotional distress (Hewitt & Flett, 1991; Stoeber & Otto, 2006). Maladaptive perfectionism is characterized by overconcern with mistakes, doubts about actions, and conditional self-worth (Frost *et al.*, 1990; Levinson *et al.*, 2013). These perfectionistic tendencies are often accompanied by heightened anxiety, as individuals become preoccupied with avoiding imperfection or criticism (Sherry *et al.*, 2016; Lloyd *et al.*, 2015).

Procrastination has been defined as the voluntary delay of intended tasks despite awareness of negative outcomes (Steel, 2007). Far from reflecting simple laziness, procrastination often serves as a coping mechanism that momentarily reduces the discomfort generated by perfectionistic pressure and performance anxiety (Metin *et al.*, 2016; Petrie, 2014). Individuals delay tasks not because of indifference, but to avoid the psychological tension of potential failure or imperfection.

1.1. ANXIETY–PERFECTIONISM–PROCRASTINATION DYNAMICS

The relationship between anxiety and perfectionism is cyclical: high self-imposed standards increase the fear of failure, while anxiety reinforces rigid self-evaluative tendencies. Studies among students and professionals indicate that perfectionistic individuals interpret feedback as global personal judgment, which intensifies both anxiety and self-criticism (Vitasari *et al.*, 2010; Von der Embse *et al.*, 2018). Over time, this cycle creates emotional exhaustion and cognitive interference that can impair motivation and task engagement (Flett *et al.*, 2016; Stoeber, 2018).

Empirical evidence indicates that perfectionism and anxiety operate as powerful antecedents of procrastination (Steel & Klingsieck, 2016). Those with self-critical or rigid perfectionism tend to postpone work when convinced that their efforts will not meet internalized ideals (Flett *et al.*, 2016; Alblwi *et al.*, 2021). Similarly, anxious arousal generates cognitive interference – rumination, self-doubt, and avoidance of evaluation – that further reinforces delay behaviors (Alharthi *et al.*, 2019; Metin *et al.*, 2016).

From a cognitive-behavioral perspective, this triadic relationship can be viewed as a self-perpetuating loop: perfectionism creates unrealistic standards, anxiety amplifies the fear of not meeting them, and procrastination offers temporary relief that ultimately maintains both anxiety and self-criticism (Steel & Klingsieck, 2016). The avoidance brings short-term comfort but long-term costs, lower performance, guilt, and emotional strain (Steel, 2007; Petrie, 2014).

Recent studies emphasize that these processes are particularly salient in younger generations exposed to constant evaluative pressure (Bornioli *et al.*, 2019; Sam *et al.*, 2020). Yet, comparative evidence across generations remains limited, and little is known about how the specific dimensions of perfectionism contribute to procrastination when anxiety is considered. The present study therefore aims to explore these predictive relationships through an intergenerational analytical framework, clarifying the joint role of anxiety and perfectionism in shaping procrastination.

2. RESEARCH DESIGN

2.1. RESEARCH OBJECTIVE

The main research objective of this study is to investigate the relationships among anxiety, perfectionism, and procrastination, focusing on intergenerational differences and predictive associations between these variables. Drawing on previous literature that emphasizes the link between maladaptive perfectionism, evaluative anxiety, and avoidance behavior (Flett *et al.*, 2016; Steel & Klingsieck, 2016; Metin *et al.*, 2016), the study seeks to clarify how specific facets of perfectionism and anxiety contribute to procrastination across Generation Z, Generation Y, and Generation X.

2.2. HYPOTHESES

Prior research suggests that younger individuals face stronger social and academic pressures, leading to increased self-criticism, performance concerns, and avoidance behaviors (Bornioli *et al.*, 2019; Sam *et al.*, 2020). Exposure to constant evaluation and unrealistic standards contributes to heightened anxiety and perfectionistic striving (Lloyd *et al.*, 2015; Von der Embse *et al.*, 2018). Based on this theoretical rationale, we propose the following hypothesis:

H1: *There are significant generational differences in the levels of perfectionism, anxiety, and procrastination, with Generation Z expected to report higher scores than Generation Y and Generation X.*

The cognitive-behavioral model of avoidance explains procrastination as a response to emotional discomfort and evaluative fear (Steel, 2007; Petrie, 2014). Considering that anxiety, particularly in performance contexts was shown to lead to task delay as a short-term strategy to regulate distress, even though it reinforces avoidance in the long term (Metin *et al.*, 2016; Steel & Klingsieck, 2016), we propose the following hypothesis:

H2: *Higher levels of anxiety (social and performance) will positively predict procrastination.*

Perfectionism has been conceptualized as a multidimensional construct encompassing rigid, self-critical, and narcissistic forms (Smith *et al.*, 2016). Among these, self-critical perfectionism, characterized by doubts about actions and concern over mistakes, is most consistently associated with avoidance and self-regulatory failure (Frost *et al.*, 1990; Flett *et al.*, 2016). Additionally, narcissistic perfectionism, including facets such as entitlement and hypercriticism, may contribute to procrastination by undermining responsibility and persistence (Sherry *et al.*, 2016). Therefore, we propose the following hypothesis:

H3: *Specific perfectionism dimensions, particularly self-critical and narcissistic perfectionism, will significantly predict procrastination beyond the effect of anxiety.*

3. METHOD

3.1. PARTICIPANTS

The study included 155 Romanian participants aged between 18 and 64 years ($M = 40.67$, $SD = 10.31$), of whom 74.8% were women. Participants were categorized into three generational cohorts based on their year of birth: Generation X (1965–1980; 40%), Generation Y (1981–1996; 46%), and Generation Z (1997–2012; 14%).

Most respondents (71.6%) were employed at the time of data collection (March–April 2025), while 21.9% were students. Among those employed, 49.7% identified as entrepreneurs, 14.8% worked in the private sector, and 13.5% in the public sector. Regarding education, 41.9% held university degrees, 30.0% postgraduate qualifications, 23.2% had completed high school, and 4.5% held doctoral degrees. Data were collected online through a self-report questionnaire distributed via snowball sampling, ensuring confidentiality and voluntary participation.

3.2. INSTRUMENTS

Perfectionism was assessed using the Big Three Perfectionism Scale (BTPS; Smith *et al.*, 2016), a 45-item instrument rated on a 5-point Likert scale (1 = strongly disagree, 5 = strongly agree). The scale measures three major dimensions of perfectionism. The first, Rigid Perfectionism ($\alpha = .92$), includes two facets: Self-Oriented Perfectionism ($\alpha = .85$) and Self-Worth Contingencies ($\alpha = .87$). The second dimension, Self-Critical Perfectionism ($\alpha = .94$), comprises Concern over Mistakes ($\alpha = .85$), Doubts about Actions ($\alpha = .91$), Self-Criticism

($\alpha = .92$), and Socially Prescribed Perfectionism ($\alpha = .78$). The third dimension, Narcissistic Perfectionism ($\alpha = .90$), includes Other-Oriented Perfectionism ($\alpha = .87$), Hypercriticism ($\alpha = .71$), Entitlement ($\alpha = .73$), and Grandiosity ($\alpha = .64$). Overall, the BTPS demonstrated excellent internal consistency ($\alpha = .96$).

Anxiety was measured using the Liebowitz Social Anxiety Scale (Liebowitz, 1987), a 24-item instrument rated on a 4-point Likert scale (1 = not at all, 4 = severe). The scale evaluates two related constructs: Performance Anxiety (13 items, $\alpha = .92$) and Social Anxiety (11 items, $\alpha = .92$). The total score reflects a global measure of social-evaluative anxiety, with excellent internal consistency ($\alpha = .96$).

Procrastination was assessed using the Procrastination at Work Scale (PAWS; Metin *et al.*, 2016), consisting of 12 items rated on a 5-point Likert scale (1 = never, 5 = always). The instrument includes two subscales: Soldiering-intentional task delay (8 items, $\alpha = .90$)-and Cyberslacking, referring to non-work-related internet use (4 items, $\alpha = .75$). The total score indicated high internal reliability ($\alpha = .89$).

3.3. RESULTS

The collected data will be interpreted through a comprehensive analytical framework that integrates descriptive, correlational, and predictive approaches to clarify the interplay between perfectionism, anxiety, and procrastination, as well as to illuminate meaningful generational patterns and the relative weight of specific psychological mechanisms underlying procrastination behaviour.

Descriptive analyses showed that participants reported moderate levels of perfectionism and anxiety, accompanied by relatively elevated levels of procrastination. Global procrastination had a mean of 3.24 (SD = 1.20), with cyberslacking (M = 3.54, SD = 1.47) exceeding soldiering (M = 2.94, SD = 1.30). Social anxiety displayed a lower average level (M = 1.93, SD = 0.72).

Table 1

Associations between procrastination and anxiety at the facet level (N = 155)

Variable	1	2	3	4
1 Performance anxiety	-			
2 Social anxiety	.83**	-		
3 Soldiering	.38**	.37**	-	
4 Cyberslacking	.13	.17*	.49**	-

Notes: * $p < .05$, ** $p < .01$

As shown in Table 1, zero-order correlations revealed a strong, positive association between performance anxiety and social anxiety ($r = .83$, $p < .01$), reflecting substantial overlap between evaluative and interpersonal forms of anxiety. Both anxiety dimensions were moderately associated with global procrastination: performance anxiety correlated at $r = .31$ ($p < .01$), and social anxiety at $r = .33$ ($p < .01$). Soldiering, the intentional postponement of tasks, showed stronger

correlations with both anxiety types (performance anxiety: $r = .38, p < .01$; social anxiety: $r = .37, p < .01$). For cyberslacking, only social anxiety showed a significant but weak relationship ($r = .17, p < .05$), whereas performance anxiety was not significantly associated ($r = .13, ns$).

Table 2 shows the procrastination associations with study variables, including anxiety facets and multiple facets of perfectionism. Within self-critical perfectionism, concern over mistakes correlated at $r = .36 (p < .01)$, doubts about actions at $r = .44 (p < .01)$, and self-criticism at $r = .33 (p < .01)$. Among narcissistic perfectionism facets, hypercriticism ($r = .26, p < .01$) and entitlement ($r = .32, p < .01$) showed moderate associations with procrastination, while grandiosity demonstrated no meaningful correlation ($r = .02, ns$). These results support the hypothesis that higher levels of anxiety and perfectionism, particularly the most evaluative and self-critical facets, are linked to increased procrastination.

To test generational differences, Kruskal–Wallis analyses were conducted across the three cohorts: Generation X (1965–1980), Generation Y (1981–1996), and Generation Z (1997–2012). Significant effects emerged for all study variables. Differences were largest for self-critical perfectionism ($\chi^2 = 25.10, p < .001, \eta^2 = .146$) and rigid perfectionism ($\chi^2 = 23.32, p < .001, \eta^2 = .135$), indicating medium-to-large effects. Social anxiety also differed significantly between generations ($\chi^2 = 22.58, p < .001, \eta^2 = .130$), while narcissistic perfectionism showed a smaller but significant effect ($\chi^2 = 10.05, p = .018, \eta^2 = .047$). Procrastination varied between generations as well, although with a smaller effect size ($\chi^2 = 8.59, p = .035, \eta^2 = .037$).

Table 2

Associations between procrastination and all study variables at the facet level (N = 155)

Variable	Procrastination
1 Performance anxiety	.31**
2 Social anxiety	.33**
3 Self-oriented perfectionism	.22**
4 Self-worth contingencies	.28**
5 Concern over mistakes	.36**
6 Doubts about actions	.44**
7 Self-criticism	.33**
8 Socially prescribed perfectionism	.30**
9 Other-oriented perfectionism	.16*
10 Hypercriticism	.26**
11 Entitlement	.32**
12 Grandiosity	.02

Notes: * $p < .05$, ** $p < .01$. Variables 3–12 represent the facet-level components of the three higher-order dimensions of perfectionism. *Rigid perfectionism* includes (3) self-oriented perfectionism and (4) self-worth contingencies. *Self-critical perfectionism* comprises (5) concern over mistakes, (6) doubts about actions, (7) self-criticism, and (8) socially prescribed perfectionism. *Narcissistic perfectionism* includes (9) other-oriented perfectionism, (10) hypercriticism, (11) entitlement, and (12) grandiosity.

Examination of mean ranks revealed a consistent pattern across generations. Generation Z showed the highest levels on all variables, with mean ranks of 105.35 for rigid perfectionism, 111.80 for self-critical perfectionism, 118.27 for social anxiety, and 96.93 for procrastination. In contrast, Generation X consistently displayed the lowest levels, with mean ranks of 57.32 for rigid perfectionism, 58.31 for self-critical perfectionism, 66.53 for social anxiety, and 66.14 for procrastination. Generation Y occupied an intermediate position, with mean ranks of 86.51 for rigid perfectionism, 83.99 for self-critical perfectionism, 76.08 for social anxiety, and 81.70 for procrastination. For narcissistic perfectionism, the pattern was more nuanced: Generation Y showed the highest mean rank (88.61), followed by Generation Z (80.34), while Generation X reported the lowest values (63.84). These results support the hypothesis that meaningful generational differences exist in perfectionism, anxiety, and procrastination, with younger adults generally reporting higher levels across these psychological dimensions.

Table 3

Hierarchical multiple regression predicting procrastination
from social anxiety and perfectionism facets (N = 155)

Model	Predictor	β	t	p	R	R^2	ΔR^2	F_{change}
1	Social anxiety	.33	4.33	<.01	.33	.11	.11	18.76**
2	Social anxiety	.16	1.94	.06	.46	.21	.10	19.33**
	Doubts about actions	.36	4.40	<.01				
3	Social anxiety	.16	1.99	.05	.51	.26	.05	9.67**
	Doubts about actions	.31	3.76	<.01				
	Entitlement	.23	3.11	<.01				

Notes: * $p < .05$, ** $p < .01$

To determine the extent to which anxiety and perfectionism predict procrastination, a stepwise multiple regression analysis was performed. Results are shown in Table 3. Social anxiety entered the model first and emerged as a significant predictor, accounting for 11% of the variance in global procrastination, $F(1, 153) = 18.76$, $p < .001$. Its regression coefficient indicated a positive relationship ($B = 0.52$, $\beta = .33$).

The addition of doubts about actions, a self-critical perfectionism facet, significantly improved the model. With this variable included, the explained variance increased to 21%, and doubts about actions became the strongest predictor ($B = 0.40$, $\beta = .36$, $t = 4.40$, $p < .01$). Social anxiety remained statistically non-significant ($p = .06$), suggesting its predictive value diminished once perfectionistic self-doubt was taken into account.

Entitlement, a narcissistic perfectionism facet, entered at the next step and added another 5% to the explained variance. In this expanded model, doubts about actions remained the strongest predictor ($B = 0.34$, $\beta = .31$), entitlement was also significant ($B = 0.34$, $\beta = .23$), and social anxiety remained only marginally significant ($p = .05$).

The final model demonstrated that procrastination is best predicted by specific perfectionism facets, particularly doubts about actions and entitlement, while social anxiety plays a secondary and diminishing role once perfectionistic cognitions are considered. Performance anxiety and the global perfectionism dimensions did not enter the model, indicating that broad perfectionism factors are less predictive than their specific cognitive components.

Overall, these results provide partial support for the hypothesis: both anxiety and perfectionism contribute to procrastination, but the most robust predictors emerge from the self-critical and entitlement-related facets of perfectionism.

3.4. FINDINGS

The present study examined the interrelations among anxiety, perfectionism, and procrastination across three generational cohorts (Generation X, Generation Y, and Generation Z). Overall, the results align with previous literature indicating that evaluative concerns and maladaptive self-regulation processes tend to co-occur and reinforce one another (Flett *et al.*, 2016; Stoeber, 2018). The findings offer additional evidence that these associations are robust across age groups, though their magnitude varies inter-generationally. Importantly, the generational differences observed were not merely quantitative but also suggest qualitative variations in how evaluative pressures are internalized and translated into avoidance behavior.

Consistent with theoretical accounts emphasizing the cognitive–emotional burden of high self-imposed standards (Hewitt & Flett, 1991; Stoeber & Otto, 2006), Generation Z reported the highest levels of rigid and self-critical perfectionism, as well as elevated anxiety and procrastination. This pattern suggests that younger adults may be more sensitive to evaluative pressures and comparison-based environments, which increase vulnerability to self-doubt and avoidance. In contemporary digital and highly competitive contexts, performance is often public, permanent, and easily comparable, potentially amplifying doubts about adequacy and intensifying self-monitoring tendencies. Generation X, by contrast, exhibited the lowest overall scores on these constructs, potentially reflecting more stable self-regulatory strategies or reduced exposure to contemporary achievement-oriented norms. Although the study did not assess cultural or technological factors directly, the generational gradient observed echoes previous work linking younger cohorts to heightened psychological strain in performance contexts (Vitasari *et al.*, 2010). These differences may also reflect developmental variations in identity consolidation, coping flexibility, and tolerance of ambiguity.

Regarding associations among variables, both performance anxiety and social anxiety correlated positively with global procrastination, supporting models proposing that avoidance of feared evaluation reinforces task delay (Steel & Klingsieck, 2016; Metin *et al.*, 2016). Notably, anxiety showed stronger associations with the soldiering facet (intentional delay) than with cyberslacking, suggesting that emotionally driven avoidance is more directly tied to decisional postponement than to task-unrelated online behaviour. This distinction may indicate that soldiering reflects a deliberate self-regulatory attempt to reduce emotional discomfort, whereas cyberslacking may be influenced more strongly by contextual or habitual digital engagement patterns. From a generational perspective, cyberslacking may represent a more normalized form of disengagement among Generation Z, embedded in everyday digital multitasking, whereas for older cohorts intentional delay may be more closely associated with performance apprehension and cognitive indecision.

The regression analysis further clarified the predictive value of perfectionism-related cognitions. Among the perfectionism facets, doubts about actions emerged as the strongest predictor of procrastination, consistent with findings that indecisiveness and fear of error undermine task initiation (Flett *et al.*, 2016). The entitlement facet of narcissistic perfectionism also predicted procrastination, indicating that individuals who hold inflated expectations about outcomes or personal standards may be more prone to disengage when performance risks inconsistency with those standards. Taken together, these findings suggest the existence of at least two partially distinct pathways to procrastination: a self-critical pathway driven by fear of inadequacy and chronic self-doubt, and a self-protective or self-enhancing pathway, in which disengagement serves to preserve a positive self-image in the face of potential failure. Although social anxiety initially contributed to the prediction model, its effect became non-significant when perfectionistic doubts and entitlement were included, suggesting that the influence of anxiety on procrastination may be partly mediated or cognitively reframed through perfectionistic beliefs. This pattern supports the interpretation that cognitive appraisals associated with perfectionism may constitute a more proximal mechanism underlying task delay than emotional distress *per se*.

These findings carry several theoretical implications. First, they underscore the importance of examining perfectionism at the facet level, given that different components display distinct associations with procrastination and may operate through partially divergent motivational pathways. Second, the generational differences identified in the present study suggest that age-related socio-emotional development and differential exposure to evaluative contexts may influence the way self-regulatory vulnerabilities manifest. Third, the results reinforce the perspective that procrastination is less indicative of deficient motivation and more reflective of a coping strategy aimed at managing perceived inadequacy, anticipated failure, or threats to self-image (Petrie, 2014). The intergenerational patterns observed further suggest that the meaning and functional role of procrastination may shift subtly across developmental stages and socio-cultural environments.

The present findings also extend existing research by clarifying how specific facets of perfectionism and anxiety interact in relation to procrastination across generational cohorts. Rather than emerging as a straightforward consequence of emotional discomfort, procrastination appears to represent a self-regulatory response shaped by evaluative concerns, internalized standards, and perceived threats to self-worth. The prominence of doubts about actions as the strongest predictor indicates that indecision and persistent self-questioning constitute central mechanisms underlying task postponement. The additional contribution of entitlement suggests that procrastination may stem not only from fear of inadequacy but also from tensions between elevated personal standards and the possibility of failing to meet them. Together, these patterns emphasize the multifaceted psychological processes underlying procrastination.

4. DISCUSSIONS AND CONCLUSIONS

From an applied perspective, the results point toward meaningful implications, particularly within educational contexts. Interventions aimed at reducing rigid self-criticism, fear of mistakes, and cognitive inflexibility may help attenuate maladaptive delay patterns, especially among younger adults who reported higher levels of vulnerability. Approaches that combine cognitive restructuring with self-compassion practices, emotional regulation strategies, and adaptive goal-setting may be particularly effective. At the institutional level, learning environments that prioritize mastery, constructive feedback, and reduced normative comparison may buffer the impact of excessive evaluative pressure. Given the generational gradient observed, prevention and intervention programs may benefit from being developmentally sensitive and responsive to the intensified performance visibility characteristic of contemporary academic and digital spaces.

Beyond individual-level processes, the generational differences identified in this study may also reflect broader psychosocial dynamics. Younger cohorts have frequently been described as experiencing heightened social comparison and increased feelings of loneliness despite constant connectivity. Within such climates, procrastination may take on a relational dimension: postponing action can serve as a temporary shield against anticipated criticism, rejection, or negative evaluation. In this sense, delay behavior may represent not only avoidance of tasks, but also avoidance of exposure. When performance becomes closely tied to identity and social validation, refraining from engagement may offer short-term protection of self-worth, even at the cost of long-term efficacy.

Several limitations must be acknowledged. The relatively small sample size (N= 155) constitutes the primary constraint of the present research and limits the generalizability of the findings. The uneven distribution across generations, including the absence of a Baby-Boomer cohort, further restricts the scope of

developmental comparisons. Accordingly, the study should be regarded as a pilot investigation offering preliminary evidence regarding intergenerational variations in anxiety, perfectionism, and procrastination. The cross-sectional design precludes causal inference, and reliance on self-report measures may introduce response biases. Additionally, generational classifications are necessarily broad and may not fully capture intra-cohort diversity shaped by socio-economic, cultural, or contextual factors.

Future research would benefit from employing larger and more heterogeneous samples, longitudinal designs, and multi-method assessments capable of capturing dynamic processes over time. Such approaches would allow for a more precise examination of mediation and moderation pathways, clarifying whether perfectionistic cognitions function as intermediary mechanisms linking anxiety to procrastination and whether these associations vary across generational cohorts. Further expanding the relational and value-based dimensions of analysis may deepen understanding of how identity formation, perceived social worth, and performance-contingent self-evaluation interact with self-regulatory vulnerabilities. In particular, future studies may explore the hypothesis that generational differences in procrastination are partially explained by the extent to which self-worth is experienced as contingent upon achievement and external evaluation.

In light of these considerations, the present pilot study contributes to a more differentiated conceptualization of procrastination as a self-protective strategy embedded within both intrapersonal dynamics and generationally shaped evaluative climates. By showing that procrastination is more strongly linked to self-evaluative doubt and entitlement than to anxiety alone, the findings suggest that models of self-regulation may be refined through a more explicit integration of internal standards and perceived self-worth as central mechanisms shaping procrastination across generations.

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TRANSLATION AND VALIDATION OF THE ROMANIAN VERSION
OF THE ITC-SENSE OF PRESENCE INVENTORY (ITC-SOPI)
FOR ASSESSING PRESENCE IN VIRTUAL REALITY

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Abstract

Presence is a multidimensional construct for VR psychological interventions, but no standardized Romanian measure existed. This study translated and culturally adapted the ITC–Sense of Presence Inventory (ITC-SOPI) and tested its psychometric properties on a sample of 241 Romanian-speaking participants, after a VR experience. Confirmatory factor analysis supported the intended structure with four factors. ML estimation indicated a modest fit (CFI = .762, TLI = .749, RMSEA = .070, SRMR = .080), while an ordinal sensitivity analysis WLSMV showed an improved fit (CFI = .906, TLI = .900, RMSEA = .058, SRMR = .097). Subscales showed acceptable-to-excellent internal consistency ($\alpha=.765-.893$). Presence-related subscales correlated positively with IPQ dimensions ($r=.33-.64$). Negative Effects subscale correlated strongly with simulator sickness (SSQ; $r=.48-.71$) and was the main predictor of total SSQ ($\beta=.711$). Overall, the Romanian ITC-SOPI appears reliable and valid for VR research.

Keywords: presence, virtual reality, ITC-SOPI, Romanian adaptation, validation.

In recent years, the accelerated development of virtual reality (VR) technologies has made it possible to create increasingly complex and multisensory digital environments (Melo *et al.*, 2022), which facilitate users' real-time interaction while at the same time providing them with a sense of control over their own actions (Saran, 2025). VR applications are used in a wide range of domains, from mental health (Dellazizzo *et al.*, 2020) to professional training (Steen *et al.*, 2024), education (Kamińska *et al.*, 2019), and entertainment (Lai *et al.*, 2019). The evaluation of these applications does not target only the degree of symptom reduction or the increase in skill level, but also users' subjective experience, which has been emphasized as an important outcome dimension in VR-based interventions (Weber *et al.*, 2021).

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One of the most important concepts for understanding how people experience VR is presence (sense of presence). In the existing literature, multiple definitions have been proposed. For example, Kisker *et al.* (2021) define presence as the subjective feeling of existing inside a virtual world, with awareness of real surroundings and hardware gradually fading away. Regardless of the diversity of definitions that have emerged over time, most authors converge on the view that presence should be understood as a multidimensional construct that calls for cross-disciplinary perspectives and user-focused assessments of subjective psychological experience (Baños *et al.*, 2008; Williams *et al.*, 2025). Theories of presence are often grouped into descriptive and structural models (Diemer *et al.*, 2015). Descriptive models, such as that of Schubert *et al.* (2001), refer to the components of presence experienced in virtual reality, such as spatial presence, involvement, and realness. At the same time, structural models, such as that of Wirth *et al.* (2007), propose that spatial presence can be described as the outcome of a two-step process in which users first construct a mental model of the mediated environment – a stage influenced by attention allocation and visual spatial imagery – and then, in a second stage, involvement and suspension of disbelief, together with this model, give rise to the subjective feeling of being located in the virtual space. Therefore, in line with this distinction, descriptive models focus on the phenomenological components of the presence experience, whereas structural models focus on the underlying cognitive and perceptual processes in the user’s mind that generate this experience of presence (Diemer *et al.*, 2015).

Empirical findings on the role of presence as a mechanism of change in VR-based interventions are mixed. Pavic *et al.*’s (2023) results demonstrate that, when eliciting positive emotions, spatial and social presence mediate the relationship between immersion and subjective arousal. Other studies have shown that there is a relationship between presence and emotions, suggesting that both highly anxious and highly relaxing virtual environments are associated with increased presence, and that presence significantly predicts a range of emotional outcomes, particularly in anxious environments such as virtual park scenarios (Riva *et al.*, 2007). In contrast to these findings, some studies indicate that presence does not consistently predict clinical improvement, with cognitive variables such as self-efficacy and changes in dysfunctional beliefs emerging as stronger predictors of symptom reduction (Price & Anderson, 2007; Côté & Bouchard, 2009).

Given the central role of presence in modulating users’ emotional and subjective responses in VR (Riva *et al.*, 2007; Diemer *et al.*, 2015), valid and reliable instruments for measuring presence are essential. Although physiological and behavioral indices of presence have also been explored in the literature (Cummings & Bailenson, 2016), empirical studies most commonly assess presence

using self-report questionnaires (Halbig & Latoschik, 2021). Among the numerous instruments developed to assess presence, the Witmer and Singer Presence Questionnaire (PQ; Witmer & Singer, 1998), the Slater–Usoh–Steed questionnaire (SUS; Slater *et al.*, 1994; Usoh *et al.*, 2000), the IGroup Presence Questionnaire (IPQ; Schubert, Friedmann, & Regenbrecht, 2001), the ITC–Sense of Presence Inventory (ITC-SOPI; Lessiter, Freeman, Keogh, & Davidoff, 2001), the Multimodal Presence Scale (MPS; Makransky, Jensen, & Aaby, 2017), and the Temple Presence Inventory (TPI; Lombard, Ditton, & Weinstein, 2009) stand out as some of the most frequently used and recommended measures in contemporary VR research (Hein *et al.*, 2018; Bareišytė *et al.*, 2024; Kukshinov *et al.*, 2025).

Among all these instruments, the ITC–Sense of Presence Inventory (Lessiter *et al.*, 2001) is the most relevant for the purposes of the present study. It was designed to assess presence across different media contexts and to facilitate the comparability of results between different laboratories (Lessiter *et al.*, 2001). The inventory developed by Lessiter *et al.* (2001) includes a total of 44 items. These are organized into four distinct dimensions: Engagement, Negative Effects, Ecological Validity/Naturalness, and the Sense of Physical Space. Sense of Physical Space reflects the feeling of being spatially located in the mediated environment and being able to act within it; Engagement captures interest and psychological involvement with the content; Ecological Validity/Naturalness refers to the perceived realism and naturalness of the environment and events; whereas Negative Effects includes symptoms such as dizziness, nausea, or discomfort that may accompany media exposure (Lessiter *et al.*, 2001). The ITC-SOPI has also been translated and validated for other language populations. For example, Vasconcelos-Raposo *et al.* (2019) adapted the ITC-SOPI for Portuguese-speaking participants, confirming the four-factor structure and reporting acceptable internal consistency for all subscales. They concluded that the instrument is suitable for assessing presence in VR research with Portuguese samples (Vasconcelos-Raposo *et al.*, 2019).

To our knowledge, no standardized and psychometrically validated presence scale is currently available in Romanian. Existing VR studies conducted with Romanian-speaking samples either rely on instruments developed in other languages or use ad hoc translations that have not undergone a formal process of cultural adaptation and validation. In this context, we chose the ITC-SOPI as the basis for developing a Romanian measure of presence that can be applied across different media configurations, so that it can be used in a wide range of study designs (e.g., experimental laboratory studies, clinical interventions, educational applications). The aim of the present study is therefore to translate and culturally adapt the ITC-SOPI for the Romanian population and to examine its factorial structure, reliability, and construct validity in a sample of Romanian VR users. In the current study, the scale was initially adapted for use in virtual reality contexts, and we intend to extend its application to other media environments in future research.

METHODS

PARTICIPANTS

A total of 241 Romanian-speaking participants took part in the study and completed all measures included in the present analyses. Participants were recruited from the Faculty of Psychology and Educational Sciences in Cluj-Napoca. The sample included 223 women (92.53%) and 18 men (7.47%). All participants had been exposed to a virtual reality (VR) scenario before completing the questionnaires. Regarding prior experience with VR, 39.49% of the participants reported that they had previously used VR in an experimental context. In terms of VR-related knowledge, participants rated themselves on a four-point scale as follows: none (48.55%), basic (46.89%), intermediate (3.73%), and expert (0.83%). Unfortunately, participants' age was not collected in the present study.

MEASURES

ITC–Sense of Presence Inventory (ITC-SOPI)

The participants' sense of presence was measured using the ITC-SOPI (Lessiter *et al.*, 2001). The inventory is a cross-media instrument, in which the items (44; 5-point Likert scale) are grouped into four subscales: Sense of Physical Space, Engagement, Ecological Validity/Naturalness, and Negative Effects. It is divided into Part A (6 items) and Part B (38 items). Following the authors' recommendation, we removed item B6, as it was not included in the scoring of any subscale and was specific to the television context (Lessiter *et al.*, 2001). Thus, the final version of the ITC-SOPI adapted for the Romanian population within the present study contains 43 items.

Regarding the back-translation procedure used to adapt the scale for the Romanian population, established multi-step guidelines were followed (Sousa & Rojjanasrirat, 2011; Zhao *et al.*, 2024). First, two bilingual psychologists, Romanian native speakers, independently translated the inventory items (forward translations) from English into Romanian. In line with recommendations to prioritize conceptual equivalence over literal equivalence, the two translations were then compared item by item and reconciled into a single preliminary version through discussion and consensus. Next, this preliminary Romanian version underwent a blind back-translation into English conducted by a third bilingual translator. This translator was not involved in the forward translation process and was kept "blind" to both the original instrument and the study objectives (Sousa & Rojjanasrirat, 2011). The back-translated version was thereafter compared with the original English ITC-SOPI by all three translators acting as an expert committee. Discrepancies in meaning, nuance, and wording were discussed until consensus

was reached, and minor adjustments were made to the Romanian phrasing to improve conceptual and semantic equivalence.

Igroup Presence Questionnaire (IPQ)

Another questionnaire used to measure the sense of presence and to use it in the study analyses is the IPQ (Schubert *et al.*, 2001). It includes 14 items rated on a 7-point Likert-type scale, having 3 subscales (Spatial Presence, Involvement, Experienced Realism) and one General Presence item, being one of the most established presence questionnaires (Tran *et al.*, 2024).

Simulator Sickness Questionnaire

The SSQ (Kennedy *et al.*, 1993) was administered to quantify cybersickness following VR exposure. Participants rated 16 symptoms on a 4-point scale, and responses were combined to obtain Nausea, Oculomotor Disturbance, and Disorientation scores, as well as an overall SSQ total score (Kennedy *et al.*, 1993). The SSQ is widely used in the VR literature to assess simulator sickness (Bimberg *et al.*, 2020).

PROCEDURE

Google Forms was the platform used to recruit participants. In the first stage, participants signed up to take part in the study by indicating the day on which they wished to participate. Subsequently, in the second stage, those who registered received a second Google Forms link in which they selected the time slot at which they wanted to attend the experimental session in the laboratory. In the third stage, after completing informed consent and meeting the eligibility criteria, participants completed a short familiarization session with the virtual reality application, Nature Treks VR (Greener Games, 2019), conducted in the research laboratory. The exclusion criteria included minors, residents of foreign countries, non-Romanian speakers; a history of mental disorders, photophobia; and an epilepsy diagnosis or the presence/experience of epileptic seizures.

After participants familiarized themselves with the virtual reality application, they engaged in the experimental task. This task consisted of exploring a virtual reality environment called White Winter (Björling *et al.*, 2022), in which participants could observe various changes in the landscape and interact with characters within the application. During this task, Meta Quest 2 virtual reality headsets were used (Meta Platforms, Inc., 2020). After completing the experimental task, participants filled out the set of instruments selected for the statistical analyses. Only one session took place, and the total duration of the entire procedure was approximately 30 minutes.

ETHICAL CONSIDERATIONS

Ethical approval was granted by the Ethics Committee of the Faculty of Psychology and Educational Sciences, Babeş- Bolyai University. All participants provided written informed consent prior to participation. The ethical approval number is No. 16.704/13.10.2025.

DATA ANALYSIS

To validate the four-factor structure of the ITC-SOPI Inventory – comprising Engagement, Sense of Physical Space, Ecological Validity, and Negative Effects – we conducted a Confirmatory Factor Analysis (CFA). The four-factor model was specified a priori based on the original ITC-SOPI structure. First, we used maximum likelihood (ML) estimation as the primary CFA estimator because the ITC-SOPI items use a 5-point response format that is commonly treated as approximately continuous in SEM, and ML facilitates comparability with prior validation studies. Item distributions were examined via skewness and kurtosis to assess univariate departures from normality and to inform the interpretation of ML-based model fit. To assess the model's quality, we reported the SRMR and RMSEA (with a 90% confidence interval), alongside the chi-square test. Furthermore, the comparative indices TLI and CFI were calculated, following the standard SEM guidelines proposed by Brown (2015) and Bentler and Bonett (1980). Second, as a sensitivity analysis, we re-estimated the CFA treating the items as ordinal indicators using the WLSMV estimator with robust standard errors. The aim was to assess whether conclusions regarding the factor structure were robust to the estimation method. Third, in an exploratory post-hoc step within the CFA, we sought to improve model parsimony by removing items with standardized factor loadings below .40, even though loadings as low as .30 are often considered acceptable in initial adaptations (Brown, 2015). Cronbach's alpha coefficient was used to calculate internal consistency regarding the ITC-SOPI scores (Cronbach, 1951). Convergent validity was evaluated through Pearson correlations between the ITC-SOPI subscales and the IPQ subscales, as well as between the Negative Effects subscale and the SSQ scores (Campbell & Fiske, 1959). Furthermore, we employed a multiple linear regression model, defining the four inventory factors as predictors and the total SSQ score as the primary dependent variable. All these analyses were conducted in order to examine the extent to which experiences related to presence and negative effects explain the variance of simulator sickness (Cohen, Cohen, West, & Aiken, 2003). All analyses were based on complete cases ($N = 241$). The comprehensive statistical workflow, spanning from the initial CFA to the final regression models, was carried out using the JASP (Version 0.18; JASP Team, 2023).

RESULTS

CONFIRMATORY FACTOR ANALYSIS

CFA was used to test the original four-factor structure of the ITC-SOPI. Prior to CFA, item distributions were inspected for univariate normality (skewness and kurtosis). Several items showed pronounced skewness ($|\text{skewness}| > 2$), including Item_6_A (skewness = -3.806) and Item_26_B (skewness = 3.719). Kurtosis was elevated for a small number of items, with three items exceeding $|\text{kurtosis}| > 10$ (Item_5_A = 10.55; Item_6_A = 15.97; Item_26_B = 14.94). These departures were considered when interpreting ML-based fit indices. The initial model, which included all items, yielded a statistically significant chi-square value, $\chi^2(854) = 1858.23$, $p < .001$, as expected given the relatively large number of items and parameters (Bentler & Bonett, 1980). Overall model fit was modest, with CFI = .762 and TLI = .749. The RMSEA was .070, with a 90% CI [.066, .074], and the SRMR was .080. Standardized factor loadings for the initial (full) CFA model are reported in Table 1.

Table 1

Standardized factor loadings for the Romanian ITC-SOPI: initial (full) four-factor CFA model (ML)

Factor	Indicator	Std. estimate	Std. Error	p
Sense of Physical Space	Item_4_B	0.568	0.048	< .001
	Item_7_B	0.591	0.051	< .001
	Item_9_B	0.604	0.053	< .001
	Item_12_B	0.597	0.052	< .001
	Item_13_B	0.457	0.059	< .001
	Item_18_B	0.765	0.037	< .001
	Item_19_B	0.507	0.067	< .001
	Item_22_B	0.381	0.047	< .001
	Item_23_B	0.569	0.047	< .001
	Item_24_B	0.427	0.066	< .001
	Item_25_B	0.594	0.055	< .001
	Item_28_B	0.618	0.050	< .001
	Item_29_B	0.356	0.055	< .001
	Item_31_B	0.478	0.064	< .001
	Item_33_B	0.462	0.055	< .001
	Item_34_B	0.680	0.046	< .001
	Item_35_B	0.545	0.050	< .001
Item_36_B	0.739	0.034	< .001	
Item_38_B	0.737	0.048	< .001	

Table 1 (continued)

Factor	Indicator	Std. estimate	Std. Error	p
Engagement	Item_1_A	0.413	0.067	< .001
	Item_3_A	0.411	0.060	< .001
	Item_4_A	0.677	0.055	< .001
	Item_5_A	0.390	0.094	< .001
	Item_6_A	0.591	0.071	< .001
	Item_1_B	0.606	0.067	< .001
	Item_2_B	0.669	0.048	< .001
	Item_3_B	0.385	0.066	< .001
	Item_8_B	0.740	0.046	< .001
	Item_16_B	0.544	0.054	< .001
	Item_17_B	0.449	0.075	< .001
	Item_30_B	0.396	0.058	< .001
	Item_32_B	0.667	0.053	< .001
Ecological validity	Item_5_B	0.751	0.044	< .001
	Item_11_B	0.785	0.044	< .001
	Item_15_B	0.806	0.030	< .001
	Item_20_B	0.409	0.066	< .001
	Item_27_B	0.537	0.062	< .001
Negative Effects	Item_2_A	0.422	0.060	< .001
	Item_10_B	0.779	0.039	< .001
	Item_14_B	0.769	0.043	< .001
	Item_21_B	0.740	0.051	< .001
	Item_26_B	0.515	0.067	< .001
	Item_37_B	0.769	0.043	< .001

Note. λ = standardized factor loading; SE = standard error; N = 241. Items labeled “_A” and “_B” correspond to Parts A and B of the inventory. The CFA model includes 43 items; item B6 was excluded following the original authors’ scoring recommendations.

As a confirmatory sensitivity analysis of the same a priori four-factor model, the four-factor model was re-estimated treating the 5-point Likert items as ordinal using the WLSMV estimator with robust standard errors. This ordinal solution indicated improved global fit: $\chi^2(854) = 1536.54$, $p < .001$; CFI = .906; TLI = .900; RMSEA = .058, 90% CI [.053, .062]; SRMR = .097. Overall, the pattern supported the intended four-factor structure.

EXPLORATORY POST-HOC REFINEMENT (ITEM TRIMMING)

As an exploratory post-hoc refinement, several items displayed relatively weak standardized loadings (*i.e.*, $< .40$) on their intended factors. Therefore, we specified an additional, more parsimonious four-factor solution in which these

low-loading items were excluded. This trimming was treated as an exploratory refinement to reduce model complexity, acknowledging that in early adaptations factor loadings around .30 may still be considered acceptable (Brown, 2015). The revised model yielded slightly better fit indices, $\chi^2(660) = 1485.24$, $p < .001$, CFI = .787, TLI = .773, RMSEA = .072 (90% CI [.067, .077]), and SRMR = .077. In the trimmed model all retained indicators loaded at .40 or higher on their intended factors. However, the improvement in global fit was marginal and did not alter the substantive interpretation of the four-factor structure. Therefore, we retained the full 43-item version to preserve content validity and comparability with the original instrument. Although the overall fit remained below commonly cited “good-fit” criteria (e.g., CFI/TLI $\geq .95$; Hu & Bentler, 1999), overall, both models were broadly consistent with the original four-factor ITC-SOPI structure, although global fit was modest.

INTERNAL CONSISTENCY

Internal consistency of the Romanian ITC-SOPI subscales was examined using Cronbach’s alpha with 95% confidence intervals. The Sense of Physical Space subscale showed excellent reliability ($\alpha = .893$, 95% CI [.87, .917]). Engagement also demonstrated good internal consistency ($\alpha = .805$, 95% CI [.77, .841]). The Ecological Validity/Naturalness subscale showed acceptable to good reliability ($\alpha = .765$, 95% CI [.713, .817]), and the Negative Effects subscale likewise displayed good internal consistency ($\alpha = .819$, 95% CI [.787, .852]). The internal consistency for all subscales shows good results. Reliability analyses revealed good internal consistency across all four subscales of the inventory.

CONVERGENT VALIDITY

Convergent validity was evaluated by correlating the ITC-SOPI subscales with the dimensions of the Igroup Presence Questionnaire (IPQ) (Table 2).

The three subscales (Sense of Physical Space, Engagement, and Ecological Validity/Naturalness) of the ITC-SOPI correlated positively with the presence measured by the IPQ dimensions (Spatial Presence, Involvement, Experienced Realism) ($r = .33-.64$, $p < .001$). The highest association was observed between Sense of Physical Space and Spatial Presence ($r = .64$, $p < .001$) and between Engagement and General Presence ($r = .54$, $p < .001$). In contrast, the Negative Effects subscale showed small and largely non-significant associations with the IPQ dimensions ($r = -.18$ to $-.03$). This suggests that this subscale is partially distinct from the presence measured in the other scales, as expected (Lessiter *et al.*, 2001). At the same time, convergent validity with simulator sickness (SSQ) was examined by correlating the Negative Effects subscale from the ITC-SOPI with the SSQ subscales and the SSQ total score (Table 3).

Table 2

Pearson correlations between ITC-SOPI subscales and IPQ dimensions (convergent validity)

ITC-SOPI subscale	IPQ dimension	Pearson's r	p
Sense of Physical Space (ITC-SOPI)	– General Presence (IPQ)	0.593	< .001
Sense of Physical Space (ITC-SOPI)	– Spatial Presence (IPQ)	0.641	< .001
Sense of Physical Space (ITC-SOPI)	– Involvement (IPQ)	0.436	< .001
Sense of Physical Space (ITC-SOPI)	– Experienced Realism (IPQ)	0.448	< .001
Engagement (ITC-SOPI)	– General Presence (IPQ)	0.543	< .001
Engagement (ITC-SOPI)	– Spatial Presence (IPQ)	0.509	< .001
Engagement (ITC-SOPI)	– Involvement (IPQ)	0.466	< .001
Engagement (ITC-SOPI)	– Experienced Realism (IPQ)	0.332	< .001
Ecological Validity/Naturalness (ITC-SOPI)	– General Presence (IPQ)	0.496	< .001
Ecological Validity/Naturalness (ITC-SOPI)	– Spatial Presence (IPQ)	0.517	< .001
Ecological Validity/Naturalness (ITC-SOPI)	– Involvement (IPQ)	0.394	< .001
Ecological Validity/Naturalness (ITC-SOPI)	– Experienced Realism (IPQ)	0.540	< .001
Negative Effects (ITC-SOPI)	– General Presence (IPQ)	-0.179	.005
Negative Effects (ITC-SOPI)	– Spatial Presence (IPQ)	-0.095	.142
Negative Effects (ITC-SOPI)	– Involvement (IPQ)	-0.032	.624
Negative Effects (ITC-SOPI)	– Experienced Realism (IPQ)	-0.057	.375

Note. Pearson correlations are reported (two-tailed). N = 241. ITC-SOPI = ITC–Sense of Presence Inventory; IPQ = Igroup Presence Questionnaire.

Table 3

Pearson correlations between ITC-SOPI Negative Effects and SSQ scores

ITC-SOPI subscale	SSQ subscale	Pearson's r	p
Negative Effects (ITC-SOPI)	– Nausea (SSQ)	0.484	< .001
Negative Effects (ITC-SOPI)	– Oculomotor (SSQ)	0.708	< .001
Negative Effects (ITC-SOPI)	– Disorientation (SSQ)	0.625	< .001
Negative Effects (ITC-SOPI)	– Total (SSQ)	0.689	< .001

Note. Pearson correlations are reported (two-tailed). N = 241. ITC-SOPI = ITC–Sense of Presence Inventory; SSQ = Simulator Sickness Questionnaire

As we can observe, the Negative Effects subscale correlated positively with all SSQ dimensions (*e.g.*, Disorientation, $r = .63$, $p < .001$) and the SSQ Total score ($r = .69$, $p < .001$). Overall, this pattern is consistent with the intended conceptualization of the Negative Effects subscale as an indicator of adverse reactions/discomfort (Lessiter *et al.*, 2001) and with its convergence with simulator sickness, as measured by the SSQ.

MULTIPLE LINEAR REGRESSION PREDICTING TOTAL SSQ SCORES FROM ITC-SOPI SUBSCALES

A multiple linear regression analysis was conducted with the Total SSQ score as the dependent variable and the four ITC-SOPI subscales as simultaneous predictors. The overall model was significant, $F(4, 236) = 59.01$, $p < .001$, and explained 50.0% of the variance in simulator sickness, $R^2 = .50$, adjusted $R^2 = .492$. Standardized regression coefficients indicated that Negative Effects was the strongest predictor of Total SSQ ($B = 26.727$, $SE = 1.766$, $\beta = .711$, $t = 15.137$, $p < .001$), whereas the other subscales showed mixed associations with simulator sickness when controlling for the remaining predictors (Engagement: $B = 10.311$, $\beta = .188$, $p = .004$; Ecological Validity/Naturalness: $B = -4.898$, $\beta = -.146$, $p = .037$; Sense of Physical Space: $B = -2.840$, $\beta = -.064$, $p = .415$). In the multiple regression model, Negative Effects emerged as the dominant predictor of Total SSQ, while the remaining ITC-SOPI subscales contributed weakly or not at all after controlling for the other predictors.

DISCUSSION

The present study aimed to translate, culturally adapt, and psychometrically evaluate the Romanian version of the ITC–Sense of Presence Inventory (ITC-SOPI) for its use in studies involving virtual reality applications. A comprehensive analytical strategy was employed, including Confirmatory Factor Analysis (CFA), reliability analyses, convergent validity correlations, and multiple regression. Overall, the results provide solid preliminary evidence that the Romanian-language ITC-SOPI yields reliable subscale scores and supports the intended four-factor conceptualization in a Romanian VR sample.

Regarding structural validity and model evaluation, the CFA results were, broadly, in line with the original four-factor structure proposed by Lessiter *et al.* (2001), comprising Sense of Physical Space, Engagement, Ecological Validity/Naturalness, and Negative Effects. As is common in the case of complex, multidimensional instruments with many indicators and parameters, the overall model fit was modest in the initial model ($CFI = .76$, $TLI = .75$). Although the CFA was initially estimated using ML (treating 5-point items as approximately continuous), the ordinal WLSMV sensitivity analysis provided converging evidence for the four-factor structure. Future studies are encouraged to prioritize ordinal estimators (*e.g.*, WLSMV) when validating the Romanian ITC-SOPI in broader Romanian samples.

A more parsimonious CFA model, conducted by excluding items with standardized loadings $< .40$, produced only marginal changes in global fit, indicating that item reduction did not significantly alter the underlying structure (CFI = .787, TLI = .773).

Even though the ML-estimated CFA models (both the initial full-item model and the trimmed ML model) do not meet the most stringent conventional fit thresholds (*e.g.*, Hu & Bentler, 1999), Marsh *et al.* (2004) argue that such “golden rules” are often overly restrictive when applied to real-world psychological data and complex, multidimensional instruments. According to their critique, rigid cutoffs can lead to the unfair rejection of theoretically sound models, especially in item-level analyses where inherent complexity naturally leads to lower fit indices such as the CFI and TLI. It is worth noting that while the comparison-based indices (CFI/TLI) were modest, the absolute fit indices in our study (RMSEA = .070; SRMR = .080) remained within acceptable ranges (Brown, 2015), providing additional support for the model's viability.

Given that the more parsimonious model did not provide a substantial improvement in fit, it was concluded that the modest indices reflect the inherent complexity of the construct rather than specific psychometric failures of the items. Consequently, we decided to retain the original 43-item version of the inventory to preserve its full theoretical breadth and content validity, since the reduction did not significantly change the underlying structural performance.

The reliability analyses further support this decision, as the internal consistency for all subscales showed good to excellent results. The high Cronbach's alpha values (ranging from .765 to .893) indicate that the Romanian version of the ITC-SOPI is a highly dependable instrument for measuring the different dimensions of the VR experience.

Beyond the structural evaluation, the convergent validity results of the inventory provide solid evidence for the psychometric robustness of the Romanian version of the ITC-SOPI. The positive correlations between the Romanian ITC-SOPI inventory and the IPQ dimensions confirm that the instrument accurately captures the core constructs of the presence experience. The particularly high association between Sense of Physical Space and Spatial Presence (IPQ) aligns with the conceptual framework proposed by Lessiter *et al.* (2001), confirming that the Romanian version effectively captures the spatial dimension of presence as intended by the original authors.

Another important result of this study concerns the relationship between the Negative Effects subscale and simulator sickness. Negative Effects showed a strong and specific positive correlation with all dimensions of the Simulator Sickness Questionnaire (SSQ), especially with the SSQ Total score ($r = .69$). In addition, the fact that this subscale showed non-significant or weak associations with the IPQ dimensions further supports its discriminant validity and conceptual distinctiveness, as adverse reactions represent a separate dimension of the VR experience from presence itself.

The predictive power of the instrument was further demonstrated through the multiple linear regression analysis. The model explained 50% of the variance in total simulator sickness, a substantial effect for psychological research. The emergence of Negative Effects as the dominant predictor ($\beta = .711$) validates the role of this subscale as a primary indicator of user discomfort, suggesting that simulator sickness is primarily associated with Negative Effects rather than with the other presence-related subscales when considered simultaneously.

Regarding practical implications, a validated Romanian version of the ITC-SOPI provides a standardized presence measure that supports the cross-laboratory comparability highlighted for the original instrument. In Romania, this can facilitate systematic comparisons of presence and negative effects across VR studies and VR setups (Lessiter *et al.*, 2001). In applied VR development, presence is commonly measured – most often with validated questionnaires – to document user experience and interaction in both research and industry (Schwind *et al.*, 2019). Because a subjective sense of “being there” is one component contributing to realistic responding in immersive virtual environments (Slater, 2009) and presence has been described as a consensual metric for evaluating virtual-environment effectiveness (Vasconcelos-Raposo *et al.*, 2021), routine presence assessment can inform evaluation and iterative design decisions for VR applications where immersion is a key goal.

Taken together, the pattern of results – modest ML-based global fit, improved fit under ordinal WLSMV estimation, and strong reliability and criterion-related validity – provides initial evidence supporting the use of the Romanian ITC-SOPI in Romanian VR research.

LIMITS AND FUTURE DIRECTIONS

The first limitation of this study is the fact that the data were collected at a single measurement point, after a single exposure to VR, which prevented the evaluation of test–retest reliability. We recommend that future studies examine stability across several repeated sessions. The second limitation is the lack of formal testing of the instrument in a pilot study before the main data collection in order to evaluate face validity and comprehensibility (Sousa & Rojjanasrirat, 2011). This stage could help identify potentially ambiguous items. Third, convergent validity was assessed using the IPQ, and not an additional Romanian-language presence instrument. Although a local measure has been mentioned in Romanian contexts (Oprîș, 2012), it is not currently available as a fully peer-reviewed, standardized, and publicly documented instrument, which limited local comparability across instruments. It is recommended to examine convergence with Romanian presence measurement instruments, if and when they become validated and accessible. The fourth limitation is the predominantly female sample, which may limit generalizability to other demographic groups. Replication on more heterogeneous samples may strengthen external validity. Unfortunately, participants’ age was not

collected, which prevented us from describing the age distribution of the sample and from examining potential age-related differences in presence or cybersickness. Accordingly, the present findings should be considered an initial validation of the Romanian ITC-SOPI. Future studies should recruit samples with a more even gender distribution, collect age data, and test measurement invariance across gender and age to evaluate whether the scale operates equivalently across demographic groups.

CONCLUSIONS

In conclusion, although the global fit indices reflect the inherent complexity of modeling a multidimensional instrument with 43 items, the overall psychometric profile of the Romanian-language ITC-SOPI is highly favorable. The combination of high internal consistency, strong convergent validity with the IPQ, and robust prediction of simulator sickness justifies the use of the full 43-item version. Overall, the present study provides initial validation evidence that the Romanian ITC-SOPI yields reliable scores and shows construct validity in a Romanian VR sample, supporting its use in Romanian VR research while future work should confirm invariance across gender and age.

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FAMILY ARCHITECTURE – THE MATRIX OF RELATIONSHIPS AND SOCIO-EMOTIONAL COMPETENCIES IN ADOLESCENCE

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Abstract

Background: Research on family diversity frequently emphasizes structural forms, yet adolescents' socio-emotional development is more proximally shaped by relational organization (attachment, emotional communication, and support for autonomy). To integrate these layers, this paper proposes the construct of family architecture – an integrative matrix linking family structural configuration with relational processes and adolescents' socio-emotional competences.

Method: Cross-sectional survey of Romanian adolescents (N = 250; 16–18 years) from Olt County: two-parent families (n = 160), single-parent families (n = 68), and alternative-care settings (institutional/professional foster care; n = 22). Measures included the Relationship Scales Questionnaire (RSQ), Inventory of Parent and Peer Attachment (IPPA; mother, father, peers), Emotional Intelligence Inventory (INEM; self-report), Friedmann Emotional Maturity Scale, and Attachment Styles and Self-Esteem Inventory (ASSI). Nonparametric tests (Kruskal–Wallis, χ^2) and multiple regression were used. Effect sizes and 95% confidence intervals (CIs) are reported where applicable.

Results: Family type was associated with attachment style distribution ($\chi^2(6) = 17.34$, $p = 0.008$, Cramer's V = 0.19, 95% CI [0.05, 0.25]). Significant differences by family type emerged for maternal trust (H = 24.40, $p < 0.001$, $\varepsilon^2 = 0.09$) and communication (H = 20.21, $p < 0.001$, $\varepsilon^2 = 0.07$), paternal trust (H = 14.35, $p = 0.001$, $\varepsilon^2 = 0.05$) and communication (H = 12.79, $p = 0.002$, $\varepsilon^2 = 0.04$), and maternal total attachment (H = 12.74, $p = 0.002$, $\varepsilon^2 = 0.04$); peer attachment did not differ by family type ($ps > 0.10$). Trait emotional intelligence (INEM global and subscales) did not differ by family type (all $ps \geq .075$; $\varepsilon^2 \leq 0.01$), whereas emotional maturity showed a small difference (H = 8.39, $p = 0.015$, $\varepsilon^2 = 0.03$). Attachment style was related to global self-esteem levels ($\chi^2(6) = 12.72$, $p = 0.048$, V = 0.16, 95% CI [0.00, 0.22]). In regression models controlling for parental attachment, peer attachment uniquely predicted global trait EI ($R^2 = 0.067$, $\beta = 0.162$, $p = 0.012$, B = 0.026, 95% CI [0.006, 0.046]) and understanding others' emotions ($R^2 = 0.066$, $\beta = 0.242$, $p < 0.001$, B = 0.041, 95% CI [0.019, 0.063]).

Conclusions: Findings support the family architecture framework: structural family type relates to parental attachment quality, while peer attachment may partially compensate in the domain of perceived socio-emotional competences. Given the cross-sectional, self-report design and the small alternative-care subsample, interpretations are associative and context-bound.

Keywords: family architecture, family type, attachment, adolescence, socio-emotional competences, emotional intelligence.

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1. INTRODUCTION

Adolescence is a developmental period marked by accelerated biological changes, increasing social complexity, and a progressive shift from parental dependence to broader relational networks. Across this transition, socio-emotional competences – such as emotional understanding, regulation, and relational functioning – are central resources for adaptation, academic engagement, and long-term well-being (Arnett, 2000). These competences develop within relational contexts that include caregivers and peers, and their formation depends not only on family structure but also on how relationships are organized in daily life.

Family research has historically described family diversity through structural indicators (*e.g.*, number of caregivers in the household, divorce, or out-of-home placement). However, structural indicators alone often have limited explanatory power because adolescents' outcomes are more proximally shaped by relational processes: emotional availability of caregivers, trust, communication patterns, and autonomy support. For example, meta-analytic evidence suggests that parenting quality and relational processes are robust correlates of adolescent adjustment, whereas the effects of structural transitions (*e.g.*, divorce) are heterogeneous and contingent on contextual mechanisms (van Dijk *et al.*, 2020).

To integrate structure and process, we propose the construct of family architecture. Family architecture refers to an integrative matrix connecting: (a) family structural configuration (family type and stability/continuity of caregiving), (b) relational organization (attachment representations and relationship quality, including trust, communication, and alienation), and (c) socio-emotional competences (*e.g.*, emotional intelligence, emotional maturity, and self-evaluative outcomes). This construct aims to move beyond a purely descriptive-structural view and foreground the relational mechanisms through which family forms become developmentally meaningful.

Adolescence is also characterized by a functional rebalancing toward peers. Peer attachment becomes increasingly salient as a secure base, a source of emotional disclosure, and a context for social learning (Delgado *et al.*, 2022). Within the family architecture framework, peers may play a compensatory role, particularly when parental relationships are strained or caregiving is disrupted.

2. FAMILY ARCHITECTURE – A CONCEPTUAL AND OPERATIONAL FRAMEWORK

In this article, family architecture is defined as the configuration of structural and relational features through which adolescents experience their family environment. Structurally, family architecture captures the type of family/care arrangement (two-parent, single-parent, or alternative care). Relationally, it encompasses attachment-related representations (secure, preoccupied, dismissing, fearful), and relationship-specific dimensions of trust, emotional communication, and alienation

toward caregivers and peers. Conceptually, family architecture also includes autonomy support and emotion socialization practices; in the present dataset, autonomy support was not directly measured and is therefore treated as a theoretical component to be operationalized in future work. Socio-emotional competences are modeled as outcomes that may be shaped by both structural factors (*e.g.*, caregiver continuity) and relational mechanisms (*e.g.*, trust and communication), with peers acting as a potential compensatory resource during adolescence.

3. OBJECTIVES AND HYPOTHESES

The study aims to empirically test key links in the family architecture framework in a Romanian adolescent sample. We examined (a) how family type is associated with attachment style distribution (RSQ) and with relationship-specific attachment dimensions (IPPA), (b) whether family type is associated with adolescents' socio-emotional competences (trait emotional intelligence and emotional maturity), (c) whether attachment style is associated with self-esteem indicators, and (d) whether attachment to peers predicts emotional intelligence beyond attachment to parents.

Based on attachment theory and process-oriented family research, we formulated the following hypotheses:

H1: Family type will be associated with relational organization. Adolescents from two-parent families will report higher trust and communication (and higher total attachment) toward mother and father than those from single-parent or alternative-care settings.

H2: Family type will be associated with socio-emotional competences (emotional maturity and trait emotional intelligence). Given the increasing salience of peers, differences in emotional intelligence by family type may be attenuated.

H3: Attachment style will be associated with global self-esteem, with secure attachment linked to higher self-esteem relative to insecure styles.

H4: Peer attachment will uniquely predict trait emotional intelligence (global and interpersonal components) beyond parental attachment.

4. METHODOLOGY

4.1. PARTICIPANTS AND SETTING

Participants were 250 Romanian adolescents aged 16–18 years ($M = 16.98$, $SD = 0.85$) from Olt County. The sample included 131 boys (52.4%) and 119 girls (47.6%), with 109 adolescents from urban areas (43.6%) and 141 from rural areas (56.4%). With respect to caregiving arrangement, 160 participants (64.0%) lived in two-parent families, 68 (27.2%) in single-parent families, and 22 (8.8%) in alternative-care settings (institutional care and/or professional foster care; AMP). Data were collected during the 2023–2024 school year from 250 high school

students, 22 of them benefiting of child protection services. Inclusion criteria were: age 16–18 years and ability to understand and complete the questionnaires. Exclusion criteria included severe intellectual disability or inability to read/understand the items, as well as temporary inability to participate (*e.g.*, absence on the day of assessment).

4.2. ETHICAL CONSIDERATIONS AND PROCEDURE

The study adhered to the ethical principles governing psychological research and professional conduct in Romania. Participation was voluntary and anonymous. For adolescents recruited from schools, informed consent was obtained from both adolescents and their parents/legal guardians. For adolescents in alternative-care settings, consent procedures involved the legal guardian and institutional gatekeepers, in line with child protection regulations. Questionnaires were administered in Romanian in group settings in schools and, for the child protection subsample, under supervised conditions appropriate to the care context. Participants were informed about the purpose of the research, confidentiality, and their right to withdraw without negative consequences.

4.3. MEASURES

Relationship Scales Questionnaire (RSQ). Attachment style was assessed with the RSQ, a self-report measure capturing adult/adolescent attachment representations aligned with four prototypical styles: secure, preoccupied, dismissing, and fearful (Griffin & Bartholomew, 1994). The RSQ yields a categorical classification based on dominant style indices.

Inventory of Parent and Peer Attachment (IPPA). Relationship-specific attachment quality was assessed using the IPPA (Armsden & Greenberg, 1987) for mother, father, and peers. IPPA dimensions include Trust, Communication, and Alienation, with higher total scores reflecting more secure attachment-related relationship quality (high trust and communication, low alienation). In this manuscript, we report subscale scores and total scores for each relationship target.

Emotional Intelligence Inventory (INEM). Trait emotional intelligence (self-reported emotional competences) was assessed with the INEM (Constantin, 2008). The INEM provides a global index and component scores covering intrapersonal and interpersonal emotional competences, including understanding and regulating own emotions, understanding and regulating others' emotions, and social desirability.

Friedmann Emotional Maturity Scale. Emotional maturity was assessed using the Friedmann scale (Friedmann, 1992), which provides a level-based index of emotional maturity related to affective balance, frustration tolerance, and emotional stability.

Attachment Styles and Self-Esteem Inventory (ASSI). Self-esteem indicators were assessed using the ASSI (Constantin, 2005), which yields categorical self-esteem levels (low, medium, high) as well as additional indices (self-devaluation and infatuation) used in the dataset.

4.4. DATA ANALYSIS AND TRANSPARENCY OF INFERENCE

All analyses were conducted at $\alpha = 0.05$ (two-tailed). Preliminary screening indicated deviations from normality for most scale variables, and several variables showed pronounced skewness/kurtosis, therefore between-group comparisons relied primarily on nonparametric tests.

Family type comparisons: Kruskal–Wallis tests were used to compare family type groups (two-parent, single-parent, alternative care) on attachment dimensions (IPPA) and socio-emotional outcomes (INEM, Friedmann). Effect size was indexed with epsilon-squared (ϵ^2) for Kruskal–Wallis tests. Given the small size and heterogeneity of the alternative-care group, results are interpreted cautiously and primarily at the level of overall group differences.

Associations between categorical variables: χ^2 tests were used for associations between family type and attachment style (RSQ), and between attachment style and categorical self-esteem levels. Effect size was indexed with Cramer's V, and 95% CIs for V were derived from noncentral χ^2 confidence limits.

Multiple testing control: To reduce Type I error inflation, Benjamini–Hochberg false discovery rate (FDR) correction ($q = 0.05$) was applied within families of related tests (*e.g.*, the set of IPPA family-type comparisons). Both unadjusted p-values and FDR-adjusted q-values are reported when relevant.

Predictive analyses: Multiple linear regression (enter method) was used to test whether peer attachment predicts emotional intelligence outcomes beyond attachment to mother and father. Multicollinearity was examined via tolerance and variance inflation factors (VIF). Other regression assumptions (normality and homoscedasticity of residuals; influential cases) should be examined in future replications with full diagnostic reporting.

Missing data: The analyzed variables contained complete data (Valid N, listwise = 250), therefore no imputation procedures were required.

5. RESULTS

5.1. ATTACHMENT STYLE DISTRIBUTION (RSQ) AND ASSOCIATION WITH FAMILY TYPE

Table 1 summarizes the distribution of attachment styles (RSQ) in the sample and their composition by gender, residence, and family type. Because the alternative-care group was small ($n = 22$), percentages for this subgroup should be interpreted cautiously.

Table 1

Distribution of attachment styles (RSQ) by gender, residence, and family type (%)

RSQ	%	Boys	Girls	Urban	Rural	Two parent	Single parent	Inst./AMP
Secure	39.6	56	44	46	54	77	17	6
Preoccupied	22.4	46	54	46	54	61	28	11
Dismissive	25.2	58	42	38	62	52	34	14
Fearful	12.8	50	50	42	58	52	45	3

Note. Inst./AMP = institutional care and/or professional foster care (asistent maternal profesionist).

A χ^2 test indicated that attachment style distribution differed across family types, $\chi^2(6) = 17.34$, $p = 0.008$, Cramer's $V = 0.19$, 95% CI [0.05, 0.25]. The two-parent group contained a higher proportion of secure and preoccupied styles, whereas dismissing attachment was relatively more frequent in the alternative-care group; given the small size of the alternative-care subsample, these patterns should be interpreted as descriptive rather than definitive.

5.2. FAMILY TYPE AND RELATIONAL ORGANIZATION (IPPA)

Table 2 reports descriptive statistics and Kruskal–Wallis tests comparing family type groups on attachment-related relationship quality toward mother, father, and peers (IPPA). To control for multiple testing within this family of comparisons, Benjamini–Hochberg FDR-adjusted q-values are also reported.

Table 2

Family type differences in attachment to mother, father, and peers (IPPA)

IPPA dimension	Two-parent M (SD)	Single-parent M (SD)	Inst./AMP M (SD)	H	p	ϵ^2	q (FDR)
Mother attachment – Trust	42.01 (6.75)	36.46 (11.60)	33.23 (9.80)	24.401	< .001	0.091	0.004
Mother attachment – Communication	33.28 (7.09)	28.40 (10.13)	25.91 (8.69)	20.21	< .001	0.074	0.004
Mother attachment – Alienation	12.68 (5.39)	14.28 (6.29)	14.23 (5.90)	3.735	0.155	0.007	0.232
Mother attachment – Total	99.56 (16.99)	88.00 (25.12)	87.95 (21.57)	12.737	0.002	0.043	0.005
Father attachment – Trust	40.08 (7.92)	34.31 (12.53)	32.91 (11.78)	14.347	< .001	0.05	0.004
Father attachment – Communication	30.64 (8.18)	26.44 (10.19)	25.32 (8.83)	12.793	0.002	0.044	0.005
Father attachment – Alienation	13.95 (5.63)	15.22 (6.83)	15.91 (6.65)	2.298	0.317	0.001	0.38

Table 2 (continued)

IPPA dimension	Two-parent M (SD)	Single-parent M (SD)	Inst./AMP M (SD)	H	p	ϵ^2	q (FDR)
Father attachment – Total	94.14 (18.49)	83.21 (28.61)	85.64 (24.85)	6.533	0.038	0.018	0.076
Peer attachment – Trust	41.39 (7.55)	39.65 (8.80)	38.50 (11.36)	1.739	0.419	0	0.419
Peer attachment – Communication	42.10 (7.66)	41.07 (8.42)	39.59 (10.34)	1.032	0.597	0	0.597
Peer attachment – Alienation	14.28 (4.48)	16.81 (13.48)	14.41 (6.16)	4.268	0.118	0.009	0.202
Peer attachment – Total	94.88 (14.33)	91.53 (17.16)	88.50 (17.95)	3.033	0.219	0.004	0.292

Note. Kruskal–Wallis tests ($df = 2$) compared three family type groups (two-parent, single-parent, institutional/AMP). ϵ^2 = epsilon-squared effect size for Kruskal–Wallis. q-values are Benjamini–Hochberg FDR-adjusted within the set of 12 IPPA tests. Inst./AMP group size was small ($n = 22$), therefore estimates involving this group may be less stable.

Consistent with H1, family type was associated with multiple indicators of parental attachment quality. Adolescents from two-parent families reported higher trust and communication toward mother and father than adolescents from single-parent families and, most markedly, those in alternative care. These differences were robust after FDR correction for maternal trust and communication, paternal trust and communication, and maternal total attachment (all $q \leq 0.005$). In contrast, attachment to peers (trust, communication, alienation, and total) did not differ by family type (all $ps \geq 0.118$), suggesting that peer attachment may be less structurally constrained in late adolescence within this sample.

5.3. FAMILY TYPE AND SOCIO-EMOTIONAL COMPETENCES (INEM)

Table 3 presents descriptive statistics and Kruskal–Wallis tests for emotional intelligence (INEM) outcomes by family type. Across global and component scores, differences by family type were small and nonsignificant.

Table 3

Family type differences in emotional intelligence outcomes (INEM)

Outcome (INEM)	Two-parent M (SD)	Single-parent M (SD)	Inst./AMP M (SD)	H	p	ϵ^2
Global trait EI (INEM)	2.33 (1.18)	2.07 (1.05)	2.36 (1.14)	2.341	0.31	0.001
Intrapersonal competence	2.27 (1.18)	2.12 (1.11)	2.23 (1.19)	0.724	0.696	0
Understanding own emotions	1.70 (0.60)	1.63 (0.61)	1.64 (0.49)	0.205	0.903	0
Regulating own emotions	1.72 (0.63)	1.65 (0.58)	1.59 (0.59)	1.193	0.551	0
Interpersonal competence	2.39 (1.14)	2.09 (1.02)	2.50 (1.11)	3.304	0.192	0.004

Table 3 (continued)

Outcome (INEM)	Two-parent M (SD)	Single-parent M (SD)	Inst./AMP M (SD)	H	p	ϵ^2
Regulating others' emotions	1.55 (0.63)	1.43 (0.59)	1.59 (0.61)	3.119	0.21	0.004
Recognizing others' emotions	1.56 (0.66)	1.46 (0.63)	1.41 (0.50)	1.585	0.453	0
Social desirability	1.56 (0.68)	1.34 (0.54)	1.55 (0.74)	5.191	0.075	0.013

Note. Kruskal–Wallis tests ($df = 2$) compared family type groups on INEM outcomes. Scores represent level-based indices used in the dataset. ϵ^2 = epsilon-squared effect size.

Contrary to H2, family type was not associated with global trait emotional intelligence or with the examined INEM components (all $ps \geq 0.075$; $\epsilon^2 \leq 0.013$). A trend-level difference emerged for the social desirability indicator ($p = 0.075$), suggesting possible differences in response style across groups; however, this did not reach conventional significance.

5.4. EMOTIONAL MATURITY (FRIEDMANN) BY FAMILY TYPE

Table 4 summarizes emotional maturity scores by family type. A small but statistically significant difference was observed across groups.

Table 4

Emotional maturity (Friedmann) by family type

Family type	M (SD)	H	p	ϵ^2
Two-parent	2.52 (0.78)	8.387	0.015	0.026
Single-parent	2.24 (0.73)			
Inst./AMP	2.27 (0.55)			

Note. Friedmann scores represent level-based indices in the dataset. Kruskal–Wallis test ($df = 2$) compares family types. ϵ^2 = epsilon-squared effect size.

Adolescents from two-parent families showed slightly higher emotional maturity scores than those from single-parent families and alternative care (Table 4). Given the cross-sectional design and the small alternative-care subsample, this finding should be interpreted as an association rather than an effect of family structure.

5.5. ATTACHMENT STYLE AND SELF-ESTEEM (ASSI)

Table 5 presents the distribution of self-esteem levels across attachment styles. Differences were evaluated with a χ^2 test of independence.

Table 5

Self-esteem (ASSI) by attachment style RSQ (percentages)

RSQ	N	Low	Medium	High	Self-devaluation	Infatuation
Secure	100	21.0%	48.0%	31.0%	13.0%	21.0%
Preoccupied	57	31.6%	45.6%	22.8%	24.6%	19.3%
Dismissive	64	29.7%	35.9%	34.4%	26.6%	25.0%
Fearful	29	51.7%	27.6%	20.7%	41.4%	17.2%

Secure attachment was associated with a higher proportion of medium-to-high self-esteem, whereas fearful attachment showed the highest proportion of low self-esteem (Table 5). Associations between attachment style and the ASSI self-devaluation and infatuation levels were not statistically significant in this sample (self-devaluation: $\chi^2(6) = 7.95$, $p = 0.242$; infatuation: $\chi^2(6) = 3.06$, $p = 0.801$).

5.6. PEER ATTACHMENT AS A PREDICTOR OF EMOTIONAL INTELLIGENCE (MULTIPLE REGRESSION)

To test the compensatory role of peer attachment proposed in the family architecture framework, we estimated multiple regression models predicting (a) global emotional intelligence and (b) understanding others' emotions, from attachment to mother, father, and peers (IPPA total scores). Table 6 reports model summaries, coefficients, and 95% CIs.

Model: Global emotional intelligence (IE_{global}) – $R^2 = 0.067$; Adj. $R^2 = 0.056$; $F(3,246) = 5.93$; $p = 0.001$

Table 6

Multiple regression models predicting emotional intelligence outcomes from parental and peer attachment (IPPA total scores)

Predictor (IPPA total)	B	SE B	β	t	p	95% CI for B	VIF
Attachment to mother	0.015	0.008	0.121	1.796	0.074	[-0.001, 0.031]	1.256
Attachment to father	0.008	0.008	0.026	0.399	0.69	[-0.008, 0.024]	1.256
Attachment to peers	0.026	0.01	0.162	2.527	0.012	[0.006, 0.046]	1.256

Peer attachment (IPPA peers total) was the only significant predictor after controlling for attachment to mother and father, supporting H4. Effect sizes were small ($R^2 \approx 0.07$), consistent with the expectation that socio-emotional competences in adolescence are shaped by multiple determinants beyond attachment alone.

6. DISCUSSION

This study set out to operationalize and test the family architecture framework in a Romanian adolescent sample by linking family type to relational organization (attachment) and to socio-emotional competences. Consistent with H1, family type was associated with key parental attachment dimensions: adolescents from two-parent families reported higher trust and communication with both mother and father, and higher total attachment to mother, than peers from single-parent families and alternative-care settings. These findings converge with process-oriented accounts suggesting that structural transitions and caregiving disruptions may be expressed in adolescents' perceived availability and emotional accessibility of caregivers, particularly through trust and communication channels (van Dijk *et al.*, 2020).

Contrary to H2, family type was not associated with global trait emotional intelligence or its components as assessed via self-report (INEM). One interpretation, aligned with the family architecture model, is that peer relationships in late adolescence may attenuate structural differences by providing alternative sources of emotional learning and support. This interpretation is consistent with evidence that peer attachment and peer functioning become increasingly salient for adolescent well-being and may partly compensate for lower perceived parental support (Delgado *et al.*, 2022). At the same time, emotional maturity (Friedmann) showed a small group difference, suggesting that some aspects of socio-emotional development may remain sensitive to caregiving context even when perceived emotional intelligence does not differ.

H3 received partial support: attachment style was associated with global self-esteem, with secure attachment linked to more favorable self-evaluations, whereas fearful attachment showed the highest prevalence of low self-esteem. This pattern is broadly consistent with attachment-based models in which internal working models organize self-worth and relational expectations. Notably, attachment style was not significantly associated with self-devaluation or infatuation levels in this dataset, indicating that attachment-related differences may be more pronounced for global self-esteem than for specific self-evaluative tendencies captured by ASSI.

Support for H4 was clear: peer attachment uniquely predicted both global trait emotional intelligence and understanding others' emotions beyond parental attachment. Although effect sizes were small, the pattern is theoretically meaningful in adolescence, when daily emotion work increasingly occurs in the peer context. This finding also illustrates a key claim of family architecture: structural family type relates strongly to parental relationship quality, whereas socio-emotional competences may be partly scaffolded by non-familial attachments when peers provide trust and emotional communication opportunities.

Interpretation and methodological considerations. The present evidence should be interpreted within several methodological constraints. First, the cross-sectional

design precludes causal inference. Associations between family type, attachment, and socio-emotional outcomes may reflect selection effects and unmeasured confounders such as socioeconomic status, cumulative adversity, or placement history. Second, the alternative-care group was small ($n = 22$) and likely heterogeneous (institutional and foster contexts), reducing statistical power and stability of subgroup estimates. Third, all constructs were assessed via self-report, increasing the risk of common method variance and socially desirable responding; the trend-level family-type differences in social desirability underscore this possibility. Fourth, psychometric properties in the current sample were not comprehensively documented, which limits assessment of measurement reliability and validity.

A further interpretive point concerns the measurement of emotional competence. The INEM assesses trait emotional intelligence via self-report, which captures perceived emotional competences and emotion-related self-concepts rather than ability-based performance. Systematic reviews have emphasized that ability-based and trait-based measures can yield different patterns and may not be interchangeable (Bru-Luna *et al.*, 2021). Therefore, the absence of family-type differences in INEM outcomes should not be interpreted as evidence that actual emotion-processing abilities are unaffected by caregiving context; rather, it suggests that adolescents' self-perceived emotional competences are not strongly differentiated by family type in this sample.

Despite these limitations, the results have practical implications when framed cautiously as directions for intervention and future research. Relationship-focused programs that strengthen caregiver–adolescent trust and communication (in intact families, single-parent families, and alternative-care contexts) may support adolescents' socio-emotional development. In educational settings, socio-emotional learning and counseling programs could incorporate structured peer mentoring and group-based emotional skills training, leveraging the observed predictive role of peer attachment. For child protection services, interventions that stabilize caregiving relationships and enhance emotionally attuned communication may be particularly relevant. Future research should use longitudinal, multi-informant designs (adolescents, caregivers, teachers) and include ability-based measures of emotional competence, as well as key contextual covariates, to test mediation pathways proposed by the family architecture model.

In future operationalizations of family architecture, it will be important to directly assess autonomy-supportive parenting and related controlling practices. Theory and recent reviews suggest that autonomy support is beneficial for adolescent adjustment (Benito-Gomez *et al.*, 2020), whereas psychologically controlling parenting is linked to less adaptive emotion regulation (Beliveau *et al.*, 2023). Incorporating such constructs would allow stronger tests of the proposed process pathways from family architecture to socio-emotional competences.

Moreover, adolescents with histories of institutional care or multiple placements may have experienced early adversity and caregiving instability, which

can influence attachment and socio-emotional development (Zeanah *et al.*, 2005). Because placement history variables were not available in the current analyses, these contextual influences could not be disentangled.

In addition, the psychometric properties of the instruments in the current sample (*e.g.*, internal consistency, construct validity, and cultural adaptation details) were not fully reported, which limits the evaluation of measurement quality and should be addressed in future replications and manuscript versions.

7. CONCLUSIONS

Family architecture offers a process-oriented lens for integrating family structural configuration with relational organization and adolescent socio-emotional development. In this Romanian sample, family type was associated with parental attachment quality (trust and communication), while perceived emotional intelligence was not structurally differentiated; instead, peer attachment emerged as a consistent predictor of emotional intelligence outcomes. Attachment style was linked to global self-esteem, reinforcing the relevance of internal working models for adolescent self-evaluation. Overall, the findings support the conceptual value of family architecture while underscoring the need for cautious interpretation and for more rigorous longitudinal and multi-method research to establish mechanisms and causal pathways.

8. LIMITATIONS AND FUTURE DIRECTIONS

The study is cross-sectional, which limits causal inference. Future research should integrate indicators of family climate (communication, cohesion, conflict), longitudinal designs, and multi-level analyses (family-school-community) to refine the family architecture model.

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