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Development, validation, and psychometric properties of the Italian and English version of the Boredom Intolerance Scale (BIS)

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ABSTRACT

The present paper proposes developing and validating the Boredom Intolerance Scale (BIS) – the first and only measure assessing the degree to which individuals are able to stand the experience of boredom. Across three studies ($N = 1397$), the psychometric properties of the BIS are presented. Exploratory factor analysis, implemented in Study 1, suggested a unidimensional and 6-item structure with high reliability. Study 2 ratified the emerged structure by using a confirmatory factor analysis. Corroboration of the measure's robustness was provided by a multigroup CFA, which yielded evidence for the gender invariance of the BIS Italian version. Study 3 validated the English version of the BIS, indicating a robust factor structure with high reliability and invariance across participants' gender. Study 3 also proved the BIS's invariance across English and Italian versions. Construct validity was examined across Studies 2 and 3, yielding significant associations of the BIS with measures of trait and state boredom, relaxation sensitivity, neuroticism, anxiety, anger, impulsiveness, depression, life satisfaction, and purpose in life. These findings suggest that the BIS is a psychometrically sound measure with possible implications for researchers and practitioners.

1. Introduction

Boredom, a ubiquitous yet often overlooked emotional state, has profound implications for individual well-being and behavior. It has received little attention from the psychological literature, making it difficult to define for many years. However, scholars have increasingly focused on boredom in the last two decades (e.g., van Tilburg et al., 2024). Fahlman et al. (2013) offered a comprehensive trans-theoretical definition of the experience of boredom, synthesizing conceptualizations from several theoretical approaches (e.g., Csikszentmihalyi, 2000; Fahlman et al., 2009; Wangh, 1975). Such approaches agree that boredom is primarily defined by the hostile experience of wanting to engage in something challenging and satisfying but being unable to do so. It is often seen as an aversive state of under-arousal due to a lack of meaningful environmental stimuli. However, it can also manifest as high-arousal states such as restlessness and frustration. Some theorists highlight that boredom alternates between agitation and lethargy. Fiske and Maddi (1961) and Bernstein (1975) note that tiredness and restlessness coexist in boredom. Hamilton (1981) explains that high arousal in boredom may be a compensatory response for self-stimulation, and Thackray (1981) adds that it may occur during monotonous tasks requiring high vigilance. Boredom also distorts the perception of time, making it seem slower and causing difficulty concentrating (Zakay, 2014). Based on these distinct theorizing, Fahlman et al. (2013) frame boredom as an aversive experience characterized by a lack of engagement, negative affect with low and high arousal, and the sensation of time passing slowly with difficulty focusing.

Although boredom does not necessarily imply discomfort, abundant research has highlighted its potential negative correlates. Empirical studies have linked boredom to various adverse mental health conditions such as depression and anxiety (Farmer & Sundberg, 1986), negative affect (Vodanovich et al., 1991), hostility and anger (Rupp & Vodanovich, 1997), alexithymia (Eastwood et al., 2007), somatization disorders (Sommers & Vodanovich, 2000), overeating and binge eating

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(Stickney & Miltenberger, 1999), pathological gambling (Mercer & Eastwood, 2010), marijuana use (Lee et al., 2007), alcohol abuse (Wiesbeck et al., 1996), job dissatisfaction (Kass et al., 2001), and poor academic performance (Jarvis & Seifert, 2002). Boredom has also been linked to lower levels of life meaning (Fahlman et al., 2009), self-realization (McLeod & Vodanovich, 1991), and life satisfaction (Farmer & Sundberg, 1986). Hence, boredom appears to be associated with significant social, psychological, and physical challenges.

Because of its potential adverse effects, numerous scholars have attempted to produce tools to detect boredom in various forms. Until the early 2000s, most of these instruments addressed boredom by focusing only on specific contexts (e.g., Job Boredom Scale, Lee, 1986; Free Time Boredom Scale, Ragheb & Merydith, 2001) or representing it as a sub-dimension of scales that measured other constructs (e.g., Boredom Coping Scale, Hamilton et al., 1984; Boredom Susceptibility Scale, Zuckerman, 1979). The only large-scale measure of boredom that has been widely used in empirical research is the Boredom Proneness Scale (BPS; Farmer & Sundberg, 1986), which has been the subject of numerous revisions over the years (e.g., Struk et al., 2017). The Boredom Proneness Scale measures the dispositional tendency to experience boredom in various situations. Although initially showing some limitations, this measure has proven to be a solid and robust tool across its revisions despite some difficulties in defining its dimensionality (Melton & Schulenberg, 2009). More recently, Fahlman et al. (2013) proposed another instrument to detect boredom: the Multidimensional State Boredom Scale. This scale also proved to be a solid, robust, and reliable measure, representing an excellent and exhaustive tool for assessing state boredom.

Surprisingly, despite the efforts of numerous scholars in developing boredom assessment tools, literature has not yet provided an instrument to measure the individual tendency to tolerate this emotion. The capacity to tolerate boredom may vary significantly across individuals, potentially impacting their mental health, productivity, and overall life satisfaction (Masland et al., 2020). Based on the exhaustive definition of boredom provided by Fahlman et al. (2013), we framed boredom intolerance as the impossibility that *individuals experience in enduring the feeling of disengagement from interesting or significant activities*. Boredom intolerance may emerge as a critical construct in psychological research and practice. High boredom intolerance could be associated with various maladaptive outcomes, such as increased susceptibility to anxiety, depression, and anger. Individuals who struggle to tolerate boredom may engage in risky behaviors or seek immediate gratification, often at the expense of long-term goals and purposes, thus potentially compromising their general well-being and life satisfaction. Notwithstanding its potential relevance, the measurement of boredom intolerance is lacking in the psychological literature. As previously mentioned, existing tools mainly focus on the dispositional tendencies to experience boredom in several situations. However, some scholars advanced criticisms about measures of trait boredom, underlining the risk that they may broadly reflect situational variance in activities rather than individual differences (Westgate & Steidle, 2020). An intolerance-based measure may instead help consider the context in which emotion is experienced, providing a more nuanced understanding of dispositional individuals' emotional responses. This may allow a deeper exploration of boredom-related factors influencing emotional regulation and coping strategies (Masland et al., 2020). Thus, the present paper introduces the Boredom Intolerance Scale (BIS), a novel instrument designed to robustly, validly, and reliably assess individuals' intolerance to boredom. The BIS aims to fill the gap in current boredom assessment tools by providing a robust measure that can be utilized across diverse contexts. The development and validation of the BIS involved several rigorous steps, including a qualitative approach for item generation, the exploration of the BIS's factorial structure and its confirmation across English and Italian samples, and participants' gender. The scale was then subjected to broad psychometric evaluation to establish its reliability and validity.

2. Study 1a: qualitative approach for items generation

Study 1a aimed to develop a list of multiple items capable of accurately and comprehensively capturing the semantic dimension of the experience of intolerance to boredom. Thus, we adopted a preliminary qualitative approach to ensure that item generation was initially guided by the content used by individuals to describe their experience of boredom intolerance to increase the potential ecological validity of the created items.

2.1. Method

2.1.1. Participants and procedure

Two expert psychotherapists (second and third Authors) conducted semi-structured interviews involving 20 Italian adult participants (10 Females, $Mage = 40.80$, $SDage = 13.47$) randomly selected from a non-clinical population. As for education, 45 % of participants had a high school diploma, while the remaining 55 % had a master's or bachelor's degree. Participants were recruited online based on their gender (50 % female and 50 % male) and invited to participate in a face-to-face interview. Written informed consent was obtained on their arrival. Seven trainee psychotherapists who were unaware of the interviewee's details transcribed the twenty interviews verbatim. The contents of the interviews were first coded and analyzed by the second Author, following the thematic analysis guidelines developed by Braun and Clarke (2006). Then, they were discussed in a focus group in which five experts on the boredom construct and psychometric techniques for evaluating measurement instruments participated (four expert Psychotherapists and a psychometrics Professor).

2.2. Results

From the focus group, it emerged that the interviews allowed the identification of semantic content consistent with the theoretical definitions of boredom and intolerance. Boredom emerged as the perceived lack of engagement in desired satisfying activities characterized by little interest or absence of personal meaning. This perception was accompanied by states of irritability, unpleasantness, frustration, stress, discomfort, and monotony, which determined the degree of tolerance and endurance of the situation by individuals. The interviews, therefore, highlighted that people shared a standard description of the experience of intolerance to boredom, which can be framed as the *inability and impossibility to tolerate the feeling of disengagement from interesting or significant activities*.

3. Study 1b: exploratory factor analysis (EFA) of the Boredom Intolerance Scale

Study 1b aimed to explore the factor structure of the BIS. Based on themes from Study 1a and a literature review, the five experts have drawn up their list of items capable of reflecting the semantic content of the definition of boredom intolerance. This led to the generation of 50 items examined and revised in a second focus group. After revisions, 20 items were considered consistent, non-redundant, and semantically representative of the boredom intolerance construct. Each of the 20 items was constructed with a 5-point Likert-type response scale that asked participants to express the degree of disagreement/agreement with the relevant statement. Seventeen items were formulated so that agreement with the statement reflected a greater intolerance to boredom, while the remaining three were in reverse format. The 20 items were retained with the consensus of all five experts and represented the initial pool implemented in the EFA.

3.1. Method

3.1.1. Participants and procedure

Lacking specific power analyses for EFA, we determined the sample size based on the guidelines from various studies (e.g., Howard, 2016), advocating for a participant-to-variable ratio of at least 5-to-1. We distributed an original set of 20 items among 506 participants, yielding a 25.3 ratio exceeding the typical recommendation.

Participants were recruited through Prolific Academic and received monetary compensation (£6/hour) for completing a short questionnaire and thus being enrolled in the present study. We prescreened participants based on their nationality (Italian), country location (Italy), first language (Italian), and gender (approximately 50 % female and 50 % male). Participants were first presented with a brief introduction describing the general aims of the research and asked to consent to participate. Then, they were presented with a questionnaire asking them to provide demographic information and to fill in the initial pool of 20 items generated for the EFA of the BIS. At the end of the questionnaire, participants were asked to answer an "instructed item" as an attention check (DeSimone et al., 2015). An incorrect response to this item was considered an indicator of insufficient effort in following instructions and providing answers to complete the survey. Thus, in such cases, the questionnaire completion was interrupted, participants were thanked for their participation, and their answers were not recorded.

The sample consisted of 261 male and 245 female Italian participants with a mean age of 30.71 ($SD = 9.07$). The sample was quite balanced regarding the geographical area of origin, with 45.7 % coming from the north, 26.5 % from the center, 20.4 % from the south, and 7.5 % from the islands. A majority (62.8 %) of the sample were non-student adults. Among them, 79.3 % were employed, while 20.7 % were retired, homemakers, or unemployed. The remaining 37.2 % of the sample were college students. As for education, 1.8 % had a lower secondary school diploma, 43.3 % a high school diploma, 48.8 % a master's or bachelor's degree, and 6.1 % had a post-graduate qualification.

3.2. Results

3.2.1. EFA and reliability

Analyses were run with the *RStudio* graphical interface (R Core Team, 2024) and its *psych* (Revelle & Revelle, 2015) and *paran* (Dinno & Dinno, 2018) packages. Before examining the BIS's factorial structure, the normality of distribution for the initial pool of 20 items was tested. Skewness and kurtosis values within ±2.00 have been considered acceptable (Curran et al., 1996). As shown in Table 1, all 20 items had values below the thresholds considered and did not present normality issues in their distributions. Thus, the EFA was conducted on a polychoric correlation (Fabrigar et al., 1999). To evaluate whether the obtained correlation matrix was factorable, we implemented Bartlett's test of sphericity and the Kaiser-Meyer-Olkin (KMO) measure of sampling adequacy. Bartlett's test showed that the correlation matrix was not random, $\chi^2(190) = 550.139$, $p < 0.001$. The KMO statistic was equal to 0.96, an excellent value and higher than the minimum standard for proceeding with the analysis (Howard, 2016). Therefore, the correlation matrix could be considered appropriate for factor analysis. Once we determined that the correlation matrix was factorable, we investigated the factor structure of the BIS by implementing an EFA with the extraction method of principal axis factoring (PA). We opted for the PA method, which proved more robust than other extraction methods (e.g., Norris & Lecavalier, 2010). Factor retention utilized the Kaiser criterion, visual scree test (Cattell, 1966), and parallel analysis (Glorfeld, 1995; Horn, 1965). Analysis revealed two factors with an eigenvalue >1. However, only the first factor showed a high eigenvalue of 9.65, while the second was just above 1 (i.e., 1.03). Furthermore, the first factor explained a large portion of the variance (i.e., 48 %), compared to the second extracted factor, which explained only 5 %. These results provided a preliminary indication in support of our theoretical expectation of a unidimensional structure of the BIS. This expectation was also corroborated by the analysis of the scree plot, where the eigenvalue curve flattened out after the first factor, and by the parallel analysis,[1] where the line of the simulated random eigenvalues intersected that of the estimated (adjusted) eigenvalues again after the first factor (Fig. 1).

Once we established the unidimensional structure of the BIS, we concentrated on the analysis of the items. The entire pool of 20 items showed good factor loadings ranging from 0.54 to 0.87 (Table 1). Thus, we selected and included only the items that showed the highest loadings in a new EFA, as they were more representative of the factor. Considering the EFA results relating to the entire pool of 20 items, we arbitrarily set our loading cutoff to ≥0.75. Following this criterion, we selected six items and again performed the EFA to test the final unidimensional solution of the BIS. Bartlett's test indicated that the correlation matrix was not random, $\chi^2(15) = 79.141$, $p < 0.001$, accompanied by an excellent KMO value of 0.92. The EFA showed a single factor with an eigenvalue of 4.09, which explained 68 % of the variance, corroborated by the scree plot and parallel analysis (Fig. 1). Factor loadings of the final six items were high and ranged from 0.78 to 0.86 (Table 1).

Finally, we also examined the reliability of BIS's 6-item and unidimensional final solution ($M = 3.36$, $SD = 0.85$). Given recent criticism about the *alpha* coefficient (Deng & Chan, 2017), we assessed reliability through *Cronbach's* α and *McDonald's* ω. Analyses revealed excellent values of 0.90 (*95 % CI* = 0.885, 0.917) and 0.90 for α and ω coefficients.

4. Study 2

Study 2 aimed to validate the Italian version of the BIS. Therefore, a confirmatory factor analysis was first conducted to evaluate the factor structure of the BIS that emerged from the exploratory analysis of Study 1. The measurement invariance of the Italian version of the BIS was also investigated across the participants' gender by implementing multigroup CFA models. Specifically, we first aimed to test whether the specified factor structure of the BIS and the related pattern of factor-indicator relationships could be considered equal across female and male participants (i.e., configural invariance). Once configural invariance was established, we aimed to investigate whether items' factor loadings could be regarded as invariant across participants' gender (i.e., metric invariance). Finally, once we established metric invariance, we aimed to examine whether intercepts of items' regressions on the latent variable were invariant across the two groups (i.e., scalar invariance) (see Vandenberg & Lance, 2000, for a review). We expected to find corroboration on all three typologies of invariances for the BIS factorial structure.

Study 2 also aimed to evaluate the BIS's psychometric properties. Thus, the potential association of the BIS with other measures of a convergent conceptual nature was first investigated to test its construct validity. Specifically, we examined the associations of the BIS with the trait measure of Boredom Proneness (Vodanovich & Kass, 1990) and the state boredom measure by Fahlman et al. (2013). We predicted that positive connections between both measures and the BIS would be found. As a further test of the BIS's convergent validity, we also examined the association of the BIS with relaxation sensitivity. Relaxation sensitivity refers to the fear or discomfort associated with relaxation, often due to the belief that relaxing can lead to adverse outcomes or heightened awareness of distressing thoughts and feelings (Luberto et al., 2021). Individuals experiencing boredom intolerance may find relaxing challenging because their minds actively seek something more engaging or stimulating. Similarly, people with high relaxation sensitivity might avoid relaxing because they associate it with negative emotional states or physical sensations. We thus expected that

[1] The *paran* package allows for the specification of the number of simulated random datasets (i.e., iterations) and the centile value. Following Glorfeld's suggestions (1995), we opted for robust values of 5000 iterations and 95 centiles for the presented parallel analysis.

Table 1
Descriptive statistics, loadings (λ), and communalities (h^2) of the items of the initial 20-item and final 6-item unidimensional solution of the BIS.

Label	Content	M	SD	Skewness	Kurtosis	Initial solution		Final solution	
						λ	h^2	λ	h^2
BIS1	**Tollero poco la noia** **I have little tolerance for boredom**	**3.24**	**1.08**	**−0.06**	**−0.97**	**0.87**	**0.75**	**0.86**	**0.74**
BIS2	**Considero la noia spiacevole** **I think boredom is unpleasant**	**3.62**	**0.95**	**−0.70**	**−0.02**	**0.79**	**0.62**	**0.80**	**0.63**
BIS3	**Considero la noia irritante** **I think boredom is irritating**	**3.42**	**1.04**	**−0.48**	**−0.53**	**0.83**	**0.69**	**0.86**	**0.74**
BIS4	Cerco in ogni modo di combattere la noia I try in every way to fight boredom	3.39	0.98	−0.28	−0.64	0.69	0.48		
BIS5	**Considero la noia un'emozione poco tollerabile** **I think boredom is a rather intolerable emotion**	**3.12**	**1.07**	**−0.08**	**−0.90**	**0.82**	**0.68**	**0.84**	**0.71**
BIS6	Non sento la necessità di combattere la noia[a] I don't feel the need to fight boredom[a]	2.48	1.05	0.44	−0.68	0.67	0.45		
BIS7	**Considero insopportabile trovarmi in situazioni noiose** **I think it's unbearable being in boring situations**	**3.18**	**1.07**	**−0.13**	**−0.76**	**0.80**	**0.64**	**0.81**	**0.66**
BIS8	È ingiusto avere una vita noiosa It's unfair to have a boring life	3.30	1.11	−0.37	−0.53	0.55	0.30		
BIS9	Non riesco a rilassarmi quando mi annoio I can't relax when I'm bored	2.99	1.17	0.05	−0.99	0.69	0.48		
BIS10	Considero la noia facilmente tollerabile[a] I find boredom easily tolerable[a]	2.73	1.01	0.07	−0.92	0.72	0.51		
BIS11	In situazioni noiose mi sento a disagio I feel uncomfortable in boring situations	3.17	1.03	−0.12	−0.90	0.70	0.49		
BIS12	Una persona dovrebbe sempre evitare di trovarsi in situazioni noiose A person should always avoid being in boring situations	3.11	1.04	−0.07	−0.69	0.54	0.29		
BIS13	**È frustrante trovarsi in situazioni noiose** **It's frustrating finding myself in boring situations**	**3.61**	**0.98**	**−0.71**	**0.08**	**0.79**	**0.62**	**0.78**	**0.61**
BIS14	Quando mi annoio mi sento spento e questo è insopportabile When I'm bored, I feel drained and this is unbearable	3.28	1.13	−0.30	−0.84	0.74	0.54		
BIS15	Posso annoiarmi anche per ore senza problemi[a] I can be bored for hours without any problems[a]	2.46	1.09	0.36	−0.75	0.64	0.41		
BIS16	La noia non mi permette di vivere la mia vita pienamente Boredom doesn't allow me to live my life fully	2.75	1.14	0.15	−0.94	0.55	0.30		
BIS17	Quando la noia dura molto diventa per me intollerabile When boredom lasts for a long time, it becomes intolerable to me	3.56	1.02	−0.61	−0.13	0.70	0.48		
BIS18	Non sopporto di trovarmi in situazioni ripetitive e monotone I can't stand being in repetitive and monotonous situations	3.28	1.06	−0.23	−0.71	0.59	0.34		
BIS19	L'idea di una vita noiosa mi fa sentire a disagio The idea of a boring life makes me feel uncomfortable	3.71	1.07	−0.74	−0.16	0.65	0.43		
BIS20	Trovo insopportabile passare il mio tempo senza fare nulla I find it unbearable to spend my time doing nothing	3.43	1.11	−0.35	−0.71	0.68	0.46		

Note: Bold lines indicate the items maintained in the BIS's final solution derived from the EFA; The table shows the initial 20 items developed and administered in Italian. We have included a translated version to facilitate understanding of each item's content. However, it must be considered that only the translation of the six items of the final solution was obtained through the back translation procedure.
[a] Item reverse.

individuals who experience boredom intolerance may also exhibit higher levels of relaxation sensitivity. This may be because both states involve a form of discomfort: boredom from a lack of stimulation and relaxation sensitivity from a fear of the internal experiences that relaxation brings about. Moreover, as previous research shows the convergence of both state and trait boredom with the personality trait of emotional stability/neuroticism (Culp, 2006; Fahlman et al., 2013), we also aimed to investigate the potential association of the BIS with such a personality trait. Consistent with the literature, we expected BIS to present positive relationships with the neuroticism trait.

In Study 2, we further assessed the BIS's construct validity by testing its associations with several negative emotional states and general well-being correlates. The association between boredom and negative affect is a well-documented psychological phenomenon, with boredom often linked to various detrimental emotional states (Raffaelli et al., 2018). Research has shown that boredom is associated with a general dimension of subjective distress (Vodanovich et al., 1991), anger expression and impulsiveness (Dahlen et al., 2004), pronounced anxiety (Csikszentmihalyi, 2000), and depression (LePera, 2011). Based on this literature, we expected that greater boredom intolerance would be positively correlated with general negative affect measure, trait anxiety, anxiety sensitivity, worry, anger, impulsivity, and depressive symptomatology. Moreover, since previous research on boredom has shown that it is related to reduced life satisfaction and the lack of meaning and purpose to pursue (Fahlman et al., 2013), we examined the potential relationship of the BIS with satisfaction with life and purpose in life expecting to find negative associations. These last two relations were investigated as a further test of the construct validity to highlight BIS's potential impact on the general well-being of individuals.

Finally, Study 2 aimed to test the discriminant validity of the new proposed measure of boredom intolerance. Rönkkö and Cho (2022) advised that "two measures intended to measure distinct constructs have discriminant validity if the absolute value of the correlation between the measures after correcting for measurement error is low enough for the measures to be regarded as measuring distinct constructs" (pp. 11). Thus, based on this recently generalized definition of discriminant validity, we statistically evaluated the associations of the BIS with measures of boredom proneness (Vodanovich & Kass, 1990) and state boredom (Fahlman et al., 2013), expecting to find correlations sufficiently low that the variables could be considered to represent distinct constructs.

4.1. Method

4.1.1. Participants and procedure

The sample size was established using a-priori power analysis designed for structural equation models (Moshagen, 2020). Following

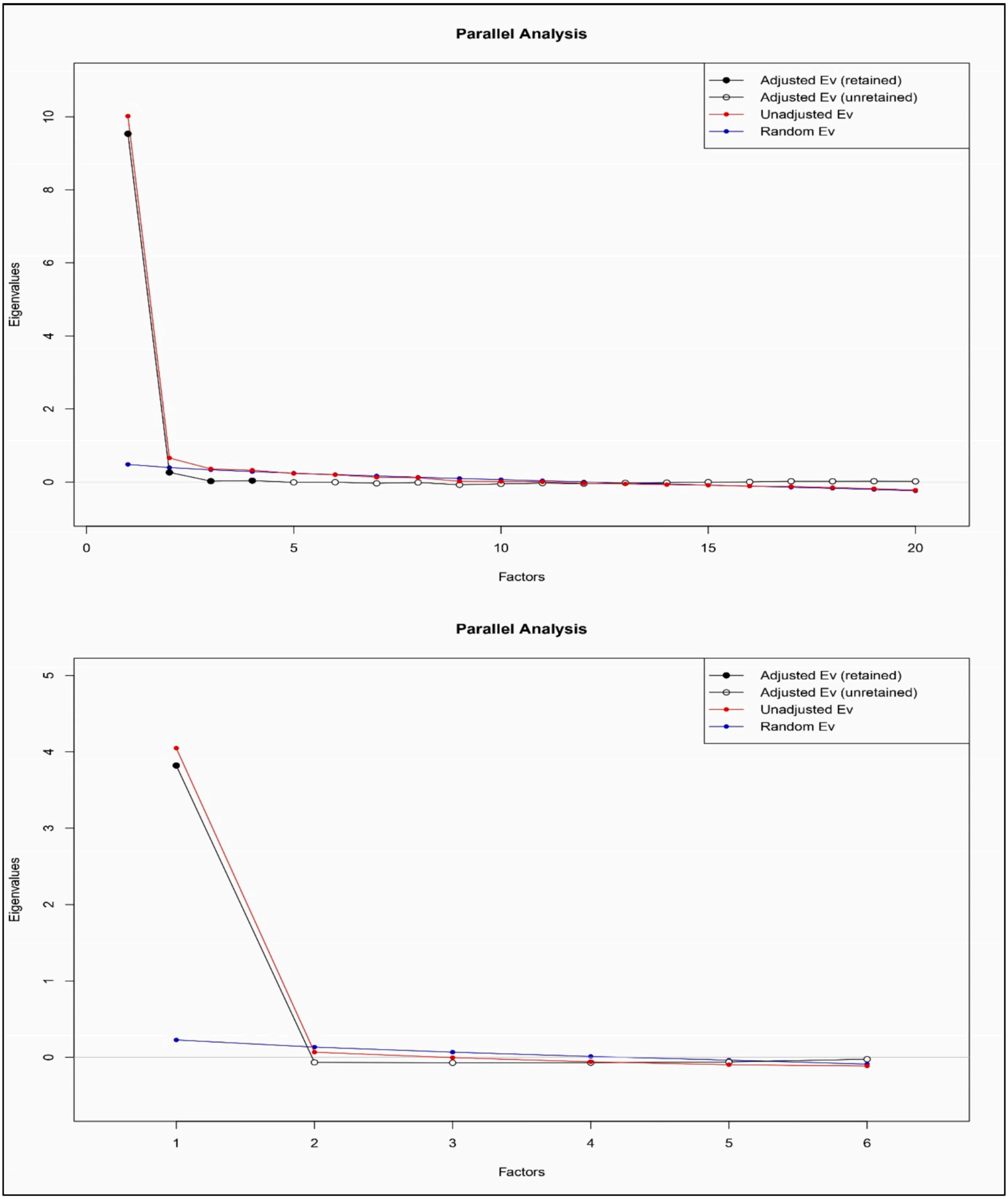

Fig. 1. Scree plot and parallel analysis of the initial 20-item (upper) and final 6-item (lower) solution of the BIS.
Note: The black line represents the eigenvalues retained (black dots) and unretained (white dots) produced for factors that are adjusted for the sample error-induced inflation (Horn's adjustment; 1995); The red line represents the eigenvalues of the observed data (unadjusted); The blue line represents the estimated eigenvalues from iterations number (i.e., 5000) of random data sets.

the indication of Moshagen and Erdfelder (2016), we set a minimum threshold for RMSEA of 0.08, an *alpha* level of 0.05, a conventional power threshold of 0.80, and 9 model degrees of freedom. Analysis indicated a minimum sample size of 274 participants to achieve the desired power. To err on the safe and conservative side, we recruited a community sample of 485 Italian respondents (328 female, *M*age = 36.91, *SD*age = 10.68) who completed an online survey. Participants were presented with a short introduction describing the general aims of the study and asked to provide informed consent. Hence, they completed a questionnaire rating demographic information, the Boredom Intolerance Scale, and other measures described below. Even for this Study, participants were asked to answer an "instructed item" as an attention check (DeSimone et al., 2015). In case of incorrect responses to this item, the participant's answers were not recorded.

Data were collected through a snowball sampling procedure. Bachelor's degree candidates were instructed to recruit up to five individuals within their bachelor's thesis program, prioritizing non-student adult respondents. Participants were pretty balanced regarding the geographical area of origin, with 23.3 % coming from the north, 46.6 % from the center, 14.8 % from the south, and 15.3 % from the islands. Regarding educational level, participants were distributed as follows: 4.5 % had a lower secondary school diploma, 30.1 % a high school diploma, 43.7 % a master's or bachelor's degree, and 21.2 % had a postgraduate qualification. As for employment conditions, a large majority (92.4 %) of the sample were non-student adults. Among them, 90.7 % were employed, 3.1 % were retired or homemakers, and 5.4 % were unemployed. The remaining 7.6 % of the sample were college students.

4.1.2. Measures

Participants answered the *Boredom Intolerance Scale* (BIS), *Boredom Proneness Scale* (BPS; Vodanovich & Kass, 1990; Craparo et al., 2013), *Multidimensional State Boredom Scale* (MSBS; Fahlman et al., 2013; Craparo et al., 2017), emotional stability/neuroticism dimension from the *10-item Big Five Inventory* (Guido et al., 2015; Rammstedt & John, 2007), *Relaxation Sensitivity Index* (RSI; Luberto et al., 2021), *Negative Affect* (NA; Watson et al., 1988; Terraciano et al., 2003), *Anxiety Sensitivity Index-3* (ASI; Taylor et al., 2007; Ghisi et al., 2016), *Trait Anxiety* (STAI-T; Spielberger, 1983; Pedrabissi & Santinello, 1989), *Penn State Worry Questionnaire* (PSWQ; Meyer et al., 1990; Morani et al., 1999), *Trait Anger* (STAXI-T; Spielberger & Reheiser, 2004; Comunian, 1992), *Barratt Impulsiveness Scale* (BIS-11; Patton et al., 1995; Fossati et al., 2001), *Center for Epidemiologic Studies Depression Scale* (CES-D; McDowell & Newell, 1996; Fava, 1983), *Satisfaction With Life Scale* (SWLS; Diener et al., 1985; Di Fabio & Gori, 2016), *Purpose in Life Scale* (PILS; Crea, 2018). Detailed descriptions of all measures (including integral versions, descriptive statistics, procedure for the scoring, and goodness-of-fit indicators) are presented in the supplemental materials.

4.2. Results

4.2.1. Confirmatory factor analysis (CFA) and reliability

CFA was conducted using *lavaan* (Rosseel, 2012), an *R* package for Structural Equation Modeling, using the *RStudio* graphical interface (2024). Analysis was performed with a *Robust Maximum Likelihood method* (MLM). The model fit was evaluated following the benchmarks provided by Hu and Bentler (1999). Given the sensitivity of the *chi-square* (χ^2) statistic to sample size (Chen, 2007), we mainly based on values above 0.95 for the Comparative Fit Index (CFI) and Tucker-Lewis index (TLI) and on values below 0.08 for the Root Mean Square Error of Approximation (RMSEA) and Standardized Root Mean Square Residual (SRMR). Before proceeding with the CFA, we assessed the normality of the items' distribution. Items did not present normality issues (Table 2).

CFA revealed an excellent model fit of the 6-item and unidimensional factorial structure (Fig. 2). Besides a significant robust *chi-square* statistic ($\chi^2 = 23.237$, $df = 9$, $p = 0.006$), the considered incremental fit indices were over the threshold of 0.95 (CFI = 0.99; TLI = 0.98). Regarding the absolute fit indexes, the analyses showed a value of 0.023 for the SRMR and 0.057 for the RMSEA (*90 % CI* = 0.034, 0.081). These findings suggested that the model's fit to the observed data was excellent. As can be seen in Table 3, items' factor loadings were high and significant, highlighting coefficients between 0.72 and 0.82 in their standardized version. Moreover, the confirmed 6-item and unidimensional factorial structure of the BIS showed excellent reliability. We again examined α and ω, obtaining high coefficients of 0.90 (*95 % CI* = 0.885, 0.916) and 0.90, respectively.

4.2.2. Measurement invariance across participants' gender

A multigroup CFA was conducted to test the robustness of the BIS's confirmed structure and its potential generalization across the gender of participants. Specifically, the measurement invariance was tested on the two subsamples of male and female participants. We first examined the model's fit separately for each subsample. Then, by performing hierarchically nested multigroup CFAs with a robust estimator (i.e., MLM), we tested the BIS's configural invariance (by estimating the CFA model simultaneously across groups), metric invariance (by constraining the items' factor loadings to be equal across groups), and scalar invariance (by constraining items' factor loadings and intercepts of items' regressions on the latent variable to be equal across groups). The invariance was assessed on $\Delta\chi^2$ computed with the Satorra-Bentler scaled chi-square difference test (Satorra & Bentler, 2010), developed for *robust chi-square* estimation. However, given the sensitivity of $\Delta\chi^2$ to sample size (Chen, 2007), we assessed invariance by also computing and mainly considering the criteria of ΔCFI and $\Delta RMSEA$ <0.01 and 0.015, respectively (Chen, 2007).

As previously indicated, the invariance test was conducted on two subsamples of male and female participants. The CFA implemented separately on each sample revealed excellent model's goodness-of-fit statistics both for males ($N = 157$; $\chi^2 = 7.764$, $df = 9$, $p = 0.56$; CFI = 0.99; TLI = 0.99; SRMR = 0.031; RMSEA = 0.001, *90 % CI* = 0.001, 0.066) and females ($N = 328$; $\chi^2 = 18.762$, $df = 9$, $p = 0.03$; CFI = 0.99; TLI = 0.99; SRMR = 0.022; RMSEA = 0.058, *90 % CI* = 0.025, 0.089). The factor loadings ranged from 0.66 to 0.79 for males and 0.73 to 0.84 for females. They were also all significant across the two groups. Table 4 shows the goodness-of-statistics for nested multigroup CFA models about gender. Comparisons across nested multigroup models indicated configural, metric, and scalar invariance of the BIS Italian version

Table 2
Items' descriptive statistics of the Italian and English versions of the BIS.

Label	Italian version	M	SD	Skew.	Kurt.	English version	M	SD	Skew.	Kurt.
BIS1	Tollero poco la noia	3.12	1.02	−0.16	−0.97	I have little tolerance for boredom	3.21	1.06	−0.23	−0.72
BIS2	Considero la noia spiacevole	3.37	0.95	−0.48	−0.61	I think boredom is unpleasant	3.62	1.04	−0.67	−0.21
BIS3	Considero la noia irritante	3.11	1.07	−0.22	−0.96	I think boredom is irritating	3.53	1.06	−0.66	−0.22
BIS4	Considero la noia un'emozione poco tollerabile	2.93	1.00	0.10	−1.00	I think boredom is a rather intolerable emotion	3.06	1.08	−0.06	−0.74
BIS5	Considero insopportabile trovarmi in situazioni noiose	2.85	0.99	0.25	−0.77	I think it's unbearable being in boring situations	3.03	1.15	−0.09	−0.92
BIS6	È frustrante trovarsi in situazioni noiose	3.20	1.00	−0.27	−0.74	It's frustrating finding myself in boring situations	3.45	1.05	−0.56	−0.33

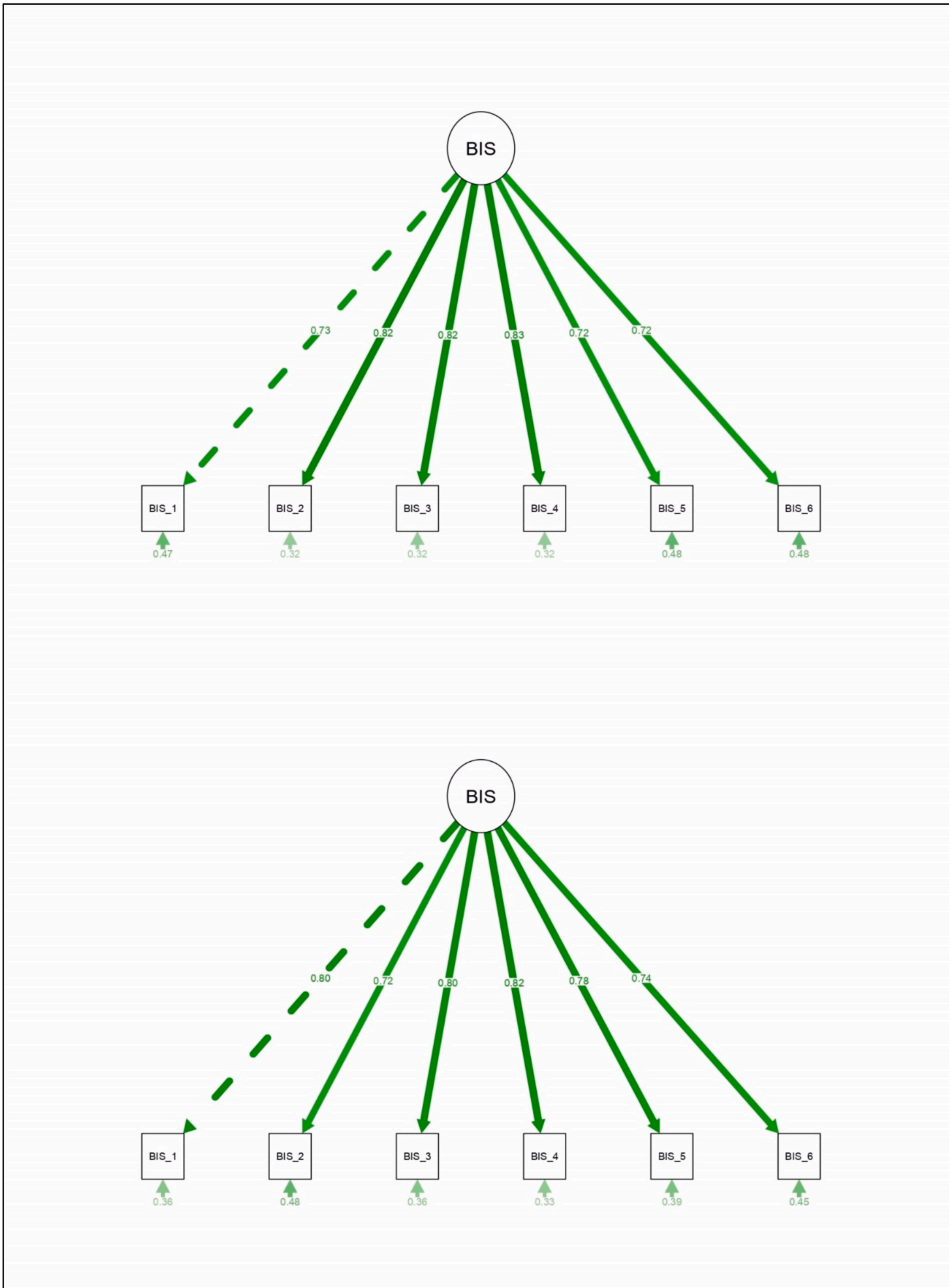

Fig. 2. Factor structure of the Italian (lower) and English (upper) versions of the BIS.

Table 3
Items' factor loadings of the CFAs for the Italian and English versions of the BIS.

Item	Italian version						English version					
	β	*se*	*z*	*p*	*95 % CI*		*β*	*se*	*z*	*p*	*95 % CI*	
					Lower	Upper					Lower	Upper
BIS1	0.80	0.023	34.55	<0.001	0.754	0.844	0.73	0.031	23.32	<0.001	0.665	0.787
BIS2	0.72	0.026	28.00	<0.001	0.672	0.774	0.82	0.026	32.28	<0.001	0.774	0.874
BIS3	0.80	0.021	38.06	<0.001	0.756	0.838	0.82	0.023	35.16	<0.001	0.776	0.868
BIS4	0.82	0.022	36.66	<0.001	0.774	0.862	0.83	0.021	38.48	<0.001	0.785	0.869
BIS5	0.78	0.022	35.37	<0.001	0.739	0.826	0.72	0.028	26.21	<0.001	0.667	0.775
BIS6	0.74	0.025	29.43	<0.001	0.694	0.794	0.72	0.034	21.40	<0.001	0.653	0.784

Table 4
Goodness-of-fit statistics of nested multigroup CFA models for participants' gender of the Italian version of BIS.

Model	χ^2 (*df*)	χ^2/df	*p*	*CFI*	*RMSEA (90 % CI)*	Model comparison	$\Delta\chi^2$	Δdf	*p*	ΔCFI	$\Delta RMSEA$
Model 1 Configural Invariance	25.622 (18)	1.42	0.11	0.995	0.042 (0.000, 0.070)	–	–	–	–	–	–
Model 2 Metric Invariance	32.256 (23)	1.40	0.10	0.994	0.041 (0.000, 0.067)	2 vs. 1	6.377	5	0.27	0.001	0.001
Model 3 Scalar Invariance	35.346 (28)	1.26	0.16	0.995	0.033 (0.000, 0.059)	3 vs. 2	1.900	5	0.86	0.001	0.008

Note. Male: N = 157; Female: $N = 328$. The goodness-of-fit statistics and the related comparison are based on a robust estimator.
$\Delta\chi^2$ has been computed using the Satorra-Bentler scaled chi-square difference test. Model 2 imposes equality constraints on loadings and Model 3 on intercepts.

concerning participants' gender.

4.2.3. Convergent validity

To test the construct validity of the Italian version of the BIS, we first computed correlations between it and other theoretically convergent measures. More in detail, we investigated the associations of the BIS with measures of the trait (i.e., BPS) and state (i.e., MSBS) boredom. As expected, the BIS was positively correlated with the trait boredom measure of the BPS ($r = 0.27$, $p < 0.001$, *95 % CI* = 0.182, 0.348) and with the pertaining subdimensions of Apathy ($r = 0.26$, $p < 0.001$, *95 % CI* = 0.170, 0.336) and External Stimulation-Challenge ($r = 0.34$, $p < 0.001$, *95 % CI* = 0.254, 0.412), while it was unrelated to Internal Stimulation-Creativity ($r = -0.01$, $p = 0.88$, *95 % CI* = −0.096, 0.082). Correlation analysis also showed a positive association of the BIS with the investigated state boredom measure of MSBS ($r = 0.28$, $p < 0.001$, *95 % CI* = 0.197, 0.361) and with the related subdimensions of Disengagement ($r = 0.27$, $p < 0.001$, *95 % CI* = 0.183, 0.348), High Arousal ($r = 0.26$, $p < 0.001$, *95 % CI* = 0.172, 0.338), Low Arousal ($r = 0.26$, $p < 0.001$, *95 % CI* = 0.180, 0.345), Inattention ($r = 0.17$, $p < 0.001$, *95 % CI* = 0.081, 0.254), and Time Perception ($r = 0.20$, $p < 0.001$, *95 % CI* = 0.115, 0.286). Then, we also tested correlations of the BIS with another conceptually convergent measure about the fear of relaxation-related events (RSI) and with the personality trait of neuroticism. As predicted, the BIS showed to be positively and significantly associated with RSI ($r = 0.22$, $p < 0.001$, *95 % CI* = 0.129, 0.299) and the BIF-10 subdimension of neuroticism ($r = 0.25$, $p < 0.001$, *95 % CI* = 0.166, 0.333). Thus, these results provide empirical support for the convergent validity of the BIS, corroborating the construct robustness of our new proposed instrument.

To further test the BIS's construct validity, we computed bivariate correlations with negative emotional states potentially related to an increased boredom intolerance. We first investigated the association of the BIS with a general measure of negative affect (i.e., NA). Analysis revealed a positive and significant correlation of the BIS with NA ($r = 0.18$, $p < 0.001$, *95 % CI* = 0.092, 0.265). Then, we investigated the potential association of the BIS with emotional negative correlates of anxiety, worry, anger, impulsivity, and depressive states. As for anxiety, we found significant positive associations of the BIS with the ASI ($r = 0.35$, $p < 0.001$, *95 % CI* = 0.270, 0.427), STAI-T ($r = 0.30$, $p < 0.001$, *95 % CI* = 0.221, 0.383), and PSWQ ($r = 0.32$, $p < 0.001$, *95 % CI* = 0.232, 0.393). As for anger and impulsivity, we found a positive association of the BIS with the STAXI-T ($r = 0.27$, $p < 0.001$, *95 % CI* = 0.183, 0.349) and with the BIS-11 ($r = 0.13$, $p = 0.005$, *95 % CI* = 0.038, 0.213). Regarding depressive symptomatology, correlation analysis revealed a positive association of the BIS with the CES-D ($r = 0.23$, $p < 0.001$, *95 % CI* = 0.148, 0.316). Finally, as a further test of the construct validity, we were also interested in investigating BIS's potential association with measures of general well-being and a boredom-related concept of the meaning of life. Specifically, we expected and found a negative association of the BIS with participants' life satisfaction (SLWS: $r = -0.10$, $p = 0.021$, *95 % CI* = −0.016, −0.192) and participants' purpose in life (PILS: $r = -0.15$, $p < 0.001$, *95 % CI* = −0.057, −0.231). Taken together, these results provided robust empirical evidence in favor of the construct validity of the BIS. For a complete overview of the correlations between all the variables investigated in Study 2, see Table 3 in the supplemental materials.

4.2.4. Discriminant validity

To test the discriminant validity of the BIS, we used the CI_{CFA} (*cut*) technique suggested by Rönkkö and Cho (2022). It implies estimating a CFA model that incorporates all the scales being assessed for discriminant validity and comparing correlations among latent factors. Rather than fix the first factor loading to 1, the latent variables are scaled by setting their variances to 1 so that the obtained covariance between them corresponds to their correlation. The technique inspects the upper limits of the *95 % CIs* of the estimated factors' correlation by comparing the value obtained in the baseline model against a model built by constraining such correlation to a specified cutoff. Rönkkö and Cho (2022) found that correlations below 0.8 were seldom problematic, suggesting this value may indicate the absence of discriminant validity problems. The analysis returns factors' correlation estimates, with their confidence intervals, and a likelihood ratio test for nested models where a significant χ^2 statistic indicates support for discriminant validity.

Thus, specifying the abovementioned cutoff of 0.8, we first compared correlations among BIS, the aggregate scale level of state boredom, and each subdimension. Analysis revealed that BIS was distinct from the MSBS aggregate score ($r = 0.31$, *95 % CI* = 0.215, 0.397, $\chi^2_{diff} = 17.09$, $df = 1$, $p < 0.001$), the subdimension of Disengagement ($r = 0.29$, *95 % CI* = 0.198, 0.388, $\chi^2_{diff} = 34.77$, $df = 1$, $p < 0.001$), High Arousal ($r = 0.29$, *95 % CI* = 0.195, 0.391, $\chi^2_{diff} = 37.65$, $df = 1$, $p < 0.001$), Low Arousal (r

= 0.30, *95 % CI* = 0.205, 0.387, χ^2_{diff} = 40.10, *df* = 1, *p* < 0.001), Inattention (*r* = 0.20, *95 % CI* = 0.100, 0.306, χ^2_{diff} = 46.65, *df* = 1, *p* < 0.001), and Time Perception (*r* = 0.23, *95 % CI* = 0.138, 0.316, χ^2_{diff} = 53.78, *df* = 1, *p* < 0.001). Discriminant validity analysis also revealed that the BIS was empirically distinct from the measure of BPS aggregate score (*r* = 0.35, *95 % CI* = 0.250, 0.440, χ^2_{diff} = 17.10, *df* = 1, *p* < 0.001), the subdimension of Apathy (*r* = 0.30, *95 % CI* = 0.194, 0.396, χ^2_{diff} = 37.97, *df* = 1, *p* < 0.001) and External Stimulation-Challenge (*r* = 0.42, *95 % CI* = 0.329, 0.517, χ^2_{diff} = 28.58, *df* = 1, *p* < 0.001). As for the dimension of Internal Stimulation-Creativity, the analysis was redundant as it was not related to the BIS (*r* = −0.05, *95 % CI* = −0.159, 0.067, χ^2_{diff} = 36.44, *df* = 1, *p* < 0.001). These results provided empirical corroboration for the discriminant validity of the BIS Italian version, highlighting its property of measuring a distinct construct from other boredom measures.

5. Study 3

Study 3 aimed to validate and assess the psychometric properties of the English version of the BIS. The scale was translated from Italian into English according to the parallel back translation procedure (Brislin, 1986), in which two bilingual persons independently translated the scale from its original language to the language under study. Then, a committee comprised of individuals who participated in the translation process assessed the new scale obtained. They prepared the scale format and the instructions identically to the original version. Then, another bilingual individual, unfamiliar with the original scale but well-versed in the psychological lexicon, translated this version back to the original language. Finally, the new Italian version obtained was sent to the original authors to verify the concordance between the original scale and the translation.

Once the English version of the scale was established, we implemented a confirmatory factor analysis to evaluate the factor structure of the BIS in an English sample. As in Study 2, we investigated the BIS's gender invariance through multigroup CFA. More importantly, we also assessed the measurement invariance by comparing the BIS's factor structure across Italian (sample of Study 2) and English (sample of the present Study) participants.

Finally, we inspected the BIS's convergent validity by probing its potential associations with boredom proneness, relaxation sensitivity, neuroticism, anxiety sensitivity, trait anxiety, and depressive symptomatology. Discriminant validity was tested by statistically evaluating whether the BIS was empirically distinct from boredom proneness and its subdimensions. Given that data for this study was collected through Prolific Academic and participants were paid based on the time taken to complete the questionnaire, we administered fewer measures than in Study 2 for reasons linked to the availability of our economic resources.

5.1. Method

5.1.1. Participants and procedure

The sample size was established using the same procedure of power analysis implemented in Study 2, which indicated a minimum sample size of 274 participants to achieve the conventional power of 0.80. Thus, we recruited a community sample of 406 English native respondents (328 female, Mage = 36.91, SDage = 10.68) to err on the conservative side. Participants were recruited through Prolific Academic and received monetary compensation (£6/hour) for enrolling in the study. We prescreened participants based on their nationality (English), country location (England), first language (English), ethnicity (White Caucasian), and gender (approximately 50 % female and 50 % male). Participants were first presented with a brief introduction describing the general aims of the research and asked to consent to participate. They were then presented with a questionnaire asking them to provide demographic information and to fill in the Boredom Intolerance Scale and other measures described below. In the present Study, we included three "instructed items" as an attention check (DeSimone et al., 2015). One was presented at the beginning of the questionnaire (after the BPS scale), one in the middle (after the STAI-T scale), and the other at the end, as in the previous Studies. In case of incorrect responses to one of these items, the completion of the questionnaire was interrupted, participants were thanked for their participation and their answers were not recorded.

The participants were all from England and were of White Caucasian ethnicity. Regarding educational level, participants were distributed as follows: 22.9 % had a secondary education, 27.1 % had a high school diploma, 48.8 % had a master's or bachelor's degree, and 1.2 % had a post-graduate qualification.

5.1.2. Measures

Participants answered the *Boredom Intolerance Scale* (BIS), *Boredom Proneness Scale* (BPS; Vodanovich & Kass, 1990), emotional stability/neuroticism dimension from the *10-item Big Five Inventory* (Rammstedt & John, 2007), *Relaxation Sensitivity Index* (RSI; Luberto et al., 2021), *Anxiety Sensitivity Index-3* (ASI; Taylor et al., 2007), *Trait Anxiety* (STAI-T; Spielberger, 1983), *Center for Epidemiologic Studies Depression Scale* (CES-D; McDowell & Newell, 1996). Detailed descriptions of all measures (including integral versions, descriptive statistics, scoring, and goodness-of-fit indicators) are presented in the supplemental materials.

5.2. Results

5.2.1. Confirmatory factor analysis and reliability

The CFA was implemented with the *lavaan* package (Rosseel, 2012) and performed with a *Robust Maximum Likelihood method* (MLM). The model fit was assessed as in Study 2. Items did not present normality issues (Table 2).

Confirmatory Factor Analysis revealed a satisfactory model fit of the 6-item and unidimensional English version of the Boredom Intolerance Scale (Fig. 2). Besides a significant robust *chi-square* statistic (χ^2 = 40.139, *df* = 9, *p* < 0.001), we found good values of 0.97 and 0.95 for the CFI and TLI, respectively. Similarly, we found an excellent value of 0.033 for the SRMR and an acceptable value of 0.092 (*90 % CI* = 0.069, 0.117) for the RMSEA. The analysis thus revealed that the tested model adhered fairly to the observed data, showing satisfactory goodness-of-fit statistics. As can be seen in Table 3, items' factor loadings were high and significant, highlighting standardized coefficients between 0.72 and 0.83. The confirmed 6-item and unidimensional English version of the BIS proved excellent reliability, showing high coefficients of 0.90 (*95 % CI* = 0.877, 0.916) and 0.90 for α and ω, respectively.

5.2.2. Measurement invariance across participants' gender and nationality

We conducted two distinct multigroup CFAs to test the BIS's confirmed structure's robustness and potential generalization. As in Study 2, we assessed for configural, metric, and scalar invariance across the subsample of male and female participants. Moreover, and importantly for generalization purposes, we tested the BIS's configural, metric, and scalar invariance across participants' nationalities by merging Studies 2 and 3 samples. CFAs were based on a robust estimator (i.e., MLM). Measurement invariance was assessed by computing and mainly considering the parameters presented in Study 2.

The first invariance test compared the BIS's factorial structure across two subsamples of male and female participants. The CFA revealed acceptable model's goodness-of-fit statistics both for males (*N* = 205; χ^2 = 27.684, *df* = 9, *p* = 0.001; CFI = 0.97; TLI = 0.95; SRMR = 0.034; RMSEA = 0.101, *90 % CI* = 0.066, 0.137) and females (*N* = 201; χ^2 = 24.432, *df* = 9, *p* = 0.004; CFI = 0.97; TLI = 0.94; SRMR = 0.039; RMSEA = 0.092, *90 % CI* = 0.058, 0.128). The factor loadings ranged from 0.72 to 0.85 for males and 0.70 to 0.83 for females. They were all significant across the two groups. Table 5 shows the goodness-of-fit statistics for nested multigroup CFA models about gender. Comparisons across nested models indicated configural, metric, and scalar invariance of the BIS English version across participants' genders.

Table 5
Goodness-of-fit statistics of nested multigroup CFA models for participants' gender of the English version of the BIS.

Model	χ^2 (*df*)	χ^2/df	*p*	*CFI*	*RMSEA (90 % CI)*	*Model comparison*	$\Delta\chi^2$	Δdf	*p*	ΔCFI	$\Delta RMSEA$
Model 1 Configural Invariance	51.922 (18)	2.88	0.001	0.966	0.096 (0.072, 0.122)	–	–	–	–	–	–
Model 2 Metric Invariance	58.944 (23)	2.56	0.001	0.964	0.088 (0.064, 0.112)	2 vs. 1	2.027	5	0.85	0.002	0.008
Model 3 Scalar Invariance	64.813 (28)	2.31	0.16	0.964	0.080 (0.058, 0.103)	3 vs. 2	3.671	5	0.60	0.000	0.008

Note. Male: N = 205; Female: *N* = 201. The goodness-of-fit statistics and the related comparison are based on a robust estimator.
$\Delta\chi^2$ has been computed using the Satorra-Bentler scaled chi-square difference test. Model 2 imposes equality constraints on loadings and Model 3 on intercepts.

The second measurement invariance test concerned the participant's nationality. This last test was particularly relevant for the potential generalization of our new scale. To carry out the invariance tests, we combined the data relating to the Italian sample of Study 2 ($N = 485$) and those of the present study on English participants ($N = 406$). The peculiar model's fit can be seen in the pertaining section of the manuscript. Table 6 shows the results of the measurement invariance test, which yielded configural, metric, and scalar invariance of the BIS's factorial structure across participants' nationalities.

5.2.3. Convergent validity

As in Study 2, we computed correlations between the BIS and other theoretically convergent measures to test the convergent validity of its English version. We first investigated the associations of the BIS with the trait boredom measure of BPS and its subdimensions. As expected, the BIS related positively with the BPS ($r = 0.27$, $p < 0.001$, *95 % CI* = 0.179, 0.359), as well as with the pertaining subdimensions of External Stimulation ($r = 0.38$, $p < 0.001$, *95 % CI* = 0.296, 0.462), Affective Responses ($r = 0.16$, $p < 0.001$, *95 % CI* = 0.062, 0.252), and Constraint ($r = 0.50$, $p < 0.001$, *95 % CI* = 0.418, 0.566). We did not find a significant association of the BIS with the dimension of Perception of Time ($r = 0.08$, $p = 0.13$, *95 % CI* = −0.022, 0.172) and, as in Study 2, Internal Stimulation ($r = -0.05$, $p = 0.36$, *95 % CI* = −0.143, 0.052). Correlation analysis also showed a positive association of the BIS with the conceptually convergent measure of the fear of relaxation-related events ($r = 0.17$, $p < 0.001$, *95 % CI* = 0.071, 0.261) and with the personality trait of neuroticism ($r = 0.18$, $p < 0.001$, *95 % CI* = 0.082, 0.271). Overall, these results provided empirical support for the convergent validity of the BIS English version, corroborating its construct soundness.

To further assess the construct validity of the English version of the BIS, we computed bivariate correlations between it and a series of negative emotional correlates of anxiety and depressive states. As for anxiety, we found significant positive associations of the English version of the BIS with the ASI ($r = 0.24$, $p < 0.001$, *95 % CI* = 0.143, 0.326) and STAI-T ($r = 0.22$, $p < 0.001$, *95 % CI* = 0.122, 0.307). As for depressive symptomatology, correlation analysis revealed a positive association of the BIS with the CES-D ($r = 0.20$, $p < 0.001$, *95 % CI* = 0.100, 0.287). These results provided further support for the construct validity of the BIS English version. For a complete overview of the correlations between all the variables investigated in Study 3, see Table 6 in the online supplemental materials.

5.2.4. Discriminant validity

We used the CI_{CFA} (cut) technique (Rönkkö & Cho, 2022), with a cutoff of 0.8, to test the discriminant validity even for the BIS English version. Discriminant validity analysis revealed that the BIS was empirically distinct from the measure of BPS aggregate score ($r = 0.31$, *95 % CI* = 0.210, 0.413, $\chi^2_{diff} = 11.34$, $df = 1$, $p < 0.001$), the subdimension of External Stimulation ($r = 0.44$, *95 % CI* = 0.330, 0.551, $\chi^2_{diff} = 14.31$, $df = 1$, $p < 0.001$), Affective Responses ($r = 0.22$, *95 % CI* = 0.101, 0.331, $\chi^2_{diff} = 33.05$, $df = 1$, $p < 0.001$), and Constraint ($r = 0.61$, *95 % CI* = 0.518, 0.707, $\chi^2_{diff} = 101.51$, $df = 1$, $p = 0.001$). As for the dimensions of Internal Stimulation ($r = -0.05$, *95 % CI* = −0.178, 0.081, $\chi^2_{diff} = 29.49$, $df = 1$, $p < 0.001$) and Perception of Time ($r = 0.05$, *95 % CI* = −0.071, 0.152, $\chi^2_{diff} = 47.13$, $df = 1$, $p < 0.001$), the analysis was redundant as they were not related to the BIS. These results provided corroboration for the discriminant validity of the BIS English version, highlighting its ability to assess a distinct construct from other boredom measures.

6. General discussion

The present research aimed to develop, validate, and evaluate the psychometric properties of the Boredom Intolerance Scale, a new tool for assessing boredom intolerance. To pursue these aims, we conducted three studies. Study 1 focused on exploring and establishing the factor structure of the BIS. An initial qualitative approach ensured that the items were ecologically valid and representative of real-life experiences of boredom intolerance. Quantitative analyses confirmed that these items cohered into a single factor, yielding a robust and reliable measure. EFA returned a single 6-item factor, aligning with theoretical expectations of boredom intolerance as a unidimensional construct. The emerged factor structure was accompanied by high factor loadings and reliability coefficients, endorsing the BIS as a concise and potentially effective tool for measuring boredom intolerance.

Study 2 aimed to validate the Italian version of the BIS and evaluate its psychometric properties. The primary objective was to confirm the unidimensional structure identified in Study 1. CFA corroborated the soundness of the 6-item unidimensional BIS, highlighting robust goodness-of-fit indicators, high factor loadings, and internal consistency. Measurement invariance indicated that the BIS structure was invariant across genders, confirming the BIS's ability to assess boredom intolerance consistently in men and women. Moreover, the BIS showed significant positive correlations with trait and state boredom, relaxation

Table 6
Goodness-of-fit statistics of nested multigroup CFA models for the Italian and English versions of the BIS.

Model	χ^2 (*df*)	χ^2/df	*p*	*CFI*	*RMSEA (90 % CI)*	Model comparison	$\Delta\chi^2$	Δdf	*p*	ΔCFI	$\Delta RMSEA$
Model 1 Configural Invariance	63.414 (18)	3.58	0.001	0.982	0.075 (0.059, 0.092)	–	–	–	–	–	–
Model 2 Metric Invariance	80.038 (23)	2.56	0.001	0.977	0.088 (0.064, 0.112)	2 vs. 1	15.916	5	0.007	0.005	0.013
Model 3 Scalar Invariance	112.795 (28)	4.03	0.001	0.966	0.093 (0.075, 0.111)	3 vs. 2	3.671	5	0.001	0.011	0.005

Note. Male: N = 205; Female: N = 201. The goodness-of-fit statistics and the related comparison are based on a robust estimator.
$\Delta\chi^2$ has been computed using the Satorra-Bentler scaled chi-square difference test. Model 2 imposes equality constraints on loadings and Model 3 on intercepts.

sensitivity, and neuroticism, supporting its convergent validity. Significant positive correlations were also found between the BIS and different measures of negative affect, including anxiety, anger, impulsivity, and depressive symptoms. Correlation analysis also indicated that the BIS was associated with reduced life satisfaction and reduced purpose in life. Moreover, discriminant validity analysis showed that associations of BIS with state and trait boredom were well below an acceptable threshold, highlighting the BIS's property of measuring a distinct construct from other boredom-related measures. Consistent with previous literature on boredom, these findings underscored the BIS's construct validity, underlining the potential implications for understanding the impact of boredom intolerance on mental health and well-being.

Finally, Study 3 aimed to validate and assess the psychometric properties of the BIS English version. The scale was translated from Italian using parallel back translation procedures, ensuring conceptual and linguistic equivalence. CFA revealed a satisfactory model fit for the 6-item unidimensional structure, accompanied by high factor loadings and reliability coefficients. Even in the English sample, multigroup CFA showed the gender invariance of the BIS. Importantly, through the combination of data from Studies 2 and 3, multigroup CFA also showed measurement invariance of the BIS across participants' nationalities. This result was particularly relevant for potentially generalizing the BIS structure, suggesting that the BIS maintains its measurement properties across English and Italian versions. Generalization purposes were also supported by construct validity analysis. Study 3 replicated the findings from Study 2, highlighting consistent associations of the BIS English version with theoretically related constructs, though not too strong for being considered overlapped (i.e., discriminant validity). Again, we found positive associations of the BIS with the trait boredom, relaxation sensitivity, neuroticism, anxiety sensitivity, trait anxiety, and depressive symptomatology, reinforcing the BIS's construct validity. Besides replicating findings from Study 2, these results aligned with existing literature on the psychological relevance of boredom.

Overall, the three studies provided solid empirical support for developing and validating our novel proposed measure of boredom intolerance, demonstrating its reliability and psychometric properties. The findings indicated that the BIS's Italian and English versions were reliable and valid instruments for assessing boredom intolerance.

6.1. Theoretical and practical implications

The development of the Boredom Intolerance Scale may have theoretical implications advancing the understanding and measurement of boredom as a psychological construct. Available boredom measures, such as the BPS and the MSBS, focus on the dispositional tendency to experience boredom (Vodanovich & Kass, 1990) or the boredom experienced when it is detected (Fahlman et al., 2013). The BIS introduces the idea that boredom intolerance may be a separate and significant individual difference that influences people's emotional responses. This may represent a broadening of the boredom conceptualization and, more specifically, of its operationalization, allowing to capture the individual capacity to tolerate it and emphasizing its relevance. The positive relationships we found between the BIS and several adverse emotional criteria underscore the potential role that boredom intolerance may play in individual mechanisms of emotion regulation. People with high boredom intolerance may engage in maladaptive behaviors to cope with the discomfort felt in experiencing this emotion. Future research could profitably investigate the potential adaptive (e.g., practicing meditation or engaging in physical activity) and maladaptive (e.g., avoidance or risky behaviors, substance abuse, sensation seeking) coping strategies associated with high boredom intolerance and their potential repercussions on pathological behavior. Furthermore, as our findings on the negative association of BIS with general well-being criteria underline, high levels of boredom intolerance could have repercussions on life satisfaction, meaning, and goal pursuit. Those who have difficulty tolerating boredom could experience a potential impairment of multiple life domains (e.g., work, relationships, etc.) with repercussions on general well-being and the pursuit of long-term goals. This highlights the need to address boredom intolerance also in therapeutic contexts. For instance, cognitive-behavioral frameworks focusing on coping strategies and emotional resilience could benefit from using such a tool.

6.2. Limitations and future research

Some limitations need to be acknowledged. First, our measure development was conducted using same-source, cross-sectional data collection. Future research based on data from other sources and objective outcomes could help bolster our results. This may also help future more in-depth tests of the criterion validity of the BIS, investigating its impact on external criteria and outcomes. Another limitation may be restricting the English-speaking population (Study 3) to Caucasians in England. England is roughly 82 % Caucasian, thus limiting our sample and preventing us from culturally generalizing our proposed instrument. Future research may profitably investigate the invariance of the BIS across cultures.

7. Conclusion

In conclusion, the present paper proposed the conceptual framework underlying boredom intolerance, the methodology employed in the scale's development, and the findings from validation studies. By offering a consistent measure of boredom intolerance, the BIS has the potential to advance research in this field and inform interventions aimed at enhancing individuals' capacity to manage boredom effectively. Through this manuscript, we aim and hope to contribute to the growing body of literature on boredom and its impacts, providing researchers and practitioners with a potentially valuable tool for assessing and addressing boredom intolerance. The implications of this work strive to extend beyond academic inquiry, highlighting the importance of fostering resilience to boredom in promoting mental health and well-being.

CRediT authorship contribution statement

Valerio Pellegrini: Writing – review & editing, Writing – original draft, Software, Methodology, Investigation, Formal analysis, Data curation, Conceptualization. **Estelle Leombruni:** Writing – review & editing, Investigation, Conceptualization. **Stefania Iazzetta:** Writing – review & editing, Investigation, Conceptualization. **Marco Saettoni:** Supervision, Conceptualization. **Andrea Gragnani:** Writing – review & editing, Supervision, Resources, Data curation, Conceptualization.

Consent for publication

The manuscript has been seen and reviewed by all authors, and all authors agree to the submission of the manuscript in its current form.

Ethics approval

All procedures performed in studies involving human participants were conducted in accordance with the ethical standards of the institutional and national research committee and the 1964 Declaration of Helsinki and its later amendments or comparable ethical standards. The article does not refer to any studies with animals performed by any of the authors.

All Studies performed in the research received ethical approval from the ethics committee of the Associazione Scuola di Psicoterapia Cognitiva (SPC), Viale Castro Pretorio, 116, 00185, Rome, Italy, with the following protocol numbers: N. Pr. 06/24 and N. Pr. 07/24.

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Declaration of competing interest

We formally declare that the manuscript has not been previously published in any form. It is neither under consideration nor in press with another publisher. The authors declare no conflicts of interest.

Appendix A. Supplementary data

Supplementary data to this article can be found online at https://doi.org/10.1016/j.paid.2025.113151.

Data availability

Data and materials are made available on Open Science Framework at the following link: https://doi.org/10.17605/OSF.IO/7F3J9

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