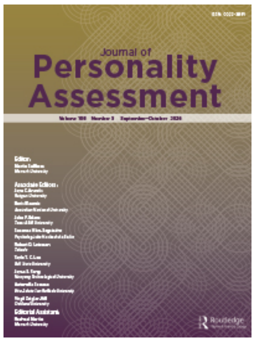

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Measurement Invariance of the Brief State Rumination Inventory Across Administration Methods, National Samples, and Sex

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ABSTRACT

State rumination is a temporary activation of repetitive, negative, self-focused thinking in response to specific situational triggers or mood states. We present the results of two studies that examined the psychometric properties and measurement invariance of the Italian version of the Brief State Rumination Inventory (BSRI). Study 1 tested its unidimensionality and measurement invariance across administration methods (paper-and-pencil, $N=216$, vs. online, $N=222$) and national samples. Study 2 ($N=675$) tested its measurement invariance across sex and its construct validity. Results supported the BSRI's unidimensional structure. Configural and weak (metric) invariance were supported across administration methods and national samples, whereas partial strong invariance was established across sex. The Italian BSRI showed construct-validity evidence consistent with theoretical expectations, correlating more strongly with brooding than reflection and more strongly with depression than anxiety or stress. Neuroticism and limited access to emotion regulation strategies emerged as other significant correlates. These findings support the Italian BSRI as a robust tool for assessing state rumination.

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Rumination constitutes a maladaptive form of self-focused cognition (Mor & Winquist, 2002; Nolen-Hoeksema et al., 2008), defined by repetitive and passive contemplation of the causes and consequences of negative affect (Nolen-Hoeksema, 1991). Extensive evidence links rumination to deleterious outcomes, including elevated risk for depression (Nolen-Hoeksema, 2000; Nolen-Hoeksema et al., 1994), prolonged depressive episodes (Kuehner & Weber, 1999; Nolen-Hoeksema et al., 1993), and comorbid psychopathologies such as anxiety disorders, substance misuse, and bulimic behaviors (Kocovski et al., 2005; Nolen-Hoeksema et al., 2007). Conceptualized as a habitual, trait-like response to negative mood states (Lyubomirsky et al., 2015; Nolen-Hoeksema et al., 2008), rumination has been operationalized through several validated instruments, including the Ruminative Response Scale (RRS; Nolen-Hoeksema & Morrow, 1991), the Rumination-Reflection Questionnaire (RRQ; Trapnell & Campbell, 1999), the Repetitive Thinking Questionnaire (RTQ; Mahoney et al., 2012), and the Perseverative Thinking Questionnaire (PTQ; Ehring et al., 2011).

Rumination has recently been conceptualized as a state-dependent cognitive response, triggered by situational factors such as perceived discrepancies between goals and current circumstances (Michel-Kröhler et al., 2023; Smith & Alloy, 2009; Tamm et al., 2024). Experimental research demonstrated that state rumination exerts detrimental effects on mood, cognition, problem-solving, goal-directed behavior, and psychopathology (Lyubomirsky et al., 2015). Despite growing interest, its assessment remains challenging; prior approaches often relied on poorly defined item sets, inconsistent temporal frames, or overly specific contexts. To address these limitations, Marchetti et al. (2018) introduced the Brief State Ruminative Inventory (BSRI), a concise self-report measure designed to capture momentary fluctuations in rumination. Items are rated on a 100-mm visual analog scale in the paper-and-pencil format or a 100-point slider online, with instructions referencing the present moment ("right now"). The initial 13-item pool, derived from established trait measures (RRS, PTQ, RTQ, RRQ), was refined to eight items through validation studies. Both formats exhibit a unidimensional structure, satisfactory reliability, and robust construct and criterion validity (Marchetti et al., 2018).

Since its introduction, the Brief State Ruminative Inventory (BSRI) has become a widely used tool for examining momentary rumination across diverse psychological contexts. Empirical applications spanned behavioral research - such as studies on alcohol-related rumination in hazardous drinkers (Mollaahmetoglu et al., 2021), maladaptive eating behaviors (Kornacka et al., 2021), and insomnia (Ballesio

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et al., 2020) - as well as developmental investigations involving adolescents and children (Bonduelle et al., 2023; García-Morales et al., 2024). Beyond behavioral paradigms, the BSRI has been integrated with neuroimaging techniques, including EEG and fNIRS (Long et al., 2024; Luo et al., 2024), and employed in intervention research targeting rumination reduction (e.g., Guillaumée et al., 2024; Horczak et al., 2023; Mandelbaum & Kalanthroff, 2024; Owens & Bunce, 2022; Perlman et al., 2024; Shireen et al., 2024). Its adaptability has facilitated cross-cultural use, with translations into multiple languages (Marchetti et al., 2024), although formal validation studies exist only for Turkish (Altan-Atalay et al., 2020) and Chinese (Wang et al., 2022).

The Turkish adaptation demonstrated good fit for a one-factor model (CFI/TLI > .90; RMSEA < .08) after minor residual correlations were freed, with strong internal consistency (α = .91) and evidence of construct and incremental validity. BSRI scores correlated strongly (*r*s ≥ |.50|) with trait rumination and negative affect, moderately (|.30| ≤ *r*s < |.50|) with depression and coping strategies, and predicted depressive symptoms beyond demographic and trait measures. Moreover, the BSRI captured state fluctuations following rumination induction ($d = 0.76$). Similarly, the Chinese version (Wang et al., 2022) showed high reliability (α = .93), a unidimensional structure explaining 65% of variance, and sensitivity to experimental manipulations ($d = 0.64$).

Given its theoretical relevance and widespread application, the present study aimed to adapt the BSRI for the Italian cultural context, evaluate its psychometric properties, and test measurement invariance across administration formats, national samples, and sex. This work addresses a critical gap, as most validation efforts have focused on non-Southern European populations, leaving its applicability in Italy largely unexplored.

Cultural variation has been documented in both the prevalence of trait rumination and its psychological correlates (Chang et al., 2010; Choi & Miyamoto, 2023; Grossmann & Kross, 2010; Kwon et al., 2013; Sakamoto et al., 2001). For example, European Americans reported lower rumination frequency than Asian Americans (Chang et al., 2010) and South Koreans (Kwon et al., 2013). Moreover, the association between rumination and maladjustment appeared weaker among Asian Americans compared to European Americans (Chang et al., 2010) and may even be adaptive in certain contexts, as evidenced by its positive link to problem-solving in Japanese samples (Sakamoto et al., 2001). In Italy, trait rumination has shown moderate-to-strong correlations with worry, anger rumination, metacognitive beliefs, eating disorder symptoms, depression, anxiety, emotion regulation difficulties, neuroticism, and spontaneous mind wandering (Mansueto et al., 2022; Palmieri et al., 2024; Vannucci & Chiorri, 2018). These findings underscore the likelihood that state rumination, like its trait counterpart, may exhibit culturally specific patterns.

Adapting the BSRI for Italian speakers is therefore essential for capturing these nuances and supporting conceptual comparability and the evaluation of measurement invariance across groups. A culturally sensitive instrument will enable researchers and clinicians to assess transient ruminative thought with greater accuracy, improving the validity of cross-cultural comparisons and informing interventions tailored to Italian-speaking populations.

The second aim of this work was to address two questions not previously examined: whether sex differences exist in state rumination and whether these differences vary across age. Research on trait rumination consistently reported small sex effects, with women scoring higher than men. A meta-analysis by Rood et al. (2009) found negligible differences in children ($d = 0.14$) and small differences in adolescents ($d = 0.36$), a pattern replicated in adults ($d = 0.24$, Johnson & Whisman, 2013; $d = 0.19$, Tamres et al., 2002). To date, comparable evidence for state rumination is lacking. Preliminary findings by Marchetti et al. (2018) suggest negligible, non-significant sex differences (*d*s ranging from 0.08 to 0.31), slightly favoring women.

Valid mean comparisons require evidence of measurement invariance across sex; without it, observed differences may reflect measurement bias rather than true variation. Previous research on trait rumination measures, such as the RRS (Whisman et al., 2020) and the Children's Response Style Questionnaire (Stewart et al., 2022), has demonstrated at least partial invariance, with occasional higher scores among boys. Building on this work, we tested the measurement invariance of the BSRI across sex.

Finally, given evidence that rumination peaks in young adulthood and declines across the lifespan - with sex and cohort differences favoring younger women (Lilly et al., 2023) - we examined whether BSRI latent scores vary by age and whether these trajectories differ by sex.

As part of the Italian BSRI validation, we also assessed construct validity, our third aim. Previous studies have correlated BSRI scores with trait brooding and reflection, affect, emotion regulation, depression, anxiety, stress, mindfulness, worry, perseverative thinking, and coping strategies (Altan-Atalay et al., 2020; Marchetti et al., 2018; Wang et al., 2022). Although formal tests were lacking, correlations with brooding exceeded those with reflection, and associations with depression were stronger than with anxiety or stress. In this study, we replicated these analyses using the 10-item RRS for brooding/reflection and the DASS-21 for depression, anxiety, and stress, while employing the Difficulties in Emotion Regulation Scale (DERS; Gratz & Roemer, 2004) for emotion regulation. Prior research indicated strong associations between DERS and rumination ($r = .69$; Mansueto et al., 2022), though subscale-level analyses suggest nuanced patterns, with the Limited Access to Emotion Regulation Strategies (LAERS) subscale showing the strongest link to rumination, particularly among women (Ando' et al., 2020).

To the best of our knowledge, no study has examined the relationship between state rumination and the personality traits of the Five-Factor Model (FFM). Prior research on trait rumination consistently reported strong associations with neuroticism (Lyon et al., 2021; Teasdale & Green, 2004; Trapnell & Campbell, 1999; Vannucci & Chiorri, 2018), supporting the notion that a tendency toward repetitive thoughts about past or future problems may underlie the heightened negative affect often observed in individuals high in neuroticism (Perkins et al., 2015). To explore this link at the state level, we included the Big Five Inventory (BFI; John et al., 1991) in the present study.

Based on these findings, we expected that the BSRI score would: (1) have a stronger relationship with a trait measure of brooding (a maladaptive form of trait rumination) than with a trait measure of reflection; (2) have a stronger relationship with neuroticism than with any of the other Big Five dimensions; (3) have a stronger relationship with depression than with anxiety and stress; and (4) have a stronger relationship with the LAERS subscale than with any of the other DERS subscales.

It should be noted that although the BSRI was expressly developed to assess momentary, situationally triggered fluctuations in rumination, its validation studies, including the present one, have relied on an administration of trait measures. As a result, evidence for construct validity primarily reflects between-person differences in participants' momentary rumination at the time of assessment, rather than dynamic within-person processes. This distinction is crucial, since correlations with trait measures (e.g., brooding, neuroticism) indicate that individuals who report higher state rumination at that specific moment also tend to endorse trait-like vulnerabilities, but such associations cannot be taken as evidence that the BSRI is sensitive to short-term intra-individual variability, which is the core theoretical rationale for a state measure.

To address these aims, we conducted two studies. Study 1 evaluated the unidimensionality, reliability, and measurement invariance of the Italian BSRI across administration formats (online vs. paper-and-pencil), replicating the original validation. Leveraging Marchetti et al. (2018) dataset, we also tested invariance across national samples. Study 2 employed a larger sample to confirm unidimensionality and reliability, assessed sex-based measurement invariance, and examined the construct validity of the Italian BSRI.

Method

Participants

In Study 1, we aimed to recruit at least 141 participants per administration format, following Marchetti et al. (2018) Monte Carlo analysis demonstrating that this sample size ensures adequate power for CFA parameter estimation (Muthén & Muthén, 2002). To account for potential exclusions due to careless responding (e.g., straightlining; Ward & Meade, 2023), we exceeded this target. The paper-and-pencil sample comprised 216 Italian adults (57.9% women; mean age = 34.94, SD = 14.22, range = 18–75; median education = high school), recruited opportunistically by psychology students among acquaintances as part of research training. No specific inclusion or exclusion criteria were applied, but, to be enrolled, participants had to read and sign three forms (information on the processing of personal data, informed consent, and consent for the processing of personal data for research purposes) as required by the Italian law and code of ethics for the administration of psychological tests for clinical and research purposes.

Sample 2 included 222 participants (82.1% women; mean age = 32.92, SD = 8.04, range = 21–64; median education = university degree) who completed the online version of the Italian BSRI. Participants were recruited among Italian community-dwelling adults *via* psychology students (distinct from those recruiting Sample 1) as part of research training, using social media platforms (e.g., Instagram, Twitter, Facebook), online forums, and email invitations. A snowball sampling strategy was employed, encouraging invited participants to share the study within their networks. Inclusion criteria mirrored those of Sample 1, with informed consent and data processing agreements obtained electronically.

In Study 2 we planned to test measurement invariance across sex. Assuming full weak and strong invariance and a small ($d = 0.20$) latent mean difference, we computed that we needed at least 125 participants per group to meet Muthén and Muthén (2002)'s criteria. However, based on the results by Marchetti et al. (2018, Table 3), we computed that we needed at least 600 cases to reach a power of .90 to find a significant difference between the correlation of BSRI with RRS-Brooding and the correlation of BSRI with RRS-Reflection. We recruited Italian community-dwelling adults *via* convenience and snowball sampling, using social media advertisements posted by the authors, assistants, and students. Inclusion criteria matched those for online participants in Study 1. Of 2,032 participants who accessed the survey link, 1,325 were excluded for failing to complete any items, declining consent, or refusing data processing authorization. The final sample (Sample 3) comprised 675 participants (68.6% women; mean age = 35.27, SD = 13.46, range = 18–86; median education = high school), with no evidence of careless responding.

Measures

Socio-demographic information collection form

Before completing the questionnaires, participants were asked to provide information about their sex, the sex they identified with, their age (in years), their educational level, their relationship status, their occupational status, and their perceived social status. The latter variable was measured using a self-anchoring scale (Adler et al., 2000; Kilpatrick & Cantril, 1960). More details are reported in Appendix A of the Supplementary Materials (SM).

Italian version of the Brief State Rumination Inventory

The English Brief State Rumination Inventory (BSRI; Marchetti et al., 2018) was initially translated into Italian using a forward-translation procedure (Behling & Law, 2000). Four bilingual authors independently produced Italian versions of each item, which were then reviewed collaboratively to ensure linguistic fidelity and clarity in standard Italian. The final Italian version was obtained by consensus between the translators. A back-translation of the items was then conducted using an online translation service. The conceptual equivalence between the back-translated and original items was subsequently evaluated by an independent native English-speaking expert in clinical psychology, who confirmed their equivalence (for more details, see Appendix B of the SM). To assess comprehensibility, 10 participants with low-to-average education (high school or less; 5 women;

mean age = 40.50, SD = 14.98, range = 21–54) rephrased the items in their own words and completed the inventory to evaluate instructions and response format. All items were judged easy to understand and score.

Big Five Inventory

The Big Five Inventory (BFI; John et al., 1991, Italian version in Ubbiali et al., 2013). The BFI is a 44-item measure of the so-called Big Five, i.e., the most global personality dimensions in the adult population: extraversion (the orientation of one's interests and energies toward the outer world; 8 items), agreeableness (the tendency to act in a cooperative, unselfish manner; 9 items), conscientiousness (the tendency to be organized, responsible, and hardworking; 9 items), neuroticism (the tendency to have strong negative emotions, such as anger, anxiety, or depression; 8 items), and openness to experience (the tendency to be open to new aesthetic, cultural, or intellectual experiences; 10 items). Each item is rated on a 5-point, Likert-type agreement scale from 1 (*disagree strongly*) to 5 (*agree strongly*).

Ruminative Response Scale

The Ruminative Response Scale (RRS; Treynor et al., 2003, Italian version in Palmieri et al., 2007) is a self-report measure of trait rumination, i.e., obsessive thinking characterized by an excessive and repetitive focus on specific thoughts or themes, and reflection, i.e., deliberate introspection to engage in cognitive problem-solving to reduce one's symptoms of depression. It comprises 10 items, 5 for each dimension, and participants are asked to rate the frequency of their responses to negative and depressed mood using a 4-point Likert-type scale, ranging from 1 (*never*) to 4 (*always*).

Depression Anxiety Stress Scales-21

The Depression Anxiety Stress Scales-21 (DASS-21; Henry & Crawford, 2005, Italian version in Bottesi et al., 2015) is a self-report questionnaire assessing the frequency of depressive, anxious, and stress symptoms. Each scale comprises 7 items, and participants are asked to rate them on a 4-point Likert-type frequency scale, ranging from 1 (*never*) to 4 (*almost always*).

Difficulties in Emotion Regulation Scale

The Difficulties in Emotion Regulation Scale (DERS; Gratz & Roemer, 2004, Italian version in Giromini et al., 2012) is a self-report questionnaire designed to assess problems related to the regulation of emotions. Its 36 items ask participants to report how often (from 1 = *almost never* to 5 = *almost always*) they resort to six different emotion regulation strategies: (1) Lack of emotional awareness (LEA), i.e., lack of awareness or inattention to emotional responses; (2) Lack of emotional clarity (LEC), i.e., the extent to which a person knows and is clear about their emotions; (3) Difficulties engaging in goal-directed behavior (DEGDB), i.e., difficulty in focusing on and/or accomplishing tasks when experiencing negative emotions; (4) Impulse control difficulties (ICD), i.e., difficulty remaining in control of one's behavior when experiencing negative emotions; (5) Nonacceptance of emotional responses (NER), i.e., tendency to have a non-accepting reaction to one's own distress; (6) Limited access to emotion regulation strategies (LAERS), i.e., the belief that there is little one can do to regulate oneself once upset.

Procedure

In Study 1, participants completed only the demographic questionnaire and the BSRI in addition to the consent forms described in the Participants section. The completion of this task took approximately 5–10 min. Participants in Study 2 were also administered the battery of measures described in the previous section. The presentation order was randomized for each participant to control for order and sequence effects. The completion of the battery took approximately 15–20 min.

All participants volunteered to participate, received no compensation, and were treated in accordance with the Declaration of Helsinki (World Medical Association, 2013) and the Ethical Principles of Psychologists and Code of Conduct of the American Psychological Association (2017) and of the Italian Psychological Association (Associazione Italiana di Psicologia, 2022).

Data analysis strategy

We first inspected item score distributions. Based on prior research and the response format, we anticipated deviations from normality, with skewness and kurtosis potentially exceeding the [−1, +1] range recommended for standard maximum likelihood (ML) estimation (Muthén & Kaplan, 1985). Consequently, we employed robust maximum likelihood estimation (MLR) in confirmatory factor analyses, providing standard errors and fit indices adjusted for nonnormality.

We then tested the plausible factor structure of the scale using dimensionality analyses, i.e., the scree-test (Cattell, 1966) and parallel analysis (Horn, 1965; Longman et al., 1989). Although we expected a single-factor measurement model, before testing it with CFA we sought empirical evidence of unidimensionality.

The measurement model of the BSRI was tested by fitting a single-factor CFA model in each separate sample. Following Marsh et al. (2004), we considered as indices of adequate and excellent fit the following: values of robust Comparative Fit Index (CFI) and Tucker-Lewis Index (TLI) larger than .90 and .95, respectively, and values of robust Root Mean Square Error of Approximation (RMSEA) smaller than .08 and .06, respectively. In case of a lack of adequate fit, we planned to investigate its sources through the modification indexes. Post hoc modifications to covariance structure models are usually discouraged, since they may capitalize on chance characteristics of the sample at hand (MacCallum et al., 1992). However, in the particular case of single-factor models, the sources of misfit are usually unmodeled residual covariances (or correlated uniquessesses, CUs) among items. Such an issue is common in empirical settings. It reflects systematic variance, often due to idiosyncratic item characteristics (e.g., similar wording), and is independent of the covariance due to the factor (Podsakoff et al., 2003).

To minimize capitalization on sample-specific characteristics when applying post hoc modifications, we implemented a resampling-based procedure. For each iteration, 1,000 bootstrap samples were drawn, the model was refitted, and modification indices were computed. Correlated uniquenesses (CUs) meeting inclusion criteria in at least 80% of replications were added to the respecified model. Model fit was then reassessed, and the procedure repeated until adequate fit was achieved. The inclusion criteria for the CUs were a modification index (i.e., the expected χ^2 value by which model fit would improve if a particular parameter was added) larger than 10 and a standardized expected parameter change (SEPC, i.e., the standardized estimated value of a fixed parameter if it were added to a model and freely estimated) larger than .20 (Whittaker, 2012).

Measurement invariance was tested using the same approach of Marchetti et al. (2018). We started by testing for factor structure invariance, also known as *configural invariance*, which requires that the same items in each sample serve as indicators of the same latent factor. We then tested *weak invariance*, which requires factor loadings to be invariant across samples. This model assumes that the scores of identical items will change by the same amount between samples with the same degree of increase on the latent factor (i.e., equality of scaling units). As in Marchetti et al. (2018), we considered weak invariance sufficient for this study, as it indicates that respondents across groups interpreted the latent construct similarly. Given the substantial socio-demographic differences among the available samples, we limited invariance testing to this level for comparisons between Italian participants completing the paper-and-pencil versus online versions, and between Italian and non-Italian participants. As shown in Appendix A of the SM (Table A1), the Italian samples differed markedly: although mean age was similar, the online sample included more women (82% vs. 58%) and was far more educated (85% held a master's degree vs. 18%). Comparable imbalances were present in cross-country comparisons: in Marchetti et al. (2018), nearly all participants were female university students, with a mean age about ten years younger than the Italian samples. Tests of weak invariance are relatively robust to differences in covariates because they do not assume equality of item intercepts or observed scores. A strong invariance model, in contrast, imposes equality constraints on both factor loadings and item intercepts. Intercepts are directly tied to observed scores and are highly sensitive to group differences in covariates. When samples are unbalanced on relevant characteristics, intercept equality may fail for reasons unrelated to measurement bias, such as genuine differences in item affectivity or response tendencies driven by demographic or cultural factors. Forcing intercept equality under these conditions risks misrepresenting the data and invalidating conclusions about latent mean differences. For these reasons, we tested strong invariance only when comparing women and men in Study 2.

A *strong invariance* model (i.e., invariance of item intercepts) evaluates differential item functioning (DIF) across populations and tests latent factor mean differences. Intercept invariance ensures that identical factor levels yield equivalent average item scores across groups, meaning observed score differences reflect true differences in latent means rather than group bias. Since it is common that strong invariance cannot be supported, we considered the possibility of resorting to partial invariance. Byrne et al. (1989) argued that mean comparisons remain valid if at least two items per latent trait satisfy weak and strong invariance, making full invariance unnecessary for further analyses. Freeing intercepts excludes non-invariant items from latent mean estimation, without loss of accuracy (Cheung & Rensvold, 2002). To identify sources of non-invariance, we applied the iterative procedure described above to free specific CUs.

Invariance models are typically compared using a chi-square difference test. However, this test is highly sensitive to sample size (Bentler & Bonett, 1980; Marsh et al., 1988). Consequently, we used other ad hoc criteria to determine when discrepancies in fit were substantial enough to dismiss a more parsimonious model (i.e., a model with more constraints and fewer estimated parameters) in favor of a more complex model. As proposed by, e.g., Marsh et al. (2010), we considered a change in CFI of less than .01 (Chen, 2007; Cheung & Rensvold, 2001) or a change in RMSEA of less than .015 (Chen, 2007) as sufficient support for the more parsimonious model.

The association with age was tested by adding age as a predictor of factor scores in the strong invariance model, in which we added a parameter constraint to test whether the regression coefficients of women and men differed.

The hypotheses about construct validity were tested using the omnibus test and the related post-hoc tests for the equality of dependent correlations by Meng et al. (1992). A false discovery rate (FDR) correction of the *p*-values was applied to post-hoc tests to control the inflation of Type I error due to multiple comparisons.

The analyses were performed using Mplus 7.0 (Muthén & Muthén, 2012) and several *R* functions, as listed in the Appendix C of the Supplementary Materials.

Results

Study 1a: Testing invariance across different administration methods

As a first step in the analyses, we computed the descriptive statistics and plotted the distributions of the BSRI item scores for all samples. With few exceptions, the distributions were all essentially symmetrical (skewness values < |1.00|), albeit somewhat flat (kurtosis values generally negative), as was to be expected given the particular response scale of the items (Figure SM1.1, Section 1 of the SMs). These results suggested the use of MLR as estimation method in CFAs.

We tested the plausible factor structure of the scale using the scree-plot and parallel analysis. Results indicated that a single factor was optimal across all samples, as the eigenvalue curve flattened after the second factor and only one observed eigenvalue exceeded the 95th percentile of its simulated distribution (Figure SM2.1, Section 2 of the SMs).

We then used CFA to test the fit of a single-factor model. According to the criteria described above, the fit was not acceptable in either sample (Paper-and-pencil sample: $\chi^2(20) = 71.685$, $p < .001$, scaling correction factor [SCF] = 1.553, CFI = .914, TLI = .880, RMSEA = .109 [.088, .132]; Online sample: $\chi^2(20) = 123.216$, $p < .001$, SCF = 1.291, CFI = .852, TLI = .793, RMSEA = .152 [.130, .176]). Given the results of the dimensionality analyses, the lack of fit could not be due to a lack of specification of additional factors and, therefore, there should be unmodeled residual covariances between the items. We thus used the procedure described above to identify the most replicable substantial CUs. In the paper-and-pencil sample, CUs that met the criteria for additional specification were those between items 1 and 2,7 and 8, and 2 and 7 (for more details, see Section 3 of the SMs, Table SM3.1). In the online sample, CUs that met the criteria for additional specification were those between items 2 and 3, 1 and 2, 1 and 3, and 2 and 7 (Table SM3.2). We then specified a CFA model that included the most replicable CU and tested the model again. If the fit was adequate, no additional parameters were specified; otherwise, we added the second most replicable CU, and so on. The additional specification of the first two CUs mentioned above to the model for the paper-and-pencil sample led to an acceptably fitting model ($\chi^2(18) = 41.558$, $p < .001$, SCF = 1.505, CFI = .961, TLI = .939, RMSEA = .078 [.053, .103], Table SM3.3). The additional specification of the first three CUs mentioned above to the model for the online sample led to an optimal fit ($\chi^2(17) = 26.124$, $p = .015$, SCF = 1.224, CFI = .987, TLI = .979, RMSEA = .049 [<.001, .081], Table SM3.4).

We then used these models to test the measurement invariance of the Italian BSRI across administration methods. The results are reported in Table 1 and show support for full weak invariance.

Finally, we computed the BSRI total score, its descriptive statistics, and its reliability in the two Italian samples. In the paper-and-pencil sample, the mean total score was 36.46 (*SD* = 19.18, range 0–77), the internal consistency (Cronbach's α) was .90 [.88, .92] and the reliability of the total unit-weighted sum score (McDonald's omega) was .91 [.89, .92]. In the online sample, the mean total score was 24.89 (*SD* = 17.34, range 0–74), Cronbach's α was .90 [.88, .92] and McDonald's omega was .90 [.88, .92].

In conclusion, this study showed that the BSRI had a replicable factor structure across administration methods and that the relationships between the underlying latent variables and their observed indicators were equivalent across groups.

Study 1b: Testing invariance across different national samples

Next, we tested whether the factor structure was invariant across national samples. Therefore, we used the data from the paper-and-pencil and online samples in the original study (Marchetti et al., 2018) to perform a test of invariance between Italian and Belgian participants for the paper-and-pencil version and Italian and US participants for the online version. The results are reported in Table 1. Full weak invariance was supported in the comparison between the Italian paper-and-pencil sample vs Belgian paper-and-pencil sample 1, while when we compared the Italian paper-and-pencil sample with the Belgian paper sample 2, the Satorra-Bentler Scaled Test Statistic was significant (see Table 1). However, the bootstrapping procedure for identifying the most replicable non-invariant factor loadings revealed that none exceeded the 80% threshold - the most replicable non-invariant parameter was the factor loading of item 1 (63.7%, Table SM4.1). Since the drop in CFI and TLI and the increase in RMSEA were less than .01, we did not consider these results as sufficient evidence for concluding that the full weak invariance model was not supported. When we compared the Italian online sample with the US online sample, we again found support for the full weak invariance model (Table 1).

In conclusion, this study showed that the BSRI had a replicable factor structure across national samples and that the relationships between the underlying latent variables and their observed indicators were equivalent across groups.

Study 2a: Testing invariance across sex

The descriptive statistics and score distributions for the BSRI items are shown in Figure SM1.1. As in Study 1, the dimensionality analyses suggested that the optimal number of factors was one (Figure SM2.1). We thus tested the single-factor model, which showed adequate fit without requiring additional parameters ($\chi^2(20) = 88.789$, $p < .001$, SCF = 1.217, CFI = .969, TLI = .957, RMSEA = .075 [.061, .089]). The mean total score was 43.77 (*SD* = 22.25, range 1–100), Cronbach's α was .91 [.90, .92] and McDonald's omega was .91 [.90, .92].

The sizes of the women and men subsamples allowed us to test the measurement invariance of the Italian BSRI across sex. When tested separately, the one-factor model fit adequately for the women sample but failed to meet adequacy criteria in the men subsample. (Table 2). The resampling analyses revealed that the most replicable substantial CU was the one between items 4 and 8 (94.1% of bootstrap replicates; for more details, see Section 6 of the SMs, Table SM5.1). After the specification of this additional parameter the fit was acceptable.

We then proceeded with multi-group analyses. The results are reported in Table 2. The configural invariance model showed an adequate fit, while the weak invariance model had a significantly worse fit. However, the resampling procedure did not identify any additional parameter in over 80% of replications (Table SM5.2). Considering the negligible differences in model fit indices, we chose not to modify the model. When we tested the strong invariance model, we again found a significant difference in fit. This time, the resampling procedure identified item 7 ('hard to shut off negative thoughts about oneself')'s intercept as non-invariant in 99.4% of replications (Table SM5.3). After freeing the invariance constraint for this item, model fit was adequate (Table 2). The

Table 1. Summary of goodness-of-fit statistics for invariance models across administration methods (paper sample: $n=216$; online sample: $n=222$) across national samples (Italian paper sample: $n=216$; Belgian paper sample 1: $n=155$; Belgian paper sample 2: $n=155$; Italian online sample: $n=222$; US online sample: $n=199$).

Model	χ^2	df	p-value	SCF	CFI	TLI	RMSEA	$\Delta\chi^2$	Δdf	Δp-value	ΔCFI	ΔTLI	ΔRMSEA
Italian paper sample vs Italian online sample													
Configural (no invariance constraints)	69.069	35	.001	1.368	.974	.958	.067 [.047, .086]	–	–	–	–	–	–
Weak invariance (all FLs equal)	80.260	43	<.001	1.275	.971	.963	.063 [.044, .082]	9.037	8	.339	−0.003	.005	−0.004
Italian paper sample vs Belgian paper sample 1													
Configural (no invariance constraints)	79.244	38	<.001	1.470	.960	.941	.076 [.057, .096]			–	–	–	–
Weak invariance (all FLs equal)	91.165	46	<.001	1.393	.956	.947	.073 [.054, .091]	10.239	8	.249	−0.004	.006	−0.003
Italian paper sample vs Belgian paper sample 2													
Configural (no invariance constraints)	72.499	38	.001	1.381	.966	.950	.071 [.050, .092]						
Weak invariance (all FLs equal)	88.809	46	<.001	1.294	.958	.949	.072 [.052, .092]	16.801	8	.032	−0.008	−0.001	.001
Italian online sample vs US online sample													
Configural (no invariance constraints)	65.412	37	.003	1.350	.978	.966	.060 [.039, .081]						
Weak invariance (all FLs equal)	77.275	45	.002	1.277	.975	.968	.058 [.038, .078]	11.043	8	.199	−0.003	.002	−0.002

Note: df: degrees of freedom; SCF: Scaling Correction Factor; CFI: Comparative Fit Index; TLI: Tucker-Lewis Index; RMSEA: root-mean-square error of approximation; $\Delta\chi^2$: Satorra-Bentler scaled chi-square difference; Δdf: degrees of freedom difference; Δp-value: p-value for the Satorra-Bentler scaled chi-square difference test. FL: factor loading.

parameter estimates showed that, once controlled for the level of state rumination, the intercept of item 7 was higher in men (42.53 [Standard error = 1.77]) than in women (34.32 [Standard error =1.37]). This model also allowed us to test sex differences in latent scores, which were not statistically significant ($d=0.14$, $p=.113$).

Finally, we specified a strong invariance model in which the latent scores were regressed onto age values. The results showed that both the regression coefficient of women and the regression coefficient of men were statistically significant and suggested a decrease of scores with age (women: $\beta = -0.31$, $p < .001$; men: $\beta = -0.34$, $p < .001$). The regression coefficients did not differ ($z=0.60$, $p=.552$). The data showed no evidence of non-linearity (see Figure SM6.1 in Section 6 of the SMs).

In conclusion, this study demonstrates that the Italian BSRI measures the same construct consistently across sexes, enabling valid comparisons of latent means.

Study 2b: Exploring construct validity

To explore construct validity, we computed the correlations of the BSRI total score with measures of personality traits, rumination, anxiety, depression, stress, and emotional regulation (Table 3). Across analyses, the pattern of results supported our predictions. We found strong correlations ($r \geq |.50|$) with neuroticism, brooding, depression, stress, non-acceptance of emotional responses, and limited access to emotion regulation strategies. Moderate correlations ($|.30| \leq r < |.50|$) were observed with conscientiousness, reflection, anxiety, reduced emotional clarity, difficulties engaging in goal-directed behavior, and impulse control difficulties.

The correlation with brooding was significantly stronger than the correlation with reflection ($z=6.03$, $p < .001$, $r=0.23$). The pattern of correlation with the Big Five dimensions was not uniform ($X2(4) = 116.99$, $p < .001$, $r=0.41$), and the correlation with neuroticism, in absolute value, was stronger than the correlation with all other traits, with small to moderate effect sizes (extraversion: $z=8.16$, FDR adjusted $p < .001$, $r=0.30$; agreeableness: $z=8.43$, FDR adjusted $p < .001$, $r=0.31$; conscientiousness: $z=4.92$, FDR adjusted $p < .001$, $r=0.19$; openness: $z=9.30$, FDR adjusted $p < .001$, $r=0.34$).

The pattern of correlation of BSRI with measures of anxiety, depression, and stress was not uniform, although with a small effect size ($X2(2) = 20.30$, $p < .001$, $r=0.17$). The correlation with depression was significantly stronger than the ones with anxiety ($z=4.49$, FDR adjusted $p < .001$, $r=0.17$) and stress ($z=2.55$, FDR adjusted $p=.016$, $r=0.10$), while the correlation with anxiety was not statistically different from the one with stress ($z=1.94$, FDR adjusted $p=.053$, $r=0.07$).

The pattern of correlation of BSRI with DERS subscales was not uniform ($X2(5) = 240.07$, $p < .001$, $r=0.59$). The correlation with the LAERS subscale was stronger than the correlations with the other subscales (LEA: $z=14.32$, FDR adjusted $p < .001$, $r=0.48$; LEC: $z=4.76$, FDR adjusted $p < .001$, $r=0.18$; DEGDB: $z=5.64$, FDR adjusted $p < .001$, $r=0.21$; ICD: $z=6.44$, FDR adjusted $p < .001$, $r=0.24$; NER: $z=2.24$, FDR adjusted $p=.025$, $r=0.09$). None of these correlations differed between women and men (see Section 7 of the SMs, Table SM7.1).

Table 2. Summary of goodness-of-fit statistics for confirmatory factor analysis total, women, and men samples and invariance models across sex in Study 2 (women: $n = 463$; men: $n = 212$).

Model	χ^2	df	p-value	SCF	CFI	TLI	RMSEA	$\Delta\chi^2$	Δdf	Δp-value	ΔCFI	ΔTLI	ΔRMSEA
Total sample	88.789	20	<.001	1.217	.969	.957	.075 [.061, .089]						
Women	52.898	20	<.001	1.215	.978	.970	.063 [.044, .082]						
Men	52.231	20	<.001	1.224	.958	.941	.091 [.063, .119]						
Men (with CU_{48})	32.738	19	.026	1.187	.982	.973	.061 [.025, .092]						
Women sample vs men sample													
Configural (no invariance constraints)	85.848	39	<.001	1.201	.980	.971	.063 [.046, .079]						
Weak invariance (all FLs equal)	103.929	47	<.001	1.131	.975	.970	.063 [.048, .078]	18.283	8	.019	−0.005	−0.001	.000
Strong invariance (all FLs and INTs equal)	135.724	54	<.001	1.113	.964	.963	.070 [.056, .084]	33.803	7	<.001	−0.011	−0.007	.007
Strong invariance (all FLs and INTs equal except INT_7)	114.805	53	<.001	1.115	.973	.971	.062 [.047, .076]	10.561	6	.103	−0.002§	.001§	−0.001§

Note: df: degrees of freedom; SCF: Scaling Correction Factor; CFI: Comparative Fit Index; TLI: Tucker-Lewis Index; RMSEA: root-mean-square error of approximation; $\Delta\chi^2$: Satorra-Bentler scaled chi-square difference; Δdf: degrees of freedom difference; Δp-value: p-value for the Satorra-Bentler scaled chi-square difference test. CU: correlated uniqueness; FL: factor loading; INT: intercept.

Table 3. Correlations of BSRI scores with measures of personality traits, rumination, anxiety, depression, stress, and emotional regulation, and descriptive statistics and reliability of scale scores ($n = 675$).

	1	2	3	4	5	6	7	8	9	10	11	12	13	14	15	16	17
1. BSRI																	
2. BFI-Ext	−0.17																
3. BFI-Agr	−0.17	.24															
4. BFI-Con	−0.33	.24	.31														
5. BFI-Neu	.53	−0.26	−0.22	−0.48													
6. BFI-Ope	−0.13	.28	.15	.20	−0.08												
7. RRS-Bro	.53	−0.20	−0.26	−0.36	.51	−0.14											
8. RRS-Ref	.32	−0.14	−0.07	−0.22	.44	.20	.44										
9. DASS-Anx	.44	−0.07	−0.17	−0.38	.49	.01	.44	.40									
10. DASS-Dep	.56	−0.22	−0.27	−0.43	.56	−0.13	.65	.47	.62								
11. DASS-Str	.50	−0.11	−0.29	−0.38	.61	.02	.49	.42	.69	.65							
12. DERS-LEA	.04	−0.12	−0.14	−0.12	.03	−0.38	.10	−0.26	.02	.05	−0.07						
13. DERS-LEC	.43	−0.26	−0.25	−0.32	.31	−0.29	.41	.16	.30	.36	.27	.51					
14. DERS-DEGDB	.40	−0.24	−0.26	−0.52	.54	−0.13	.44	.33	.40	.46	.44	−0.01	.26				
15. DERS-ICD	.36	−0.06	−0.29	−0.45	.48	−0.15	.50	.26	.47	.43	.51	.15	.43	.55			
16. DERS-NER	.51	−0.21	−0.10	−0.37	.47	−0.09	.59	.42	.41	.51	.43	.02	.42	.47	.54		
17. DERS-LAERS	.58	−0.31	−0.25	−0.49	.61	−0.10	.62	.44	.46	.66	.53	−0.03	.40	.66	.60	.66	
M	43.77	26.01	32.89	31.75	25.75	37.54	10.63	9.76	11.25	12.69	14.92	15.32	10.20	14.10	12.26	12.74	17.35
SD	22.25	5.31	4.88	6.31	5.67	6.12	3.01	2.83	3.48	4.02	4.06	4.73	3.52	4.61	4.37	4.75	6.04
alpha	.91	.79	.65	.84	.78	.81	.77	.73	.82	.86	.84	.79	.79	.85	.83	.84	.87
omega	.91	.79	.64	.85	.78	.82	.77	.75	.82	.86	.85	.80	.80	.86	.83	.84	.87

Note: critical correlation values: .08 at $p = .05$; .10 at $p = .01$; .13 at $p = .001$; BSRI: Brief State Rumination Inventory; BFI: Big Five Inventory; Ext: Extraversion; Agr: Agreeableness; Con: Conscientiousness; Neu: Neuroticism; Ope: Openness; RRS: Ruminative Responses Scale; Bro: Brooding; Ref: Reflection; DASS: Depression Anxiety Stress Scales; Anx: Anxiety; Dep: Depression; Str: Stress; DERS: Difficulties in Emotion Regulation Scale; LEA: Lack of Emotional Awareness; LEC: Lack of Emotional Clarity; DEGDB: Difficulties Engaging in Goal-Directed Behavior; ICD: Impulse Control Difficulties; NER: Nonacceptance of Emotional Responses; LAERS: Limited Access to Emotion Regulation Strategies; M: mean; SD: standard deviation; alpha: Cronbach's alpha; omega: McDonald's omega.

Discussion

In this article, we report two studies examining the psychometric properties of the Italian BSRI: (1) its unidimensionality, reliability, and measurement invariance across administration methods (online vs. paper-and-pencil) and national samples; and (2) its measurement invariance across sex and construct validity.

Consistent with prior validation studies of the BSRI (Altan-Atalay et al., 2020; Marchetti et al., 2018; Wang et al., 2022), results supported the unidimensionality of the Italian version of the scale. Dimensionality and confirmatory factor analyses (CFA) in both studies supported an essentially unidimensional structure, with high internal consistency and satisfactory fit after accounting for minor residual covariances. These results align with the theoretical view of state rumination as a coherent cognitive-affective construct and replicate prior evidence of the BSRI's structural robustness across languages and cultures.

Measurement invariance testing provided important insights into the generalizability and replicability of the BSRI measurement model. The scale demonstrated configural and weak (metric) invariance across administration methods (paper-and-pencil vs. online) and national samples (Italy vs. Belgium/USA), indicating that the underlying construct was represented similarly and that factor loadings were comparable across groups. Across sex, evidence supported partial strong invariance, permitting meaningful comparisons of latent means after accounting for the non-invariant intercept of Item 7.

Study 2 contributed novel evidence regarding sex differences in state rumination. While prior research has consistently documented small-to-moderate sex differences in trait rumination (e.g., Johnson & Whisman, 2013; Rood et al., 2009), our findings suggest that such differences do not extend robustly to state rumination. After establishing partial strong invariance across sex, latent mean differences in BSRI scores were not statistically significant, consistent with Marchetti et al. (2018) initial observation of small-to-negligible sex effects in state rumination. Although further research is needed to clarify this finding, we hypothesize that trait rumination may be shaped over time by societal expectations, similar to other personality traits (e.g., neuroticism; Feingold, 1994; see also Eagly et al., 2000). In contrast, state rumination, being situational, is likely less influenced by long-term socialization and instead reflects immediate reactions that are more consistent across women and men. It has also been shown that women may be more likely to use emotion-focused coping, which includes rumination, while men may lean toward problem-focused coping (see, e.g., Ptacek et al., 1992, 1994). These preferences develop over time, affecting trait rumination, but in acute situations (as is the case for state rumination), both sexes may ruminate similarly depending on the stressor. In any case, this dissociation underscores the importance of distinguishing between trait and state conceptualizations of rumination in both research and clinical settings.

Although the differential item functioning (DIF) observed for item 7 in the partial strong invariance model may reflect sample-specific characteristics, its content ('hard to shut off negative thoughts about oneself') and higher scores among men warrant attention. While women generally report greater rumination, partly due to gendered norms surrounding emotional expression (Nolen-Hoeksema & Jackson, 2001), men often face cultural pressures to maintain emotional stoicism and self-sufficiency (Mahalik et al., 2003). Engaging in rumination, a behavior socially coded as feminine, may therefore elicit heightened self-criticism and fear of judgment in men (Borinca et al., 2021; Şirin et al., 2004). Masculine norms that discourage emotional disclosure and favor aggression as a coping strategy can exacerbate depressive symptoms when rumination occurs (Panayiotou & Papageorgiou, 2007; Simonson et al., 2011). These sociocultural dynamics shape emotional processing, particularly in contexts of humiliation and self-esteem, with evidence suggesting men respond more adversely to such experiences, intensifying self-referential rumination (Coleman et al., 2009). Thus, while overall rumination patterns may appear similar across sexes, cultural constructions of masculinity likely amplify negative self-assessments among men (Chen-Xia et al., 2022; Bosak et al., 2018).

The construct validity of the Italian BSRI was supported by its pattern of associations with other constructs. As hypothesized, BSRI scores showed stronger associations with brooding than with reflection, and stronger correlations with depression than with anxiety or stress, replicating previous findings (Altan-Atalay et al., 2020; Wang et al., 2022). Moreover, BSRI scores exhibited the strongest association with neuroticism among the Big Five personality traits, aligning with prior research on trait rumination (Teasdale & Green, 2004; Vannucci & Chiorri, 2018). These findings suggest that, although state rumination is situational, its underlying cognitive and emotional mechanisms (e.g., attention to negative thoughts, emotional reactivity) are influenced by stable personality traits. High neuroticism predisposes individuals to negative emotions and both trait and state rumination, while low extraversion and low conscientiousness are associated with less adaptive coping and greater rumination (Connor-Smith & Flachsbart, 2007). As a result, when someone experiences a stressful event, their personality traits shape how they respond, even in the moment.

Interestingly, the BSRI demonstrated a distinct pattern of associations with emotion regulation strategies, showing the strongest correlation with limited access to emotion regulation strategies (LAERS). This result is consistent with earlier findings by Ando' et al. (2020) and highlights the interplay between momentary rumination and perceived regulatory efficacy. Note, however, that, differently from Ando' et al. (2020), we did not find any substantial difference between women and men in the correlations of BSRI with any DERS subscale.

As anticipated in the Introduction, a central implication of our findings is that the construct-validity evidence obtained here should be understood as indexing stable between-person differences in momentary rumination, rather than verifying the BSRI's capacity to track within-person fluctuations over time. Because participants completed the BSRI only once, observed correlations with trait rumination, personality traits, and emotion regulation strategies reflect how individuals differ from one another in their momentary cognitive-affective state under an uncontrolled context, not how their rumination rises and falls in response to situational triggers. While this pattern is consistent with prior validations and provides valuable information about the BSRI's nomological network, it highlights an important direction for future research. Designs incorporating repeated measurements (such as ecological momentary assessment or short-term longitudinal protocols) would allow for the estimation of within-person reliability (e.g., generalizability coefficients, multilevel omega) and would more directly test the BSRI's theorized sensitivity to situationally driven changes in rumination.

Age-related analyses revealed a negative association between state rumination and age, replicating recent longitudinal findings suggesting that rumination tends to peak in young adulthood and decrease thereafter (Lilly et al., 2023). The age-related decline was evident across both men and women, further confirming the generalizability of this trajectory.

Despite the study's strengths - including large and diverse samples, the use of robust analytic techniques, and cross-cultural comparisons - several limitations should be acknowledged. First, reliance on convenience sampling limits the generalizability of findings. Participants were predominantly young and more-than-average educated. While they can be considered representative of the sampling frame, they do not reflect the demographics of the general population. Furthermore, the lack of random assignment to the paper-and-pencil and online samples for Italian participants and the lack of balance on background characteristics of the national samples prevented us to validly test the invariance of item intercepts and latent means. Second, although the translation process included independent forward translations, expert review, an automated back-translation, and cognitive debriefing, it did not fully adhere to more comprehensive cross-cultural adaptation procedures. Specifically, the back-translation was generated using an online translation service rather than an independent bilingual translator, and the process did not involve a formal multidisciplinary expert committee to reconcile all translation stages. Accordingly, although the evidence collected supported the comprehensibility and conceptual adequacy of the Italian items, future adaptations could strengthen methodological rigor by incorporating independent human back-translators and a structured committee-review process. Third, the cross-sectional design precluded inferences about the temporal dynamics of state rumination, which may benefit from ecological momentary assessment approaches. Fourth, unlike previous studies (Altan-Atalay et al., 2020; Marchetti et al., 2018; Wang et al., 2022), the present one did not assess the instrument's sensitivity to rumination induction, which would have provided additional evidence of its criterion validity. Lastly, future research should examine the BSRI's diagnostic sensitivity and specificity in clinical samples: for instance, we expect that individuals with depression will score higher on the BSRI than individuals without a diagnosis.

In conclusion, the present findings support the Italian BSRI as a psychometrically sound tool for assessing state rumination. The scale demonstrated configural and weak (metric) invariance across administration methods and national contexts, and partial strong invariance across sex, supporting comparable measurement of the construct across these groups while allowing latent mean comparisons only for sex. These results not only advance the cross-cultural applicability of the BSRI but also enrich our understanding of state rumination as a context-sensitive process with distinct correlates and developmental patterns. Future studies may build upon this work by exploring the temporal dynamics of state rumination and its role in intervention outcomes in both clinical and non-clinical populations.

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Disclosure statement

No potential conflict of interest was reported by the author(s).

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