Studia Psychologica, Vol. 67, No. 3, 2025, 264-281
https://doi.org/10.31577/sp.2025.03.923

Psychometric Evaluation of the Slovak Version of the Self-Reflection and Insight Scale (SRIS)

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The aim of the present study is to evaluate the psychometric properties of the Slovak version of the Self-Reflection and Insight Scale (SRIS), including both the original 20-item version and the recently introduced 12-item short version. Two samples, totaling 440 participants, were used to examine the factorial structure, convergent validity, and test-retest reliability of the scale in the first phase. In the second phase, the factorial structure of the short version was replicated on 1132 participants. The results indicated that the long version of the scale did not meet the required criteria. However, our findings provide initial support for the short version of the scale. Despite its limitations, this study serves as an important first step in the validation process for the Short Self-Reflection and Insight Scale and potentially provides a basis for future research using this version of the scale in the Slovak language.

Key words: Self-Reflection and Insight Scale (SRIS), psychometric properties, self-awareness

Introduction

A considerable amount of recent research literature has focused on constructs related to mechanisms of focusing attention on and being aware of oneself, i.e., self-awareness. However, to the best of our knowledge, a Slovak version of the scale for measuring self-awareness is not available. This study aims to fill this gap by adapting the Self-Reflection and Insight Scale (SRIS) to the Slovak language and providing initial evidence of various types of validity for both the original full and short versions of the scale. In the following text, we will describe self-awareness as a construct and the Self-Reflection and Insight Scale (SRIS) as a measurement tool.

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Online Supplements are available at https://journals.savba.sk/index.php/studiapsychologica/article/view/3282/version/3228
Data, materials and analytic code are available on the project's Open Science Framework page: https://osf.io/mvtcg/

Received August 21, 2024

Self-Awareness

Self-awareness is generally referred to as one's tendency to be attentive to or to cognitively explore one's internal and external responses to a certain situation, namely one's inner thoughts, feelings, and behaviors (Hamaideh et al., 2024; Kreibich et al., 2020). It can be understood as a targeted exploration of oneself that can provide important insight into functioning and shape one's perception of oneself. As such, self-awareness can play an important role in self-regulating strategies used by an individual, which are crucial for personal development, but also any type of motivated action, including achieving goals, facing obstacles, or experiencing an action crisis (Kreibich et al., 2020).

The construct of self-awareness has been studied extensively; however, there are some differences in the definition of the construct based on the researchers' approach. The initial definitions described self-awareness simply as a state of heightened attention to some aspects of the self (Duval & Wicklund, 1972). However, interindividual differences in practicing self-awareness were emphasized very soon, and various other aspects were also considered. This led to the differentiation between private and public self-awareness based on whether the attention was on one's internal states and processes or the aspects of the visible self, such as gestures and overall looks (Feningstein et al., 1975). Some authors also emphasize the metacognitive perspective of having an objective understanding of some internal events, especially thoughts, which enables a person to monitor and control their internal states and processes, improving emotional regulation and reducing mental distress (Jankowski & Holas, 2014). Moreover, authors like Vago (2014) stressed the reflective properties of the awareness that enable individuals to observe their inner states and processes from an objective, detached point of view of the self.

Based on these developments, we understand self-awareness as a meta-cognitive ability to focus on – i.e., monitor and reflect upon – aspects of one's self in terms of internal states and processes that enable one to self-regulate, pursue goals, etc. In fact, monitoring one's actions, progress, and outcomes as an integral part of self-awareness distinguishes this construct from related constructs, such as mindfulness. The main difference is that mindfulness is a form of awareness that does not evaluate or judge and accepts everything being observed as it is (Evans et al., 2009). We will focus on the measurement of self-awareness in the next section.

Measurement of Self-Awareness and the Self-Reflection and Insight Scale

As alluded to in the previous section, self-awareness can be approached as a trait (i.e., as a relatively consistent attentive pattern) or a state (also often referred to as self-consciousness) (Kreibich et al., 2020). Grant and his colleagues (2002) focused on the first-mentioned approach to self-awareness. Rooted in the principles of coaching, personal development frameworks, and metacognitive theories, and informed by earlier studies on private self-consciousness, they focused on the distinction between two main metacognitive factors of self-awareness: self-reflection and insight. While *self-reflection* refers to inspecting and evaluating one's internal processes, *insight* encompasses the clarity of understanding these processes (Grant et al., 2002).

Grant and his colleagues (2002) proposed a generic model of self-regulation and goal attainment in which both aspects of self-awareness – self-reflection and insight – play an important role. More specifically, after setting a

goal and developing an action plan, the action is carried out while being monitored (requiring self-reflection) and evaluated (associated with insight) to determine what steps must be taken next. This model expands one of the main theories proposed by Carver and Scheier (1998), in which self-focused attention produces effects on behavior that seem to reflect the engagement of regulating feedback processes. The importance of these processes led Grant and his colleagues (2002) to attempt to develop a measurement – The Self-Reflection and Insight Scale (SRIS).

An important aspect of understanding the broader context of the scale involves pondering how it was developed. The work of Grant et al. (2002) was based on the strengths of the Private Self-Consciousness Scale (PrSCS) created by Fenigstein and colleagues in the mid-seventies. While public self-consciousness reflects attending to how the self appears to others, private self-consciousness focuses on attention directed to one's self. The PrSCS was originally comprised of two sub-scales, which led to some psychometric issues. Also, revision of items, especially those capturing self-reflection, was necessary. Grant and his colleagues (2002) conducted three studies to develop a new measurement tool. The final version of their scale consisted of 20 items, each accompanied by a six-point agreement scale. The factorial analyses revealed two factors labelled "self-reflection" (SRIS-SR) and "insight" (SRIS-IN). While the insight scale consists of 8 items, the self-reflection scale has more items as it can be further differentiated into two more specific facets – "need for self-reflection" (6 items) and "engagement in self-reflection" (6 items). Later on, however, the literature regarded the scale mainly as a two-factor measure, mostly due to inconsistencies in the factorial structure of the scale (e.g., Havrdová et al., 2020; Silvia, 2021).

According to Grant et al. (2002), the scales for both factors performed better than previous measurements. They had good internal consistency (i.e., coefficient alpha ranging from .71 to .91 for the SRIS-SR and from .82 to .87 for the SRIS-IN), as well as test-retest correlation over seven weeks (correlation coefficient .77 for the SRIS-SR and .78 for the SRIS-IN). The two factors were not correlated in the first study. However, they were negatively correlated in their third study, which is in accordance with the ambiguity of this relationship reported by some previous authors (Grant et al., 2002). Other authors (e.g., Havrdová et al., 2020; Silvia, 2021) did not find a significant relationship between the two factors. One explanation for this discrepancy may lie in the extent of the conscious engagement in self-reflection or other internal processes (Grant et al., 2002), which deserves more research attention. Further explanations consider various biases, such as acquiescence bias, as outlined below.

Several studies also explored the new scale's convergent validity. For example, SRIS-IN was negatively correlated with various well-being-related variables, such as depression, anxiety, and stress, and positively correlated with cognitive flexibility and self-regulation (Grant et al., 2002). Subsequent research studies have found additional relationships between these factors and other constructs, such as mindfulness or personality traits (e.g., Havrdová et al., 2020; Silvia et al., 2023). The observed differences in relationships with other constructs underscored the necessity to distinguish between self-reflection and insight. While self-awareness is often expected to be connected to self-regulation and self-evaluation (Nie et al., 2014), this is not always supported when considering the specific aspects of self-awareness.

On one hand, insight was repeatedly found to positively correlate with well-being, life satisfaction (e.g., Harrington et al., 2016), or resilience (e.g., Cowden & Meyer-Weitz, 2016),

and to negatively correlate with rumination (Bucknell et al., 2022), depression and anxiety (Grant et al., 2009). On the other hand, self-reflection was found to be related to rumination (e.g., Lyke, 2009) and often positively associated with indicators of both good mental health and ill health (Bucknell et al., 2022). Rumination, referring to repetitive thoughts of negative aspects of one's life, seems to be a part or a result of self-reflective activities (Bucknell et al., 2022); thus, it is also related.

Over the following years, the SRIS scale became highly popular and effective in assessing individual differences in self-awareness. Despite the many strengths of the scale, multiple researchers have noted inconsistencies, especially regarding factor analytical findings (e.g., Havrdová et al., 2020; Silvia, 2021). This led Silvia (2021) to examine the psychometric properties of the SRIS scale and identify poorly performing items that were eliminated. His efforts resulted in a more concise 12-item scale (6 items per each factor). The internal consistency of the short SRIS scale was found to be acceptable (Cronbach's alpha was .87 for SRIS-SR and .83 for SRIS-IN). Although Silvia's work might have some limitations, such as the use of young college-aged participants, the brief form of the popular SRIS scale might be much appreciated in applied research. In fact, a short version of the scale was not only created to provide superior psychometric properties by preserving items with high discrimination levels (Silvia, 2021), but Silvia et al. (2023) also showed that it exhibited a very similar pattern of relations with external criteria and profiles as did the original version. Thus, concern about the limited construct coverage is not supported.

The Present Study

The SRIS scale is, to this day, a widely used tool, adapted in other languages (e.g., Aşkun & Cetin, 2017; Chen et al., 2016; Havrdová et al., 2020). However, to our knowledge, the scale has not been adapted to the Slovak language. Likewise, despite creating and adapting short versions with psychometric and practical benefits into various languages (Banner et al., 2023; Silvia, 2021; Silvia et al., 2023), the Slovak language has not been addressed. Considering the need to look more closely into self-awareness's role in self-regulating goal attainment, we find it important to adapt the proper measurement to help researchers achieve that goal in the Slovak language as well. This paper, therefore, focuses on the initial phases of adapting the SRIS scale to the Slovak language and provides an initial examination of the psychometric properties of (A) the original 20-item version (Grant et al., 2002) and (B) more recent short 12-item version (Silvia, 2021) of the scale. In the next section, we will focus on Phase 1 of the adaptation process.

Phase 1 (Studies A and B)

In the first phase of the validation process, we conducted a backward translation of the SRIS and examined its factorial validity, internal consistency, convergent validity, and test-retest reliability of the scale across two studies.

First, we investigated the factor structure and associated psychometric properties of the SRIS in its full and abbreviated forms. We investigated the default two-factor solution in the main text but looked at a three-factor solution in the online appendix. Due to convergence with the research literature and its hypothesized reflective nature, we used a Covariance-based Confirmatory Factor Analysis. We also analyzed internal consistency and acquiescence bias.

Next, we examined the convergent validity. We selected rumination, resilience, and life satisfaction based on previous research.

Table 1 *Overview of pattern of relationships based on previous research*

Construct	Expected direction of relationships	Effect size (r) derived from the literature	Reference for expected direction and effect size
Rumination	Positive with self-reflection	.27	Harrington & Loffredo (2010)
	Negative with insight	-.41	Bucknell et al. (2022)
Resilience	Positive with self-reflection	.40	Bucknell et al. (2022)
	Positive with insight	.48	Cowden & Meyer-Weitz (2016)
Life satisfaction	Positive with insight	.44	Harrington et al. (2016)

Table 1 summarizes the expected pattern of relationships in terms of direction and effect size. The literature review indicates a positive medium-to-large relationship between self-reflection and rumination and large-to-very-large relationships between self-reflection and resilience, a negative relationship between insight and rumination, and a positive relationship between insight and resilience and life satisfaction. We also examined test-retest reliability.

Methods

Participants and Procedure

The sample consisted of 440 participants in total across two studies: 230 people with a mean age of 27.85 years (*SD* = 9.33; Median = 24; Mode = 23), of which 64% were women in the first study, and 210 people with a mean age of 29.99 years (*SD* = 11.77; Median = 24; Mode = 24), of which 74% were women in the second study.

We conducted two separate studies to examine the robustness of the results, as this approach allows the examination of how the scales behave in different contexts and this procedure also decreases the burden for participants (as scales used for convergent validity were distributed across two studies, decreasing the time necessary to fill in the questionnaires).

This study aimed to provide pilot results given resource constraints (e.g., time and other resources). The research was conducted via convenience sampling, using an anonymous online questionnaire via social media. Two phases, each lasting a week, were carried out with a one-month gap between them to assess test-retest reliability. The university's Institutional Ethics Board approved the study.

Instruments

Examined scales

The Self-Reflection and Insight Scale (SRIS) had two versions: A, the original full version, and B, the short version.

The *original full form of SRIS* (Grant et al., 2002) consists of 20 items accompanied by a 6-point agreement scale. The first factor, "insight" (SRIS-IN), consists of 8 items, e.g., "I am usually aware of my thoughts". The second factor, "self-reflection" (SRIS-SR), is measured by 12 items. This factor can be further divided by a "need for self-reflection" scale (6 items,

e.g., "It is important for me to evaluate the things I do") and an "engagement in self-reflection" scale (6 items, e.g., "I frequently examine my feelings").

The short version of SRIS (Silvia, 2021) consists of 12 items – 6 items for each factor. This brief version did not make any changes to the items included. Silvia (2021) suggested to work with 12 items and 8 problematic items from the original version were eliminated. As a result, SRIS-SR does not differentiate between the need and the engagement in self-reflection, although these components are equally represented by 3 items. Silvia (2021) did not use a 6-point agreement scale. However, to compare the psychometric properties of both versions of the SRI's scale, we have used the original 6-point response scale as other authors underwent similar goals (e.g., Banner et al., 2023).

A back-translation (English – Slovak) was conducted by four independent translators (two persons translated items into the Slovak language, two back to English). Translations were then compared, and a reconciliation procedure was conducted to address any problematic items, while taking into account cultural nuances for a better understanding of the items.

Other scales

The first part of the Rumination-Reflection Questionnaire (RRQ; Trapnell & Campbell, 1999) was used to measure rumination. Rumination was examined by 12 statements (e.g., "I tend to 'ruminate' or dwell over things that happen to me for a really long time afterwards") accompanied by a 5-point agreement scale. The test of internal consistency showed good reliability (ωt = .93).

The Brief Resilience Scale (BRS; Smith et al., 2008) contains 6 items examining resilience (e.g., "It does not take me long to recover from a stressful event") accompanied by 5-point agreement scale. The test of internal consistency showed acceptable reliability (ωt = .88).

The Satisfaction with Life Scale (SWLS; Diener et al., 1985) measures life satisfaction through 5 items (e.g., "In most ways my life is close to my ideal") accompanied by a 7-point agreement scale. The test of internal consistency showed acceptable reliability (ωt = .89).

Analysis

Confirmatory factor analysis was used to examine the evidence for factorial validity. The MLR estimator (i.e., Maximum Likelihood with Robust Standard Errors) was used, and accordingly, scaled and robust statistics are reported. As the chi-square test can highlight even minor discrepancies, model fit indices were also considered. The model was evaluated according to the following benchmarks of model-data correspondence: RMSEA optimally < .05 or at most .08; CFI > .90 or preferably > .95; SRMR < .08 (Gana & Broc, 2019). Please note that fit indices could be understood as the effect size quantifying the degree of mismatch from various viewpoints rather than a binary decision.

Results

Evidence for the Factorial Validity of the Full Scale

In the first step, we conducted a CFA for the original scale (20 items; Grant et al., 2002) with a default two-factor solution.

The CFA model did not provide an adequate fit, as indicated by the significant chi-square test ($\chi^2(169) = 546.99$, $p < .001$). Additionally, continuous approximate fit indices suggested poor model-data correspondence, with the Robust Comparative Fit Index (CFI) at 0.864,

the robust Root Mean Square Error of Approximation (RMSEA) at 0.080 (90% CI [0.073, 0.088]), and the Standardized Root Mean Square Residual (SRMR) at 0.096, all outside recommended thresholds for acceptable fit.

Next, we looked at the factor loadings in two samples. All items were loaded into their respective factors; however, some had borderline and suboptimal factor loadings.

For the self-reflection (SR) factor, item 10 (i.e., "I frequently examine my feelings") demonstrated the highest factor loadings in both groups (-.83). Items 16 ("I frequently take time to reflect on my thoughts"; -.75) and 19 ("I often think about the way I feel about things"; -.75) also showed strong loadings. Item 15 ("I have a definite need to understand the way that my mind works"; -.73) and item 18 ("It is important to me to try to understand how my thoughts arise"; -.71) provided lower, but still adequate factor loadings. Similarly, item 7 ("I am very interested in examining what I think about"; -.69) and item 12 ("It is important to me to try to understand what my feelings mean"; -.64) showed moderate loadings. Items 2 ("I am not really interested in analyzing my behavior"; .49), 13 ("I don't really think about why I behave in the way that I do"; .50), 8 ("I rarely spend time in self-reflection"; .45) and 1 ("I don't often think about my thoughts"; .44) showed less optimal and borderline loadings. However, item 5 ("It is important for me to evaluate the things that I do"; -.33) provided suboptimal factor loading.

For the insight (I) factor, item 17 ("Often I find it difficult to make sense of the way I feel about things") showed the highest factor loadings (-.81). Items 11 ("My behavior often puzzles me"; -.77) and 4 ("I'm often confused about the way that I really feel about things"; -.76) also demonstrated strong loadings. Item 9 ("I'm often aware that I'm having a feeling, but I often don't quite know what it is"; -.70) and item 14 ("Thinking about my thoughts makes me more confused"; -.65) provided adequate factor loadings. Item 20 ("I usually know why I feel the way I do"; .62) and item 6 ("I usually have a very clear idea about why I've behaved in a certain way"; .47) showed less optimal but still acceptable loadings. Item 3 ("I'm usually aware of my thoughts"; .33) provided suboptimal factor loading.

The covariance between the SR and I factor was low and insignificant (i.e., .08, p = .175). Please note that the relationship was positive in this specific CFA; however, it is important to remember that the sign of the factor is determined by the order and wording of items (e.g., the first item in the SR factor should be reversed). Thus, the sign of correlation between factors is negative in reality.

Next, we examined the local fit using standardized residuals consistency. There were several problems in the model. For example, item 20 had large residuals with multiple items, particularly items 16 and 6. We also examined the modification indices. These suggested that items 1 and 2, items 3 and 6, and especially items 6 and 20, be correlated in terms of error terms. Considering cross-loadings, items 3 ("I'm usually aware of my thoughts") and 20 ("I usually know why I feel the way I do") might also load to the self-reflection factor. We did not implement these suggestions as they did not align with the original scale, and our goal was not to modify it.

The internal consistency, assessed via McDonald's omega, was sub-optimal – below the recommended range of .70 (i.e., in a .63 range). The average variance extracted was also below the recommended range of 50% (in the .40 to .47 range), indicating a high proportion of measurement error compared to the variance explained by the construct, thus showing very limited internal consistency and convergent validity, respectively. As such, it

can be concluded that we did not find initial support for factor validity, internal consistency and convergent validity of the examined Slovak version of the original 20-item scale (Grant et al., 2002).

Evidence for Factorial Validity of the Short Version of the Scale

In contrast to the original full version of the scale (Grant et al., 2002), the CFA with a short scale and default two-factor solution (Silvia, 2021) provided a reasonable model-data correspondence with relatively adequate fit according to the approximate fit indices when less stringent criteria were applied. Also, the local fit was better than in the original version.

More specifically, although the chi-square test of model fit was statistically significant ($\chi^2(53) = 127.83$, $p < .001$), the approximate fit indices were in an acceptable range for the default two-factor solution of the short version of the scale, indicating reasonable model-data correspondence. The Robust Comparative Fit Index (CFI) at 0.96 was good, the Robust Root Mean Square Error of Approximation (RMSEA) at 0.063 (90% CI [0.049, 0.077]) was good, and the Standardized Root Mean Square Residual (SRMR) at 0.072 was acceptable according to the less stringent criteria.

We have also tested invariance by examining configural, metric, and scalar invariance across two groups to ensure that the factor structure is equivalent. The metric invariance model did not significantly worsen the model fit compared to the configural model ($\Delta\chi^2(10) = 8.32$, $p = .597$), supporting metric invariance. The scalar invariance model also did not significantly worsen the model fit compared to the metric model ($\Delta\chi^2(10) = 12.91$, $p = .229$), indicating that the scalar invariance holds in two samples. This suggests that the factor structure, loadings, and intercepts are roughly equivalent across the two sampling groups.

Next, we examined factor loadings across two samples. All items were loaded into their respective factors, with varying factor loadings.

More specifically, for *the self-reflection (SR) factor*, item 10 ("I frequently examine my feelings") demonstrated the highest factor loadings (.87). Items 19 ("I often think about the way I feel about things"; .78) and 16 ("I frequently take time to reflect on my thoughts"; .74) also showed strong loadings. Item 7 ("I am very interested in examining what I think about"; .67) and item 12 ("It is important to me to try to understand what my feelings mean."; .61) provided adequate factor loadings. However, item 5 ("It is important to me to evaluate the things that I do") showed suboptimal loading (.30).

For the insight (I) factor, item 17 ("Often I find it difficult to make sense of the way I feel about things") showed the highest factor loadings (.83). Items 4 ("I'm often confused about the way that I really feel about things"; .77) and 11 ("My behavior often puzzles me"; .77) also demonstrated strong loadings. Item 9 ("I'm often aware that I'm having a feeling, but I often don't quite know what it is"; .71) and item 14 ("Thinking about my thoughts makes me more confused"; .66) provided adequate factor loadings. Item 20 ("I usually know why I feel the way I do") showed negative loading (-.58), which is consistent with the reverse-scored nature of this item. A more nuanced model with differentiated groups is depicted in Online Supplement 2.

The covariance between the SR and I factors was significant and small to modest (.17, $p = .003$). It was positive in this specific CFA, but it is important to consider the wording and order of the items (e.g., the majority of the items in the I factor, including the first item, should be reversed). Thus, the sign of the correlation is negative in reality.

Next, we examined local fit. Examination of standardized residuals indicated fewer and

less severe problems compared to the full version of the scale. However, some problems were identified, mainly with item 20, showing large residuals with other items, namely 16 and 5. Modification indices had only several suggestions, mainly related to item 20 being part of the self-reflection factor. Other suggestions were group-specific. As in the previous case, we did not implement these suggestions as this was not our goal.

The internal consistency, as assessed via McDonald's omega (total), was good (i.e., ~ωt =.84 for self-reflection and ωt =.79 for insight subscales). Also, the average variance explained was borderline/above the recommended 50%, but still borderline (i.e., .49 and .53; .50 and .55 in one group and .47 to .51 in another group for self-reflection and insight, respectively).

Sensitivity Analysis Dedicated to the Acquiescence Bias

We did not find initial support for the factorial validity of the original long version of the scale (Grant et al., 2002), but we found some initial support for the factorial validity of the short version (Silvia, 2021). However, as the reviewer argued, one reason could be related to acquiescence bias. Therefore, we conducted a sensitivity analysis focusing on this issue.

We examined the acquiescence method factor via the model controlling acquiescence bias (see, e.g., Havan et al., 2025). The model fit indices of the full-scale model improved ($\chi^2(336) = 588.51$, $p < .001$; Robust CFI = 0.92, robust RMSEA = 0.064 (90% CI [0.055, 0.072]), robust SRMR = 0.079). There were consistent but relatively small loadings to the acquiescence factor and associated variance, indicating not severe but consistent presence of the acquiescence bias (loadings around .23 to .29). Thus, some misfits could be attributed to the acquiescence method factor. However, global and local fit are still not optimal, and several items emerged as problematic (e.g., item 5 is still problematic due to lower factor loading, and items 8 and 3 are borderline).

The model also provided superior properties when applying the same analysis to the short version of the scale. The model fit indices improved ($\chi^2(104) = 157.94$, $p < .001$; Robust CFI = 0.97, robust RMSEA = 0.053 (90% CI [0.035, 0.069]), robust SRMR = 0.062). Loadings for the acquiescence factor are similar in a range of .28 to .30). Although item 5 still has a low factor loading (.20), other factor loadings are not problematic. Thus, although not without problems, we prefer the shortened scale version due to fewer problems and practicality, as discussed below. All models are summarized in Online Supplement 1.

Evidence for Convergent Validity

In the second step of the first phase of the validation process, we analyzed the relationships between the SRIS and theoretically related constructs. Here, we report only a short version of the scale, as the original version did not exhibit acceptable psychometric properties (but it is worth noting that differences between correlations to selected variables between the short and long-scale versions were inconsequential (in a range of (r_{diff}=) .01 to .03 as can be seen in Online Supplements-containing the extensive correlation matrix, means, and SD for all variants of the scale and factors for the interested reader). We present latent correlations between latent self-reflection, insight, and related variables to provide a more realistic estimation below. Model with resilience and life satisfaction demonstrated acceptable global fit: robust CFI = 0.95, robust RMSEA = 0.047 (90% CI [0.034, 0.058]) and SRMR = 0.069, although the scaled chi-square test was significant ($\chi^2(224) = 315.38$, $p < .001$). A model with rumination exhibited less

optimal fit: robust CFI = 0.87, robust RMSEA = 0.076 (90% CI [0.067, 0.085]), and SRMR = 0.076 with the chi-square statistic being significant ($\chi^2(249) = 549.50$, $p < .001$).

The analysis of correlation between latent factor indicated that insight (with recoded items) was negatively associated with rumination ($r = -0.36$, $p < .001$, 95% CI [-0.49, -0.22]) and positively linked to resilience ($r = 0.59$, $p < .001$, 95% CI [0.46, 0.72]) and life satisfaction ($r = 0.42$, $p < .001$, 95% CI [0.26, 0.58]). Self-reflection exhibited a positive link to rumination ($r = 0.46$, $p < .001$, 95% CI [0.32, 0.61]) and life satisfaction ($r = 0.18$, $p = .023$, 95% CI [0.03, 0.34]), but shown non-significant negative association with resilience ($r = -0.13$, $p = .103$, 95% CI [-0.29, 0.03]) (i.e., we were not able to reject the null-hypothesis, however, as the magnitude of the effect was very small, it could be due to limited statistical power). The relationship between self-reflection and (recoded) insight varied by model. It was negative in one study ($r = -0.23$, $p = .002$, 95% CI [-0.38, -0.08]) but non-significant in another study ($r = -0.13$, $p = .161$, 95% CI [-0.31, 0.05]), which could also be a matter of statistical power. The sensitivity analysis results with simple Pearson correlations (between mean scores) provided similar results with slightly lower effect sizes[1].

Test-Retest Reliability

In the last step of the first phase, we examined the test-retest reliability of the scale's short version (Silvia, 2021) after 1 month. As in the previous case, the correlation matrix with different scale versions can be found in the online appendix. Latent correlation between two-time time points was $r = .86$, $p < .001$, 95% CI [.76, .95] for insight, $r = .82$, $p < .001$, 95% CI [.72, .93] for self-reflection, respectively, suggesting strong temporal stability even when measurement error is accounted for. However, the model's fit was not optimal and showed some mismatch: robust CFI = 0.84, robust RMSEA = 0.085 (90% CI [0.070, 0.099]), and SRMR = 0.089. These values fell below established thresholds. The scaled chi-square test further indicated significant misfit, $\chi^2(246) = 442.53$, $p < .001$. The simple correlation coefficient conducted as a sensitivity analysis provided similar results, but with a smaller effect size[2].

Phase 2 (Study 2)

In the first phase, we found initial support for the short version of the SRIS scale (Silvia, 2021) across two pilot studies. Although study one employed two samples, both were severely limited in diversity and size (post hoc-power analysis can be found in the online appendix). Also, although the results of the first phase indicated that a shortened version of the scale provided acceptable psychometric properties, some issues were observed, such as lower factor loading for some items. This is worth further examination.

In the second phase of the validation process, we aimed to replicate and further examine the factorial validity of the shortened scale version in a bigger and more diverse sample of participants. Also, we shifted the rating scale from a 6-point scale to a more standard 5-point Likert scale. The main reason was 1) negative feedback about this response

[1] The simple Pearson correlations indicated that insight correlated negatively with rumination ($r = -.33$, $p < .001$, 95% CI [-.44, -.21]) and was positively related to resilience ($r = .51$, $p < .001$, 95% CI [.40, .60]) and life satisfaction ($r = .40$, $p < .001$, 95% CI [.28, .51]). Self-reflection was not statistically significantly related to resilience ($r = .08$, $p = 0.263$, 95% CI [-.21, .06]), but it was positively related to rumination ($r = .39$, $p < .001$, 95% CI [.27, .49]) and life satisfaction ($r = .19$, $p = .005$, 95% CI [.06, .32].

[2] The simple Pearson correlation coefficient between the scores obtained from the same participants at two-time time points was $r = .74$, $p < .001$, 95% CI [.63, .81] for insight and $r = .71$, $p < .01$, 95% CI [.60, .80] for self-reflection, respectively.

scale in the pilot phase, 2) observation that difficult response categories can lead to difficulties for respondents in discriminating between options, potentially introducing noise and 3) the fact that a short version was not created using a 6-point scale in the first place.

Participants and Procedure

The sample consisted of 1132 participants with a mean age of 52.92 years (*SD* = 14.70; Median = 53; Mode = 68; minimum = 18, maximum = 88 years), of which 641 (i.e., 62% were women). The sample was collected by a research agency for a larger research project dedicated to goal pursuit and health-related obstacles. We need to emphasize that the sample is not fully representative regarding various population quotas. However, it is much bigger and more diverse and, therefore, suitable for our goal to replicate the factorial structure in a bigger and more age-diverse group. The university's Institutional Ethics Board approved the study.

Instruments

Examined scales

The short version of the Self-Reflection and Insight Scale (SRIS) (Silvia, 2021) has been used similarly to the previous study. However, in contrast to Phase 1, only items for the short version were included. A 5-point Likert-type agreement scale has been used, as discussed above.

Analysis

As in Phase 1, confirmatory factor analysis was used to examine the evidence for factorial validity with the MLR estimator (i.e., Maximum Likelihood with Robust Standard Errors). Scaled and robust statistics are reported.

Results

Factorial Validity

In accordance with Phase 1, the CFA for a shortened scale provided a reasonable model-data correspondence. More specifically, although the chi-square test of model fit was statistically significant ($\chi^2(53) = 304.136$, $p < .001$), indicating some misfit, the approximate fit indices, quantifying the degree of misfit, were in an acceptable range. More specifically, the robust Comparative Fit Index (CFI) at 0.95 was good, the robust Root Mean Square Error of Approximation (RMSEA) at 0.078 (90% CI [0.070, 0.087]) was acceptable, and the Standardized Root Mean Square Residual (SRMR) at 0.079 was acceptable according to the less stringent criteria.

The items loaded significantly into their factors. The items had acceptable loadings (above 0.5) except for one item. More specifically, for *the self-reflection (SR) factor*, Item 5 (i.e., item 10 in the original scale) ("I frequently examine my feelings") demonstrated the highest factor loading (.85). Item 11 (item 19 in the original; i.e., "I often think about the way I feel about things"; .77) and item 3 (item 7 in the original scale; "I am very interested in examining what I think about"; .75) also showed strong loadings. Item 9 (item 16 in the original scale; "I frequently take time to reflect on my thoughts"; .71) and 7 (item 12 in the original scale "It is important to me to try to understand what my feelings mean"; .66) provided adequate factor loadings. Item 2 (item 5 in the original scale) ("It is important to me to evaluate the things that I do"; .43) showed a less optimal but acceptable factor loading, better than in Phase 1.

For the insight (I) factor, item 1 (item 4 in the original scale; "I am often confused about the way that I really feel about things")

showed the highest factor loading (.87). Items 6 (11 in the original scale; "My behavior often puzzles me"; .83), item 10 (17 in the original scale; "Often I find it difficult to make sense of the way I feel about things"; .83), and 8 (14 in the original scale) ("Thinking about my thoughts makes me more confused"; .82) also demonstrated strong loadings. Item 4 (9 in the original scale; "I'm often aware that I'm having a feeling, but I often don't quite know what it is"; .75) provided an adequate factor loading. However, item 12 (20 in the original scale; "I usually know why I feel the way I do"; -.18) showed negative and weak loading. This is consistent with its reverse-scored nature but may indicate potential issues with this item. The model is depicted in Online Supplement 2.

The covariance between the SR and I factors was significant and relatively large (i.e., .63, $p < .001$). As in the previous case, it was positive in this specific CFA, but it is important to consider that the majority of the items in the I factor, including the first item, should be reversed. Thus, the sign of the correlation is negative in reality.

Examination of local fit indicated that item 12 (item number 20 in the original scale, "I usually know why I feel the way I do"; -.18) is problematic due to large residuals with many variables.

The internal consistency, as assessed via McDonald's omega (total), was good (i.e., ~ωt = .86 for self-reflection and ωt = .81 for insight subscales). The average variance explained was above the recommended 50% (i.e., .52 and .59 for self-reflection and insight, respectively).

Sensitivity Analysis Dedicated to the Acquiescence Bias

Due to the suspicious correlation between the factors and the problematic item, we also examined the acquiescence bias as a form of sensitivity analysis similarly as in the previous case. A model accounting for potential acquiescence bias provided significant chi-square ($\chi^2(52) = 199.907$, $p < .001$), but improved model-to-data correspondence in terms of the approximate fit indices (the robust CFI = .97, the robust RMSEA = 0.052 (90% CI [0.051, 0.069]), and the SRMR = 0.069. The acquiescence factor was moderate. Interestingly, acquiescence bias was higher than in study 1 (loadings ~0.4). When acquiescence bias is accounted for, the correlation between the two factors decreased (0.53) but remained significant and strong. Here, item twelve (i.e., item 20 in the original scale; "I usually know why I feel the way I do") was less problematic (-0.43), but the second item (i.e., item 5 in the original scale: "It is important to me to evaluate the things that I do") showed low factor loading (0.17) when acquiescence bias is accounted for.

Re-examination of Silvia (2021)

Due to the interesting pattern of results when acquiescence bias is taken into account, we also re-examined the results of Silvia (2021), accounting for acquiescence bias. Full results can be found in the online appendices; here, we will comment only on three observations. First, both models had significant χ^2 (i.e., $\chi^2(53) = 380.987$, $p < .001$ and $\chi^2(52) = 388.960$, $p < .001$, respectively), but acceptable to borderline model-to-data correspondence with approximate fit indices CFI = .925, RMSEA = 0.083, 90% CI [0.075, 0.091], SRMR = .065 and CFI = .926, RMSEA = 0.083, 90% CI [0.076, 0.091], SRMR = .062, respectively (all examined models are summarized in Online Supplement 1). Factor loadings were slightly below .5 for item 8 in both models, worsening in the model accounting for acquiescence bias (acquiescence bias was comparable to

our studies – loadings of ~.3). Correlation between latent factors changed from -.07 (p = .068) to -.29 (p = .031) when acquiescence bias was accounted for.

Discussion

Previous research has highlighted the importance of self-awareness in various areas, such as conscientious self-regulatory activities aimed at achieving goals. The concept of self-awareness has been studied for decades using various approaches and methods. However, one of the most widely used measurements is the Self-Reflection and Insight Scale (SRIS) created by Grant et al. (2002). The measure distinguishes between two related but different aspects of self-awareness – self-reflection and insight. Silvia (2021) recently crafted a short version of the scale, providing several benefits over the original scale, including superior psychometric qualities and brevity. However, to our knowledge, no adaptation of SRIS was conducted to the Slovak language, hindering research on the Slovak-speaking population. Reflecting on this gap, this paper aimed to adapt and examine the psychometric properties of the Slovak version of the SRIS scale – both the original (20-item) version and the short (12-item) version – in two consequent phases.

The results of the Confirmatory Factor Analysis indicated that the original full (i.e., 20-item) version of the SRIS scale (Grant et al., 2002) did not provide adequate model-data correspondence. For example, the global fit was sub-optimal and further analysis of the local fit indicated several problems. This pattern of results is in line with other studies, such as Banner et al. (2023) and Aşkun and Çetin (2017). Furthermore, other problems, such as relatively low internal consistency and average variance explained, were observed, as well as problematic factor loadings for some items. For example, item 5 ("It is important to me to evaluate the things that I do") from the self-reflection scale, or item 3 ("I'm usually aware of my thoughts") from the insight factor provided lower or inconsistent factor loadings. Some items were already problematic in the original scale, such as item 3, which had inconsistent factor loadings in two subsequent studies carried out by Grant and colleagues (2002), leading some authors to exclude problematic items (e.g., Havrdová et al., 2020).

The short version provided by Silvia (2021) attempted to ameliorate the problems while providing the brevity needed in some research settings. In our study, fit indices for the short version indicated better and reasonable model-data correspondence. Factor loadings were also better. Internal consistency was good. As such, the shorter version seems to be preferable over the original version.

One can argue that acquiescence bias is responsible for these results (see Havan et al., 2025 for a discussion on this phenomenon). Therefore, we also examined this possibility. In fact, data-to-model improved when acquiescence bias is accounted for. However, the short version is still preferable even when acquiescence bias is accounted for. Although the lower number of reversed items could complicate the interpretation of acquiescence bias in the short version, as discussed further, the short version exhibited fewer problems in all comparisons.

Although preferable, the shorter version is not without potential problems. For example, item five, "It is important to me to evaluate the things that I do", had a lower factor loading (.33; more specifically, .39 in one and only .21 in another sample). As this problem could be sample- or method-specific, we decided to examine the short version of the scale with a bigger, more diverse sample and shift the rating scale. Consequently, factor loading was not optimal but more acceptable (.43).

However, another item ("I usually know why I feel the way I do") provided sub-optimal factor loading (-.18). This can be caused by the specificity of the sample (e.g., people with health-related issues could understand this item differently than healthy young population). However, there are other explanations as well. Acquiescence bias can be another explanation. When acquiescence bias is accounted for, the loading is improved, partially supporting the concern. However, when bias is accounted for, the fifth item, "It is important to me to evaluate the things that I do" (problematic also in phase 1), shows sub-optimal factor loading.

Why does the item "It is important to me to evaluate the things that I do" seem to be problematic but was not problematic in the original study? The discrepancy can be caused by many factors, such as problems with the translation of the item or a shift in the meaning of the term evaluation across culture and time. In the Slovak language, this term might have a negative connotation. Also, the apparent relation to the construct of self-reflection could be more nuanced than expected by the original authors. Thus, the relevant question is if the problem with this item is also present in more recent data from another country. When data from Silvia (2021) are re-examined, the factor loading (.39) for the most problematic item, "I usually know why I feel the way I do", is lower than all other loadings (average .65). However, it is still higher than in our sample. Other studies found similarly lower factor loading for this item (see, e.g., Banner et al., 2023, who found factor loading for this item in a range of .31 to .36 in long and short versions, respectively), but still higher than we found. Therefore, we can speculate that the interaction of specific content and translation nuances or also sample specificity can exacerbate the problem. In both cases, further examination is needed.

Practically, while a short version (Silvia, 2021) in the Slovak language is more recommended than a more extended version (Grant et al., 2002) due to superior psychometric properties and brevity, researchers should be cautious with regard to problematic item. Our goal was not to create a new scale. Thus, we did not eliminate this item. Yet, it is one of the possibilities. Another concern about the short version of the scale is that a lower number of items can, at least theoretically, underrepresent the construct. However, a short version was created applying the item response theory (Silvia, 2021). As such, underperforming items were eliminated. Also, a recent study, conducted by Silvia et al. (2023), showed that shorter versions exhibited a very similar correlation with external criteria and created similar profiles. Similarly, our results indicate that both versions had almost the same patterns of correlations, diverging in terms of the effect size only marginally.

Also, it is worth noting that the correlation between the two factors is inconsistent – from non-significant (sample 1A) to significant but small (sample 1B) to relatively large (sample 2). However, this aligns with a pattern also observed in the research literature. For example, while some authors observed a negligible or very small correlation (±0.10), others found a much higher correlation. There are various explanations, such as how people purposefully engage in self-reflection (e.g., by keeping a diary), the capacity to move from self-reflection to insight, or further differentiation of various types of self-reflection (Grant et al., 2002). However, other reasons could be important as well. When data from Silvia (2021) are re-examined, the relationship is non-significant. However, it is significant when acquiescence bias is accounted for, indicating that this bias can play some role in solving the puzzle. However, even if the bias is accounted for in our sample, divergent results remain. Also,

when not many items should be recoded (as in the case of the short version), it is questionable if the captured variance reflects acquiescence. Thus, the puzzle seems more nuanced, and other types of bias or specificities of samples could play a role here. For example, it is possible that people in our second study had a different constellation of unattainable goals, and self-focused vs. problem-focused self-reflection worked differently. Grant et al. (2002), for example, argued that the levels of insight could increase while levels of self-focused self-reflection could decrease after systematically attaining a goal that was previously considered unattainable, mainly due to over-extensive engagement in self-focused self-reflection hindering progress. Alternatively, people in study two could have engaged in self-focused self-reflection due to health-related problems. Also, as pointed out by the reviewer, the factors of self-reflection and insight can be understood as the frequency of self-reflecting and the evaluation of one's success at self-reflecting. Items with cross-loadings start with words related to frequency and are related to self-reflection. Such reframing could also explain instability in correlations, as correlation differences could be caused by differences between samples (e.g., age). However, future research is needed to examine all these possibilities.

The examination of the convergent validity of the short version showed that insight was positively linked to resilience and life satisfaction, while negatively associated with rumination. Insight was negatively correlated with rumination and resilience and positively correlated with life satisfaction. Self-reflection was correlated with rumination[3]. The pattern of relations regarding direction and effect size was generally in line with previous research findings (e.g., Bucknell et al., 2022; Cowden & Meyer-Weitz, 2016; Harrington et al., 2016; Lyke, 2009). For example, insight was previously found to have a positive relationship with well-being (e.g., Harrington et al., 2014; Harrington et al., 2016) and a negative relationship with rumination (Bucknell et al., 2022). Self-reflection was previously associated with various health indicators, especially rumination (Bucknell et al., 2022; Harrington & Loffredo, 2010). However, self-reflection was not related to resilience. This is not in line with Cowden and Meyer-Weitz (2016). However, a pattern of a relationship may be more complex than expected. For example, Bucknell et al. (2022) found that self-reflection was associated with perceived resilience indirectly via insight.

[3] Although not of main interest in the present study, when the subscales are analyzed, engagement in self-reflection did not correlate with resilience, or life satisfaction, but it was positively related to rumination. The need for self-reflection was positively related to rumination, and life satisfaction but not to resilience.

Limitations, Implications, and Future Directions

The main aim of the presented study was achieved; however, some limitations need to be acknowledged. Firstly, while four independent translators conducted a back-translation and reconciliation procedure, some items might still need to be revised for a better concept representation, as shown by problems with one item of the short version of the scale. However, lower factor loadings could also be caused by the distributional properties of the item in the examined sample (Sellbom & Tellegen, 2019). Thus, examination of the factorial structure in a representative sample can be beneficial for future research with invariance testing across genders and various groups of people to ensure that these groups understand the construct consistency and can be meaningfully compared. Also, the new scale can be crafted in case the problems with problematic items will be prevalent. Re-

latedly, a multi-country validation study of the short version of the scale could be beneficial for future research, alongside the addition of a more person-oriented approach and analysis to inform content validity. Also, although we tried to account for acquiescence bias via the unmeasured latent variable technique (Havan et al., 2025), the question of what exactly the factor represents remains, especially with regards to the short version as it is possible that acquiescence bias also captures some construct-relevant variance and this should be interpreted with caution (see Podsakoff et al., 2024).

Notwithstanding some limitations, this study presents unique findings that can inform future studies and facilitate the research not only in Slovakia. Namely, the psychometric properties of the short Slovak version of the SRIS scale were initially supported, providing the basis for using this scale to measure self-reflection and insight in future research. As this is an adaptation study, practical implications are limited, but future studies focused on self-reflection and insight using a short version of the scale can bring several practical implications across various psychological sub-disciplines.

Conclusions

The present study provided initial evidence for the factorial validity, convergent validity, and test-retest reliability of the Slovak short (i.e., 12-item) version of the Self-Reflection and Insight Scale (SRIS) (Silvia, 2021). We did not find support for the original version of the scale (20 items; Grant et al., 2002) due to problems with factorial validity and internal consistency. However, we found some evidence that the problem could be partially caused by acquiescence bias. This bias could also be one piece of the puzzle explaining the inconsistent pattern of correlations between insight and self-reflection across studies. Even when acquiescence bias is accounted for, a short version of the scale suggested by Silvia (2021) seems preferable, but it is not without potential challenges, as one item could function sub-optimally, as indicated by lower factor loading.

Acknowledgement

This work was supported by the Scientific Grant Agency of the Ministry of Education, Science, Research and Sport of the Slovak Republic and Slovak Academy of Sciences (contract no. 1/0145/23) and Slovak Research and Development Agency under Contract No. APVV-23-0548.

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