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Psychometric Evaluation of the Short Self-Regulation Questionnaire across Three European Countries

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The aim of this study was to extend the psychometric evaluation of the Short Self-Regulation Questionnaire (SSRQ) by assessing the factor structure across three countries from Central and Eastern Europe. The sample included 1809 students from Slovakia, Lithuania and Hungary. Based on an initial confirmative factor analysis, a 2-factor structure by Neal and Carey (2005) was confirmed in the Lithuanian sample. Next, exploratory factor analyses were used on the Slovak and Hungarian subsamples separately. For both national subsamples, a very similar four factor solution was found, which was confirmed by confirmatory factor analyses on the rest of the data. Despite the reduced number of items, the abridged scale did not suffer in terms of its internal reliability and thus provides an adequate approximation of self-regulation levels as the entire scale or as the scale with the proposed 4-factor solution.

Key words: SSRQ, self-regulation, psychometric evaluation, university students

Introduction

From a social cognitive perspective, self-regulation is the ability of an individual to manage his own behavior through observation, evaluation, and consequation. It involves generating thoughts, feelings, and actions that are planned and adapted to the attainment of personal goals (Zimmerman, 2000). Self-regulation refers to the regular exercise of control over oneself in order to adapt (Zimmerman, 2000) and bring oneself in line with preferred standards (Carver & Scheier, 1998; Vohs & Baumeister, 2004). Self-regulation correlates with various aspects of life. For example two longitudinal, prospective studies of middle school students found that self-regulation helps students to study, complete homework, behave positively in the classroom, get better grades, and school attendance (Duckworth, Quinn, & Tsukayama, 2012;

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Duckworth & Seligman, 2005). In the context of health and interpersonal relationships, life goals have been found to be associated with higher levels of well-being and better mental health (Martos & Kopp, 2012; Kasser & Ahuvia, 2002; Sheldon, Klinesmith, Houser-Marko, Osbaldiston, & Gunz, 2007). Hofer, Bush, and Kärtner (2011) found among a sample of university students that those with higher self-regulatory capabilities had higher levels of well-being. Other studies found self-regulation related to lower depression, anxiety and stress among university students (e.g., Park, Edmondson, & Lee, 2012) or low psychopathological symptoms and better interpersonal relationships (Tangney, Baumeister, & Boone, 2004).

Self-regulation skills also emerged as the predictors of avoiding problematic use of the internet and alcohol (Sebena, Orosova, & Benka, 2013; Seay & Kraut, 2007). Lower levels of self-regulation functions were found to be a risk factor for experiencing alcohol-related consequences and for reductions in alcohol use and consequences over time for heavier drinking college students (Hustad, Carey, Carey, & Maisto, 2009).

Self-regulation can also be seen as relevant with respect to its theoretical implications. It is a key concept for understanding what the human self is and how it operates. Self-regulation is an important function of the human self, one that helps define self and is relevant with the self's executive function, which is related to self-control, control of the environment (has some relevance to self-knowledge and to interpersonal belonging), self-directed behavior or decision-making and choosing (Baumeister, Schmeichel, & Vohs, 2007).

This study is based on Miller and Brown's (1991) theory, which proposed seven dimensions of self-regulation: 1) informational input, 2) self-monitoring current progress towards a personal goal, 3) motivation for change, 4) commitment to reaching the goal, 5) development of a plan to reach the personal goal, 6) work according to the plan and 7) re-evaluation of the plan. Miller and Brown's model implies that deficits in any one stage can lead to self-regulation difficulties and so individuals may have problems to regulate their behavior and to achieve the desired outcomes or goals.

In order to capture the dimensions of the mentioned Miller and Brown model (1991), the Self-Regulation Questionnaire (SRQ), a 63-item instrument was developed. After a psychometric evaluation of the SRQ, the 7 factors of the proposed self-regulation theory were not confirmed and the total sum score was recommended as a measure of general self-regulation skills (Brown, Miller, & Lewandowski, 1999). Carey, Neal, and Collins (2004) extended the psychometric evaluation of the SRQ by evaluating its factor structure. However, the results did not confirm the 7-factor scale but rather a single factor. From this, 31 of the 63 items loaded significantly. As a result, a short form of the SRQ (SSRQ – Short Self-Regulation Questionnaire) consisting of these 31 items was produced (Carey, Neal, & Collins, 2004). The next verification of the SSRQ found a 2-factor solution: impulse control and goal setting factors (Neal & Carey, 2005).

To the best of our knowledge, no valid and reliable measurement tool for the self-regulation construct exists in the Slovak Republic. It was decided to choose a previously generated and tested general measure of self-regulation rather than a more specific one, so it could be used in various domains of human functioning. The aim of this study is to extend the psychometric evaluation of the SSRQ by assessing the factor structure across three countries from Central and Eastern Europe (Slovakia, Lithuania and Hungary).

The study specifically aimed to:

1) confirm the previous 7-factor theoretical model by Miller and Brown (1991), the 1-factor

model by Carey, Neal, and Collins (2004) and the 2-factor model by Neal and Carey (2005).

2) if confirmation failed, it aimed to provide a exploratory factor analysis across the three national samples.

Method

Participants

Data were used from the first wave of the Student Life Cohort in Europe (SLiCE), a multinational longitudinal study among first year university students from several European countries. This study is based on data collected in 2011. This study has been developed from the previous activities of the Cross-National Student Health Survey (El Ansari et al., 2007). The multinational cohort was planned for the whole period of university education. The collaborating universities were selected according to the personal contacts of the researchers. This analysis is based on the data from first year students in Hungary, Lithuania and the Slovak Republic, where over 500 participants were recruited from each country. In the three countries, nine universities took part in the study; four universities in Kaunas, Lithuania (Lithuanian University of Health Sciences, Kaunas University of Technology, Aleksandras Stulginskis and Vytautas Magnus University), two universities in Hungary (Eötvös Loránd University and the University of Miskolc) and three universities in Kosice, Slovak Republic (P. J. Safarik University, the University of Veterinary Medicine, and the Technical University).

In each location, students were asked to complete self-administered online questionnaires. The strategies of recruiting respondents differed at each place because of the structural differences in the participating countries. The universities in Lithuania and the Slovak Republic provided access to the e-mails of all enrolled first year students. The project was introduced to students during their lectures and seminars and an invitation e-mail to participate was subsequently sent out. The Hungarian students were informed using university newsletters and other formal as well as informal methods. Following this, they registered on the SLiCE website and filled in the form. In total, the sample included 1809 students. The overall response rate was 22.69% (20.03% in Slovakia, 23.05% in Lithuania, and 25% in Hungary) when comparing respondents to all enrolled university students in the given year at the respective universities. Generally, the universities included in this study represented biomedical, social, physical, and technical sciences.

Characteristics of the Sample

The mean age of the students was 20.15 years ($SD = 3.38$). The description of the samples from the different countries in terms of sex and age are shown in Table 1.

Table 1 *Description of the sample by country*

	Slovakia N = 649	Lithuania N = 582	Hungary N = 578	p-value
Sex				.034*
Female [%]	73.4	70.3	75.4	
Male [%]	26.6	29.7	24.6	
Age [mean (SD)]	19.61(3.64)	20.00(2.83)	20.84(3.64)	<.001

* Chi-square for comparison of sex proportionality between the three countries

Student participation in the study was voluntary and anonymous. Students were informed that by completing the questionnaire they were providing their informed consent to participate. They were also informed that they could terminate their participation at any point while filling out the questionnaire. No incentives were provided. Permission to conduct the study was granted by the ethical commissions of the participating institutions. Initially, the questionnaire was compiled in English and subsequently translated into the local languages using two independent forward translations for each language. The research team reviewed any cases of disagreement and the authors familiar with the respective languages, usually native speakers, made the final decisions.

Measures

The measures used in this study were part of a multi-topical questionnaire assessing health and health behaviors among university students.

Demographic Variables

Students' sex and age were based on individuals' self-reports in the survey.

The Short Self-Regulation Questionnaire. The SSRQ is a 31-item scale that was designed to assess self-regulation skills. Items are scored on a 5-point scale from 1 – strongly disagree to 5 – strongly agree. A previous study by Neal and Carey (2005) indicates that the SSRQ has two distinct factors; an impulse control and a goal-setting factor. Questions on the impulse control factor include for example: "It's hard for me to notice when I've had enough (alcohol, food, sweets)," or "I am able to resist temptation." Questions on the goal setting factor include for example: "Once I have a goal, I can usually plan how to reach it" or "I am able to accomplish goals I set for myself".

Statistical Analysis

Two methods were used to assess the factor structure of the SSRQ. Principal component analysis was used to identify the factors in the whole sample and Structural Equation Modeling was used to further confirm the results across the countries.

Firstly, a confirmatory analysis approach was employed to test the factor structure of the three competing models. Secondly, exploratory factor analyses were conducted. Principal component analysis with a direct oblique rotation (Oblimin) and a covariance matrix as the input for the 2/3 national subsamples separately were used. The Goodness of Fit Index (GFI) was computed, which showed how closely the model replicated the observed covariance matrix. In order to examine the results of the factor analysis further, an analysis of the individual items was conducted to assess the reliability and the convergent properties of the reduced SSRQ subscales. Then, we tested whether each item was individually correlated with the sum of the items in the same factor (item-test correlation) and the aggregate of the remaining items in that subscale (item-rest correlation). Finally, the internal consistency (Cronbach's Alpha) was computed for each factor, and for each factor minus one item in order to determine whether dropping an item would increase the overall internal consistency in a meaningful fashion.

To verify and confirm the factor structure of the model, Confirmatory Factor Analyses (CFA) were used. In order to define the good fit, several fit indices were applied. A satisfactory degree of fit requires the comparative fit index (CFI) to be close to 0.95, and the model should be rejected when these indices are below 0.90 (Brown, 2006). The next fit index was the root mean squared error of approximation (RMSEA). A RMSEA below 0.05 indicates an excellent fit, a value around 0.08 indicates an adequate fit, and a value above 0.10 indicates a poor fit.

The PCLOSE measure, which goes with the RMSEA and provides the probability of a hypothesis test that the RMSEA is no greater than 0.05, should be above 0.05 (Byrne, 2010).

The data analysis was performed using the statistical program PASW for Windows, version 18.0 and Amos 16.

Results

Confirmatory Factor Analyses of Previous Models

First of all, given that previous studies have tested the factor structure of the SSRQ, three competing models using the confirmatory analysis approach were initially tested on data that had no missing values. Based on the theorized model (Brown, Miller, & Lewandowski, 1998), the 7-factor solution (Receiving, Evaluating, Triggering, Searching, Formulating, Implementing, Assessing) was tested on each national sample. However, this initial model, did not fit the data well in each selected country (see Table 2).

The results of Carey, Neal, and Collins (2004) revealed a 1-factor model on which all 31 items was loaded. However, testing the single factor model on each national sample did not provide an adequate fit to the data either (Table 3). The modification indices were examined within each factor structure model, which indicated that a number of substantial changes would need to be made in order to achieve a good fitting model. After that it was decided to test the bi-factorial structure by Neal and Carey (2005), considering the Impulse control factor and Goal setting factor. Neither of these factor structures fitted the data well for the Slovak and Hungarian national samples (see Table 4). However, for the Lithuanian sample, the modification indices in-

Table 2 *CFA of the 7-factor theorized model by Brown, Miller and Lewandowski (1998)*

	N	*χ2 (p)*	*GFI*	*CFI*	*RMSEA*	*PCLOSE*
Slovak Republic	649	2537(p<.001)	.764	.664	.086	<.001
Hungary	578	3433(p<.001)	.689	.646	.109	<.001
Lithuania	582	2291(p<.001)	.763	.720	.086	<.001

Table 3 *CFA of the one-factor model by Carey, Neal, & Collins (2004)*

	N	*χ2 (p)*	*GFI*	*CFI*	*RMSEA*	*PCLOSE*
Slovak Republic	649	1566 (p<.001)	.784	.701	.085	<.001
Hungary	578	2313 (p<.001)	.746	.740	.092	<.001
Lithuania	582	2007 (p<.001)	.788	.746	.084	<.001

Table 4 *CFA of the two-factor model by Neal and Carey (2005)*

	N	*χ2 (p)*	*GFI*	*CFI*	*RMSEA*	*PCLOSE*
Slovak Republic	649	1192 (p<.001)	.835	.732	.091	<.001
Hungary	578	1126 (p<.001)	.824	.740	.093	<.001
Lithuania	582	397 (p<.001)	.927	.93	.05	.062

dicated that three covariances between the errors of indicators within the same factor would need to be added in order to improve the model's fit. Two items ("I have trouble following through with things once I've made up my mind to do something." (Impulse Control) and "When I'm trying to change something, I pay a lot of attention to how I'm doing." (Goal Setting)) were removed because of a low factor loading. After this change, the goodness of fit for the Lithuanian sample was $\chi^2 = 397$ (141 df, $p \leq .001$), χ^2/df = 2.82, SRMR = 0.05, CFI = 0.93, RMSEA = 0.05, PCLOSE = 0.06. In terms of reliability, Cronbach's Alpha for the total score was $\alpha = .87$, $\alpha = .80$ in the Goal setting and $\alpha = .81$ in Impulse control dimensions. This factor structure was accepted for the Lithuanian sample.

As the next step it was decided for the Slovak Republic and Hungarian national samples to move towards an exploratory factor analysis approach on the SSRQ items.

Exploratory Factor Analyses

A principal component exploratory analysis was conducted on 2/3 randomly chosen separate Slovak and Hungarian subsamples. For both subsamples, four factors were revealed with eigenvalues of 6.3, 2.1, 1.6 and 1.5 that cumulatively explained 50.1% of the total variance for the subsample from Slovakia and with eigenvalues of 8.7, 1.8, 1.4 and 1.3 that cumulatively explained 55.2% of the total variance for the Hungarian subsample. Parallel analysis verified that the eigenvalues of the four factors of each national subsamples were greater than what was expected by chance, given the number of items and sample size. The 4-factor solution was rotated (Direct Oblimin) to improve its interpretability; 23 items were loaded onto the rotated factors at .435 or higher for the Slovak Republic and 24 items were loaded onto the rotated factors at .573 or higher for the Hungarian subsample. The KMO-measure of sampling adequacy was .88 for the Slovak and .912 for the Hungarian subsample. The Bartlett's test of sphericity was significant ($p < .001$) across both national samples, supporting the factorability of the data. For the Slovak subsample, from the SSRQ-31 items, 23 items were classified as single-loading (loading > .4 on one factor and < .2 on the other), 8 were classified as cross-loading items (loading > .4 on one factor and > .2 on the other) or non-loading items.

7 items loaded significantly onto the first factor (self-discipline), 5 loaded significantly onto the second factor (goal-setting), 4 onto the third factor (learning from mistakes) and 7 loaded significantly onto the fourth factor (impulse control).

The study found very similar results for the factor structure for the Hungarian subsample, where 24 items were classified as single-loading and 7 as cross-loading or non-loading items. 8 items loaded significantly onto the first factor (self-discipline): "If I wanted to change, I am confident that I could do it." 5 loaded significantly onto the second factor (goal-setting): "I set goals for myself and keep track of my progress." 4 onto the third factor (learning from mistakes): "I usually only have to make a mistake one time in order to learn from it." 7 loaded significantly onto the fourth factor (impulse control): "Often I don't notice what I'm doing until someone calls it to my attention." The results of the exploratory factor analysis for both national subsamples are provided in Table 5.

Item Analyses, Reliability Analyses

For the item analyses to assess the reliability and convergent properties of the reduced SSRQ subscales, the study tested items belonging to each factor separately. Firstly, the items that had negative factor loadings were reversed to maintain consistency in the analyses.

In the Slovak subsample for factor 1, the item-test correlations ranged from .64 to .77, item-

Table 5 *Factor Loadings from Revised Principal Component Analysis with Direct Oblimin Rotation*

	Slovak Republic ($N = 433$)				Hungary ($N = 388$)			
	F1	*F2*	*F3*	*F4*	*F1*	*F2*	*F3*	*F4*
I am able to accomplish goals I set for myself.	.603				.741			
I have a lot of willpower.	.790				.678			
Once I have a goal, I can usually plan how to reach it.	.729				.702			
If I wanted to change, I am confident that I could do it.	.684				.750			
I can stick to a plan that's working well.	.650				.663			
I know how I want to be.					.670			
I can usually find several different possibilities when I want to change something.					.567			
I am able to resist temptation.	.729							
I have trouble making plans to help me reach my goals.	-.580				-.638			
As soon as I see a problem or challenge, I start looking for possible solutions.		.591				.508		
When I'm trying to change something, I pay a lot of attention to how I'm doing.		.750				.804		
If I make a resolution to change something, I pay a lot of attention to how I'm doing.		.725				.788		
I usually keep track of my progress toward my goals.		.698				.606		
I set goals for myself and keep track of my progress.		.746				.594		
I learn from my mistakes.			.796				.760	
I usually only have to make a mistake one time in order to learn from it.			.787				-.804	
I don't seem to learn from my mistakes.			-.796				-.735	
I usually think before I act.			.457				.621	
I put off making decisions.				-.670				-.593
It's hard for me to notice when I've "had enough" (alcohol, food, sweets).								-.612
I tend to keep doing the same thing, even when it doesn't work.				-.435				-.801
Most of the time I don't pay attention to what I'm doing.				-.462				-.796
Often I don't notice what I'm doing until someone calls it to my attention.				-.565				-.632
When it comes to deciding about a change, I feel overwhelmed by the choices.				-.678				
I have trouble making up my mind about things.				-.728				
I get easily distracted from my plans.								-.573
I have trouble following through with things once I've made up my mind to do something.				-.627				-.652

Note. Values less than .40 are not displayed. Revised Factor Analysis included only the items that loaded on either of the four (SVK, HU) rest correlations ranged from .63 to .75, (overall alpha = .76). For the second factor, item-test correlations ranged from .65 to .77, item-rest correlations ranged from .62 to .74, (overall alpha = .79). For the third factor, item-test correlations ranged from .76 to .81, item-rest correlations ranged from .73 to .79, (overall alpha = .69). For the fourth factor, item-test correlations ranged from .56 to .70, item-rest correlations ranged from .53 to .67, (overall alpha = .73).

In the Hungarian subsample for factor 1, the item-test correlations ranged from .64 to .77, item-rest correlations ranged from .61, to .75, (overall alpha = .84). For the second factor, item-test correlations ranged from .69 to .81, item-rest correlations ranged from .65 to .78, (overall alpha = .81). For the third factor, item-test correlations ranged from .64 to .86, item-rest correlations ranged from .60 to .83, (overall alpha = .77). For the fourth factor, item-test correlations ranged from .62 to .75, item-rest correlations ranged from .60 to .72, (overall alpha = .81).

The total internal consistency was also acceptable for the Slovak Republic (Cronbach's Alpha = 0.87 (23 items)) and Hungary, (Cronbach's Alpha = 0.92 (24 items)).

The correlation between the full 31-SSRQ version and the shortened version was .99 among both national subsamples.

Confirmatory Factor Analyses of SSRQ by Structural Equation Modeling

To verify the factor structure of the models from two countries, structural equation modeling was used. The analyses were made on the final 1/3 of the data, which were not used in the exploratory factor analyses. The model fit was evaluated in terms of chi-square, standardized root mean square residuals (SRMR), and various goodness of fit indices.

First, we had a look at the factor structure of the Slovak Republic model. The modification indices indicated that adding three covariance between the errors of indicators within the same factor would need to be made in order to improve the model's fit. After this change, the goodness of fit for the model was χ^2 =338.97 (199 df, $p \leq .001$), χ^2 /df = 1.70, SRMR = 0.05, CFI = 0.91, RMSEA = 0.05, PCLOSE = 0.22. The Hungarian model, derived from the exploratory factor analyses, showed very similar factor structure to the one derived from the Slovak Republic subsample. The modification indices indicated that two covariances would need to be added between the errors of indicators in order to improve the model's fit. After this change, the goodness of fit for the model was χ^2 =292.61 (144 df, $p \leq .001$), χ^2 /df = 2.03, SRMR = 0.05, CFI = 0.92, RMSEA = 0.05, PCLOSE = 0.12. Although the χ^2 was significant in both models here, other indexes showed that the model still fitted very well. Thus, it is possible to say that both models describe the data very well. The reason for this undesirable χ^2 significance is the large sample size.

Discussion and Conclusion

The aim of this study was to provide a psychometric evaluation of the SSRQ -31 (Neal & Carey, 2005) based on a university student sample from Slovakia, Lithuania and Hungary.

Exploratory Factor Analyses

Firstly, it was attempted to confirm the factor structure of the 7-factor theoretical model by Miller and Brown (1991) as well as the 1- and 2-factor model (Carey, Neal, & Collins, 2004; Neal & Carey, 2005).

The results have supported the findings from other studies, namely, that the SSRQ does not follow the steps in the self-regulation theory as described by Miller and Brown (1991). Regarding the Slovak and Hungarian data, the study also failed to confirm the structure of the other two models (Carey, Neal, & Collins, 2004; Neal

& Carey, 2005); none of them provided an adequate fit to the data in either country. However, we confirmed the 2-factor – impulse control and goal setting factors model by Neal and Carey (2005) in the Lithuanian sample. This two-factor model was also demonstrated in Portugal by Dias and Garcia del Castillo (2014).

For the Slovak and Hungarian data, it was decided to move towards an exploratory factor analysis, which showed that the best fitting model for both national datasets was the 4-factor model in comparison to the alternative models. The factor structure was similar in both national subsamples and from all 31 items, the final Slovak and Hungarian models revealed 20 items in common for both models.

The first two factors are made up of positively connoted items that are very similar to Neal and Carey's (2005) goal setting factor and seems to be very important in accomplishing the planned goals. This first factor was labeled in this study as *Self-discipline*. It contains items related to one's self-confidence, self-discipline and willpower to reach the goals: ("If I wanted to change, I am confident that I could do it.").

The second factor (*Goal setting*) consists of five items related to the ability to plan, set and keep track of a person's progress towards goal attainment. In both national samples, this factor contains the same 5 items, and all of them were part of Neal and Carey's (2005) goal setting factor (e.g., "I set goals for myself and keep track of my progress.").

The last two factors in both national samples are made up of 11 items that were part of Neal and Carey's (2005) impulse control factor. In the context of self-regulation, it is the ability to resist temptation, urges or impulses that may disrupt the goal directed behavior.

The third factor is labeled as *Learning from mistakes* and consists in both samples of four items related to learning from previous mistakes ("I usually only have to make a mistake one time in order to learn from it.").

The last factor *(Impulse control)* contains seven items that represent the awareness of a person's own thoughts and actions. All these items are negatively formulated with the aim of identifying one's automatic or mindful actions ("Most of the time I don't pay attention to what I'm doing.").

Verification of Slovak and Hungarian Factor Structure Model of SSRQ by Structural Equation Modeling

In the next step, the study tried to verify the Slovak and the Hungarian factor structure models by Structural Equation Modeling (SEM). In the analyses, only the items that loaded on the four factors were included. It can be concluded that the 4-factor solution showed a good fit in the structural equation modeling in both national samples.

For all national samples we tested, it can be concluded that the validation of the SSRQ led to a satisfactory factor structure in all the national samples. The Goal Orientation and Impulse control factors seem to be the core of the SSRQ and self-regulation capacity and have also been confirmed in other studies (Neal & Carey, 2005; Dias & Garcia del Castillo, 2014).

Despite the limited number of items, the abridged scales did not suffer in its internal reliability, the correlation between the full 31-SSRQ version and the shortened versions showed a strong positive correlation and thus provides an adequate approximation of self-regulation levels as the entire scale or as the scale with the proposed 2 (Lithuania) or 4 factor solutions (Slovak Republic and Hungary).

It is believed that this shortened measure could provide valuable information about a person's self-regulation level. However, given the slightly differential results obtained in this study, it is recommended that there be a further examination of these factor structures as well as an assessment of the divergent and convergent validity. Its structure needs to be confirmed by other populations in order to make conclusive statements.

Possible Implementations

The potential usage of the SSRQ appears to be strong. The measure of general self-regulation capacity may be predictive of a wide range of behaviors. Self-regulation refers to the global process used to achieve goals. Thus, this measure could be used in predicting goal directed behaviors such as academic achievement, gambling, financial difficulties and behavior change process. In the future, we would like to evaluate this measure of self-regulation in relation to alcohol consumption and alcohol related problems.

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Cognitive and Social Sources of Adolescent Well-being: Mediating Role of School Belonging

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The aim of the study was to explore direct and indirect pathways from cognitive factors (academic and social self-perception) to adolescent well-being through school belonging. The direct and indirect effects of cognitive factors were investigated on both concurrent and prospective well-being indicators. The first data collection was held in the beginning of the school year, the study sample consisted of 139 freshmen of three high schools in Nitra (53 boys, 86 girls, M_{age} = 15.63, SD = 1.15). The second data collection was held six months later from 109 respondents (40 boys, 69 girls, M_{age} = 15.16). Self-report questionnaires were utilized. Results showed that school belonging mediated the association between social and academic competence and students' concurrent optimism, connectedness and happiness. School belonging also mediated the relationship of social competence to prospective optimism, connectedness and happiness. Findings suggest that a developed sense of connection to school in the transition period may promote overall well-being in adolescents.

Key words: self-competence, school belonging, well-being, mediation analysis

Studying well-being, defined as positive characteristics and indicators of satisfied and happy life, is a topic of great theoretical and practical significance. However, examining well-being in youths has lagged behind the studies of adults. Adolescents spend a considerable amount of time at school. Hence, it is not surprising that school experience plays an important role in their health and successful lifelong development (Jose, Ryan, & Pryor, 2012; Phan, Ngum, & Alrashidi, 2016). Particularly, the period of school transitions holds considerable importance, as it is when parallel changes in social and cognitive factors of adolescent functioning occur (Lester, Waters, & Cross, 2013). In the past years, there have been calls for schools to expand the aims of education beyond learning outcomes to promote well-being (Gajdošová, 2016). In the present study, we explore the role of both cognitive and social resources of adolescent school life after the transition, in promoting concurrent and prospective positive psychological functioning, examined via the EPOCH model of subjective well-being. We investigate whether the effect of cognitive factors (self-perception) on adolescents' well-being is mediated by social domain (school belonging).

Subjective well-being is a meta-construct, consisting of not only the absence of psychological symptoms and disorders but also of the presence of positive feelings, thoughts and behavior (Kern, Benson, Steinberg, & Steinberg, 2016; Seligman, 2014). In recent years, conceptualization of subjective well-be-

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ing has expanded beyond feeling happy to broader concepts of psychological, personal, cognitive and motivational attributes that define optimal human functioning. Numerous models have been suggested, for example Ryff's theory of Psychological well-being (1989), Antonovsky's Salutogenic theory (1993), Self-determination theory (Deci & Ryan 2000) or in positive psychology Seligman's (2014) multidimensional model of well-being called PERMA.

Based on Seligman's theoretical perspective, Kern et al. (2016) introduced a theoretical and psychometrical model of optimal functioning in adolescents. The model called EPOCH describes five positive characteristics (engagement, perseverance, optimism, connectedness, happiness), which, according to the authors, support flourishing in adulthood. Engagement is described as a capacity to be immersed in and focused on one's activity and involvement and interest in life tasks. Perseverance is defined as a capacity to achieve one's goals regardless of difficulties or obstacles. Optimism is understood as a perception of the future as bright and favorable, being hopeful and confident. An optimistic person considers negative events as temporary, external and specific to a situation. Connectedness represents one's belief that he/she is loved, cared for, valued, supported, while also providing love, support, and care for others. Finally, happiness refers to stable states of positive mood and emotions regarding one's life. All of the five dimensions are non-developmental (normative immaturity does not lead to lower levels of well-being) and not specified in particular context. The EPOCH model represents the general capacity of youth to be engaged, perseverant, optimistic, connected and happy, regardless of the context specificity.

Well-being in adolescents is understood as a multi-dimensional and multi-causal phenomenon, influenced by both individual and contextual variables (Danielsen, Samdal, Hetland, & Wold, 2009; Rodriguez-Fernandez et al., 2016; Ronen, Hamama, Rosenbaum, & Mishley-Yarlap, 2014). Therefore, a good way to study adolescent positive functioning is to take into account multiple predictors, multiple pathways and multiple well-being outcomes (Friedman & Kern, 2010). Moreover, recent research has suggested that there are not only direct but also indirect (mediated) pathways from predictors to well-being (Blatný & Šolcová, 2016). However, most research has focused on the mediating role of cognitive factors (Danielsen et al., 2009; Rodriguez-Fernandez et al., 2016), whereas social factors as potential mediators of the influence of cognitive factors on positive functioning have not been given significant attention. This is why we find it necessary to explore whether cognitive factors directly predict various indicators of adolescent well-being and whether the effect of cognitive factors is (fully or partially) explained via social variables.

The direct effect of cognitive factors (such as self-efficacy, expectancy, effort) on well-being is well documented. In the present study, we focused on one of the cognitive factors – self-concept, that we defined as a multifaceted phenomenon composed of various aspects, roles, perspectives or selves. Multidimensional approach of the self assesses how people evaluate their competence, adequacy and knowledge they have in various domains of their lives (Harter, 2012). In the school life of adolescents, scholastic learning and social interactions represent two main activities. Accordingly, the present study focuses on self-perception of academic and social competencies. Academic competence refers to adolescent perception of cognitive competence applied to schoolwork (e.g., perception how well one is doing at schoolwork, at figuring out the answers, perception of one's intelligence). Social competence is described as one's perception of knowing how to make friends, how to become popular, having skills to get others to like oneself.

According to Harter (2012), domain specific self-perceptions have their unique and separate relations to various indicators of well-being. As previous research suggested, perception of academic competence is more predictive of academic related indicators of well-being, such as persistence, optimism or academic engagement (Phan et al., 2016; Prokopčáková, 2015; Mih & Mih, 2013), whereas perception of social competence more strongly contributes to non-academic indicators of well-being, such as life satisfaction, quality of life and positive relations with others (Danielsen et al., 2009; Phan et al., 2016; Rodriguez-Fernandez et al., 2016). According to Olsson, McGee, Nada-Raja, & Williams (2013) social and academic aspects of adolescent development are two important but different pathways to well-being in adulthood.

Indirect effects of cognitive dimensions on well-being through social factors have not been frequently studied. Few studies have suggested that positive self-perception may be indicative of greater engagement of adolescents at school, which in turn may lead to students' positive outcomes (Mih & Mih, 2013; Phan et al., 2016). It follows that students' positive feelings, experiences and relationships within the school (emotional engagement) may serve as a mediator of the effect of students' individual cognitive factors on optimal functioning.

In the present study we posit school belonging (conceptualized as a factor of emotional engagement) as a central mediator of the relation between self-perception and adolescent well-being. School belonging is defined as "the extent to which students feel personally accepted, respected, included and supported by others in the school social environment" (Goodenow, 1983) and its importance in adolescent development is explained through Maslow hierarchy of needs (1968), Stage-environment fit theory (Eccles et al., 1993) and Self-determination theory (Deci & Ryan, 2000). We built our theoretical model on the empirical support suggesting that cognitive factors (such as self-competence, self-esteem, self-efficacy) are significant sources of students' sense of school belonging (Allen, Kern, Vella-Brodrick, Hattie, & Waters., 2016; Pittman & Richmond, 2007; Vaz et al., 2015), which in turn is concurrently and also longitudinally associated with adolescent behavioral, emotional, social and academic outcomes (Pittman & Richmond, 2007; O'Neel & Fuligni, 2013; Pečjak & Pirc, 2017; Jose et al., 2012; Lester et al., 2013). The theoretical foundation of the mediational model can be found in the Expectancy-value theory (Eccles & Wigfield, 1995), according to which individual's expectation for success (perceived competence) influences school engagement, which in turn predicts important outcomes in adolescents. Nevertheless, research simultaneously analyzing multiple domains of self-perception and school belonging in relation to various well-being indicators in non-educational settings is limited.

Considering the growing consensus on the conceptualization of well-being as a multi-dimensional and multi-causal phenomenon and the need to study complex relational patterns between its predictors, we explore both direct and indirect pathways from cognitive factors (self-perception) through social factors (school belonging) to multiple indicators of adolescent well-being (engagement, perseverance, optimism, connectedness, happiness). Taking into account different pathways of multiple predictors leading to adolescent well-being, we focus on both academic and social self-perception and investigate their potentially different contributions to various aspects of adolescent well-being. Filling the gap in previous research, the aim of the study was to test whether the effect of perceived self-competence on context unspecified indicators of well-being can be mediated by social factors of adolescents school life (school belonging).

Empirical findings also determined longitudinal associations between cognitive and social variables and adolescent well-being, suggesting their different predictive effect on later well-being (Jose et al., 2012; Olsson et al., 2013). Considering these findings, the second aim of the study was to test whether self-perception directly or indirectly (through school belonging) predicts later well-being (six months later).

Present research focused on the sample of high school freshmen shortly after the school transition. Studying students' self-perception and school belonging shortly after a transition has a particular meaning, as school transitions have significant effect on psychological, social and cognitive functioning of adolescents (Harter, 2012). Longitudinal studies suggest that the level of connection to school during the transition period strongly predicts students' later emotional health and academic values (Lester et al., 2013; O'Neel & Fuligni, 2013). Hence, in the present study we specifically aim to identify crucial factors acting during the period of students' transition, which can promote adolescent concurrent and prospective well-being.

Method

Sample and Procedure

The first data collection started in the beginning of the school year (during October and November 2016) among students attending three high schools in Nitra. The initial sample consisted of 176 students, attending first, second and fourth year of the study. In order to elaborately examine the effect of transition, we decided to include only freshmen in our sample. The final sample consisted of 139 freshmen, 53 boys (38.1%) and 86 girls (61.9%) with mean age 15.16 (SD = 0.49). After approval of the survey by the school principals, questionnaires were administered to students during one class period (45 minutes). At the beginning, participants were met by researchers who briefly acquainted them with the purpose of the study and accustomed them to research context. The questionnaires were completed on a voluntary and anonymous basis. Respondents completed all the questionnaires included in the study.

The second data collection was held six months later in April 2017. Sample consisted of 109 respondents (40 boys and 69 girls, M_{age} = 15.16) who completed the EPOCH questionnaire. 19 freshmen from the first data collection did not participate in the second data collection because of an absence during the testing day and 11 students were lost to follow-up.

Materials

Self-perception profile for adolescents (Harter, 2012) – measures the global self-esteem and competence in eight specific domains of adolescent's life. In the present research two subscales – academic and social competence were used. Each of the scales consists of five statements that are formulated in a "structured alternative format" and individual chooses from two alternative statements (e.g., *"Some teenagers do very well at their classwork or Other teenagers don't do well at their classwork"*) and then rates if the statement is "Really true for me" or "Sort of true for me". The authors of the original questionnaire confirmed good internal consistency (α = .85 – .91). Self-perception profile for children (in which social and academic competence scales are identical with the adolescent version) was validated in the Slovak context, documenting satisfactory validity and reliability (Babinčák, Mikulášková, & Kovalčíková, 2012). In the present study, internal consistency of academic competence scale was α = .70, for social competence α = .83.

School belonging – questionnaire is a part of the Student questionnaire developed by OECD and utilized in international testing PISA

(2003). The questionnaire consists of six statements (e.g., *"I feel like I belong at school"*). Students are asked to indicate how they feel about each item using a four-point scale from "strongly disagree" to "strongly agree". The questionnaire displayed satisfactory level of reliability, validity and cross-cultural applicability (OECD, PISA, 2003). In the present study, internal consistency of the scale was $\alpha = .83$.

EPOCH (Kern et al., 2016) – measures five dimensions of adolescent well-being (engagement, perseverance, optimism, connectedness, happiness). The questionnaire consists of 20 items – 4 items for each dimension (e.g., *E: "I get completely absorbed in what I am doing" P: "I finish whatever I begin", O: "In uncertain times, I expect the best", C: "There are people in my life who really care about me", H: "I have a lot of fun"*). Items are rated on a 5-point scale ranging from 1 (almost never) to 5 (almost always). The authors documented high internal consistency ($\alpha = .77 – .83$) and test-retest reliability ($r = .37 – .49$) of the scale among adolescents from USA and Australia. Convergent and divergent validity of the scale was supported by correlations with similar and dissimilar constructs. The questionnaire was translated into Slovak by two independent translators, following back translation. Confirmatory factor analysis supported five-factor solution of the Slovak version of EPOCH (χ^2 (160) = 229.388, $p < .001$; GFI = .87, CFI = .93, RMSEA = .06, CI 90% [.039, .072], SRMR= .066). In the present study, internal consistency of the EPOCH subscales ranged between $\alpha = .73 – .86$ at T1 and between $\alpha = .69 – .87$ at T2.

Results

Preliminary Analyses

Prior to analysis, all the variables were checked for missing data. Since the pattern of missing values across variables was random and did not exceed 5%, cases with missing data were replaced with mean of the given variable (Tabachnik & Fidel, 2001). There were no significant differences in the results of the analysis, when comparing deleting missing cases and imputation of mean values for missing data.

Firstly, the normality of distribution of all variables was calculated. Values of skewness and kurtosis were converted to *z*-scores. School belonging and connectedness were negatively skewed, however, considering the large sample size, we utilized parametric statistics (Tabachnik & Fidel, 2001). Bivariate correlations were calculated to display interrelationships among predictors, mediator and outcome variables and to test the presence of multicollinearity. None of the correlation coefficients exceeded .90, which confirmed that multicollinearity among study variables was not severe. The results of correlation and descriptive analysis are presented in Table 1. Academic and social competence scales displayed significant positive correlations with well-being indicators (first testing – T1) and school belonging, social competence having stronger associations. School belonging showed significant strong positive correlations with well-being (T1), except for a small correlation with engagement and perseverance scales. There were no significant correlations between engagement and perseverance scales at second testing (T2) with perceived competence and school belonging. Optimism, connectedness and happiness at T2 showed significant positive, moderate to strong correlations with social competence and school belonging. The strength of correlation was interpreted in line with Cohen's (1992) recommendation.

To determine gender differences in the study variables, independent samples *t*-tests were calculated. Boys reported higher levels of academic competence than girls ($t(137) = 2.97$, $p = .004$, $d = 0.52$), whereas girls reported higher level of connectedness ($t(137) = -2.16$, $p = .003$, $d = 0.37$) than boys did. The effect size of differ-

Table 1 *Descriptive characteristics, reliability and correlations among study variables in sample (T1) N = 139, T2 in sample n = 109*

	M(SD)	α	1	2	3	4	5	6	7	8	9	10	11	12
1. AC	13.37(2.93)	.70	-											
2. SC	13.91(3.31)	.83	.36**	-										
3. SB	19.37(3.01)	.83	.39**	.73**	-									
4. ET1	12.82(3.11)	.76	.28**	.26**	.22**	-								
5. PT1	13.40(3.13)	.73	.27**	.20**	.27**	.46**	-							
6. OT1	13.50(3.52)	.77	.40**	.53**	.59**	.32**	.34**	-						
7. CT1	16.90(3.00)	.80	.22**	.50**	.50**	.21*	.29**	.38**	-					
8. HT1	14.74(3.66)	.86	.33**	.64**	.70**	.30**	.27**	.64**	.57**	-				
9. ET2	11.81(2.98)	.78	.13	.12	.10	.51**	.36**	.14	.23*	.20*	-			
10. PT2	12.97(2.89)	.69	.17	.11	.05	.28**	.78**	.23*	.30**	.24*	.31**	-		
11. OT2	12.67(3.40)	.77	.12	.34**	.36**	.14	.13	.60**	.42**	.53**	.15	.12	-	
12. CT2	17.04(3.03)	.83	.20*	.53**	.56**	-.02	.17	.14	.71**	.34**	-.02	.05	.38**	-
13. HT2	14.18(3.67)	.87	.19	.46**	.50**	.14	.14	.45**	.54**	.74**	.19*	.13	.65**	.53**

Note. AC = academic competence, SC = social competence, SB = school belonging, E = engagement, P = perseverance, O = optimism, C = connectedness, H = happiness
$^{*}p < .05$, $^{**}p < .01$

ences was medium and small, respectively (Cohen, 1992). There were no statistically significant gender differences in other tested variables.

Mediation Analyses at T1

In order to examine the mediating role of school belonging (M) in the relationship between perceived social and academic competence (IV) and indicators of well-being (DV: engagement, perseverance, optimism, connectedness, and happiness), the mediation analysis was performed. Also, multiple regression analysis was utilized (Baron & Kenny, 1986). Direct (effect of IV on DV controlling for M) and indirect effect (effect of IV on DV through M) were calculated using the statistical program called PROCESS developed by Hayes (2013). We tested the significance of the indirect effect using the Sobel test (*Z*) and bootstrapping procedure. Standardized indirect effect was calculated for each 10 000 bootstrapped samples and the 95% bootstrapped confidence interval was computed. The effect size was calculated as a proportion of variance in dependent variable, explained by the indirect effect (R^2_{med}) (Hayes, 2013).

Overall, 10 mediational models were tested with well-being indicators as dependent variables at Time 1. We did not perform mediation analysis for engagement and perseverance at Time 2 and for academic competence (IV) because the correlations between these and other tested variables were too small and non-significant. Three mediational models with well-being indicators as dependent variables at Time 2 were performed. Results of the mediation analysis (unstandardized total, direct and indirect effects with bootstrapped 95% confidence intervals and effect size) are presented in Table 2. Significant mediational models expressed in standardized regression coefficients with total (c), direct (c') and indirect (ab) effects are presented in Figures 1 to 3.

Results of mediational analysis showed that there is only significant direct effect of academic competence on engagement and perseverance at Time 1. School belonging did not mediate the relationship between academic competence and engagement and perseverance, respectively ($Z = 1.30, p = .194$, resp. $Z = 1.99, p = .051$). This was also revealed by 95% bootstrap confidence interval that contained zero or was only slightly above zero. Moreover, R^2_{med} indicates that only 3%/4% of the variance in engagement/perseverance respectively, is attributable to the indirect effect of academic competence through school belonging.

There was a significant indirect effect of academic competence on optimism (0.24, $Z = 3.45$, $p < .001$), connectedness (0.19, $Z = 3.30$, $p = .001$) and happiness (0.32, $Z = 4.25$, $p < .001$) at Time 1, as confirmed also by a 95% bootstrap confidence interval that was entirely above zero. Given that direct effect of academic competence on optimism after controlling for mediator remained statistically significant, mediating role of school belonging was partial. The R^2_{med} indicates that 13% of the variance in optimism is attributable to the indirect effect of the academic competence through school belonging. The direct effect of academic competence on connectedness and happiness was no longer statistically significant, meaning that school belonging fully mediated the relationship between academic competence and connectedness, and happiness, respectively. The R^2_{med} indicates that 5%/10% of the variance in connectedness/happiness is attributable to the indirect effect of the academic competence through school belonging.

Next, mediational models with social competence as the independent variable were tested. Results showed no direct or indirect effect of social competence on engagement and at Time 1. There was a significant indirect effect of social competence on perseverance (0.18, $Z = 2.28$, $p = .023$). However, 95% bootstrap confidence interval was only slightly above zero and the R^2_{med} indicates that only 4% of the variance in perseverance is attributable to the indirect effect of the social competence through school belonging. There was a significant indirect effect of social competence on optimism (0.34, $Z = 4.07, p < .001$) connectedness (0.20, $Z = 2.42$, $p = .016$) and happiness (0.40, $Z = 4.29, p < .001$) at Time 1, as revealed also by a 95% bootstrap confidence interval that was entirely above zero. Given that direct effect of social competence on optimism, connectedness and happiness remained statistically significant, the mediating role of school belonging was in all cases partial. The R^2_{med} indicates that 26%/21%/37% of the variance in optimism/connectedness/happiness, respectively, is attributable to the indirect effect of the social competence through school belonging.

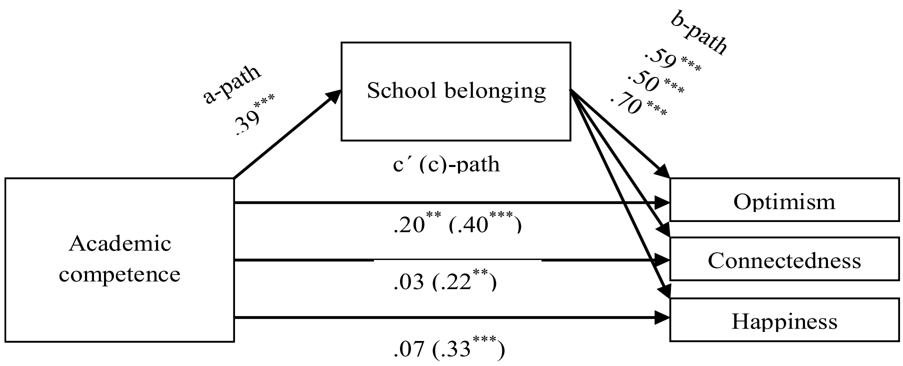

Figure 1 Mediating model of the relationship between academic competence and optimism, connectedness and happiness, respectively, with school belonging as mediator

Mediation analyses at T2

Finally, mediational models with well-being indicators at Time 2 as dependent variables were tested. There was no significant direct or indirect effect of social competence on optimism at Time 2. However, the R^2_{med} indicates that 10% of the variance in optimism is attributable to the indirect effect of the social competence through school belonging.

There were significant indirect effects of social competence on connectedness (0.24, Z = 3.01, p = .003) and happiness (0.29, Z = 2.87, p = .004), as revealed also by a 95% bootstrap confidence interval that was entirely above zero. Given that direct effect of social competence on connectedness remained statistically significant, the mediating role of school belonging was partial. The R^2_{med} indicates that 25% of the variance in connectedness is attributable to the indirect effect of the social competence through school belonging. The direct effect of social competence on happiness was, after controlling for the mediator, no longer statistically significant, meaning that school belonging fully mediated the relationship between social competence and happiness at Time 2. The R^2_{med} indicates that 19% of the variance in happiness is attributable to the indirect effect of the social competence through school belonging.

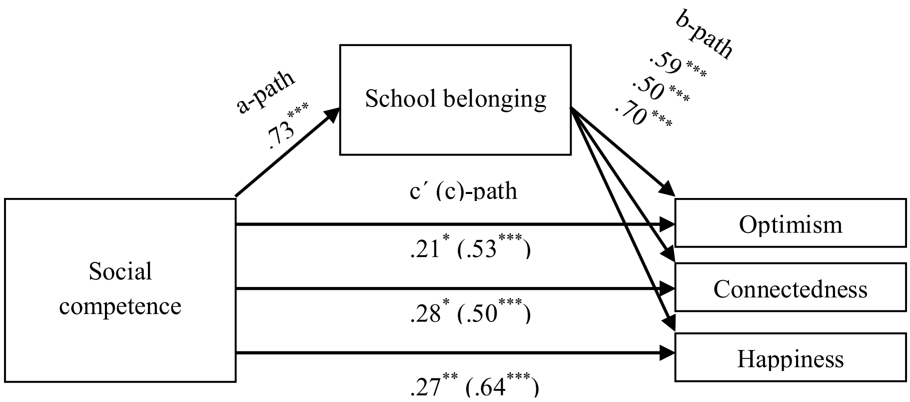

Figure 2 Mediating model of the relationship between social competence and optimism, connectedness and happiness, respectively, with school belonging as mediator

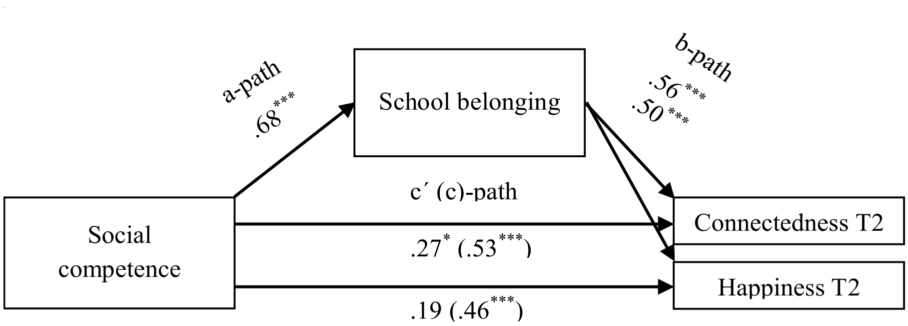

Figure 3 Mediating model of the relationship between social competence and connectedness at Time 2 and happiness at Time 2, respectively, with school belonging as mediator

Table 2 *Unstandardized direct, indirect and total effects of all tested mediational analysis*

Predictor	Mediator	Criterion	Effects [95% CI]		Effect size R^2_med
Academic competence	School belonging	Engagement	Direct Indirect Total	0.25^{*} [0.05, 0.44] 0.05 [-0.02, 0.14] 0.30^{**} [0.12, 0.47]	.03
Academic competence	School belonging	Perseverance	Direct Indirect Total	0.21^{*} [0.01, 0.41] 0.08 [0.01, 0.17] 0.29^{**} [0.11, 0.47]	.04
Academic competence	School belonging	Optimism	Direct Indirect Total	0.24^{*} [0.05, 0.43] 0.24^{***} [0.14, 0.37] 0.48^{***} [0.29, 0.67]	.13
Academic competence	School belonging	Connectedness	Direct Indirect Total	0.03 [-0.15, 0.21] 0.19^{**} [0.10, 0.32] 0.23^{*} [0.06, 0.40]	.05
Academic competence	School belonging	Happiness	Direct Indirect Total	0.09 [-0.08, 0.25] 0.32^{***} [0.19, 0.46] 0.41^{***} [0.21, 0.62]	.10
Social competence	School belonging	Engagement	Direct Indirect Total	0.20 [-0.04, 0.45] 0.04 [-0.12, 0.24] 0.24^{*} [0.10, 0.39]	.05
Social competence	School belonging	Perseverance	Direct Indirect Total	0.01 [-0.22, 0.22] 0.18^{*} [0.04, 0.37] 0.19^{*} [0.04, 0.34]	.04
Social competence	School belonging	Optimism	Direct Indirect Total	0.22^{*} [0.04, 0.40] 0.34^{***} [0.20, 0.52] 0.56^{***} [0.39, 0.73]	.26
Social competence	School belonging	Connectedness	Direct Indirect Total	0.25^{*} [0.05, 0.45] 0.20^{*} [0.06, 0.36] 0.45^{***} [0.30, 0.60]	.21
Social competence	School belonging	Happiness	Direct Indirect Total	0.30^{**} [0.09, 0.52] 0.40^{***} [0.22, 0.57] 0.70^{***} [0.56, 0.85]	.37
Social competence	School belonging	OptimismT2	Direct Indirect Total	0.16 [-0.11, 0.44] 0.18 [-0.03, 0.42] 0.35^{**}[0.12, 0.57]	.10

Table 2 continues

Table 2 continued

Predictor	Mediator	Criterion	Effects [95% CI]		Effect size R^2_med
Social competence	School belonging	ConnectedT2	Direct	0.24^{*} [0.04, 0.45]	.25
			Indirect	0.24^{**} [0.07, 0.43]	
			Total	0.49^{***}[0.28, 0.69]	
Social competence	School belonging	HappinessT2	Direct	0.21 [-0.13, 0.55]	.19
			Indirect	0.29^{**} [0.07, 0.58]	
			Total	0.50^{***}[0.28, 0.72]	

$^{*}p < .05$, $^{**}p < .01$, $^{***}p < .001$

Discussion

The aim of the study was to explore the direct and indirect (through school belonging) pathways from perceived social and academic competence to various indicators of well-being among students shortly after the transition to high school. We tested whether school belonging mediates the relationship between self-perception and well-being indicators concurrently and prospectively six months later.

The school belonging turned out to be an important mechanism linking perceived self-competence and well-being indicators. School belonging mediated the relationship between academic competence and well-being indicators at Time 1. After accounting for school belonging, relationships of academic competence and connectedness, optimism and happiness decreased substantially and became nonsignificant (except for optimism). The results suggested that perceived higher cognitive competence applied in schoolwork is associated with higher belonging to school, which in turn leads to greater general sense of connectedness, optimism and happiness in adolescents. Our results are similar to Phan et al. (2016), who documented the mediational role of engagement in the relationship between academic self-efficacy and well-being at school. Our research extends previous findings, suggesting that school belonging could be a link between academic competence and psychological adjustment of adolescents in non-educational settings. However, it should be mentioned that the mediational effect accounted only for 5-13% of connectedness variance (Fairchild, MacKinnon, Taborga, & Taylor, 2009).

School belonging did not mediate the relationship between academic competence and engagement and perseverance, respectively. When controlling for school belonging, academic competence directly predicted students' engagement, perseverance and optimism at Time 1. The results suggest that perceived higher intelligence and cognitive abilities are directly related to higher engagement and perseverance, independent of the sense of school belonging. Our findings indicate that to promote student engagement and perseverance, cognitive resources, particularly perception of academic competence, should be emphasized. These findings correspond with the previous research of Mih and Mih (2013) documenting positive relationships between academic self-concept and engagement in learning, persistence in difficult tasks and enjoyment of academic work. Similarly, Phan et al. (2016) determined that academic self-efficacy positively influenced academic engagement in secondary school students. Our results add to previous research documenting direct concurrent association of academic competence to engagement, perseverance and optimism in a non-learning context. However, our results also suggested that a significant association between academic competence and engagement, perseverance and optimism exists in young people only concurrently. The positive effect of perceived academic competence shortly after the transition (directly or through school belonging) diminishes over time.

School belonging mediated relationships between perceived social competence and well-being indicators at Time 1. The results suggested that higher perceived competence applied in social context enhance the students' sense of belonging to school, which in turn heightens their general capacity for optimism, connectedness and happiness. The effect size of mediation suggested that considerable amount of variance in optimism/connectedness/happiness was attributable to the indirect effect (Fairchild et al., 2009), suggesting that school belonging in the transition period plays a crucial role in promoting adolescent positive functioning, concurrently. Similar patterns were found in prospective relationships with well-being indicators six months later. School belonging fully mediated (except for connectedness) the relation between social competence and optimism, connectedness and happiness. Our results suggest that students who perceive themselves as socially competent shortly after the transition tend to feel greater attachment to school and school members. This in turn may support later development of relatedness to other people, hopefulness and positive emotions in general. These results are in accordance with findings of Lester et al. (2013), who confirmed significant links between school belongingness measured at the beginning of the year with decreased depression at the end of the school year. Confirming the Stage-Environment Fit Theory (Eccles et al., 1993) and Expectancy Value Theory (Eccles & Wigfield, 1995), findings indicate that the experience of acceptance and support in the context of school, as a result of an individual's perception of competence, may lead to further enhancing of adolescents' general capacity to connect to other people, to be optimistic and happy, regardless of the context.

In contrast to academic competence, social competence was directly related to optimism, happiness, and later connectedness, but not to engagement and perseverance. Our results are in line with previous research documenting stronger associations between social competence and psychological adjustment, whereas academic competence is more saliently related to achievement related outcomes (Kern et al., 2016; Olsson et al., 2012; Phan et al., 2016). Engagement and perseverance are connected with individual ability and strength and are thus more reflected in perceived academic competence. Social competence is applied in interpersonal context and thus promotes sense of connectedness and feelings of happiness.

However, it should be mentioned that association between academic and social competence and indicators of well-being was confirmed only concurrently and the relation decreased substantially after controlling for school belonging. The results suggested that, to enhance student's optimism, connectedness and happiness, developing students' cognitive resources might not be a sufficient condition. Moreover, as the tested prospective relations have suggested, to foster happy students with an optimistic view and positive relationships with others, intervention efforts should focus on individual feelings of school belonging during the transition period. Our findings confirm the belongingness hypothesis (Baumeister & Leary, 1995), stating that although attachments to parents and positive relationships with friends are important for an individual's adjustment, those who do not have a sense of connection to a larger group or community will likely experience increased stress and emotional distress. For adolescents, school is the first source of experience of connection to a larger group. Our results indicate that a developed sense of connection to school in the period of transition may promote later overall well-being in adolescents, especially socially grounded indicators of youth well-being. Therefore, the transition to high school is an important period for enhancing youth mental health by providing social and contextual support (Lester et al., 2013).

The present study has several limitations. First, only self-report questionnaires were utilized. Results can be biased by self-evaluation of the students. Secondly, the sample consisted of students from high schools in one Slovak city, which limits generalizability of the results to other types of high schools and regions. Thirdly, we examined the level of self-perception and school belonging at only one time point, which did not allow us to test changes in cognitive and social factors and their predictive effect on students' well-being. However, our research enables us to identify crucial factors, acting in the transition period, that predict actual and prospective students' well-being. Moreover, other research has suggested that school belonging across the course of high school remains remarkably stable (O'Neel & Fuligni, 2013) and its level in the beginning of the freshmen year has strong causal path leading to students' mental health in the following years (Lester et al., 2013).

Fourth, prospective relations were examined only over a six-month period using 109 respondents. The students that were absent and students lost to follow-up in the second measurement (approximately 21% of the initial sample) may have biased the results. A further longitudinal study should test respondents repeatedly during a longer period. A longitudinal study could also determine causal direction between study variables, which cannot be clearly addressed by the present study. For example, based on correlational analysis, it is unclear whether higher academic and social competence leads to better school belonging or stronger sense of school belonging leads to more positive perceptions of one's learning skills and interpersonal competence at school (Pittman & Richmond, 2007).

Fifth, despite the fact that there were no significant gender differences in the studied variables, it would be plausible to test whether the direct and indirect effect of self-competence on well-being indicators might be moderated by gender. As the present study is a part of a larger longitudinal project and the article has limited length, we aim to test whether school belonging may function differently for male than female students and younger adolescents in further research. Finally, future research should examine several domains of connectedness and other individual characteristics (e.g., self-esteem, self-efficacy, motivation) in the prediction of adolescent positive functioning.

Despite the limitation, the strength of the present study is in examining multiple factors, acting during the transition period, that can promote indicators of adolescent well-being, concurrently and prospectively six months later. So far, only a small number of studies have investigated direct and indirect effects of school-related cognitive factors (both academic and social self-competence) through social sources on positive indicators of youth development in other than the academic domain. The implication of our findings is that context unspecified optimism, connectedness and happiness may be improved indirectly and prospectively (at least six months later) by fostering positive relations and emotions within the school. Satisfying developmental need for relatedness in the context of transitioning to high school, as a result of positive self-perception, seems to provide adolescents with a substantial basis for feeling positive about themselves, their future and other people. In contrast, students' individual ability to persevere and to be engaged in life tasks can be directly promoted by emphasizing their cognitive resources, particularly their academic competence as perceived shortly after the transition to high school.

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The Moderating Effect of Hardiness on the Relationships between Problem-Solving Skills and Perceived Stress with Suicidal Ideation in Nursing Students

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Recent evidence indicates an elevated risk of suicidal ideation for undergraduate nursing students. This research was designed to enhance the understanding of suicidal ideation in nursing students by investigating the relationships between problem-solving skills, perceived stress, hardiness, and suicidal ideation, with the possibility of hardiness acting as a moderator in the relationships between problem-solving skills appraisal and perceived stress with suicidal ideation. A multi-stage cluster random sample of Malaysian nursing undergraduate students (N = 204) completed self-report questionnaires. The results of structural equation modeling revealed that poor problem-solving skills, greater levels of perceived stress, and low levels of hardiness predicted greater levels of suicidal ideation. Also, hardiness emerged as a moderator in the links between problem-solving skills appraisal and perceived stress with suicidal ideation. The findings incrementally improve our understanding about the importance of hardiness as a moderator in explaining how problem-solving skills and perceived stress affect suicidal ideation. The results of this study are obtained from Malaysian nursing students and possible generalization to other populations should be verified by further studies.

Key words: problem-solving skills appraisal, perceived stress, hardiness, suicidal ideation

Suicide has become a serious and negative occurrence throughout the world, and it has been identified as the second leading cause of death in the individuals aged 15 to 25 years old (World Health Organization, 2012). Research findings have shown that the prevalence of suicide in nurses is higher than other medical occupations (Hawton, Agerbo, Simkin, Platt, & Mellanby, 2011). One possible explanation for the higher suicide risk in nurses, compared with other medical occupations, is that their working conditions and the nature of the job make them vulnerable to burnout, anxiety, depression, and suicidal ideation (Leal & Santos, 2015). Also, nursing students may be at greater suicide risk than other university students because in addition to theoretically based class sessions, they are also exposed to stressful clinical attachment. Hence, there is a need to learn to cope with suffering and death (Pulido-Martos, Augusto-Landa, & Lopez-Zafra, 2012). This highlights the idea that undergraduate nursing students are at higher suicide risk than other

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university students. Therefore, it is important to understand suicidal ideation in nursing students when planning prevention programs to acquire better outcomes.

Maniam et al. (2014) revealed that Malaysian young adults had the highest risk for suicidal behavior compared to other age groups. There are series of significant cognitive shifts, physiological changes, and emotional changes that place young adults at risk for suicidal ideation (Maddi, 1989). The period of young adulthood is a period of rapid reorganization in all developmental aspects, and the effects of these developmental changes may influence young adults when making reasonable decisions concerning stressful life conditions, and such characteristics may contribute to their suicidal ideation (Wagner, 2012). Given such characteristics during this period, ineffective problem-solving skills may contribute to their self-destructive behavior.

Problem-solving skills have been shown to play a central role in the mental health of nursing students (Abdollahi, Talib, Yaacob, & Ismail, 2014; Campos et al., 2016). Heppner and colleagues (2004) highlighted the role of poor problem-solving skills in clarifying psychological disorders, such as depression and suicide. They coined the term *problem-solving skills appraisal,* and have defined it as one's capacity and skill to deal with problems throughout his/her life. The authors identified three styles of problem-solving skills: problem-solving confidence, approach-avoidance style, and personal control. Problem-solving confidence is defined as an estimation of one's opinion about one's own capacity to achieve goals and what is necessary to reach the final goal (Heppner et al., 2004). In their work, the authors explained that some individuals could not make a decision when encountering problems, and they might withdraw from forthright confrontation, preferring to avoid their problems, or they might prefer straight confrontation of their problems and provide responses to the problems (approach-avoidance style). The third style is personal control, which is similar to the theoretical conceptualization of control created by Rotter (1966). Heppner and colleagues (2004) explained that when individuals encounter problems, they may perceive that they have abilities to control their emotional responses during the stressful situations (personal control). Numerous studies have suggested that ineffective problem-solving skills serve as a significant correlate of suicidal ideation (Abdollahi, Talib, Yaacob, & Ismail, 2015a, 2015b). However, to the best of our knowledge, no study has investigated the relationship between problem-solving skills and suicidal ideation among Malaysian nursing students.

Stress has been considered a hindrance to concentration, problem-solving skills, and mental health in nursing students (Reeve, Shumaker, Yearwood, Crowell, & Riley, 2012). Sources of stress affecting nursing students, such as economic problems, frequent examinations, increased separation from family, adjusting to an new academic environment and culture as well as perceived pressures for academic achievement and increased workloads are commonly reported sources of nursing student stress (Jameson, 2014). Such demands create a stressful developmental period for nursing students, which may be harmful to them. In the present study, perceived stress was the degree to which a circumstance in one's life is estimated to be stressful (Cohen, 1996). Previous studies also revealed that nursing students experienced high levels of stress (Abdollahi, Abu Talib, Yaacob, & Ismail, 2014). According to Lazarus and Folkman's (1984) cognitive-transactional stress theory, appraisal and interpretation of the event is more important than the event itself. Several studies have found a direct relationship between stress and suicidal ideation among nurses (Aradilla-Herrero, Tomás-Sábado, & Gómez-Benito, 2014; Wang, Lai, Hsu, & Hsu, 2011); however, the relationship between stress and suicidal ideation among Malaysian nursing student has not been investigated.

According to Maddi and colleagues (2012), psychological hardiness as a secondary appraisal is composed of three interacting attitudes, labeled as follows: challenge – an attitude that continuing to learn from experience is more fulfilling than expecting comfort and security; control – an attitude that the struggle to influence outcomes is more advantageous than passivity and powerlessness; and commitment – an attitude that involvement is more advantageous than detachment. Maddi (2006) believes that hardiness enables individuals to feel committed, to have a greater sense of emotional control under stressful situation.

There is still a gap in the literature according to a great deal of evidence of the associations between poor problem-solving skills appraisal and perceived stress with suicidal ideation. It is unclear why some individuals with effective problem-solving skills and low levels of perceived stress experience greater amounts of suicidal ideation. There are two reasons to support the moderating role of hardiness: individuals with poor problem-solving skills are less likely to think of suicide, because hardiness enables them to handle their emotions, allows for commitment to fruitful social activities, challenges them with their stressful conditions, and provides flexibility by turning negative conditions into opportunities to grow and gain wisdom (Maddi, 2006). Hardy individuals are also less likely to evaluate stressful circumstances as frightening and unmanageable, and they appraise stressful situations as a challenge for acquiring new experiences even under stressful conditions. Therefore, they are less likely to suffer from suicidal thoughts (Abdollahi, Talib, Yaacob, & Ismail, 2015c; Jameson, 2014). Most prior studies have generally investigated linear associations between problem-solving skills and perceived stress with suicidal ideation in nursing students (Leal & Santos, 2015), but have neglected the possible moderating influence of hardiness on these associations. Therefore, the precise mechanism that accounts for problem-solving skills and perceived stress with suicidal ideation remains to be delineated. Consequently, this study is designed to examine the possible moderating effect of hardiness in the links between problem-solving skills appraisal and perceived stress in relation to suicidal ideation among Malaysian nursing undergraduates.

Therefore, the present study hypothesized as follows: 1) poor problem-solving skills and high levels of perceived stress will positively predict suicidal ideation, and 2) hardiness will moderate the links between problem-solving skills appraisal and perceived stress with suicidal ideation among nursing students.

Method

Participants

A total of 214 nursing students were recruited via a multi-stage cluster sampling approach to participate in the study. Of these, four cases were omitted from the analysis due to incomplete information and 6 cases were removed from the analysis due to outlier values. Finally, 204 nursing students were involved in the present analysis.

Procedure

Ethical endorsement was granted by the Ethical Research Committee of University Putra Malaysia (UPM) before research beginning. Before collecting data, permission to conduct research among nursing students was obtained from the universities involved. In the first stage, from the four public universities in Selangor state, two universities were selected randomly (Universiti Putra Malaysia and Universiti

Kebangsaan Malaysia). In the next stage, a class from each nursing school according to the students' grade year (freshman, sophomore, junior, and senior) was randomly selected, and data were collected during one of the regularly scheduled classes. Data were collected from January 13, 2014 to April 15, 2014. All the participants were informed regarding the research goals, and signed the consent form. They were also guaranteed confidentiality and their rights to withdraw from the study at any time. The respondents completed the self-reported questionnaires.

Measures

The 32-item Problem-Solving Inventory (PSI) (Heppner, 1988) was used to assess one's capacity and skill to deal with hardships in one's daily life (Heppner, 1988). A 6-point Likert scale was used for all questions. This measure consists of three factors, as follows: Problem-Solving Confidence (PSC), Approach-Avoidance Style (AAS), and Personal Control (PC). Heppner (1988) suggests that the factors are dependent; therefore, in this study, the sum of three factors was used. A lower score in PSI shows effective problem-solving skills. Previous studies revealed that PSI had good validity and acceptable internal consistency, with $\alpha = 0.80$ (Heppner, 1988; Heppner et al., 2004).

The 10-item Perceived Stress Scale (Cohen, Kamarck, & Mermelstein, 1983) was used to assess the degree of stressful circumstances and measure the opinions of nursing students about how circumstances are unpredictable, uncontrollable, and excessive. A 4-point Likert scale was used for all questions, with total scores ranging from 0 to 40. A higher score indicates higher levels of perceived stress. Cohen (1996) reported perceived stress scale had a good internal consistency, with $\alpha = 0.84$.

The 18-item Personal Views Survey II (Maddi et al., 2006) was used to assess one's ability and skill to handle stressful life events. Each item was rated on a 4-point Likert scale, ranging from 0 (not at all true) to 3 (very true). The sum of these three components indicates degree of hardiness (Maddi, 2006). The total score range is from 0 to 54. A higher score indicates greater hardiness. The study showed an acceptable internal consistency (Maddi et al., 2006). The hardiness literature supports the sum of the subscales of the hardiness components – commitment, control, and challenge (Cole, Feild, & Harris, 2004; Maddi, 2006). Thus, the sum of three scores was utilized in this study.

The 21-item Beck Scale for Suicidal Ideation (BSSI) (Beck, Steer, & Ranieri, 1988) was used to measure suicidal ideation, planning, and intent to commit suicide in the past week. Each item was rated on a 3-point Likert scale from 0 to 2 for all questions, with total scores ranging from 0 to 38 (Beck et al., 1988). A high BSSI score means that the respondent is at higher risk of suicide. The study confirmed concurrent validity between individuals with a high BSSI score and experience of suicide attempts (Beck et al., 1988).

Demographic information was collected to measure different features of the nursing students' backgrounds. Respondents completed a demographic survey about their gender, race, age, educational level, and marital status.

Results of Pilot Study

The questionnaires were pilot tested by convenience sampling, in which 30 nursing students were chosen. Cronbach's alpha was used for evaluating the reliability of the questionnaires (English version). The Problem-Solving Inventory had a Cronbach's alpha value of $\alpha = 0.86$, the Perceived Stress Scale had a Cronbach's Alpha value of $\alpha = 0.81$, the Hardiness Scale had a Cronbach's alpha value of $\alpha = 0.78$, and the Beck Suicidal Ideation Scale had a Cronbach's alpha value of $\alpha = 0.90$. Therefore, the reliability coefficients showed that all the questionnaires were stable and consistent.

Data Analysis and Data Preparation

Structural Equation Modeling (SEM), using AMOS version 23 (Analysis of Moment Structures), and SPSS version 23 (Statistical Package for the Social Sciences) were employed to analyze the data.

In the process of data screening, the missing data, outliers, and normality of distribution were checked and addressed. Missing data were addressed using the regression imputation method. Outliers were checked by high Mahalanobis *d*-squared with both *p1* and *p2* equal .000 and .000; 6 cases were found as outliers, and these cases were deleted. For all variables, skewness values ranged from 0.003 to 0.598, and kurtosis values ranged from -1.038 to 1.625; on that basis, the distribution was considered normal [As a rule of thumb, Byrne (2010) described that when skewness and kurtosis values are lower than ±3 and ±5, respectively, data are appropriate to assume normality].

According to Kline's (2010) recommendation, the following 6 indices were employed to assess the measurement model fit: the chi square/degree of freedom ratio (CMIN/DF), the Goodness-of Fit Index (GFI), the Comparative Fit Index (CFI), the Tucker-Lewis Index (TLI), and the Incremental Fit Index (IFI). A rule of thumb for the fit indices is that values equal or greater than .90 are an acceptable fit (Kline, 2010). Furthermore, the model may be classified as acceptable, if the Root Mean Squared Error of Approximation (RMSEA) is between .03 and .08.

Results

Respondent Profile

The age range of the respondents was from 18 to 24 years old (mean age = 21. 51, *SD* = 3.21). The sample consisted of 98 (48%) males and 106 (52%) females, and the majority of respondents (86.4%, *n* = 177) were single. Nearly half of the respondents were Malay (48%, *n* = 98), followed by Chinese (41%, *n* = 83), and Indian (11%, *n* = 22). Of the 204 nursing students, 30.2% (*n* = 61) were in their freshman year; 22.8% (*n* = 47) were in their sophomore year; 25.6% (*n* = 52) were in their junior year, and 21.4% (*n* = 44) were in their senior year.

Convergent Validity and Construct Reliability of the Instruments

The convergent validity (Average Variance Extracted; AVE) and construct reliability (Construct Reliability; CR) of the instruments (English version) were assessed in the main study. The establishment of convergent validity and construct reliability are confirmed if AVE and CR are above the recommended standard values of .50 and .70, respectively (Byrne, 2010). As seen in Table 1, the values of AVE and CR showed good internal consistency values

Table 1 *Values of Average Variance Extracted (AVE) and Construct Reliability (CR)*

Measure	AVA	CR
Problem-Solving Inventory	.71	.86
Perceived Stress Scale	.62	.76
Personal Views Survey	.59	.73
Beck Scale for Suicidal Ideation	.79	.85

Note. AVE: Average Variance Extracted and CR: Construct Reliability and convergent validity values for the instruments.

Descriptive Statistic

Table 2 shows the mean, standard deviation, actual range, and inter-correlations of the studied variables. Table 2 shows that poor problem-solving skills (r= 69, p < .01) and perceived stress (r = 53, p < .01) were positively correlated with suicidal ideation, whereas hardiness (r = -.57, p < .01) was negatively correlated with suicidal ideation. Additionally, hardiness was negatively correlated with perceived stress (r=-.69, p<.01) and problem-solving skills (r=-.63, p<.01).

Measurement Model of Study

The results showed that the measurement model fit the data well: CMIN/DF = 1.78, p < .05, GFI = .91, CFI = .92, TLI = .94, IFI = .90, and RMSEA = .04. Although the model based on significant chi-square values did not fit the data, other indices depicted good fit indices. Therefore, according to Kline (2010), the measurement model fitted the data perfectly.

Structural Model

The results, as demonstrated in Figure 1, indicated that poor problem-solving skills appraisal (β = .42, p < .01), high levels of perceived stress (β = .31, p < .01), and low levels of hardiness (β = -.32, p < .01) had significant contributions in prediction of suicidal ideation. These exogenous variables explain 71.0% of the variance in suicidal ideation.

Moderation Test of Hardiness

To investigate the moderating role of hardiness on the associations between problem-solving skills and perceived stress with suicidal ideation, multi-group analysis in AMOS (Version 23) was performed. The first step in examining the moderating role of hardiness on the associations of problem-solving skills, perceived stress, and suicidal ideation is to compare the unconstrained model (paths free to vary) versus the constrained model (paths constrained to equality). If the unconstrained model is better than the constrained model, we can conclude that hardiness has a moderation effect on the overall model. The next step in examining the moderating role of hardiness is to check for the significance of individual paths. The path is moderated by the moderating variable if 1) the beta for one group is significant, while the beta value for another group is not significant, or 2) the beta values for all groups are significant, but the correlation coefficient for the association between two variables for one group is positive and for another group is negative (Kline, 2010).

Table 2 *Correlation among study variables, and the mean, standard deviations, and observed range*

Variables	-1	-2	-3	-4
(1) Suicidal Ideation	1			
(2) Perceived Stress	.53**	1		
(3) Hardiness	-.57**	-.69**	1	
(4) Problem-Solving Skills Appraisal	.69**	.62**	-.63 **	1
Mean	5.43	17.34	29.71	81.68
Standard deviation	4.02	8.79	12.42	21.59
Observed range	0-27	4-36	10-54	32-154

Note. ** p < .01.

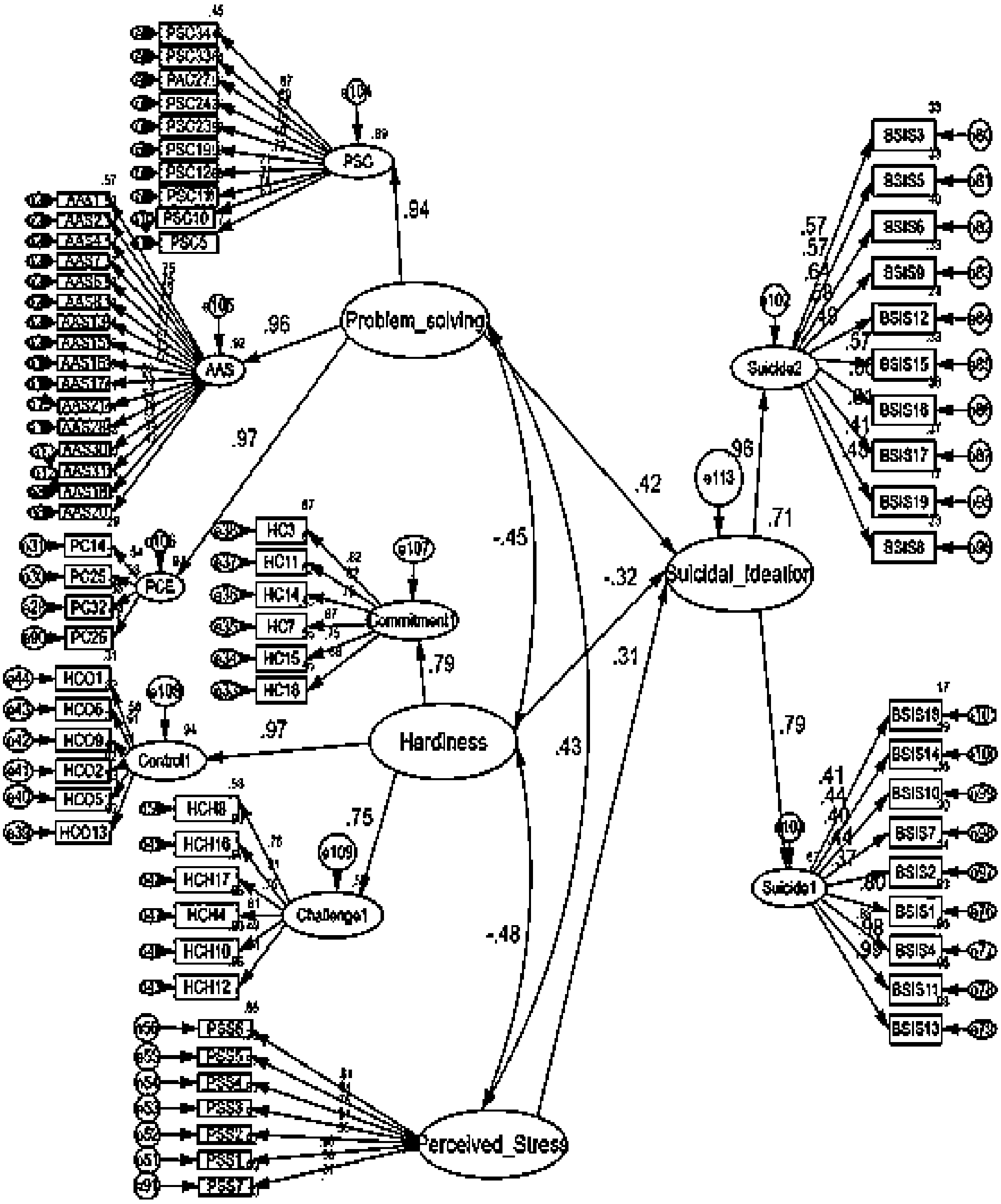

Figure 1 Structural model of suicidal ideation

Table 3 *Standardized regression weights (hardiness variant model)*

Hypothesis			Hardiness	
			Std Est	CR(SE)
Suicidal ideation	<---	Perceived stress	-.312**(.077)	-1.843(.653)
Suicidal ideation	<---	Problem-solving skills	.224*(.096)	.987(.704)

Note. SE = standard error, CR = critical ratio, Std Est = standard estimate.
Result for the low hardiness group is presented first, and results for the high hardiness group is presented in parenthesis.
*p < .05, ** p < .01

To test the moderating role of hardiness, the sample data were divided into high and low groups based on the median split in hardiness (29.71). The measurement fit indices for the unconstrained model (χ^2 = 4.212, p < 0.001, RMSEA = .041, CFI = .911, GFI = .933, NFI = .921) were better than the measurement fit indices for the constrained model (χ^2 = 5.133 p < 0.001, RMSEA = .065, CFI = .897, GFI = .880, NFI = .833), because the chi-square value for the unconstrained model was smaller than the chi-square value for the constrained model, and other fit indices for the unconstrained model were greater than the measurement fit indices of the constrained model (Kline, 2010). Therefore, there is a considerable difference in the influence of hardiness on the relationships between problem-solving skills and perceived stress with suicidal ideation.

The result, as demonstrated in Table 3, revealed that hardiness moderated the association of perceived stress and suicidal ideation; that is, the nursing students from the high hardiness group were less likely to experience suicidal ideation, regardless of their perceived stress. The moderating role of hardiness on the link between problem-solving skills and suicidal ideation was also supported; specifically, nursing students with high levels of hardiness were less likely to experience suicidal ideation, regardless of their ineffective problem-solving skills.

Discussion

The first hypothesis of the present study involved examining the relationships between problem-solving skills and perceived stress with suicidal ideation among nursing students. As anticipated, our findings provide support for positive associations between poor problem-solving skills and high levels of perceived stress with suicidal ideation. One possible explanation for the positive association between poor problem-solving skills and suicidal ideation is that individuals, who tend to use less effective problem-solving skills are more likely to engage in withdrawal, decreased assertiveness, sadness, irritability, and passivity that contribute to the risk for suicidal ideation. On the other hand, individuals with effective problem-solving skills are able to regulate their own emotions and neutralize negative thoughts, and are less likely to suffer from suicidal ideation (Abdollahi et al., 2015b; Khan, Hamdan, Ahmad, Mustaffa, & Mahalle, 2016; Siu, 2009). This finding is consistent with previous studies (Becker-Weidman et al., 2010) that found that ineffective problem-solving skills have a great influence on increased suicidality. This idea emanates from studies that have demonstrated that effective problem-solving skills reduced the risk of suicidal ideation (Speckens & Hawton, 2005). Given the chemical, physical, and psychological changes during puberty (Heim & Binder, 2012), undergraduate students need to find a new role in society and take decisions about academic, occupational, and marital affairs, which affect their future and their identities. Therefore, these demands mean that problem-solving skills are necessary for them, because they need to find logical solutions for their problems; otherwise, they may be suffering from suicidal ideation (Becker-Weidman et al., 2010).

This finding also showed that perceived stress positively predicted suicidal ideation among nursing students. This finding is in agreement with Abdollahi and colleagues (2015c), who found that stress was positively correlated with suicidal ideation. Suicidal behavior is an outcome of perceived stress, meaning that one cannot bear the stressful conditions and thinks that suicide is the best solution for escaping an unbearable situation. According to the cognitive-transactional stress theory (Lazarus & Folkman, 1984), perceived stress is produced by the perception of stressful situations and estimation of one's own ability to manage the stressors in which the effects on the individuals are more than that of the real stressors. This theory suggests that there are diversities in everyone's reaction to the same potential stressors, and these differences influence both the level of appraisal and the ability to solve problems.

The second hypothesis of this study involved examining the moderating role of hardiness on the links between problem-solving skills and perceived stress with suicidal ideation among nursing students. The present findings showed that hardiness moderated the link between problem-solving skills appraisal and suicidal ideation. That is, nursing students with high levels of hardiness are more likely to apply hardy attitudes and courage to confront problems, and they are more flexible in facing problems. The hardiness theory (Maddi, 2006) displays two types of appraisals in relation to suicidality. First, the situation is assessed by problem-solving skills as a primary appraisal. Second, hardiness as a secondary appraisal plays an important role in the relationship between problem-solving skills and suicidal ideation. This means that when the situation is assessed as a threatening and uncontrollable situation, the likelihood of suicidal ideation increased. However, when the situation is assessed as a challengeable or opportunity for learning, hardiness acts as an obstacle to engaging in suicidal ideation, meaning that hardy individuals were less likely to experience suicidal ideation even at the highest levels of ineffective problem-solving skills. The findings of this study support previous research showing that positive secondary appraisals may reduce the risk of suicidal ideation (Johnson, Gooding, Wood, & Tarrier, 2010), proving that hardiness is an important element in completing tasks and adapting to challenging situations.

The findings also revealed that hardiness acted as a significant moderator in the link between perceived stress and suicidal ideation. The findings showed that one of the influencing factors against stress is hardiness and that nursing students with strong hardiness, even those having high levels of perceived stress, were less likely to experience suicidal ideation. The findings further confirm the theorists' claims by explaining that hardiness provides courage, motivation, and effective coping strategies to enhance performance and mental health under stressful circumstances (Maddi, 2006). Hardy attitudes help individuals to have healthy life by looking for effective solutions for their daily problems and providing persevering attitude to view stressful circumstances as less stressful and more challengeable and controllable (Abdollahi et al., 2016). The findings of this study emphasize the important role of hardiness as an enhancing factor in decreasing the amount of perceived stress and suicidal ideation. Hardiness appears as a general orientation towards self, and the world conceptualized as comprising of a sense of commitment, control, and challenge. Particularly, individuals with strong hardiness are committed to what they do in diverse aspects of their lives; believe in having some control over the causes and solutions of problems, and view life changes and adjustment demands as challenges and opportunities. Therefore, hardy individuals tend to remain strong during hardship, and these characteristics appear to protect them against feeling of hopelessness

Implications for Practice

The current study contributes to the existing knowledge on suicidal ideation among nursing students in the Malaysian context, where research on suicidal ideation in Malaysia is limited. Theoretically, the findings of this study fill the gap in the suicidal research carried out in Malaysia and add to the existing literature on understanding hardiness as an influencing and enhancing factor for nursing students against suicidal ideation. Thus, the current study contributes valuable information through developing the credibility of hardiness theory and supports the portability of this theory to a Malaysian nursing student sample. Methodologically, this study adopts a more sophisticated statistical procedure in using structural equation modeling. This analysis takes one step closer to a better understanding of the hypothesized model of suicidal ideation in contrast to conventional linear regression. Using structural equation modeling and multi-group analysis provide further insight into the relationship between the studied variables. Therapeutically, clarifying the relationships between problem-solving skills appraisal, perceived stress, and hardiness with suicidal ideation among nursing students can assist in the creation of effective prevention and intervention programs to reduce suicidal ideation. For example, when practitioners or psychologists evaluate the risk of suicide in nursing students, they need to assess the amount of perceived stress, problem-solving skills, and hardiness. Although causality is unable to be determined in this study, hardiness training may decrease the likelihood of suicidal ideation.

Strengths and Limitations

The findings of the current study must be viewed in relation to its limitation. This study is restricted to nursing students in Selangor state, and possible generalization to other populations should be verified by further studies. This study utilized self-report questionnaires. Even though the questionnaires used in this research are psychometrically qualified, self-report questionnaires pose a threat to the internal validity of the data. Therefore, future studies should use a multi-method approach, including quantitative and qualitative methods that provide incremental validity to the data. As the present study was a cross-sectional survey, it constrained us to a discussion about the associations between variables rather than causality. It is necessary to replicate this study with longitudinal and experimental designs to examine the process through which the problem-solving skills, hardiness, and perceived stress affect suicidal ideation.

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Validity of the Montreal Battery of Evaluation of Amusia: An Analysis Using Structural Equation Modeling

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The Montreal Battery of Evaluation of Amusia (MBEA) is the gold standard for diagnosing amusia. We aimed to evaluate its factorial and convergent validity. Data were collected for the MBEA and a self-report Amusic Dysfunction Inventory on a non-random sample (n = 249), and the following Structural Equation Modeling (SEM) procedures were conducted: confirmatory factor analysis of the theoretical model; exploratory SEM for alternative non-restricted factor solutions; and structural models with each of these solutions as predictors of the inventory's items. The theoretical model did not prove acceptable goodness of fit, and two- and three-factor non-restricted models were better-fitted solutions for Scale, Contour and Interval tests, and Meter and Memory tests, respectively, than the theoretical one-factor model. This may reflect distinct perceptual processes related to neurocognitive demand. The non-restricted models of Scale, Meter and Memory showed to be acceptable predictors of self-reported capacity for melodic perception, vocal production, rhythmic coordination, and memory.

Key words: amusia, music cognition, structural equation modeling, validity

Introduction

Amusia is a neurocognitive impairment of music perception, memory or performance, primarily explained by a deficit in pitch processing that could be congenital- or lesion-based (Hyde & Peretz, 2004; Peretz, Champod, & Hyde, 2003), and that cannot be explained by hypoacusia, general neurocognitive functioning, education or exposure to music (Cuddy, Balkwill, Peretz, & Holden, 2005; Ayotte, Peretz, & Hyde, 2002).

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Assessment of amusia has been almost exclusively performed using the Montreal Battery of Evaluation of Amusia (MBEA) (Pfeifer & Hamann, 2015).

Though not explicitly stated as a formal measurement model in the original theoretical scheme (see Peretz et al., 2003), the MBEA evaluates cognitive processing of music through six first-order factors: scale, contour, and interval, as part of a *melodic organization* second-order factor; rhythm and meter, as part of a *temporal organization* second-order factor; and *memory-recognition*, which is conditioned by melodic- and temporal organization due to the nature of the task, which implies the previous execution of the melodic and temporal tests.

The MBEA theoretical model assumes modularity of music processing in three levels: 1) with regard to other cognitive functions (Peretz & Coltheart, 2003); 2) between second-order factors (e.g., impairment of melodic organization could exist in the presence of non-impaired temporal organization [Hyde & Peretz, 2004]; impairment of memory-recognition could exist in the presence of non-impaired melodic or temporal organization [Peretz et al., 2003]), and 3) between first-order factors (e.g., impairment of interval processing could exist in the presence of non-impaired contour processing, but not vice versa; and impairment of rhythm could exist in the absence of impaired meter [Peretz et al., 2003]).

Research using the MBEA has seen a significant rising; because of this, its validity is being rigorously tested, as it could be expected of such a relatively new and almost unique measure. Several observations can or have been made in this regard. First, the supposed modularity of music processing is in contradiction with the MBEA's composite score ultimately proposed by Peretz et al. (2003) as a unitary measure (a third-order factor) of music perception and memory functioning. Second, the only study that has undertaken the task of analyzing the items' dimensionality of the MBEA, has reported failure to replicate the assumed unidimensionality of five of its tests (Nunes-Silva & Haase, 2012). Third, one study (Henry & McAuley, 2013) has highlighted the crucial involvement of decision-making processes in the simple behavior of listening and responding to the MBEA items; these processes are not considered in the theoretical model in which the battery is founded (justifiably because of parsimonious reasons), but could have a modifying role in the structure of the theory. Fourth, little is still known regarding the relation of performance of subjects on the MBEA and actual specific amusic behaviors (e.g., ability to distinguish melodies without lyrics, minimal capacity to sing in tune or to follow the basic rhythm of a song), with several studies approaching the phenomena through the inclusion of non-representative samples of so-called "self-declared amusics" with no clear or measured diagnostic criteria other than the MBEA itself. Fifth, there is still some discrepancy concerning the use of the MBEA in terms of the number of subtests, items, scoring, cutoff scores, diagnostic accuracy and consequent estimation of prevalence (Henry & McAuley, 2013; Peretz & Vuvan, 2017; Pfeifer & Hamann, 2015).

For our consideration, in order to reach a better understanding of amusia, and all other neurocognitive features that might be addressed with, a more precise comprehension of the MBEA is needed, especially as it has been taken for granted as the gold standard for proving the existence of the very phenomenon it intends to measure. None of the studies targeting this knowledge gap (Henry & McAuley, 2013; Nunes-Silva & Haase, 2012; Pfeifer & Hamann, 2015) undertakes the task of thoroughly assessing the validity of the theoretical model underlying the MBEA. We believe that this assessment may provide information for further refinement of the theory behind the

MBEA, and could have practical implications concerning the pertinence of each of its tests and items, and the potential relation of specific music perceptual domains (e.g., perception of pitch and perception of rhythm) to other cognitive and neural processes (e.g., decision-making and frontal lobes).

Structural equation modeling (SEM) provides a way to address this issue based on systematic fit assessment procedures and estimation of relationships between latent constructs corrected for measurement error (Asparouhov & Muthén, 2009; Morin, Arens, & Marsh, 2016). Factorial validity of the MBEA can be assessed using: a) Confirmatory Factor Analysis (CFA) to empirically test the original theoretical model, assuming that specific indicators are only related to specific factors (e.g., variances in items of the MBEA's Rhythm test are caused by one and only one pre-specified factor measuring rhythm perception); and b) Exploratory SEM (ESEM) to estimate non-restricted relationships between multiple measures in the absence of a valid pre-specified model (e.g., variances in items of the MBEA's Rhythm test may be caused by two or more interrelated factors measuring different features of rhythm perception). For evaluation of the criterion validity, SEM can also be utilized to test the capacity of the best factorial models (restrained or unrestrained) to predict specific outcomes (e.g., a three factor model obtained through ESEM may have better associations with self-reported rhythm difficulties than a one factor model obtained through CFA).

The overall purpose of this study was to comprehensively assess the validity of the MBEA, aiming for three objectives: 1) to test the factorial structure of the MBEA's original theoretical model; 2) to explore an alternative factorial structure for the MBEA using ESEM; 3) to evaluate the convergent validity of the MBEA in relation to specific indicators of self-perceived amusic dysfunction.

Method

Participants

Data were collected in three Mexican cities (Mexico City, Veracruz, and Merida) during July 2014-April 2015, through convenience sampling. Inclusion criteria were: 1) age between 14 and 70 years-old (the range was established considering previous studies with this minimum age [Nunes-Silva & Haase, 2012], and also aiming to reduce the effect of age-associated cognitive decline [Deary et al., 2009]); 2) literacy; and 3) oral informed consent after assuring the participant's complete understanding of confidentiality, research purposes, presence of low risks (possible feelings of lost time, boredom, fatigue, or minimal hearing discomfort), procedures for minimizing risks (comfortable volume level, and breaks between tests), and voluntary completion of activities. Exclusion criteria were checked through the participant's self-report of: history of neuropathology (one or more of the following formal diagnoses: traumatic brain injury, stroke, epilepsy, neuroinfection, dementia, schizophrenia, or other neurodegenerative disease); compromised hearing acuity (formal diagnosis of damage to tympanic membrane or inner ear structures, or subjective complaint of disability); and formal music training (two or more years of study in a music school).

Measures

MBEA items consist of piano melodies that participants listen to through headphones. For Scale, Contour, Interval and Rhythm tests, participants are presented with 31 pairs of monophonic melodies and asked to judge whether the two melodies are the same or different. A catch trial is included randomly within each of the first four tests to assure that the participant is paying attention. For Meter test, participants are presented with a single homophonic melody, and asked to judge if the presented melody is a march or a waltz. For Memory test, a single monophonic melody is played and participants must judge whether they heard it or not during previous tests. A 30-point scale score is computed for each of the tests, and an average score can be calculated from the six subtests to obtain a global measure of music cognition (Peretz et al., 2003). Though the diagnostic accuracy of this scoring procedure has been questioned (Henry & McAuley, 2011), we decided to retain it for the descriptive analysis because the theoretical factorial structure of the MBEA was originally developed on the evidence provided by this scoring system, under the practical assumption that response to individual items could be added and averaged into broader indexes (e.g., a global score), and because this system is still widely used in research of music perception and amusia (for examples, see: Fujito et al., 2018; Peretz & Vuvan, 2017; Tang et al., 2018). For the purpose of the main analyses (CFA and ESEM), we used the participant's responses to individual items.

A 9-item self-report Amusic Dysfunction Inventory (ADI) was designed ex profeso for this study as a way to explore amusic complaints in four domains: melodic perception (item 1: *"I can tell when an instrument is out of tune"*, item 2: *"I'm capable of clearly noticing an incorrect note in a familiar melody"*); rhythmic coordination (item 3: *"I dance fluently and to the rhythm of music"*, item 4: *"It is difficult for me to follow the rhythm of a song with my hands or my feet"*); vocal production (item 5: *"People say I sing out of tune"*, item 6: *"I sing out of tune"*); and memory (item 7: *"I have difficulty in recognizing the melody of a song when it has no lyrics"*, item 8: *"I have troubles remembering melodies that I have heard several times"*, item 9: *"I can only remember lyrics of the songs and I often forget the melodies"*) (all items' statements are literally translated from Spanish). The inventory is based on the MBEA's theoretical dimensions of melodic/temporal perception and memory, and on complaints of daily-life impairments frequently reported by amusic individuals, or specific dysfunctions observed by researchers (Cuddy et al., 2005; Peretz et al., 2003). Participants were asked to rate frequency (1 – *"Never"*, *2 – "Rarely"*, *3 – "Frequently"*, *4 – "Always"*) of behaviors and situations that could be related to amusic dysfunction.

A demographic questionnaire included items about gender, age, years of education (completed levels), handedness, and a brief checklist of exclusion criteria.

Procedure

After an oral informed consent and a brief checklist for exclusion criteria, the MBEA and the ADI were computer-administered. Participants were instructed to register their responses on spreadsheets, arranged as follows: two columns for responses of each of the MBEA tests ("Yes" or "No" for all the tests except Meter for which the responses were "March" or "Waltz"), and four columns for the Likert-type responses of the ADI. All evaluations were administered in quiet settings, during one session, and carried out by psychologists trained in the procedures of the study.

All procedures were in accordance with the declaration of Helsinki, and approved by the Research Committee of the Anahuac University.

Statistical Analyses

Gender, age, years of education, and handedness, as well as sum scores of the MBEA and distribution of responses on the ADI were described as mean (standard deviation) or frequencies (percentages) for continuous and categorical data, respectively, in order to characterize the sample of participants. Statistical differences and correlations with MBEA global score were computed for each of the demographics aiming to identify significant ($p < .01$) confounders. All missing values were reported for specific variables. This analytical procedure was performed with IBM SPSS Statistics Version 22.

To assess factorial validity of the MBEA, the three one-factor models corresponding to theoretical dimensions of the MBEA were individually assessed through CFA, including: all 90 individual items from Scale, Contour, and Interval tests as indicators of melodic organization, all 60 individual items from Rhythm and Meter as indicators of temporal organization, and the 30 individual items of Memory test as indicators of the memory-recognition dimension. A one-factor model using all pooled items was also run to assess the global dimension of the MBEA. For every model, previously identified confounders were controlled. To determine the goodness of fit of the models, the following indices and correspondent cutoff values were taken into account: $p > .05$ in chi square test (χ^2); comparative fit index (CFI) > .95; Tucker-Lewis index (TLI) > .90; and root-mean-square error of approximation (RMSEA) < .05 (Hu & Bentler, 1999; Schreiber, Stage, King, Nora, & Barlow, 2015).

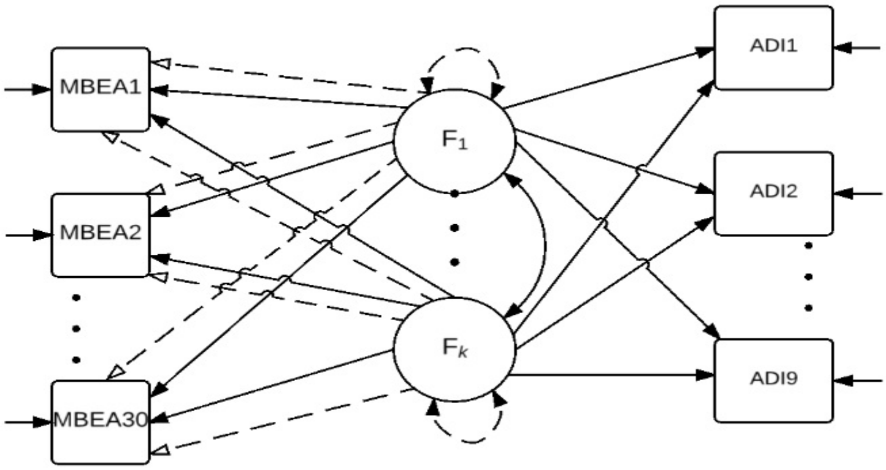

Figure 1 Prototypical representation of ESEM of the MBEA and association with ADI indicators.

Abbreviations: MBEA = Montreal Battery of Evaluation of Amusia; F = Factor; ADI = Amusic Dysfunction Inventory.

Left side of the figure represents prototypical non-restricted models with different number of factors. Right side of the figure represents path analysis, with ESEM factors as predictors of specific ADI's indicators. Confounders are not depicted in the figure.

Squares represent observed variables (indicators) and circles represent unobserved variables (factors). Full unidirectional arrows linked to indicators or factors represent the item uniqueness or factor disturbances. Full unidirectional arrows pointing to indicators represent measurement error. Bidirectional full arrows linking ovals represent factor covariances and correlations. Bidirectional dashed arrows connecting single ovals represent factor variances (Morin et al., 2016). Three-pointed vertical lines indicate ellipsis (e.g., 1, 2, 3…10).

ESEM was conducted to consecutively assess goodness of fit of different non-restricted multifactor models of melodic and temporal dimensions, as well as for individual MBEA tests (Scale-Memory) using Geomin rotation. To choose a specific model, chi square test difference with scaling factor was utilized, considering $p < .01$ to avoid type I error. Significant ($p < .05$) standardized factor loadings and between-factors covariations were identified for each of the retained Scale-Memory models. To evaluate criterion validity of the MBEA, six independent structural models were performed using these factorial solutions as predictors of ADI's indicators, as prototypically depicted in Figure 1. Identified confounders were also controlled for both analytical procedures.

For all SEM analyses, we decided to use individual items as ordered categorical indicators, instead of using scores based on the aggregation of the correct answers (MBEA scoring system); this procedure allows us to obtain latent variables that represent non-observable traits that underlie the responses of the test (Bollen & Lennox, 1991). The use of linear composites (sum of correct items) was discarded because coefficients estimated with this method may be upwardly or downwardly biased. All SEM procedures were conducted in Mplus Version 6.12 (Muthén & Muthén, 1998-2011) using the weighted least squares (WLSMV) means and variance adjusted estimator.

Results

Demographic Characteristics, MBEA Scores, and Responses to ADI

From 261 participants that concluded the assessment, six individuals were excluded from

Table 1 *Demographic characteristics and MBEA scores*

	Freq. (%) | Mean (SD)	Differences or correlations[a]
Gender		$t(247) = 1.86, p = .06$
Male	109 (43.8)	
Female	140 (56.2)	
Handedness		$F(2,248) = .14, p = .86$
Right	229 (92.0)	
Left	17 (6.8)	
Ambidexterity	3 (1.2)	
Age	40.25 (16.37)	$r = -.26, p = .00$
Years of education	15.90 (4.02)	$r = .13, p = .02$
MBEA		
Scale[b]	24.00 (4.34)	
Contour[c]	23.71 (4.11)	
Interval	22.40 (4.09)	
Rhythm[b]	24.99 (3.57)	
Meter[c]	22.93 (5.24)	
Memory[b]	24.77 (4.23)	
Global	23.80 (3.04)	

Notes. $n = 249$; [a]Differences and correlations are conducted with the MBEA global score as the dependent variable; [b]1 missing value; [c]2 missing values.

Abbreviations: MBEA – Montreal Battery of Evaluation of Amusia.

analysis due to a posteriori self-report of neuropathology history (n = 5) and compromised hearing acuity (n = 1), and six were excluded due to several discrepancies between item number and correspondent response on one or more MBEA tests. A total of 249 cases were included in the analyses. Table 1 displays the demographic characteristics and MBEA scores for this sample. Mean frequencies and standard deviations of responses on ADI were as follows: item 1 = 2.38(.87); item 2 = 2.35(.92); item 3 = 2.12(.95); item 4 = 1.70(.76); item 5 = 2.58(.99); item 6 = 2.71(1.19) [1 missing value]; item 7 = 2.02(.71); item 8 = 1.73(.64); item 9 = 1.78(.62). Age and years of education showed weak correlations with the MBEA global score (see Table 1), but only the former proved p value < .01, and thus was considered as confounder in the main analyses.

Evaluation of Factorial Validity: CFA and ESEM

As shown in Table 2, CFAs of the theoretical one-factor dimensions (melodic organization, temporal organization, memory-recognition, and global) did not prove acceptable goodness-of-fit values or convergence, and neither did the ESEM one-factor alternative solutions for each of the MBEA tests. Alternative multifactor models proved acceptable goodness-of-fit for individual subtests: two-factor solutions on Scale, Contour, Interval, and Rhythm, and three-factor solutions on Meter and Memory. Covariations between F1 and F2 were not observed for any of the two-factor models (Scale: β = .12, p = .42; Contour: β = .15, p = .33; Interval: β = -.11, p = .43;

Table 2 *Goodness-of-fit and difference testing for the MBEA theoretical and alternative model*

	Fit indices					Difference test		
	χ^2	df	$RMSEA_S$ [90% CI]	CFI_S	TLI_S	$\chi^2_{S\text{-}B}$	df	p
CFA for theoretical one-factor dimensions[a]								
Global	No convergence.							
Melodic[b]	4739.25	4004^{**}	.02 [.02, .03]	.84	.83			
Temporal[b]	No convergence.							
Memory -recognition[d]	657.58	434^{**}	.04 [.03, .05]	.76	.75			
ESEM alternative solutions for theoretical dimensions[a]								
Melodic								
1	4739.25	4004^{**}	.02 [.02, .03]	.84	.83			
2	4102.03	3914^{*}	.01 [.00, .01]	.95	.95	493.71	90	.00
3	3981.88	3825^{*}	.01 [.00, .01]	.96	.96	137.04	89	.00
4	**3869.96**	**3737**	**.01 [.00, .01]**	**.97**	**.96**	**123.27**	**88**	**.00**
5	3765.30	3650	.01 [.00, .01]	.97	.97	115.65	87	.02
Temporal								
1	2416.56	1769^{**}	.03 [.03, .04]	.67	.66			
2	1926.49	1709^{**}	.02 [.01, .02]	.89	.88	321.86	60	.00
3	1777.53	1650^{*}	.01 [.00, .02]	.93	.92	142.97	59	.00
4	1675.37	1592	.01 [.00, .02]	.95	.95	109.97	58	.00
5	**1595.92**	**1535**	**.01 [.00, .02]**	**.96**	**.96**	**85.29**	**57**	**.00**
6	1526.24	1479	.01 [.00, .02]	.97	.97	76.20	56	.03

Table 2 continues

Table 2 continued

Table 2 *Goodness-of-fit and difference testing for the MBEA theoretical and alternative model*

	Fit indices					Difference test		
	χ^2	df	$RMSEA_S$ [90% CI]	CFI_S	TLI_S	$\chi^2_{S\text{-}B}$	df	*p*
ESEM alternative solutions for tests[a]								
Scale[c]								
1	644.11	434**	.04 [.03, .05]	.85	.84			
2	**438.55**	**404**	**.01 [.00, .03]**	**.97**	**.97**	**166.00**	**30**	**.00**
3	400.54	375	.01 [.00, .02]	.98	.97	40.83	29	.07
Contour[d]								
1	562.19	434**	.03 [.02, .04]	.88	.87			
2	**433.12**	**404**	**.01 [.00, .02]**	**.97**	**.96**	**105.53**	**30**	**.00**
3	390.72	375	.01 [.00, .02]	.98	.98	45.15	29	.02
Interval								
1	651.15	434**	.04 [.03, .05]	.84	.83			
2	**451.04**	**404**	**.02 [.00, .03]**	**.96**	**.96**	**145.40**	**30**	**.00**
3	412.46	375	.02 [.00, .03]	.97	.96	41.30	29	.06
Rhythm[c]								
1	602.16	434**	.03 [.03, .04]	.86	.85			
2	**422.09**	**404**	**.01 [.00, .02]**	**.98**	**.98**	**141.59**	**30**	**.00**
3	384.80	375	.01 [.00, .02]	.99	.99	37.77	29	.12
Meter[d]								
1	591.69	434**	.03 [.03, .04]	.90	.90			
2	493.05	404*	.03 [.01, .03]	.94	.94	90.09	30	.00
3	**432.47**	**375**	**.02 [.01, .03]**	**.96**	**.95**	**60.68**	**29**	**.00**
4	389.03	347	.02 [.00, .03]	.97	.96	44.01	28	.02
Memory[c]								
1	657.58	434**	.04 [.03, .05]	.76	.75			
2	447.32	404	.02 [.00, .03]	.95	.94	166.91	30	.00
3	**395.02**	**375**	**.01 [.00, .02]**	**.97**	**.97**	**55.05**	**29**	**.00**
4	360.24	347	.01 [.00, .02]	.98	.98	36.18	28	.13

Notes. $n = 249$; [a]Controlling for age; [b]3 missing values; [c]1 missing value; [d]2 missing values; $^{*}p \leq .05$; $^{**}p \leq .001$.

Bolds indicate the best model based on scaled testing (model solution is retained when the next solution reaches $p > .01$).

Rhythm: $\beta = -.21$, $p = .08$), nor between the factors of the Memory test (F1 with F2: $\beta = .18$, $p = .22$; F1 with F3: $\beta = .26$, $p = .10$; F2 with F3: $\beta = .07$, $p = .70$). Significant associations were only found for Meter test, between F1 and F2 ($\beta = .41$, $p = .01$) but not between the rest of its factors (F1 with F3: $\beta = -.17$, $p = .21$; F2 with F3: $\beta = .10$, $p = .33$). Not controlling for the age confounder, these results showed a similar pattern, with the exception of Meter test that rendered a correlation between F1 and F3 ($\beta = .38$, $p = .01$). Factor loadings for each of the MBEA retained ESEM solutions are displayed in Table 3.

Table 3 *Standardized factor loadings for the ESEM solutions of each of the tests of the MBEA*

Items[b]	Factor loadings[a]													
	Scale		Contour		Interval		Rhythm		Meter			Memory		
	F1	F2	F1	F2	F1	F2	F1	F2	F1	F2	F3	F1	F2	F3
1	.11(.10)	**.23*(.11)**	**.62**(.06)**	.03(.11)	-.18(.13)	**.48**(.10)**	**-.29**(.09)**	**.33**(.09)**	.20(.12)	.16(.11)	**.26*(.11)**	**.33*(.16)**	**.61**(.16)**	-.04(.15)
2	**.73**(.06)**	.12(.11)	**-.19*(.10)**	**.34**(.10)**	**.82**(.05)**	.14(.13)	**.66**(.07)**	.08(.10)	**.87**(.14)**	-.22(.25)	-.02(.07)	**.78**(.09)**	.16(.13)	.00(.06)
3	**-.23**(.09)**	**.39**(.09)**	.04(.11)	.21(.11)	**.67**(.07)**	**-.31*(.12)**	**.47**(.08)**	**-.22*(.10)**	**.53**(.12)**	.03(.14)	.11(.12)	-.01(.06)	**.59**(.13)**	.26(.13)
4	-.04 (.09)	**.27*(.10)**	.09(.09)	.07(.14)	**.44**(.08)**	-.12(.10)	-.08(.14)	**.55**(.10)**	.07(.13)	.01(.10)	**.38**(.10)**	**.54**(.10)**	-.02(.11)	-.16(.14)
5	-.13(.10)	**.31**(.10)**	-.08(.10)	**.38**(.10)**	**.72**(.05)**	-.16(.10)	-.03(.15)	**.41**(.13)**	**.56**(.11)**	-.01(.03)	**.39**(.13)**	**.69**(.08)**	-.15(.12)	.00(.09)
6	**.76**(.07)**	.08(.12)	**.57**(.07)**	.04(.11)	**.27*(.11)**	**.48**(.11)**	**.75**(.05)**	.00(06)	.16(.13)	**.33**(.11)**	.13(.11)	**.64**(.08)**	.02(.06)	**-.22*(.10)**
7[c]	**.66**(.06)**	.00(.07)	**.61**(.08)**	**.32**(.09)**	.12(.10)	**.38**(.10)**	**.95**(.08)**	**.33*(.15)**	.19(.12)	**.28*(.12)**	-.05(.11)	-.06(.15)	**.64**(.11)**	-.08(.18)
8[d]	**.72**(.05)**	.01(.07)	.07(.09)	**.34**(.13)**	**.77**(.06)**	.00(.06)	-.00(.03)	**.53**(.09)**	**.33*(.14)**	**.10**(.13)**	**.27*(.11)**	.04(.14)	**.67**(.09)**	-.12(.12)
9	-.15(.14)	**.63**(.10)**	**.60**(.06)**	.02(.10)	-.04(.12)	**.45**(.10)**	**.67**(.07)**	-.01(.13)	-.01(.12)	**.31**(.11)**	.13(.11)	**.54**(.11)**	**.30**(.09)**	-.15(.10)
10	**.73**(.05)**	-.05(.10)	-.00(.02)	**.60**(.10)**	-.01(.12)	**.50**(.10)**	**.72**(.06)**	-.00(.08)	.02(.09)	**.48**(.11)**	**.28*(.12)**	.04(.10)	**.49**(.13)**	**.50**(.14)**
11	**.67**(.07)**	**-.23**(.08)**	**-.43**(.09)**	**.38**(.10)**	**.66**(.07)**	.19(.11)	**.34*(.14)**	**.50**(.11)**	**.45**(.12)**	**.39**(.14)**	.02(.11)	**.51**(.08)**	-.05(.12)	.05(.12)
12	.16(.15)	**.76**(.08)**	**.52**(.07)**	-.01(.08)	.17(.11)	**.55**(.09)**	**.77**(.06)**	-.02(.11)	-.05(.15)	.07(.14)	**.35**(.12)**	**.56**(.06)**	.03(.09)	-.08(.11)
13	**.73**(.05)**	**-.21*(.10)**	-.09(.11)	**.51**(.10)**	**.47**(.08)**	**-.19*(.09)**	**.71**(.08)**	-.12(.16)	**.29*(.13)**	**.33*(.15)**	-.20 (.11)	**.54**(.07)**	.00(.09)	.04(.12)
14	-.17(.10)	**.33**(.10)**	**.67**(.10)**	**.31**(.11)**	-.07(.12)	**.57**(.10)**	.05(.13)	**.56**(.11)**	.00(.09)	-.02(.11)	**.80(.09)**	.16(.14)	**.81**(.10)**	.00(.04)
15	.10(.13)	**.53**(.10)**	**-.25*(.12)**	**.59**(.10)**	.08(.13)	**.49**(.14)**	-.21(.11)	**.57**(.08)**	**.38**(.12)**	**.36**(.12)**	**.27*(.13)**	**.57**(.07)**	.06(.11)	.00(.10)
16	**.71**(.08)**	**.27**(.09)**	**-.47**(.08)**	.18(.09)	-.16(.10)	**.54**(.08)**	**.76**(.06)**	.02(.10)	**.46**(.12)**	**.50**(.16)**	-.15(.12)	-.14(.08)	**.33**(.09)**	.06(.12)
17	.11(.15)	**.38**(.13)**	-.02(.11)	**.28*(.12)**	**.47**(.08)**	**-.30**(.10)**	**.67**(.07)**	-.04(.11)	**.45**(.12)**	**.44**(.15)**	.18(.14)	**.55**(.09)**	.08(.14)	**.25*(.12)**
18	**.63**(.07)**	-.21(.11)	**.70**(.05)**	-.00(.06)	**.63**(.07)**	-.01(.10)	**.76**(.05)**	-.14(.09)	**.29*(.13)**	**.57**(.12)**	.05(.14)	-.06(.11)	**.40**(.13)**	.09(.15)
19	.00(.05)	**.42**(.11)**	.06(.14)	**.82**(.09)**	**.52**(.07)**	-.17(.09)	-.13(.11)	**.49**(.10)**	.22(.12)	**.58**(.12)**	-.15(.12)	**.54**(.06)**	.11(.10)	**.27*(.11)**
20	.13(.12)	**.51**(.14)**	**.60**(.07)**	.05(.14)	-.13(.10)	.23(.12)	-.05(.12)	**.51**(.09)**	**.49**(.13)**	**.39*(.17)**	-.12(.13)	.00(.05)	**.64**(.15)**	.36(.20)
21	-.10(.11)	**.42**(.10)**	.16(.12)	**.52**(.12)**	.06(.10)	**.35**(.10)**	**.68**(.07)**	-.19(.11)	-.07(.13)	**.46**(.12)**	**.25*(.12)**	.17(.10)	**.71**(.12)**	.31(.16)
22	**.69**(.07)**	-.09(.11)	**.80**(.05)**	-.04(.11)	**.72**(.06)**	.13(.10)	.17(.14)	**.62**(.10)**	-.13(.15)	**.54**(.13)**	**.46**(.12)**	-.10(.07)	**.47**(.11)**	**.42**(.13)**
23	**.83**(.06)**	**.22*(.10)**	**.69**(.05)**	-.04(.12)	-.17(.10)	**.50**(.09)**	**.74**(.07)**	-.01(.08)	-.23(.13)	**.77**(.10)**	-.00(03).	**.69**(.11)**	**-.26*(.11)**	.21(.13)
24	.12(.16)	**.78**(.08)**	**.64**(.07)**	.22(.12)	**.78**(.05)**	.11(.09)	.19(.18)	**.76**(.11)**	.21(.14)	**.46**(.12)**	.12(.12)	**.55**(.09)**	-.00(.01)	**.52**(.08)**
25	**.76**(.05)**	.00(.09)	**.63**(.07)**	-.20(.11)	.07(.12)	**.53**(.10)**	-.06(.10)	**.47**(.09)**	-.01(.10)	**.63**(.10)**	.14(.12)	-.02(.11)	**.66**(.11)**	.08(.18)
26	**.66**(.06)**	-.03(.09)	-.22(.12)	.19(.10)	**.78**(.05)**	-.00(.06)	-.03(.12)	.46(.11)	.03(.10)	**.74**(.07)**	-.02(.10)	**-.22*(.09)**	**.35**(.12)**	.06(.14)
27	**.78**(.04)**	.00(.07)	**.66**(.05)**	-.00(.07)	.00(.01)	**.63**(.08)**	**.75**(.06)**	.19(.11)	.00(.07)	**.91**(.05)**	-.12(.11)	.00(.01)	**.59**(.19)**	**.86**(.13)**
28	**.77**(.04)**	-.02(.08)	.12(.14)	**.58**(.12)**	.15(.10)	**.46**(.09)**	.07(.13)	**.28*(.12)**	-.03(.11)	**.73**(.08)**	-.00(.10)	**.76**(.07)**	-.09(.12)	.00(.19)
29	-.02(.11)	**.55**(.09)**	**.64**(.06)**	-.16(.11)	**.77**(.05)**	-.15(.12)	-.04(.09)	**.39**(.11)**	-.02(.07)	**.60**(.08)**	**-.39**(.09)**	**.52**(.12)**	-.18(.12)	**.47**(.10)**
30[e]	-.02(.12)	**.67**(.08)**	**.70**(.06)**	.14(.08)	**.68**(.05)**	.01(.09)	**.80**(.07)**	.19(.10)	.07(.12)	**.75**(.10**	**-.27*(.11)**	-.07(.09)	**.47**(.16)**	**.45**(.15)**

Notes. n = 249; [a]Age confounder is controlled for all the models; [b]Items are different for each test but follow the same numeration. Catch trials aiming to assess alertness of the listener (Scale: trial 7, Contour: trial 21, Interval: trial 28, Rhythm: trial 14) are not taken as effective items and thus are excluded from this numeration. [c]1 missing value in Rhythm; [d]Missing values in: Contour = 1, Meter = 1; [e]Missing values in: Scale = 1, Meter = 1; Memory = 1.

Shaded cells indicate the correct response to the item: "different" for Scale-Rhythm, "march" for Meter, and "No" for Memory. Blank cells indicate the other option of response.

Abbreviations: ESEM – Exploratory structural equation modeling; MBEA – Montreal Battery of Evaluation of Amusia; F – Factor.

$^{*}p \leq .05$; $^{**}p \leq .001$

Assessment of Convergent Validity: MBEA as Predictor of Self-Perceived Amusic Dysfunction

SEM analyses for testing convergent validity of MBEA alternative models with regard to specific amusic dysfunctions are displayed in Table 4. For each of the structural models, fit indices were as follow: Scale: $\chi^2(656) = 707.49$, $p = .08$; RMSEA = .01 [.00, .02]; CFI = .97; TLI = .97. Contour: $\chi^2(656) = 696.75$, $p = .13$; RMSEA = .01 [.00, .02]; CFI = .97; TLI = .97. Interval: $\chi^2(656) = 704.77$, $p = .091$; RMSEA = .01 [.00, .02]; CFI = .97; TLI = .97. Rhythm: $\chi^2(656) = 689.54$, $p = .17$; RMSEA = .01 [.00, .02]; CFI = .98; TLI = .97. Meter: $\chi^2(618) = 682.43$, $p = .03$; RMSEA = .02 [.00, .029]; CFI = .97; TLI = .96. Memory: $\chi^2(618) = 643.94$, $p = .22$; RMSEA = .01 [.00, .02]; CFI = .98; TLI = .98.

Discussion

The main purpose of this study was to evaluate the validity of the MBEA by thoroughly analyzing the original and alternative factorial structures via systematic fit assessment procedures, identification of latent constructs, and estimation of relationships between them and with regard to self-report of specific amusic-related impairments.

Factorial Validity

Overall, the findings strongly suggest that the theoretical factorial structure of the MBEA is not a good-fitted measurement model of music neurocognition. More specifically, neither the assumed three second-order factor structure (melodic and temporal organization) of the battery, nor the presumed unifactoriality of each of its tests, stood empirical proof. Instead, the results of ESEM suggest to discard melodic organization and temporal organization as composite measures, and to consider multifactoriality for each of the tests.

For the cases of Scale, Contour, Interval, and Rhythm (Scale-Rhythm), respective two-factor solutions showed better goodness-of-fit when compared to other multifactorial alternatives.

Table 4 *Structural analysis estimates for MBEA tests' ESEM solutions as predictors of ADI*

Item	Scale[a]		Contour[b]		Interval		Rhythm[a]		Meter[b]			Memory[a]		
	F1	F2	F1	F2	F1	F2	F1	F2	F1	F2	F3	F1	F2	F3
Melodic perception[c]														
1	**-.19**(.07)**	-.07(.08)	-.10(.07)	**-.16*(.08)**	-.13(.07)	-.02(.07)	-.08(.07)	.00(.07)	**-.32*(.15)**	.03(.13)	.24(.12)	.31(.19)	-.06(.21)	**-.42**(.15)**
2	**-.34**(.06)**	-.11(.08)	**-.21**(.07)**	**-.24**(.08)**	**-.25**(.07)**	-.08(.08)	**-.18*(.07)**	**-.15*(.07)**	**-.42**(.14)**	.05(.11)	**.28*(.12)**	.21(.21)	-.13(.17)	**-.57**(.16)**
Rhythmic coordination[c]														
3	-.10(.07)	**-.18*(.07)**	-.04(.08)	-.12(.08)	-.12(.07)	-.14(.07)	-.10(.07)	-.14(.08)	**-.32*(.13)**	.06(.11)	.09(.11)	.12(.12)	.02(.12)	**-.26*(.10)**
4	**-.31**(.06)**	-.04(.09)	**-.23**(.07)**	-.16(.08)	**-.24**(.07)**	-.08(.09)	-.11(.07)	-.06(.09)	**-.56**(.14)**	.02(.12)	**.31*(.13)**	.10(.15)	-.17(.11)	**-.31*(.13)**
Vocal production[c]														
5	**-.39**(.06)**	-.06(.07)	**-.29**(.07)**	**-.19*(.07)**	**-.30**(.07)**	-.07(.08)	**-.15*(.07)**	-.00(.08)	.04(.13)	-.25*(.11)	-.08(.12)	.10(.17)	-.06(.11)	-.25(.13)
6	**-.42**(.06)**	**-.17*(.08)**	**-.34**(.07)**	**-.22**(.08)**	**-.39**(.07)**	**-.22*(.09)**	**-.20**(.07)**	-.05(.08)	-.11(.13)	-.13(.11)	.10(.10)	.09(.21)	.01(.11)	**-.35**(.13)**
Memory[c]														
7	**-.21**(.07)**	-.05(.09)	**-.21**(.07)**	-.02(.09)	**-.25**(.07)**	-.02(.07)	**-.17*(.07)**	.06(.08)	-.15(.13)	-.08(.11)	.09(.12)	.18(.11)	-.19(.13)	-.25(.13)
8	**-.17*(.07)**	.09(.09)	-.09(.07)	-.05(.09)	-.12(.07)	.01(.08)	.01(.08)	.09(.09)	-.24(.15)	-.02(.12)	.09(.12)	.05(.12)	.13(.09)	.10(.12)
9	**-.22**(.07)**	.06(.08)	**-.14*(.06)**	-.03(.08)	-.11(.07)	-.03(.08)	-.05(.06)	-.04(.08)	-.19(.14)	.08(.11)	.08(.12)	.08(.13)	-.04(.09)	**-.30*(.12)**

Note. $n = 249$; [a]1 missing value; [b]2 missing values; [c]Controlling for age; $^{*}p$ d" .05; $^{**}p$ d" .001.
Abbreviations: MBEA – Montreal Battery of Evaluation of Amusia; ESEM – Exploratory Structural Equation Modeling; ADI – Amusic Dysfunction Inventory.

Observing the factor loading patterns of these four models, a clear tendency can be detected: F1 was mostly loaded by items containing altered stimuli as the second melody (scale-, contour-, interval-, and rhythm-violated conditions, according to Peretz et al. [2003]), and F2 was mainly comprised by non-violated conditions. We suggest labeling these factors: *assessment of difference* (AoD), and *assessment of sameness* (AoS), respectively. Both factors might be reflecting distinct neurocognitive processes or degrees of ability needed to perceive, retain, and compare the two melodies of each trial, as well as different degrees of decision bias partially founded on the difficulty to process each melody.

As already noted by Henry and McAuley (2013) using the signal detection theory, this response process is the result of the additive contributions of the sensitivity of the listener to correctly discriminate between same versus different melodies and his/her response bias, which in the case of the MBEA seems to originate from a systematic measurement error founded in the dichotomous options of response for each of its tests. Thus, the patterns of response may be an effect of common-method bias attributable to the MBEA (and thus a serious limitation of the battery) and not the expression of a critical feature of music cognition itself. However, taking into account the nature of the ESEM models (latent variables are free of measurement errors [Bollen, 1984]), our results suggest that the obtained models are independent of the response bias, and the source of the responses for same versus different is the consequence of two independent variations within the process of perception-retention-comparison-decision, which determine the final response of the listener. Knowing which of these processes is more critical for diagnosis of amusia falls beyond the scope of this study.

Cross-loadings between AoD and AoS further shed light on this process, as most of the covariates shared negative directions, possibly meaning that in order to produce an accurate assessment of the difference or similarity between melodies, individuals ought to execute one process of perception-retention-comparison-decision while restricting the use of the other (e.g., involvement of AoS – in theory more suitable for less neurocognitively demanding trials – for responding to trials with violated conditions might produce a higher rate of inaccurate responses).

In the case of the Memory test, a two-factor solution, though it showed acceptable goodness-of-fit, was not statistically better informative than a tree-factor model. Many of the items loaded distinctively in F1 and F2 in a similar pattern to Scale-Rhythm tests, suggesting a similar distinction of assessment processes for discriminating between recently learned and newly perceived musical material, possibly founded on our proposed assessment processes related to neurocognitive demand. Nonetheless, this three-factor model displayed several heterogeneous cross-loadings, making it harder to parsimoniously support this interpretation. Dependence of Memory test on perception and performance on the rest of the previous tests might have influenced this pattern of modeling.

Convergent Validity

Exploring convergent validity, the value of these two-factor solutions is further stretched by the fact that mostly AoD for each Scale-Rhythm test was moderately associated (β ranging from -.14 to -.42) with ADI's indicators. From all of these tests, Scale's AoD associated better with expected outcomes related to vocal production and melodic perception, whereas Contour and Interval displayed more heterogeneous covariation of AoD and AoS with all of the outcome indicators, nonetheless biased to AoD. This may be explained by the fact that Scale test seems to have better accuracy for diagnosis of amusia (Peretz & Vuvan, 2017; Goulet, Moreau, Robitaille, & Peretz, 2012) (for this reason, it has been used as a screening measure of amusia in some studies [McDonald & Stewart, 2008; Peretz et al., 2008]).

Contrary as might be expected, no association between Rhythm test's factors and ADI's indicators of rhythmic coordination was found, but only low associations with some of the rest of the ADI. Furthermore, low associations between ADI's memory items and Scale-Interval's AoD were found, but only one ADI's memory indicator (self-perceived capacity to remember song lyrics but not melodies) proved a meaningful covariation with F3 of MBEA's Memory test. Interestingly, indicators of melodic perception, rhythmic coordination, and vocal production (signing out of tune), also displayed moderate covariations with Memory's F3.

This pattern of relationships may move forward the hypothesis of AoD and AoS as two different assessment processes strongly related to task complexity, meaning that Scale's AoD and Memory's F3 items may be fairly more neurocognitively demanding, and thus could be closer to the complexity of real-life musical behaviors. Other studies have already noted the sensitivity and specificity of the Scale and Memory test to detect amusia (Peretz & Vuvan, 2017; Pfeifer & Hamann, 2015; Henry & McAuley, 2013), although one study has reported scale processing as a rather automated function of the brain (Brattico, Tervaniemi, Näätänen, & Peretz, 2006).

Special Considerations for Meter test

For this test, a three-factor solution showed a better goodness-of-fit, though with several cross-loadings between two or even all its factors, and no clear profile of the items as to characterize latent constructs. We hypothesize that, similar to the other MBEA tests, the distinction between these factors might rely on different degrees of neurocognitive demand, perhaps with regard to interaction between meter pulse and more detailed musical features of the pieces comprising the test, such as tempo, saliency of rhythmic chords, or complexity of note durations (e.g., more frequency of crotchets or quavers). Insufficient number of items and thus non-representativeness of musical features, however, prevent this analysis.

Concerning its convergent validity, Meter's F1 and F3 (both containing fewer items than F2) were significantly associated with expected ADI's indicators pertaining to rhythmic coordination; particularly, F1 displayed inverse association with self-perceived difficulty to follow rhythm with hands or feet and to dance fluidly, whereas F2 showed positive association with the latter indicator. Interestingly though, very similar patterns were also observed in relation to ADI's indicators of perception of melody. These patterns may support our hypothesis (e.g., F1 reflects a different degree of neurocognitive demand than F2); however, not having any more data to support this assumption, it is beyond the scope of this work to speculate on the matter. Thoughtful questioning about the inclusion of the Meter test in the MBEA total scoring can be stated though, as has been somewhat evidenced by other results that signal its particular behavior in contrast with the rest of the MBEA tests (Henry & McAuley, 2013; Nunes-Silva & Haasse, 2012; Paraskevopoulos, Tsapkini, & Peretz, 2010; Toledo-Fernández & Salvador-Cruz, 2015).

Limitations

First, the use of non-random sampling might limit the external validity of the results due to lack of methodological control of common confounders in neuropsychological testing, such as age and years of education. This issue was addressed in the case of age via statistical control within the tested models. We avoid statistical control of education based on its very weak correlation with MBEA's global score, and because the concept of amusia itself excludes it as a confounder (Ayotte et al., 2002).

Second, the sample size could also limit the reach of the findings because of the number of MBEA's indicators included in the analyzed models. Particularly, CFA models require considerably large sample size to attain an admissible solution when complex models are assessed (e.g., models with high-order factors such as the one proposed by Peretz et al. [2003]). However, the flexibility of ESEM models decreases due to the lack of restriction in the covariance matrices and the use of Geomin rotation (Asparouhov & Muthén, 2009).

Third, items of the ADI were not rigorously developed through a formal methodological process (e.g., items' pooling, piloting, or review by judges) but rather designed ex profeso based only on previous questionnaires, reports of amusics' common complaints of daily-life impairments and laboratory-observed dysfunctions (Cuddy et al., 2005; Ayotte et al., 2002). It is interesting to note that recent questionnaires with similar evaluation of melodic perception, vocal production, rhythmic ability and memory were developed contemporarily to our ADI (Müllensiefen, Gingras, Musil, & Stewart, 2014; Pfeiffer & Hamann, 2015), and could be used for further validation. Lastly on this matter, the MBEA being the gold standard for diagnosis of amusia, the criterion validity of the ADI could not be tested a priori either. We believe that, in view of the results and the lack of another plausible measure of self-report amusic dysfuntion, this study stands as a first examination of the validity of the ADI's items.

Fourth, dependence of observations might have influenced the associations between MBEA's and ADI's outcomes, since participants may have judged their self-perceived musical capacities predisposed by their perceived self-efficacy on the MBEA tests. This assessment procedure, however, is not rare in studies using the MBEA (Cuddy et al., 2005; Pfeifer & Hamann, 2015). Ulterior studies could employ random or alternate sequences of the assessment procedure, or use ecologically-valid behavioral measures (e.g., singing, hand coordination, dancing, etc.), aiming to avoid this bias.

Lastly, and thought not exactly a limitation of our study's design per se, it is important to highlight the possible influence of acculturation in the observed performance of our sample in the MBEA. Evidence for the effects of musical acculturation on the MBEA has been reported for the Greek population, since their musical system significantly differs from that of the Western world (Paraskevopoulos et al., 2010). Considering that most of the Mexican popular music is founded on the Western tonal system, that for centuries it has received influences from European and North American music, and that Mexicans are currently highly exposed to international music which is also based on this musical system, we believe that our findings could be transferred to other populations. Further cross-cultural studies are needed in order to test this assumption.

Conclusion

Our analyses showed that alternative good-fitted models unveil an even more complex phenomenon of music neurocognition as measured by neuropsychological testing, and the insufficiency of the current, most used measurement model for diagnosing it. Further neuropsychological research on the MBEA should be conducted using latent variable modeling to evaluate these measurement models in accurately pre-diagnosed amusic individuals, as well as in different groups with plausibly-related brain pathologies (e.g., temporal epilepsy, Alzheimer, Parkinson, schizophrenia), while rigorously testing the diagnostic precision of the alternative factorial models, their assumed dissociation with other neurocognitive functions, and their ecological validity.

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Factor Structure of Slovak Adaptation of Attentional Control Scale

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The Attentional Control Scale (ACS) is a tool developed to assess the ability to voluntarily control attentional resources. The current aim was to verify the factor structure of the scale and its hypothesized inverse relationship with measures of trait anxiety on Slovak sample. The Principal Component Analysis (PCA) suggested two correlated factors resembling the hypothesized Focusing and Shifting subscales. The factorial solution suggested by the PCA had the best fit against one-factorial and two-factorial orthogonal solutions in the Confirmatory Factor Analysis (CFA) conducted on an independent sample. The entire scale had good internal consistency (ω_t = .85). The Focusing (ω_t = .81, α_{ord} = .81) and Shifting (ω_t = .67, α_{ord} = .66) subscales reached acceptable to good values of internal consistency. The ACS showed a negative relationship with trait anxiety inventory and behavioral inhibition scale. The differences of our results compared to other studies investigating factor structure of ACS are discussed, together with limitations of the current study, validity and applicability of the scale.

Key words: Attentional Control Scale, factor structure, internal consistency, trait anxiety, self-report measures

Attentional control can be conceptualized as a unitary construct, however two dichotomous components have been proposed (Derryberry & Reed, 2002; Taylor, Cross, & Amir, 2016; Telzer et al., 2008). Focusing denotes the ability to withhold attentional focus on a relevant target while ignoring distracting, although significant stimuli. Shifting is related to the concept of cognitive flexibility and, as such, represents the ability to voluntarily disengage attention from a target that is no longer relevant to the task at hand and shift it to another target. The Attentional Control Scale (ACS) is a self-report tool designed to assess individual differences in the ability to voluntarily control attentional resources (Derryberry & Reed, 2002).

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Regarding the psychometric properties of ACS, its first version consisted of two separate scales, namely Focusing and Shifting subscales (Derryberry & Rothbart, 1988). These two scales were later compiled into one scale (ACS), which, according to the authors (Derryberry & Reed, 2002), was supposed to reflect the correlated factors of *attentional focus*, *attentional shift*, and *flexible control of thought*. However, studies examining psychometric properties of ACS are still quite recent. Considering the factor structure of the scale, Verstraeten, Vasey, Claes, and Bijttebier (2010) conducted a Confirmatory Factor Analysis (CFA) of the Dutch version of ACS on a sample of children from 8 to 18 years old. They reported the two-factor solution (with positively correlated factors) having a superior fit to the one-factor solution. Reliability estimates (Cronbach's alpha – α) for the Focusing and Shifting subscales reported in this study were α = .70 and α = .63, respectively. The Icelandic version of ACS (Ólafsson et al., 2011) also yielded two components, when submitted to the Principal Components Analysis (PCA). This two-factor model with strongly correlated factors was reasonably supported by CFA on an independent sample. Internal consistency of the subscales was comparable to the values reported in Verstraeten, Vasey, Claes, and Bijttebier (2010). The fact that Ólafsson et al. (2011) also hypothesized and tested differential dependency of ACS subscales with symptoms of anxiety and depression is noteworthy. They found the Focusing subscale to be uniquely associated with trait-anxiety, while the Shifting subscale was uniquely associated with symptoms of depression. In a study conducted by Judah, Grant, Mills, and Lechner (2014), the English version of ACS has also been found to be composed of two related factors in PCA as well as in CFA. Again, internal consistency of the subscales was $\alpha = .82$ for Focusing and $\alpha = .71$ for Shifting subscale. Recently, Abasi, Mohammadkhani, Pourshahbaz, and Dolatshahi (2017) reported two factors from an exploratory factor analysis conducted on an Iranian sample with internal consistency and test-retest reliability after two weeks for the Focusing subscale ($\alpha = .78$, $r = .80$) and for the Shifting subscale ($\alpha = .66$, $r = .72$). It is worth to mention that these studies used the α coefficient to estimate internal consistency of the whole scale as well as of its subscales, despite the evidence from either PCAs or CFAs, suggesting a clear violation of the tau-equivalency and uni-dimensionality assumptions (Graham, 2006). Finally, the reported studies differed in final number of retained items in ACS, mainly due to insufficient factor loadings of certain items.

The concept of attentional control stems from the Attentional Control theory (Posner & Petersen, 1990), which was developed in order to account for the effects of anxiety-related bias of attention towards a threat (Bar-Haim, Lamy, Pergamin, Bakermans-Kranenburg, & van IJzendoorn, 2007; Derryberry & Reed, 2002; Taylor, Cross, & Amir, 2016). Specifically, according to this theory, anxiety disrupts the balance between posterior and anterior systems, heightening the influence of the stimulus driven posterior system and lowering the influence of the goal-directed anterior system (Eysenck, 2007; see also Hermans, Henckens, Joëls, & Fernández, 2014). Fox, Russo, and Georgiou (2005) suggested that anxiety facilitates automatic processing of threat-related stimuli and hinders the influence of goal-directed processes over attention, with these effects being pronounced in anxiety-inducing environments.

Moreover, it has been reported that high-anxious individuals are more prone to exhibit this attentional bias (Bar-Haim, Lamy, Pergamin, Bakermans-Kranenburg, & van IJzendoorn, 2007; Derryberry & Reed, 2002; Koster, Crombez, Verschuere, Van Damme, & Wiersema, 2006). It has also been assumed that hyper-responsivity of amygdala, as a pre-attentive threat detection system, toward a threatening stimulus in high trait anxiety individuals is responsible for this difference between anxious and non-anxious populations (Mathews, Mackintosh, & Fulcher, 1997). However, this view has been modified by incorporating the role of prefrontal cortical mechanisms (Öhman, 2005). There have even been reports (Bishop, 2009; Eden et al., 2015; Kim & Whalen, 2009) suggesting that high trait anxiety might be generally associated with impoverished control mechanisms, regardless of whether they are applied during or in absence of threat-related stimuli exposure.

Indeed, studies examining the relationship between self-report measures of attentional control and trait anxiety have frequently reported their inverse dependency (Abasi, Mohammadkhani, Pourshahbaz, & Dolatshahi, 2017; Fajkowska & Derryberry, 2010; Judah, Grant, Mills, & Lechner, 2014; Ólafsson et al., 2011). However, most of these studies employed the State-Trait Anxiety Inventory (STAI-T)

(Spielberger, Gorusch, & Lushene, 1970) as a measure of trait anxiety, which is believed to tap more into the average anxiety level rather than the general sensitivity of anxiety system (Carver & White, 1994; Fowles, 1987). Despite this, ACS has been found to be negatively associated also with measures of trait-anxiety such as the self-report measure of the behavioral inhibition system (BIS) (Carver & White, 1994), which do not necessarily reflect only average experience of anxiety on daily basis (Fajkowska & Derryberry, 2010). Importantly however, we should not view high trait anxiety only as an indicator of a poor attentional control. On the contrary, it has been reported that highly anxious individuals with good attentional control can effectively reduce their bias towards threatening stimuli when given the appropriate amount of time (Derryberry & Reed, 2002). Considering that highly anxious persons with low attentional control are at increased risk for development of anxiety disorders, some reports suggest that effective interventions by the means of cognitive-behavioral therapy or working memory training may reduce symptoms of anxiety and related attentional bias towards threat (Bowler et al., 2012; Hadwin & Richards, 2016).

The aim of the current study is to verify the hypothesized two-factor structure of the ACS on a Slovak sample and to provide more suitable reliability estimates of the tool. We anticipated ACS to be formed of two-positively related factors and hence we were expecting two-factorial non-orthogonal solution to have the best fit from amongst one-factorial and two-factorial orthogonal solutions.

Regarding the inverse dependency of attentional control and trait anxiety, we hypothesized the ACS scores to be negatively related to self-report measures of trait anxiety (STAI-T and BIS). Finally, we also hypothesized that only the Focusing subscale is negatively related to measures of trait anxiety.

Methods

Participants

In total, 474 (354 females) subjects with average age 21.7 (SD = 2.5) years, comprised mostly of university students (majority from Comenius University in Bratislava) of various study programs (e.g., psychology, economics, law, medicine), participated voluntarily in this study. Subjects were not screened for health status or history of mental illness and hence were not selected or removed on this basis. First half of the subjects served for the PCA and the second half for the CFA (see Statistical Analyses).

Prior power analysis suggested that sample sizes for both correlation analyses (see Statistical Analyses) are sufficient for detection of moderate or stronger effects (r = .30 to .50 and more) while α = .05 and $1 - \beta$ = .80.

Self-Report Measures

Attentional Control Scale. Three scales were administered in the form of online questionnaires. Two independent translations of the ACS into the Slovak language (and back to English) were done and final edits were made after mutual consent between two translators. The ACS (Derryberry & Reed, 2002) includes 20 statements regarding the difficulty to control attention during everyday circumstances (Appendix) to which participant responds on a 4-point Likert scale (*almost never – always*) with no middle point.

State-Trait Anxiety Scale. The STAI-T subscale for trait anxiety (Spielberger, Gorusch, & Lushene, 1970) consists of 20 statements with responses organized on a 4-point Likert scale (*almost never – almost always*), asking subject to report usual frequency of anxiety related physical and mental states. Internal consistency of the scale in our sample reached very good values for research purposes (ω_t = .93, α_{ord} = .93).

Behavioral Inhibition Scale. The BIS scale as self-report measure of anxiety system sensitivity (Carver & White, 1994) is made up of 7 statements also with 4-point Likert response scale (*very true for me – very false for me*). Internal consistency of the scale in our sample reached adequate level (ω_t = .79, α_{ord} = .79).

Statistical Analyses

Principal Component Analysis. Since the responses on the ACS scale were represented in ordinal form, PCA was conducted on polychoric correlation matrix. Number of components to be extracted was determined by Parallel analysis (Horn, 1965) using 500 randomly generated matrices of equal sample size for comparison. Only components with eigenvalue higher than its 95th percentile counterpart were retained. Assumptions of sampling adequacy and reducibility of the data were tested by Barttlet's test of sphericity and Keiser-Meyer-Olkin (KMO) test of sampling adequacy. As we expected components to be correlated, extracted components were rotated by oblique "oblimin" rotation method. Sample size for PCA was N = 237 (166 females) of average age 21.5 (SD = 2.1) years. Parallel analysis and PCA were conducted with psych package in R statistical software (Revelle, 2017).

Confirmatory Factor Analysis. In CFA, the appropriateness of the model was assessed by following fit indices and their optimal values: $\chi^2/df \leq 2$, comparative fit index ($CFI \geq .95$), Tucker-Lewis index ($TLI \geq .95$), root mean square error of approximation ($RMSEA \leq .05$), and standardized root mean square residual ($SRMR \leq .05$). In addition to the significance of testing the difference between predicted and observed covariance matrices by χ^2, we followed the recommendations by Ullman (2006), who suggested the use of the ratio of χ^2 to degrees of freedom (χ^2/df) in samples larger than 200, since even small discrepancies between the compared matrices might result in a significant difference. We also checked for possible model improvements suggested by modification indices (MI). Since responses on the scale were presented in the form of ordered factors, we used diagonally weighted least squares (WLSVM) (e.g., Flora & Curran, 2004) as the estimation method, which however did not enable us to compare the tested models directly by means of the AIC or BIC criterion, in the case of non-nested scenario. Therefore, in the comparison between one and two-factorial solutions, we relied only on the quality of fit indices, while in the comparison between the two-factorial orthogonal and non-orthogonal solution, we ran a scaled chi-square difference test (Satorra, 2000). CFA was performed with the lavaan package in R (Rosseel, 2012) on a sample size of N = 237 (188 females) with average age of 21.9 (SD = 2.9) years.

Internal consistency. We used four different estimates of internal consistency: ordinal alpha (α_{ord}) (Gadermann, Guhn, & Zumbo, 2012), which is more suitable for ordinal data, McDonald's omega total (ω_t) and hierarchical (ω_h) (Zinbarg, Revelle, Yovel, & Li, 2005), and Revelle's beta (β) (Cooksey & Soutar, 2006; Revelle, 1979), based on the worst split of a test. The coefficients ω and β were utilized for possible violation of tau-equivalency and uni-dimensionality of the scale (Graham, 2006). The polychoric correlation matrix served as an input for all internal consistency estimates. Internal consistency was analyzed using the psych package (Revelle, 2017). The reliability coefficients were estimated on the whole sample.

Correlation analyses. To assess the relationship of ACS and its subscales with measures of trait anxiety, Pearson and Spearman correlation coefficients were chosen based on the prior test of bivariate normality (Henze & Zirkler, 1990) assumption. We also used partial correlation coefficients for testing unique dependencies of the ACS subscales with measures of trait anxiety. The sample for the correlation analysis between ACS and STAI-T consisted of 330 subjects (242 females) with average age of 21.3 (*SD* = 1.9) years. Relationship between ACS and BIS was estimated on a sample size of $N = 130$ (101 females) of average age 22.6 (*SD* = 3.5) years. Subjects from these samples came from the samples used in PCA and CFA.

Results

Principal Component Analysis

Looking at the correlation matrix, we noticed that three items (item 9, 16 and 20) had remarkably poor average inter-item correlation (< .08) with other items compared to the rest of the items (mean polychoric correlation $r_{poly} = .21$) and hence were removed from further analyses. Results of parallel analysis suggested an extraction of two components of which eigenvalues surpassed 95th percentile of their randomly generated counterparts.

The data were suitable for reduction ($\chi^2 = 429.9$, $df = 136$, $p < .001$; $KMO = .787$). Three items with highest loadings on first component (Table 1) were item 3 ("When I'm working hard on something, I still get distracted by events around me"), item 2 ("When I need to concentrate and solve a problem, I have trouble focusing my attention") and item 7 ("When trying to focus my attention on something, I have difficulty blocking out distracting thoughts"). Eleven items in total loaded (>. 35) on the first component and according to the content of these items, we named this component "Focusing". Two items with highest loadings on the second component (Table 1) were item 19 ("It is easy for me to alternate between two different tasks") and item 10 ("I can quickly switch from one task to another"). Seven items in total loaded (> .35) on the second component and

Table 1 *Standardized loadings (pattern matrix) of ACS items after oblique rotation from PCA*

Item	1st comp.	2nd comp.	h^2
1. Hard to concentrate for me when there are noises around	**.65**	.10	.46
2. Need to concentrate on a difficult task/trouble focusing	**.76**	-.01	.58
3. Working on something/distracted by events around me	**.81**	-.19	.37
4. My concentration is good/music in the room	**.41**	**.36**	.35
5. When concentrating/unaware of what's going on	**.51**	.06	.27
6. When reading/easily distracted if there are people talking	**.64**	.02	.42
7. Trying to focus/difficult blocking distracting thoughts	**.73**	.06	.55
8. Hard time concentrating/when excited about something	**.45**	.07	.22
10. Quickly switch from one task to another	-.10	**.76**	.56
11. It takes me a while to get really involved in a new task	**.45**	.20	.28
12. Difficult to coord. attention betw. writing and listening	**.67**	-.10	.43
13. Interested in a new topic very quickly when I need to	.19	**.36**	.19
14. Easy for me to read while I'm also talking on the phone	.23	**.56**	.42
15. I have trouble carrying on two conversations at once	.11	**.35**	.15
17. After being distracted/easily shift my attention back	**.46**	**.40**	.44
18. Distracting thoughts/shift attention away	.27	.25	.16
19. It's easy for me to alternate between two different tasks	-.07	**.79**	.61

Note. h^2 – communality of an item. Correlation of components ($r = .20$). Components explained 40% of overall variability. Bold – items retained for CFA.

we named this component "Shifting". Two items (4 and 17) loaded on both factors. Item 18 did not load sufficiently on any of the components. Cumulatively, components accounted for 40% (Focusing = 26%, Shifting = 14%) of overall variance and were weakly positively related (r = .20).

Confirmatory Factor Analysis

Please note that the sample for the CFA received only those items from the ACS which were retained after the PCA. In the first model, we let all items be saturated by one factor. This model did not yield an adequate fit $\chi^2 = 250.5$, $df = 104$, $p < .001$, $\chi^2/df = 2.4$, $CFI = .88$, $TLI = .86$, $RMSEA = .077$, $SRMR = .087$.

In the second model, we let items 1, 2, 3, 5, 6, 7, 8, 11, 12 be saturated by the first factor (Focusing), items 10, 13, 14, 15, 19 by the second factor (Shifting) and items 4 and 17 were saturated by both factors. Factors were treated orthogonally. This structure showed even lesser appropriateness than the first model with values of $\chi^2 = 290.9$, $df = 102$, $p < .001$, $\chi^2/df = 2.85$, $CFI = .84$, $TLI = .82$, $RMSEA = .088$, and $SRMR = .092$.

In the third model, we directly followed suggestions made by PCA (Table 1). Therefore, the structure of the third model was the same as in the second one with exception of the allowed covariation between factors. This model resulted in considerably better fit, compared to the first model and significantly better fit compared to the second one ($\Delta\chi^2 = 30.01$, $\Delta p < .001$, $\chi^2/df = 1.68$, $CFI = .94$, $TLI = .93$, $RMSEA = .053$, $SRMR = .072$). However, values of fit indices were still unsatisfactory.

Throughout all models, modification indices (MI) suggested substantial model improvements by letting item 4 be saturated only by the Focusing factor and item 12 be saturated by the Shifting factor. Despite the suggestion of MI to let item 17 be saturated by both factors (leading to the best fit), to ensure better interpretability of the model, we let item 17 be saturated only by the Shifting factor. Furthermore, allowing covariation between error terms of items 10 and 19 revealed redundancy of item 10 (i.e., factor loading of item 10 substantially dropped from .50 to .30) and hence, we decided to remove this item. Finally, we let error terms of items 7 and 8 covary, which did not result in considerable decrease in their factor loadings. For the convenience, we adjusted all previously tested models by these modifications and compared them again (Table 2).

Based on these results, we can conclude that acceptable fit is obtained under the assumption that items in the scale are saturated by two

Table 2 *Fit indices of three compared models*

Model	χ^2	df	p	χ^2/df	CFI	TLI	$RMSEA$	$SRMR$	$\Delta\chi^2$	Δp
1. One-factor	156	89	<.001	1.76	.94	.93	.056	.073	-	-
2. Two-factor (orthogonal)	381	89	<.001	4.28	.74	.69	.118	.113	-	-
3. Two-factor (correlated)	102	88	.137	1.18	.99	.98	.026	.059	93.2	<.001

Note. All models have been adjusted according to MIs. Reported value of $\Delta\chi^2$ in third row refers to the comparison between the second and the third model. Moving item 17 to Focusing factor did not result in considerable improvement of the two-factor correlated model ($\chi^2 = 99$, $df = 88$, $p = .198$, $\chi^2/df = 1.14$, $CFI = .99$, $TLI = .99$, $RMSEA = .023$, $SRMR = .058$). Fit indices of the 3rd model with parallel saturation of item 17 were $\chi^2 = 85$, $df = 87$, $p = .524$, $\chi^2/df = 0.98$, $CFI = .99$, $TLI = .99$, $RMSEA = .000$, $SRMR = .054$.

correlated factors. Nine items loaded on the Focusing factor and six items on the Shifting factor (Figure 1). All factor loadings were significant at p < .001 and factors strongly and positively correlated (Table 3).

Internal Consistency of ACS and its Subscales

We conducted three separate analyses of internal consistency. First one analyzed internal consistency of the whole scale and the other two examined internal consistency of each subscale. Reliability estimate for the variance, which is only due to the general factor, yielded poor level with $\omega_h = .46$. However, considering the estimate of reliability for variance, which is due to the general factor as well as specific factors, we can conclude that the scale has good internal consistency $\omega_t = .85$. Coefficient $\beta = .60$ for the whole scale.

The results showed appropriate values of internal consistency indexes $\omega_t = .81$, $\alpha_{ord} = .81$ and $\beta = .61$ for the Focusing subscale. For the Shifting subscale, we observed the following values $\omega_t = .67$, $\alpha_{ord} = .66$ and $\beta = .53$ which might be considered as acceptable. Furthermore, average value of β for both subscales is .585, which is still less than that of β coefficient for the whole scale, hence scores from both subscales may be combined into a single scale (Cooksey & Soutar, 2006).

Correlation of ACS with Measures of Trait Anxiety STAI-T and BIS Scales

Data concerning the correlation between ACS and STAI-T were bivariate normal ($HZ = .58$, $p = .65$). We observed moderate, negative and significant correlation between these two measures, $r(330) = -.48$, $p < .001$, (Figure 2). ACS and BIS data exhibited bivariate non-normality ($HZ = 1.66$, $p < .001$). BIS was also negatively and significantly related to ACS scores, $\rho(130) = -.36$, $p < .001$. Correlation coefficients were not statistically different, $Z = 1.40$, $p = .162$. Next, we partialled out shared variance of Focusing and Shifting subscales and tested their relationship to measures of trait anxiety, respectively.

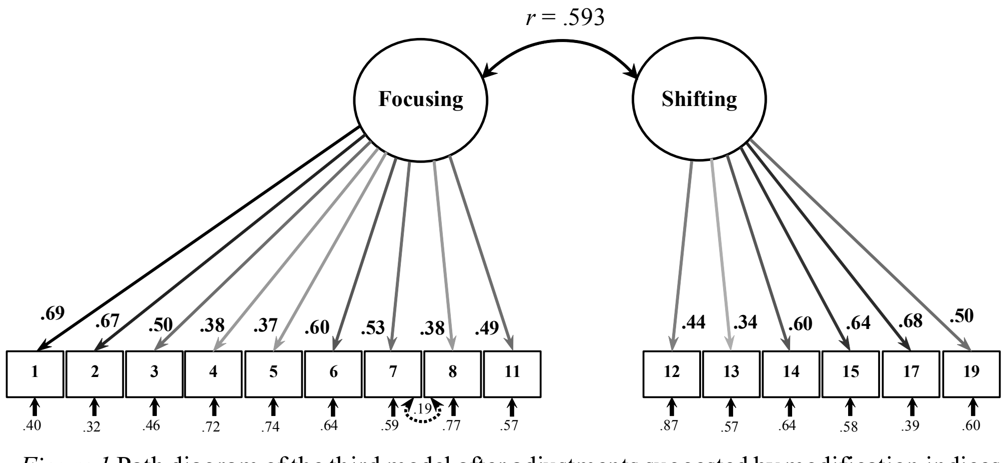

Figure 1 Path diagram of the third model after adjustments suggested by modification indices. Circles represent latent factors. Squares represent individual items of the scale. Numbers above the items denote standardized estimates and numbers below the items denote error terms.

Table 3 *Standardized estimates for two-factorial correlated model of ACS from CFA*

Item	Focusing (SE)	Shifting (SE)
1. Very hard to concentrate for me when there are noises around	.69(.04)	
2. When I need to concentrate on a difficult task/trouble focusing	.67(.05)	
3. Working hard on something/get distracted by events around me	.50(.05)	
4. My concentration is good even if there is music in the room	.38(.06)	
5. When concentrating/focus attention/unaware of what's going on	.37(.06)	
6. When reading-studying/easily distracted if there are people talking	.60(.05)	
7. Trying to focus my attention/difficult blocking distracting thoughts	.53(.05)	
8. I have a hard time concentrating/when excited about something	.38(.06)	
11. It takes me a while to get really involved in a new task	.49(.06)	
12. Difficult to coordinate my attention between writing and listening		.44(.07)
13. I can become interested in a new topic quickly when I need to		.34(.07)
14. Easy for me to read-write while I'm also talking on the phone		.60(.07)
15. I have trouble carrying on two conversations at once		.64(.05)
17. After being distracted/easily shift my attention back		.68(.06)
19. It's easy for me to alternate between two different tasks		.50(.07)

Note. SE – robust standard errors of estimates. Factors were strongly positively related (r = .593, p < .001). All estimates were significant at p < .001.

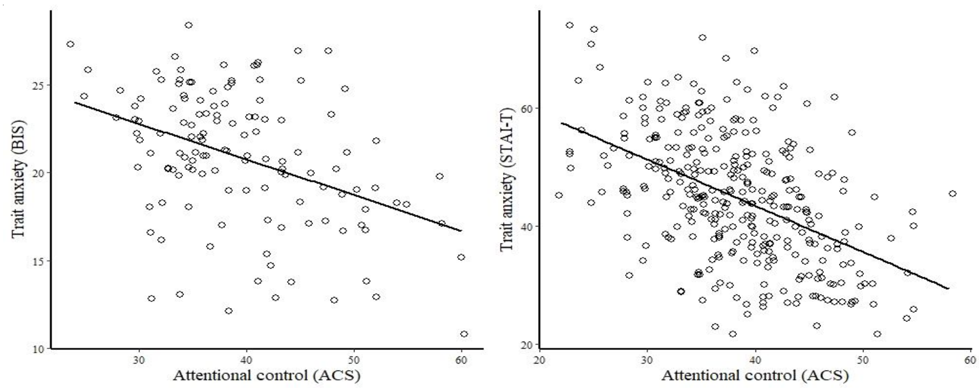

Figure 2 Left: correlation between ACS and BIS ($\rho(130) = -.36$ $p < .001$); Right: correlation between ACS and STAI ($r(330) = -.48$, $p < .001$). Small jitter (noise) was added to the plots to disperse overlaying data points.

After controlling for Shifting, the Focusing subscale was significantly, moderately and negatively related to BIS, $\rho(130) = -.31, p < .001$. On the other hand, the Shifting subscale did not show significant relation to BIS after controlling for Focusing, $\rho(130) = -.09, p = .275$. STAI-T also yielded moderate, negative, and significant relationship with Focusing subscale, $r(330) = -.34, p < .001$. While Shifting subscale did not correlate with BIS, it did show weak negative association with STAI-T, $r(330) = -.18$, $p < .001$, however we should consider the size of the sample on which this correlation was estimated.

Additionally, we tested the samples for possible gender differences in the measured variables. No differences were observed either for the BIS, $U = 1310.5, Z = -0.873, p = .383$, or for the ACS, $t(472) = 1.181, p = .238$, and STAI-T, $U = 10545.0, Z = -0.134, p = .893$.

Discussion

The current study tested psychometric properties of Slovak adaptation of the Attentional Control Scale by means of the PCA and CFA and its hypothesized relationship with trait anxiety. The PCA, using a parallel analysis, suggested two components, which indeed resembled two assumed factors of Focusing and Shifting. Importantly, some of the items that were excluded were also reported in other studies as being problematic (Judah, Grant, Mills, & Lechner, 2014; Ólafsson et al., 2011). For example, Fajkowska and Derryberry (2010) examined the content and internal validity of ACS items by including them amongst other statements referring to formal characteristics from a temperament inventory measuring endurance. Four recruited judges were in good agreement, however they judged items 9 and 20 from the ACS incorrectly and replaced them with items measuring temperamental endurance. Item 9 has been also reported as problematic in the Icelandic version of the ACS (Ólafsson et al., 2011). Furthermore, reported studies differed fairly in their criteria concerning the retention of items in the final structure. In our study, we used only one criterion regarding the PCA, namely that factor loading must be equal or higher than .35. However, in a study conducted by Judah, Grant, Mills, and Lechner (2014), the authors used rather strict constraints of factor loadings of at least .40, with minimum factor loading difference between two components being .25. Although their procedure resulted in considerably smaller number of retained items (12), the main purpose of this approach was to distinguish both factors from each other as much as possible. Nevertheless, our results from the PCA supported two-factorial structure of the ACS.

Using the CFA on an independent sample, we found that from amongst one-factorial and two-factorial orthogonal solutions, the two-factorial model with correlated factors suggested by PCA fitted the data best. This is consistent with the hypothesized structure of the ACS, and although initial configuration of this model did not yield satisfactory fit across all fit indices, subsequent adjustments suggested by modification indices proved to be sufficient to meet their required values. Please note that these adjustments are of no theoretical value regarding the factor structure of the scale (i.e., slight change of saturation pattern and allowing covariation between error terms of items is still consistent with the hypothesized structure of the scale). Noteworthy are also the results that the one-factorial model exhibited better fit to data than the two-factorial orthogonal model and its fit indices could be considered as acceptable (after MI adjustments). Despite this, only the third model reached all required values of fit indices and, also, as the only one did not result in significant χ^2 statistics. However, there were some differences in comparison to other studies. For example, we let item 12 be saturated by the Shifting factor instead of the Focusing factor and even though this item was saturated by the Focusing factor in previous studies, we suggest that the content of this item (Appendix) resembles more the ability to shift rather than to focus attention. Also in comparison to the previous studies, we found item 10 to be redundant when allowed to correlate with item 19. Again, if we look at the content of these two items, we find that these statements are indeed very similar. Therefore, we removed item 10 from the final model (3rd model). Finally, probably most different from factorial solutions obtained in the previous studies are findings concerning saturations of items 11 and 17. In the Icelandic version (Ólafsson et al., 2011), items 11 and 17 loaded on the Shifting factor. In a study by Judah, Grant, Mills, and Lechner (2014), item 11 did not reach sufficient loading, although it loaded more on the Focusing factor and item 17 was saturated only by the Shifting factor. In our study, the results of PCA and CFA (Tables 1 and 3) showed clearly greater contribution of the Focusing factor to item 11. We suspect that this is a result of specific wording in the Slovak language used in our translation. Instead of the word "involve", we used "focus". Similarly, the same might be true of item 17. This item showed approximately similar saturation by both factors in both PCA and CFA. Again, instead of "shift my attention back" we used "focus back on". Therefore, the subjects might have attributed this statement equally to the ability to reallocate attentional resources as well as to the ability to focus these resources.

In previous studies (Abasi, Mohammadkhani, Pourshahbaz, & Dolatshahi, 2017; Judah, Grant, Mills, & Lechner, 2014; Ólafsson et al., 2011) Cronbach's alpha was the only measure of internal consistency provided for the ACS. It is well documented that this coefficient provides only a lower bound and biased estimate of internal consistency (Graham, 2006; Zinbarg et al., 2005), since the tool does not measure a homogenous construct with the same precision. We observed that this is indeed the case of ACS, which does not resemble only one homogeneous construct and furthermore, as PCA and CFA showed, contains items varying in factor loadings, and thus we proceeded with alternative measures of internal consistency. For the whole scale, we found internal consistency to be at a good level, as indicated by ω_t. For individual subscales, estimates of internal consistency were fairly similar to those reported in the previous studies with Focusing subscale being the more stable construct. Shifting subscale, on the other hand, indicated a possible non-homogeneity. Indeed, certain concerns about generalization of various shifting forms have been previously raised (Ravizza & Carter, 2008). To tackle with the dissociation of different aspects of cognitive flexibility is beyond our subject, but we suspect that it is necessary to differentiate between the ability to alternate between two tasks with known and established rules (e.g., item 19 "It's easy for me to alternate between two different tasks") and the ability to adapt to a new set of rules (e.g., item 13 "I can get interested in a new topic very quickly when I need to").

We consistently observed a negative association between ACS and measures of trait anxiety. Although ACS correlated with STAI-T more strongly than with BIS, the correlation coefficients were not statistically different. However, Fajkowska and Derryberry (2010) reported similarly strong association between BIS and ACS as we evidenced between ACS and STAI-T. Importantly, one might suspect that these correlations could have been confounded by gender differences regarding the levels of trait anxiety or attentional control. However, direct comparisons of genders revealed no differences in all three measures. Moreover, we observed that only the Focusing subscale is significantly inversely related to trait anxiety. This is consistent with reports from a study by Ólafsson et al. (2011), which found lower levels of Focusing to be associated with higher levels of trait anxiety and lower levels of Shifting to be associated with symptoms of depression. Although we found a weak correlation between the Shifting subscale and STAI-T, we did not measure levels of depressive symptoms and so we could not control for their shared variance, which is generally assumed between these two constructs.

Regarding the predictive and convergent validity of ACS, we may claim that there is certainly more research to be done. However, as authors of the scale (Derryberry & Reed, 2002) demonstrated in their experiments, the scale could predict reduction of attentional bias toward threatening stimuli in highly anxious individuals. Recently, Judah, Grant, Mills, and Lechner (2014) found ACS together with its subscales to be related to performance-based measure of working memory capacity (Letter-number sequencing). More specifically, only the Shifting subscale showed positive correlation with this measure. Furthermore, utilizing the Mixed Antisaccade Task, Judah, Grant, Mills, and Lechner (2014) observed positive correlation between Focusing, antisaccade performance, and prosaccade latency. Shifting was found to be related to switch-trial performance on the Mixed Antisaccade Task. However, the scale's predictive and convergent validity should be tested against other executive control tasks, considering the multifaceted nature of executive functions as well (e.g., Miyake et al., 2000). For example, Derryberry (2002) discusses unpublished experiments, which showed high ACS scores to be related to better performance in stop-signal task or to better ability to inhibit dominant conceptual association in a priming task.

Taken together, the attentional control scale might prove useful due to its quick administration capabilities. For example, easy to use, the property of the scale might prove useful in research of stress or anxiety effects on cognitive control (i.e., controlling for individual differences in the ability to reduce attentional bias). Finally, we should address the limitations of the current investigation. Perhaps the most noticeable one is the inequality of genders in both samples, however, we did not observe any possible mediating effect of this variable on the relationship between ACS and measures of trait anxiety. Moreover, as our sample consisted mainly of university students, the structure of the ACS should be tested also on a sample from a broader general population. Problematic is also the saturation of item 17. Given that letting this item be saturated either by the Focusing, or by the Shifting factor results in comparatively satisfying solutions, it makes the interpretation of this item arbitrary in terms of whether it is reflecting more the ability to focus or to shift attention. Perhaps a change in translation of this item should be considered in future investigations utilizing the Slovak version of ACS, leaning the meaning of this item only to the shifting ability.

Summary

The structure of the first Slovak version of the ACS supports the notion of this tool as being saturated by two positively related factors. The scale can be used to assess attentional control as a unitary construct or to assess individual components of Focusing and Shifting as well. Values of internal consistency for the whole scale and separate subscales promotes the usage of this tool mainly for research purposes. The next step should be to verify its validity either in correlational or experimental studies. Further investigation of its psychometric properties should be also conducted on a sample from the general population.

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APPENDIX

Attentional Control Scale in Slovak Language

Item
1. Je pre mňa veľmi ťažké sústrediť sa na náročnú úlohu, keď je v mojom okolí hluk. (R)
2. Keď sa potrebujem sústrediť a riešiť úlohu, je pre mňa ťažké zamerať na ňu moju pozornosť. (R)
3. Keď na niečom ťažko pracujem, veci okolo mňa ma neustále vyrušujú. (R)
4. Dokážem sa dobre koncentrovať, aj keď v mojom okolí hrá hudba.
5. Keď sa sústredím, dokážem tak zamerať svoju pozornosť, že prestanem vnímať, čo sa deje naokolo.
6. Keď čítam alebo sa učím a v miestnosti sú ďalší ľudia, ktorí sa rozprávajú, ľahko ma to vyruší. (R)
7. Je pre mňa ťažké blokovať rušivé myšlienky, keď sa snažím na niečo sústrediť. (R)
8. Je pre mňa ťažké koncentrovať sa, ak som nadšený/á alebo vzrušený/á. (R)
9. Keď sa sústredím, ignorujem pocity hladu alebo smädu.
10. Dokážem rýchlo preskočiť z jednej úlohy na druhú.
11. Chvíľu mi trvá, kým sa skutočne začnem sústrediť na novú úlohu. (R)
12. Keď si počas prednášky píšem poznámky, je pre mňa náročné koordinovať moju pozornosť medzi počúvaním a písaním. (R)
13. Keď je to potrebné, dokáže ma nová úloha zaujať veľmi rýchlo.
14. Je pre mňa jednoduché niečo čítať alebo písať, aj keď popri tom telefonujem.
15. Mám problém viesť dve konverzácie naraz. (R)
16. Je pre mňa náročné prísť rýchlo s novým nápadom. (R)
17. Ak ma niečo vyruší, dokážem sa opäť ľahko sústrediť na to, čo som robil/a predtým.
18. Keď mi na um zíde rušivá myšlienka, je pre mňa ľahké odpútať od nej moju pozornosť.
19. Pri práci je pre mňa jednoduché preskakovať medzi dvoma odlišnými úlohami.
20. Pri riešení úlohy je pre mňa ťažké prestať o nej rozmýšľať jedným spôsobom a pozrieť sa na ňu z iného uhla pohľadu. (R)

Note. Only items with designated number in bold were retained after PCA and CFA. (R) – item with reversed scoring. Items are scored on a 4-point Likert scale (1 – *almost never*; 2 – *sometimes*; 3 – *often*; 4 – *always*) (Derryberry & Reed, 2002).