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Stability of the factor structure of Barrat's Impulsivity Scales for children across cultures: A comparison of Spain and Colombia

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Despite the great importance of impulsivity in many childhood and adolescence disorders, there are few self-reports on child impulsivity. Recently, a modified version of Barratt's BIS-11 questionnaire adapted for children has been developed, which may be useful in assessing this personality dimension. The present study reports an adaptation of this questionnaire in a different culture (Colombia) and assesses the degree of convergence between the factor structures of both adaptations using consensus oblimin rotation. The results indicate not only that the factor structure of the test remains stable across both adaptations, and that two of the three scales in the Colombian version show acceptable reliabilities, but also that cultural and linguistic issues are important in test adaptation even when the same language is used.

Estabilidad de la estructura factorial de la escala de impulsividad para niños de Barratt a través de culturas: una comparación entre España y Colombia. A pesar de la gran importancia de la impulsividad en múltiples patologías de la infancia y la juventud, existen pocos cuestionarios de impulsividad para niños. Recientemente se ha desarrollado una versión modificada del cuestionario BIS-11 de Barratt adaptado para niños que puede ser útil en la evaluación de esta dimensión de personalidad. Este estudio presenta una adaptación de este cuestionario a una cultura distinta (colombina) y evalúa el grado de convergencia entre las soluciones factoriales de ambas versiones del test utilizando el método de rotación consensos oblimin. Los resultados indican una elevada estabilidad de la solución factorial a través de las versiones y una fiabilidad satisfactoria para dos de las tres escalas en la adaptación colombiana. Por otra parte se pone de manifiesto la importancia de los aspectos culturales y lingüísticos en la adaptación de un test incluso cuando se trata de la misma lengua.

In recent years there has been an increase in the amount of research on impulsivity. One of the main reasons for this is that impulsivity plays a prominent role in understanding and diagnosing various forms of psychopathology, specially those linked to the lack of impulse control, such as aggression, substance abuse, etc. (Barratt & Slaughter, 1998; Whiteside & Lynam, 2001; McMurran, Blair, & Egan, 2002; Scarpa & Raine, 2002; Vigil-Colet, Morales-Vives, & Tous, 2008). The importance of impulsivity during childhood and adolescence has been established and related to a wide variety of externalising and internalising pathologies, such as hyperactivity, aggression, learning problems, anxiety disorders, depression etc. (Fink & McCown, 1993; Zaparniuk & Taylor, 1997; American Pyshological Association, 2000; Willcutt & Pennington, 2000; Summerfeldt, Hood, Antony, Richter, & Swinson, 2004; Cataldo, Nobile, Lorusso, Battaglia, & Molteni, 2005; Jensen, Youngstrom, Steiner, Findling, Meyer et al., 2007).

Despite the great importance of impulsivity in children, there is a lack of self-report measures for this factor. Impulsivity in children is often measured with rating scales completed by other individuals such as parents or teachers or with behavioural tasks. Self-reports are not as frequently used with children as with adults because it is assumed that they are less accurate at assessing their own behaviours (Fink & MacCown, 1993). Nevertheless, there is evidence to suggest that children between 8 and 12 years old are better informants than parents (Rapee, Barrett, Dadds, & Evans, 1994; Muris, Merckelbach, Van Brakel, & Mayer, 1999; Cosi, Canals, Hernández-Martínez, & Vigil-Colet, 2010). Furthermore, self-report and behavioural tasks seem to measure different components of impulsive behaviour, which indicates that both kind of measures need to be used when assessing impulsivity (Reynolds, Ortengren, Richards, & de Wit, 2005).

Few impulsivity self-report scales have been specifically designed for children. Two that have are Eysenck's I6 impulsivity scale and the children's adaptation of Dickman's impulsivity questionnaire (DII-c) (Eysenck & Eysenck, 1980; Eysenck, Easting, & Parson, 1984; Brunas-Walgstaff, Tilley, Verity, Ford, & Thompson, 1997). The I6 questionnaire was developed by Eysenck and Eysenck (1980) to measure two specific dimensions related to impulsivity (impulsiveness and venturesomeness) which have shown good internal consistencies both for the original English version, and the German and Spanish adaptations although the latter version was tested only in an adolescent sample (Eysenck & Eysenck, 1980; Silva, Martorell, & Clemente, 1987; Stadlet & Janke, 2003).

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The DII-c scales tried to replicate the dimensions of dysfunctional impulsivity (DI) and functional impulsivity (FI) proposed by Dickman (1990). This inventory, however, showed poor internal consistency with children: in particular, FI was below the commonly accepted standards for a test to be considered sufficiently reliable (Brunas-Walgstaff et al., 1997; Cosi, Morales-Vives, Canals, Lorenzo-Seva, & Vigil-Colet, 2008).

Cosi, Vigil-Colet, Canals, & Lorenzo-Seva (2008) proposed a third approach for assessing impulsivity in children using self-reports: adapting Barrat's impulsivity scales, one of the most widely used self-reports for adults in impulsivity assessment (Patton, Stanford, & Barratt, 1995). Fossati, Barratt, Acquarini, & Di Ceglie (2002) developed an adolescent's version of this scale known as BIS 11-a, consisting of short and very simple items that may also be easily understood by children. Cosi et al., (2008) adapted it into Spanish and found a three-factor structure for the Barratt impulsivity scale for children (BIS-c). This structure comprised the factors of motor (Im), lack of planning (Inp) and cognitive (Ic) impulsivity. Two of the scales showed good or sufficient reliabilities but Ic was below acceptable cut-off points, and showed a reliability of α= 0.60. Taking this into account, Cosi, Canals, & Vigil-Colet (2008) developed an improved version of the Ic scale by adding three new items, which improved the reliability (α= 0,70). They also found significant relationships between the BIS 11-c scales and measures of aggression and scholastic performance, which have been often associated with impulsivity. Some evidence of the validity of BIS 11-c was therefore given.

The aim of the present study is twofold. The first objective is to adapt the BIS 11-c to a culture other than the Spanish one and to assess its factorial structure. The second objective is to compare the factor structure obtained from the Spanish sample with the factor structure obtained from the Colombian sample. Adapting impulsivity tests so that they can be applied in places such as Colombia acquires particular importance if it is taken into account that the levels of aggression in these countries are high and that impulsivity has an important role in their origin.

The second objective addresses the question of that the extent to which both instruments measure the same constructs in exactly the same way across cultures. As Byrne (Byrne & Watkins, 2003; Byrne, 2008) has pointed out, this is a key issue in cross-cultural adaptation because it refers to the extent that a group of nested equalities between both instruments are accomplished. These equalities refer to such key aspects such as whether both adaptations have equal factor structures in the sense that the number of factors and pattern of indicator–factor loadings are identical across groups (configurational invariance), the extent to which relationships between the scores and latent variables are equivalent (equality of factor loadings, also called weak factorial invariance) for both cultures and the strong factorial invariance in the case of equal indicator intercepts (Meredith, 1993).

One way of testing whether they measure the same constructs is to use the means and covariance structures (MACS) approach proposed by Byrne (Byrne & Stewart, 2006; Byrne, 2008). This approach proposes a set of steps that first establishes the separate model fit for each group, then tests a configurational invariance model as a baseline model and finally tests different nested and more restricted models. Nevertheless, several authors have pointed out that conventional Confirmatory factor analysis (CFA) goodness of fit criteria are too restrictive when applied to most tests, especially in personality research (Ferrando & Lorenzo-Seva, 2000; Asparohov & Muthén, 2009).

One solution to this problem is to use the Exploratory structural equation modelling approach (ESEM) proposed by Asparohov & Muthén (2009). This procedure uses a rotated exploratory factor analysis (EFA) measurement model that is applied to a structural equation model, yielding the goodness of fit indexes usually obtained in SEM, and allowing a step approach to multigroup analysis such as the one proposed by Byrne (2008). Nevertheless this method has an important limitation, because the cut-off values for goodness of fit indexes in ESEM have not yet been established because ESEM studies involve a greater number of estimated parameters than the CFA approach. (Marsh et al., 2009). Furthermore, as Asparohov & Muthén (2009) pointed out ESEM is only able to estimate the measurement invariance of factors conjointly and not the specific invariance for one factor.

A second approach to the problem is based on EFA and involves the simultaneous rotation of the loading matrices obtained in two samples to show a mixture of simplicity and optimal agreement between them. The method used is Consensus Direct Oblimin with γ= 0 (Lorenzo-Seva, Kiers, & Ten Berge, 2002). In this rotation method, the loading matrices are obliquely rotated together to satisfy two criteria: simplicity and agreement. This method will allow us to assess the degree of agreement between the different factorial solutions for each factor and determine if there is any lack of agreement, which items are responsible for it. It should be mentioned that this analytical approach has been successfully used to compare the factorial structures obtained in personality tests across different cultures (i.e., Vigil-Colet, Lorenzo-Seva, Codorniu-Raga, & Morales, 2005), and similar methods have also been applied in cross-cultural adaptation using other rotational procedures such as Procustes rotation (i.e., Balluerka Gorostiaga, Alonso-Arbiol, & Aramburu, 2007). Nevertheless, being an exploratory approach, it does not provide a goodness of fit index and configurational invariance, weak factorial invariance etc. cannot be tested using nested models.

Taking into account the advantages and limitations of both approaches we will apply both in order to complement the information given by the consensus oblimin method (COM) on the overall (scales) and specific (items) congruences with the goodness of fit indexes and different nested models of the MACS approach using ESEM.

Method

Participants

The participants were 616 children (306 boys and 310 girls) aged between 8 and 16 years with a mean of 13 years (SD=2 .38) from one private and two state schools in Bucaramanga (Colombia). They were from medium-low and low social classes, respectively. The cases with missing data (n= 78) were removed from the analysis. Bucaramanga is a town of 716,000 in habitants and the capital of the Santander region.

The sample used for comparison (Cosi et al., 2008) consisted of 413 children (186 boys and 227 girls) aged between 9 and 13 years with a mean of 11 years (SD= .92). The children came from thirteen schools in Reus (Spain), which were randomly chosen from the state schools and state-subsidized private schools in the town. Reus is a medium-sized town of 100,000 inhabitants.

Procedure

The head teachers of each school were informed about the nature of the research before they authorized the test. Then, the instrument was applied to all the classrooms of the centers to volunteer students guaranteeing anonymity. The test was administered by a psychologist in a collective way to groups around 30 individuals.

Instruments

Barrat Impulsivity Scales-11 for children (Cosi et al. 2008): The BIS-11c consisted of 30 items with a 4-point response format (Never/Almost Never, Sometimes, Often, Always / Almost always). Answers were scored with 0, 1, 2 and 3, respectively. The questionnaire measures motor (Im), non planning (Inp) and cognitive (Icog) impulsivity, and shows internal consistencies of 0.80, 0.73, and 0.68, respectively.

Although test adaptation usually involves translating the test from one language to another, cultural and language differences that affect test scores are not only a translation (Hambleton, 2005). As a consequence, although Spain and Colombia share the same language, various cultural differences and dialectal variants made it necessary to analyze the items of the original Spanish scale to see if they were appropriate for administration in Colombia. With this purpose, one Colombian and two Spanish psychologists with experience in test adaptations analyzed the degree to which the BIS 11-c items were culturally and linguistically suitable for administration in a Colombian sample. After various modifications in the wording of 21 of the 26 items—for instance we changed «perder los nervios» for «me desespero con facilidad» because the first expression is not commonly used in Colombia—we sent the test to eight Colombian teachers who gave their opinion on whether the items would be understood by Colombian children and, if not, proposed an alternative version of the items. This second phase involved only slight changes such as the use of «organize» instead of «plan». Table 1 shows the Spanish version of BIS 11-c and the alternative items for the Colombian version.

Table 1

Items of the Spanish version of BIS 11-c. Between brackets alternative items for the Colombian adaptation. Mean, Standard deviation (s.d.), ítem-total correlations (r_{jx}) and Cronbach's alpha if the element is removed (α-j)

Item	Media	d.t.	r_{jx}	α-j
1. Planifico las cosas que hago (organizo las cosas que hago)	1.52	.89	.41	.69
2. Hago las cosas sin pensarlas	.85	.79	.34	.72
3. Decido las cosas rápidamente (decido rápidamente)	1.42	.92	.22	.47
4. Cuando mis amigos me preguntan algo, puedo responder rápidamente	1.69	.98	.29	.43
5. Me cuesta estar atento (me cuesta trabajo estar atento)	.88	.87	.37	.71
6. Pienso rápidamente (pienso con rapidez)	1.63	.92	.37	.38
7. Planifico mi tiempo libre (organizo mi tiempo libre)	1.57	1.11	.44	.69
8. Pierdo los nervios con facilidad (me desespero con facilidad)	1.20	1.00	.33	.72
9. Me concentro rápidamente	1.70	.98	.26	.45
10. Ahorro todo lo que puedo (ahorro lo que más puedo)	1.68	1.09	.26	.72
11. Me gusta pensar detenidamente las cosas (me gusta pensar bien las cosas)	2.19	.91	.52	.67
12. Hago proyectos para el futuro (hago planes para el futuro)	2.13	.99	.29	.70
13. Digo las cosas sin pensar (digo cosas sin pensar)	.99	.83	.45	.71
14. Soy de los primeros en levantar la mano en clase cuando el profesor hace una pregunta	.97	.87	.21	.48
15. Cambio a menudo de ideas (cambio con facilidad mi manera de pensar)	1.21	.97	.23	.73
16. Actúo impulsivamente [sin pensar] (actúo sin pensar)	.72	.80	.44	.71
17. Me distraigo con facilidad cuando tengo un problema complicado (cuando estoy haciendo algo que requiere concentración, me distraigo con facilidad)	1.00	.88	.41	.71
18. Me dejo llevar por mis impulsos	.96	.88	.37	.71
19. Me gusta pensar las cosas	2.19	.90	.54	.66
20. Cambio frecuentemente de amigos (cambio con frecuencia de amigos)	.55	.88	.15	.74
21. Compro las cosas sin pensar (compro cosas sin pensar)	.58	.84	.40	.71
22. Soluciono los problemas uno a uno (soluciono los problemas uno por uno)	1.96	1.00	.39	.70
23. Gasto más de lo que puedo (gasto más de lo que tengo)	.77	.95	.38	.71
24. Cuando pienso en algo me distraigo fácilmente (cuando estoy pensando en algo me distraigo con facilidad)	.97	.87	.43	.71
25. Estoy inquieto en el cine o en clase (me cuesta trabajo quedarme quieto en el cine o en clase)	.84	.93	.34	.72
26. Planifico mis actividades (organizo mis actividades)	1.93	.97	.55	.66

Data analysis

Consensus Oblimin Method: The data obtained in the Colombian population was factor analyzed as follows: first the polychoric correlation matrix was computed, and then Unweighted Least Squares was performed so that three factors were retained. In order to apply the rotation method described above we used two factor structures: one reported by Cosi, Canals, & Vigil-Colet (2008) obtained in a sample of 456 children, and another obtained in our own sample (i.e., the Colombian sample).

To assess the degree of similarity among factor solutions we computed the averaged Tucker's congruence index (Tucker, 1951) for items, factors and overall factor solution (for the details of the computing see, for example, Chan, Ho, Leung, Chan, & Yung, 1999). A threshold of .85 was used for assess congruence (Lorenzo-Seva & ten Berge, 2006). The degree of factor simplicity of each one of the factor solutions was described by the Loading Simplicity (LS) index (Lorenzo-Seva, 2003).

The steps of the analysis were as follows: first, all the factor solutions were simultaneously rotated by Consensus Direct Oblimin; and then the averaged congruence values for items among samples were computed, and the need to eliminate items was assessed.

Finally, the internal consistency reliabilities of the factor scales were assessed using Cronbach's index.

Factorial Invariance using ESEM: As we have described above this method involves to test the multigroup equivalence of a test using the ESEM approach developed by Asparohov & Muthén (2009). To this end we tested the 3-factor model for each group separately. We then tested the configural invariance that was used as the baseline model for other nested models which tested weak and strong factorial invariance.

We analysed the data using SPSS 17.0, Mplus 5.1 and FACTOR (Lorenzo-Seva & Ferrando, 2006). We used FACTOR for EFA as well as SPSS because it enabled us to use polychoric correlation matrices and gave complementary analyses which are not provided by SPSS.

Results

Tables 1 and 2 show descriptive statistics for the BIS 11-c items and scales. As can be seen, the Colombian sample showed higher values than the Spanish sample for all scales although the differences were significant only for Im (d= 0.22) and Inp (d= 0.43). Finally, all the scales, with the exception of Icog ($F_{(1,1120)}$ = 6.54; $p<0.05$), showed the same variance in the Spanish and Colombian samples. We found that neither sex nor age had any effect on BIS 11-c scales.

COM showed that the overall factor congruence value after the rotation was .94. If we follow the guidelines by Lorenzo-Seva and ten Berge (2006), this congruence value suggests a fair similarity among factor structures. The factor congruence for Icog was .90 (i.e., a fair similarity), whereas the congruence values between the Inp and Im factors were .95 and .97, respectively. These values suggest that these factors can be considered equivalent. Table 3 shows the factor structures in both samples after Consensus Oblimin rotation, and the congruence between items. Most of the items showed a congruence larger than .95, and six items (4, 5, 8, 21, 22 and 23) showed congruence values between .85 and .95 (i.e., a fair similarity). Finally, item 20 showed a congruence of .79, which suggests that it should be deleted. However, the inspection

Table 2
Descriptive statistics for BIS 11-c scales in Colombian and Spanish samples (Cosi et al., 2008)

Scale	Sample	Mean	Std. dev.
Cognitive	Colombia	7.41	**2.73**
	Spain	7.14	**3.09**
Non planning	Colombia	**15.21**	4.57
	Spain	**13.29**	4.50
Motor	Colombia	**12.95**	5.42
	Spain	**11.65**	5.51

p<0.01

Table 3
Factor structures in both samples after consensus. Oblimin rotation and congruence between items. The largest loading values of each item are printed in bold face

Items	Spanish sample			Colombian sample			Congru-ences
	Cognitive	Non-Planning	Motor	Cognitive	Non-Planning	Motor	
1	.07	**-.53**	.04	.11	**-.46**	-.05	.98
2	.03	.06	**.61**	.05	.15	**.44**	.97
3	**.42**	.07	.15	**.51**	.02	.24	.99
4	**.34**	-.10	-.02	**.41**	.06	-.12	.89
5	-.28	-.04	**.54**	-.04	.07	**.53**	.90
6	**.58**	-.05	.03	**.63**	.00	-.03	.99
7	.06	**-.51**	.11	-.06	**-.57**	.09	.97
8	-.09	-.21	**.45**	.06	-.09	**.49**	.93
9	**.43**	-.22	-.24	**.41**	-.15	-.24	.99
10	-.09	**-.41**	-.10	.01	**-.27**	-.12	.96
11	-.11	**-.55**	-.04	-.05	**-.61**	-.08	.99
12	.17	**-.42**	.24	.11	**-.32**	.09	.98
13	.13	.16	**.66**	-.04	.09	**.57**	.97
14	**.41**	-.16	-.17	**.34**	-.20	-.08	.97
15	-.03	-.18	**.47**	-.04	-.17	**.34**	.99
16	.05	.13	**.68**	.01	.09	**.61**	1.00
17	-.15	-.11	**.53**	-.11	-.03	**.56**	.99
18	.10	.05	**.45**	.04	.13	**.49**	.98
19	-.02	**-.60**	-.22	-.04	**-.63**	-.18	1.00
20	.03	.02	**.30**	.19	-.27	**.43**	.79
21	.15	.06	**.48**	.03	-.07	**.59**	.94
22	.07	**-.49**	-.17	.09	**-.51**	.01	.94
23	.17	.04	**.42**	-.01	-.05	**.57**	.91
24	-.16	-.06	**.44**	-.20	-.04	**.53**	1.00
25	-.09	-.10	**.52**	.00	.03	**.53**	.96
26	.10	**-.63**	.09	.03	**-.69**	.03	.99

of the loading values in the Colombian sample shows that the salient loading was in the same factor as in the Spanish sample so we finally decided to keep it. However, if this problem is replicated in other studies, it could finally be deleted. We concluded that both versions of the test were congruent among samples.

Table 4 shows the consensus inter-factor correlation values for the reduced test. As we can see Im and Inp are positively related while Ic shows a negative pattern of relationships with them. This pattern of relationships is not surprising because impulsive individuals often show inhibition deficits and do not predict consequences (both inhibition and prediction are related to Im and Inp). On the other hand, Ic is more related to quick decisions when this strategy is appropriate than to non reflexive responses.

To assess the level of factor simplicity, we computed the LS index. The values were .34 and .46 for the Spanish and the Colombian samples, respectively. Actually, the index showed that the simplest factor solution was the one obtained in the Colombian sample.

Table 5 shows the goodness of fit indexes for the different models tested with ESEM. Initial testing of the 3-factor model showed that the overall fit was acceptable: it was good in the Spanish sample and marginally good in the Colombian one. In the latter case the Root mean-square error of approximation (RMSEA) was quite good but the Comparative fit index (CFI) was slightly lower than the standard cut-off values of 0.08 and 0.9, respectively (Bentler, 1990; Browne & Cudeck, 1993).

When configurational invariance was tested, we again found a good RMSEA index (RMSEA= 0.045) and a C.F.I. value close to the 0.9 cut-off value (CFI= 0.88). Nevertheless, as we have stated in the introduction, it is not clear whether these cut-off values are appropriate for ESEM. A more restrictive model (weak factorial invariance) led to a slightly worse fit (RMSEA= 0.049; CFI= 0.85) while the more restrictive model with strong factorial invariance showed a fit that was clearly worse (RMSEA= 0.061; CFI= 0.73), which indicates that the differences in impulsivity between Colombian and Spanish children may be explained by the fact that the instrument lacks strong factorial invariance and not by true differences in impulsivity between the two populations.

Table 4
Consensus inter-factor correlation values

	Cognitive	Non-planning
Non-planning	-.24	–
Motor	-.15	.35

Table 5
Tests for Spanish, Colombian and invariance models of BIS 11-c: goodness of fit statistics using ESEM

Model	χ^2	d.f.	r.m.s.e.a	r.m.s.e.a. 90% CI	cfi
Spain	490.2	250	0.043	0.037 - 0.048	0.910
Colombia	503.37	250	0.045	0.039 - 0.051	0.86
Configurational	1029.87	500	0.045	0.041 - 0.049	0.88
Measurement	1261.8	569	0.049	0.045 - 0.052	0.85
Equal interceps	1.749	592	0.061	0.058 - 0.065	0.73

Finally, the internal consistency reliabilities of the factor scales are presented in Table 6. As can be observed, the reliabilities are systematically lower in the Colombian sample, although the values are only unacceptable for the Ic scale.

Table 6
Reliabilities of the scales between samples.
The 95th confidence intervals are printed in brackets

Factor	Spanish sample	Colombian sample
Cognitive	.68	.59
	(.64; .72)	(.55; .62)
Non-planning	.73	.72
	(.69; .76)	(.68; .75)
Motor	.80	.74
	(.77; .82)	(.70; .77)

Discussion

The results described above indicate that the factor structure obtained in the BIS 11-c is also found when it is adapted into another culture, at least in terms of configurational invariance. COM has showed that the factor congruence between the Spanish and Colombian versions is high with few items showing inappropriate or borderline congruences. On the other hand ESEM, with the cut-off limitations described above, has showed that both adaptations seem to have the same number of factors and pattern of indicator–factor loadings. Furthermore, the equality of factor loadings (weak factorial invariance) cannot be totally rejected because although authors such as Cheung and Rensvold (2002) consider that a difference in CFI greater than 0.01 is an indicator of worse fit, others such as (Little, 1997) have proposed a value of 0.05, which is higher than the ΔCFI= 0.03 obtained when the weak factorial invariance model is compared with the configurational invariance model in our data. To sum up, the results seem to confirm that both adaptations, which have the same items, have a three-factor structure and show the same relationship between latent variables and scale scores.

Finally the bad fit to the strong invariance model shows that the differences between the Colombian and Spanish sample may not show true differences in impulsivity between the two samples because these differences may also be due to test bias, which under- or overestimates impulsivity levels in one of the samples, procedure differences or differences in sample characteristics. In this sense we have to point one of the limitations of this study that is the incidental nature of the Colombian sample which also may explain the differences between founded between children of both nations. Nevertheless, this is a minor point if, as in this case, we are more interested in obtaining a measure of impulsivity that can be used to predict behaviour problems such as aggression, etc. than in cultural comparisons across nations.

These results seem to indicate that BIS 11-c may be a valuable instrument for assessing impulsivity in children and adolescents, a field in which, as we have stated above, there is a lack of psychometric measurement instruments

Furthermore, the results indicate how important it is to take into account linguistic and cultural issues in test adaptation even when cultures share the same language. In this respect we can see that the factor structure of the test is congruent after 21 of the 26 items were modified. Of course we do not know what would have happened if we had administered the unmodified Spanish version to the Colombian sample but the results would probably have been worse. Indeed this is a hypothesis that further research will have to prove.

As we have stated in the introduction, this scale showed reliability problems associated to Icog when the Spanish version was developed, which was the main reason why new items were added to the first version of the Spanish scale. On the other hand, the factors related to Ic have the lowest reliabilities in the different versions of Barrat's Impulsivity Scales (Stanford et al., 2009). Taking this into account we believe that further research is needed to improve the psychometric characteristics of this scale by adding new items at least to the Colombian adaptation. Nevertheless, it should be mentioned that Im and Inp showed good or acceptable reliabilities in both versions and that these scales are most related to the dysfunctional aspects of impulsivity, which are the best predictors of aggression, psychopathological problems, addictions, etc.

Now that the factor structure and psychometric properties of two different adaptations of BIS 11-c have been determined, future research will have to show their predictive validity and their convergent validity with other psychometric and behavioural measures of impulsivity.

Finally we should point out that ESEM has some limitations that future research should override. The first is the lack of clear cut-off values to test goodness of fit (see Marsh et al., 2009). The second is the fact that the ESEM procedure needs to restrict to zero the relationship between the first item and the *p-1* remaining factors, the relationship between the second item and the *p-2* remaining factors and so on, chosen them arbitrarily. As a result, depending upon which are the first items in the data file, ESEM might provide a different model fit. In this regard, we believe that is better to choose a set of marker items from previous exploratory studies to define the factors in a semi-restricted factor analysis, as Ferrando & Lorenzo-Seva (2000) proposed, in order to get the best fit possible. In fact, an interesting alternative to the ESEM model may be to extend their model to multigroup analysis. Finally the exploratory nature of ESEM makes it easier to achieve good model fit than with CFA, but we do not believe that this is also the case in multigroup analysis, in which CFA does not test the equivalence between the «minor» loadings for all groups because they are assumed to be zero. ESEM, on the other hand, tests the equivalence of item loadings on all factors. This means that what gives ESEM the advantage over AFC may be its weakness in multigroup analysis. Taking everything into account, although there is no doubt that new methods such as ESEM may be highly valuable in the future as alternatives to classical multigroup CFA, they must be used with caution until all the issues mentioned above have been clarified.

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