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A Confirmatory Factor Analysis of the "Autoconcepto Forma 5" Questionnaire in Young Adults from Spain and Chile

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The aim of this work is to examine the pentafactorial validity of the AF5 Self-Concept Questionnaire in Spanish and Chilean young adults. From the responses of a total of 4,383 young adults aged 17 to 22 years (1,918 Spanish, 44%, and 2,465 Chilean, 56%) it was analyzed the reliability of the instrument, the compared validity of the 5 oblique factor model proposed by the authors versus the unifactorial and the orthogonal alternative models, and was studied the invariance of one Chilean sample. The results of confirmatory factor analyses supported the authors' pentafactorial model. The multi-group factorial invariance showed that Chilean sample of the AF5 does not change neither the Spanish factor weights, nor the variances and covariances of the factors, or the error variances of items. Finally, the internal consistency of the five scales was good in the samples of both countries.
Keywords: self-concept, confirmatory factor analysis, young adults, Spain, Chile.

El objetivo de este trabajo es examinar la validez pentafactorial del cuestionario de Autoconcepto Forma 5 (AF5) con jóvenes españoles y chilenos. A partir de las respuestas de un total de 4.383 adultos jóvenes de 17 a 22 años (1.918 españoles, 44%, y 2.465 chilenos, 56%) se analiza la fiabilidad del cuestionario, se compara la validez del modelo de cinco factores oblicuos que proponen los autores versus los modelos alternativos unifactorial y ortogonal, y se estudia la invarianza de la estructura factorial entre la muestra española y la chilena. Los resultados de los análisis factoriales confirmatorios ratifican la validez del modelo teórico pentafactorial de los autores. El estudio de la invarianza factorial multigrupo indica que la muestra chilena no varía en los pesos factoriales respecto de la española, ni tampoco las varianzas y covarianzas de los factores o la varianza de error de los ítems. Finalmente, la consistencia interna de las cinco escalas fue aceptable en los dos países.
Palabras clave: autoconcepto, análisis factorial confirmatorio, jóvenes adultos, España, Chile.

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The AF5 Self-Concept Questionnaire (AF5, García & Musitu, 1999) is a 30-question test for measures five major domains of self-concept. Each one of AF5 domains is measured with 6 specific items using a 1-to-99 point scale. The AF5 was developed primarily in Spain and was normed separately by sex and age groups with a large sample of 6,500 Spanish subjects between the ages of 10 and 62. Butler and Gasson (2005, pp. 196-197) noted that Self-Concept/Self-Esteem "scales have tended to be normed on geographically limited samples, with potential problems in generalizability". For example, the Rosemberg Self-Esteem Scale (RSES, Rosenberg, 1965) was normed sampled from only one State, and the Coopersmith Self-Esteem Inventory (CSEI, Coopersmith, 1967) was normed sampled from only a University of the North of Carolina. In Spanish-speaking countries, the AF5 is one of the most widely used scales for the measure multidimensional of self-concept (e. g., Jiménez, Musitu, Ramos, & Murgui, 2009; Oliden, 2008; Pereda & Forns, 2004). The AF5 questionnaire was derived from previous AFA Self-Concept Questionnaire (AFA, Musitu, García, & Gutiérrez, 1991) with three major improvements: first, a 99-point large scale of response to discriminate between the self-concept scores high; second, a measure for the physical dimension of self-concept; and third, a well-defined five-factor structure where each dimension was measured with only 6 specific elements.

The *academic/professional* dimension refers to the perception that the subject has of the quality of their role performance as student (or worker). The dimension refers to two stages, the academic and professional, which represent different chronological periods from the contexts of work. The *Social* dimension refers to the perception that the subject has of his performance in social relations, their social network and the ease or difficulty to maintain and expand his social network, and also some important qualities in interpersonal relationships (friendly and cheerful). *Emotional* dimension refers to the perception that the subject has of his emotional state and their responses to specific situations, with some degree of commitment and involvement in their daily lives. The dimension has two sources of meaning: the first refers to the general perception of their emotional state (e.g., I'm nervous, I get scared easily) and, the second, to specific situations (e.g., when they ask me, they tell me) where the other person involved is of a higher rank (e.g., professor, head). The *family* dimension refers to the perception that the subject has of their involvement, participation, and integration in the family. Finally, the *physical* dimension refers to the perception that the subject has of their physical appearance, and physical condition (García & Musitu, 1999).

The five-factor structure of AF5 that relates each dimension with his items was developed a priori following the theoretical model proposed by Shavelson et al. (Byrne & Shavelson, 1996; Shavelson, Hubner, & Stanton, 1976). These authors suppose the hierarchical organization of the construct with an unspecific general apex, whereas the more specific base dimensions of the Self presents various aspects, non-orthogonal but enough distinguishable, so that may be more differentially between them and be more related to the different specific areas of human behavior (see Musitu et al., 1991; Palacios & Zabala, 2007; Wylie, 1979).

Some empirical studies support the validity of multidimensional five-factor structure of AF5, studies established consistent theoretical relationships between the dimensions of AF5 and other psychological constructs. AF5-dimensions scores were positively related to different measures of adjustment and negatively related with different measures of maladjustment (Byrne & Shavelson, 1996; Shavelson et al., 1976). For example, empirical studies have found that emotional-AF5 was positively related with emotional intelligence (Calvet et al., 2005) and negatively with inhibition in physical activities and sports (Bernal, 2006); social-AF5, positively related with peer perceived popularity (Košir & Pečjak, 2005); family-AF5, positively related with the optimum family functioning (García & Gracia, 2009, 2010); physical-AF5, positively related with performances of students in physical education classes (Busso, 2003) and sport practice (Etxaniz, 2005); social-family-AF5 were negatively related with family and school violence (Cava, Musitu, & Murgui, 2006), immigrant integration problems (Del Barco, Castaño, Carroza, Delgado, & Pérez, 2007), and difficulties in integrating into high-risk neighbourhoods (Fariña, Arce, & Novo, 2008). Academic-AF5 has been positively associated with pro-social, conformity and self-direction values; and physical-AF5, with self-enhancement and security values (Insa, Pastor, & Ochoa, 2001). Also, academic- family- emotional-AF5 dimensions were positively associated with the quality of life in adolescents (Gómez-Vela, Verdugo, & González-Gil, 2007).

The other hand, some studies that has been assessed AF5 one-dimensionally as global self-concept, too have observed negative relationships between global scores on the AF5 and eating disorders in young women (e.g., Gual et al., 2002; Martinez-Gonzalez et al., 2003).

Available evidence also shows adequate bivariate correlations between AF5 scales and related measures. Statistically significant and meaningful (i.e., $|r| \geq .35$, $p < .001$) relationships were generally in keeping with theoretical predictions. For example, Fuentes, García, Gracia, and Lila (2011) found strong relationships between academic-AF5 with grade point average, $r(1281) = .60$; and, social-AF5 with social competence subscale of the Perceived Competence Scale (Harter, 1982), $r(1281) = .46$; emotional-AF5 with subscale of emotional instability, $r(1281) = -.41$, of Personality Assessment Questionnaire (Rohner, 1990). Also Garaigordobil, Durá, and Pérez (2005) reported relationships between emotional-AF5 and the subscales of depression, $r(140) = -.62$, and anxiety, $r(143) = -.49$, of the Symptoms Checklist-90-Revised (Derogatis, 1983). As well, Esteve (2005) found strong relationships between physical-AF5 with subscales of coordination, $r(351) = .66$, physical activity, $r(351) = .57$, sports competence, $r(351) = .76$, physical self-concept, $r(351) = .76$, appearance, $r(351) = .64$, strength, $r(351) = .57$, flexibility, $r(351) = .54$, endurance, $r(351) = .69$, and, self-esteem, $r(351) = .59$, of Physical Self-Concept Questionnaire (Marsh, Richards, Johnson, Roche, & Tremayne, 1994).

Additionally, AF5 scales have been used as criteria to validate self-concept (e.g., Garaigordobil & Bernarás, 2009) and self-esteem measures (Martín-Albo, Núñez, Navarro, & Grijalvo, 2007; Ramos, 2008), as well as criteria to validate scales that measuring related constructs: sport motivation (Martin-Albo et al., 2007), effective personality (Pellerano, Trigo, del Buey, Palacio, & Zapico, 2006), academic motivation (Alonso, Lucas, Izquierdo, & Lobera, 2006), and a psychosocial measure for treatment of addiction (Pérez, López, Cuesta, & Caballero, 2005).

On the other hand, a number of authors have highlighted four main methodological problems in measuring the self-concept (e.g., García, Musitu, & Veiga, 2006; Tomás & Oliver, 2004). First, the distributions of the items have a pronounced negative skew on the distribution curve, that is, most individuals reported having relatively high self-concept with a minority reporting lower self-concept (García & Musitu, 1999; Roth, Decker, Herzberg, & Brähler, 2008), with two opposite trends according to which side of the median is considered: 1) a great large variance in the direction of low self-concept, and 2) a small variance in the direction of high self-concept. This problem increases seriously when the scale of response is very small (e.g., *yes* or *not*; *true* or *false*); even, when join the serious problems of skew and kurtosis, is possible that only be enables the measurement of self-concept in the low to medium range, whereas differentiations at the upper end of the range be unfeasible (García et al., 2006; Roth et al., 2008). Second, constructing self-concept questionnaires with some negatively worded items, not necessary correct tends to produce skewed distributions. In some questionnaires, even, negative worded items produce positive skewed distributions, or seem associated with a method effect: positively and negatively worded items systematically load in two different factors (Tomás & Oliver, 1999). For example, the *positive* and *negative* self-concept dimensions obtained from the factor analyses on the one-dimensional Rosenberg Scale (RSES; Rosenberg, 1965). How one individual have, concurrently, two scored high on the two dimensions: the positive and negative self-concept? Third, instruments derived theoretically and based on multidimensional constructs as the AF5, generally also presented more unclear relations of item-factor structure on hypothesized domains of self-concept between different studies. For example, the factor structure of TSCS (Tennessee Self Concept Scale, Fitts, 1965) one of the most popular measures of self-concept, never has been empirically replicated (see Alfaro-Garcia & Santiago-Negrón, 2002). Finally and fourth, when applying exploratory or confirmatory factor analysis on scales originally conceptualized as one-dimensional, the studies generally can found a very consistent and easily replicated factor structure; as easily replicated as absurd for psychological theory like the *positive* and *negative* image of the self (Tomás & Oliver, 1999; Verkuyten, 2003).

However, structural validity evidence for the AF5 is supportive. As well to the initial scale development study (García & Musitu, 1999), other studies have examined the structural validity of the AF5 using exploratory factor analysis with Spanish (e.g., Busso, 2003), Brazilian (Martínez, Musitu, García, & Camino, 2003), and Italian (Marchetti, 1997) samples. All these studies, reported that all AF5 items loaded on their assigned subscales, and that there were no complex items. Confirmatory factor analyses of the AF5 have also been conducted. An initial confirmatory factor analysis with a large sample of Spanish supported the correlated five-factor model of the AF5, and showed no method effects associated with negatively worded items (Tomás & Oliver, 1999). García et al. (2006) also showed that the correlated five-factor model of the AF5 fit the data better than other alternative models. García et al. (2006) studied moreover the invariance of a Portuguese version with a sample of adult aged 21-66 years; the multi-group factorial invariance analyses showed that Portuguese translation of the AF5 does not change neither the original factor weights, nor the variances or covariances of the factors; only change the error variances of items. Another study, analyzing the structure invariance of a Basque language version, showed the same supportive results (see Elosua & Muñiz, 2010).

Although these advantages and the accumulated knowledge on this questionnaire, still not has been established whether the AF5 factor structure is invariant across countries that share the Spanish language. Even though the instrument has been widely applied on the basis that cultural differences should not modify the meaning of items and the factor structure of AF5 (Argentina: Castañeiras & Posada, 2007; Mexico: Saldaña, Becerra, & Gasca, 2007; Peru: Calvet et al., 2005; Paraguay: Alonso, Lucas, Izquierdo, & Lobera, 2006; Colombia: Lila, Musitu, & Buelga, 2000) the fact is that, until now, there is no empirical evidence that supports this main assumption. Another important issue that arises from previous studies on the AF5 factor structure is that most studies using age-heterogeneous samples (Tomás & Oliver, 2004) or adults across a wide range of ages (García et al., 2006) and none have focused specifically on young adults (Martín-Albo et al., 2007).

The first objective of the present study is examine the good fit of the correlated five-factor model based on AF5 versus one-dimensional and five-dimensional orthogonal alternative models, in a sample of Spanish young adults. The second objective is to test the invariance of the correlated five-factor model based on AF5 with another sample of Chilean young adults. It was hypothesized that: 1) the five-factor correlated model would fit the data better than both alternative models; and, 2) adjust of Chilean young adults would meet invariant respect to Spanish young adults.

Method

Participants

The sample was 4,383 young adults, 1,918 Spanish (44%) and 2,465 Chilean (56%). Spanish ranged in age from 17 to 21 (M = 18.86, SD = 1.6), with 1,212 women (63.2%) and 706 men. Chilean ranged in age from 17 to 22 (M = 17.8, SD = .69), with 1,397 women (57.0%) and 1,068 men. All Chilean and 71% of Spanish were university students (29% of Spanish were workers). The subgroup of Spanish workers was randomly sampled from Spanish Valencia City neighborhoods, the city were stratified by quartile of average household wealth, and three neighborhoods in each stratum were randomly selected (see Gracia, García, & Lila, 2009). A convenience sample of university students was recruited from three Spanish universities (Andalusia, Valencian Community, and Basque Country) and one Chilean (Araucanía). Of all participants contacted, a total of 4,383 completed the instruments, 13% of workers and 3% of university students refused to participate.

Instrument

AF5 (García & Musitu, 1999) questionnaire is designed to measure five self-concept dimensions: academic/professional, social, emotional, family, and physical. The items are statements that the subject must rate using a response on a 99-point scale, ranging from 1: complete disagreement, to 99: complete agreement.

Analysis

For testing the first hypothesis we compared the fit of the five-factor correlated model with four alternative models (Garcia & Musitu, 1999). First, we tested a one-factor model. This model represented the primary view of self-concept as one-dimensional construct (e.g., Rosenberg, 1965; Coopersmith, 1967). Second, we tested an orthogonal five-factor model. This model define self-concept as a multidimensional construct, consisting of five AF5 dimensions –academic/professional, social, emotional, family, and physical– other than as orthogonal (separate) dimensions underlying self-concept. Third, we tested the correlated five-factor model based on the AF5. This is the same model as the previous one, but with the five dimensions correlated (Byrne & Shavelson, 1996; Shavelson et al., 1976). Finally, and fourth, we freed error covariances for the strongly correlated pairs of items in each factor of third model (Byrne & Shavelson, 1996). Research on self-concept has widely noted that the correlation between the residuals of items influences the fit of the models. This notice is applied in the present study only to the items within each factor where the content is more equivalent (e.g., Byrne & Shavelson, 1996; Klassen et al., 2009; Yin & Fan, 2003). On the other hand, once the baseline model was established in fourth model we tested the method effects associated with positively and negatively worded items (Martín-Albo et al., 2007; Verkuyten, 2003). Other previous study (Tomás & Oliver, 2004) reported no-method effect associated with negatively worded items but analyzing only the four negatively worded items of social and family dimensions. In this work we compared models that test separately for method effects associated with the 10 negatively worded items of AF5, and the 20 positively worded items of AF5, but also the negatively and positively worded items together. Additionally, we tested other time the three models of the method effects associated with positively and negatively worded items, but using as baseline model the correlated five-factor model based on the AF5, without freed error covariances to “error covariances were constrained to zero in all models, in order to avoid opportunistic fitting” (Roth et al., 2008, p. 193), also following the same procedure as a previous work (Tomás y Oliver; 2004).

Following preliminary studies (García et. al, 2006; Tomás & Oliver, 2004), we used maximum likelihood (ML) as the estimation method in the confirmatory factor analyses (Chou & Bentler, 1995; West, Finch, & Curran, 1995). ML assumes that variables have a multivariate normal distribution, but non-normality appears to have a slight impact on model parameters estimated via ML, parameters remain relatively unbiased (e.g., Curran, West, & Finch, 1996), and, in any case, always reduce the confirmatory fit index measures given that the extended tails tends to increase the standard errors (Tomás & Oliver, 2004, p. 288). Furthermore, distribution-free or robust EQS methods (Byrne, 2006) did not substantially change our conclusions obtained with maximum likelihood, so we do not present these results here. However, we find that a major decrease in confirmatory fit index measures came from the extreme scores of the tails of the asymmetric distributions, which increasing the standard error heavily (see García et al. 2006). The distribution of some items was very similar to the reaction time when intervening effects of task as fatigue, lack of motivation, or errors in previous trials (Pérez, Navarro, & Llobell, 2000). To prove the negative effect of long standard errors on confirmatory fit indexes, we repeated all analyses for the four models after transforming the 99-points scale item into a shortest dichotomous response scale (< median or ≥ median). This scale change could seem similar to using a dichotomous response format (Yes/No) directly, but while the wide 99-points scale of the AF5 allows to splitting into two equal groups, using directly a short dichotomous response format only assurance come together the serious problems of skew and kurtosis.

The dichotomous scale is a non-linear transformation which preserves order relations with the original raw scores –in the dichotomous scale, self-concept that represents 1 point is less than the self-concept that represents 2 points– and reduce the variance of error markedly, as compensation for the loss of information of change from a 99-point scale to a dichotomous scale (García, Pascual, Frías, Van Krunckelsven, & Murgui, 2008). As original 99-points scale was transformed in dichotomous, these models were tested with the Satorra-Bentler chi-squared statistic (Satorra & Bentler, 2001) and associated robust confirmatory fit index provided by EQS6.1 (Byrne, 2006).

In addition, if a large size of sample is fine for controlling the standard error of measurement (García et al., 2008), it is a problem when apply the null hypothesis as criteria to testing alternative models; overall chi-square tests of model goodness-of-fit are likely to be significant due to the oversensitivity of the chi-square statistic to sample size (Bentler & Bonett, 1980). Thus, apart from χ^2-test and the ratio χ^2/gl, it has been suggested that values 2.00-3.00 or less constitute good fit (Marsh & Hau, 1996), other recommended criteria for goodness of fit were adopted, including root mean square error of approximation (RMSEA; Hu & Bentler, 1999), it has been suggested that values < .05 constitute good fit, values in the .05 to .08 range acceptable fit, values in the .08 to.10 range marginal fit, and values > .10 poor fit (Browne & Cudeck, 1992); the goodness of fit index (GFI), its adjusted version which takes model complexity into account (the adjusted goodness of fit index, AGFI; Jöreskog & Sörbom, 1989), and comparative fit index (CFI; Bentler, 1990), guidelines have suggested for GFI, AGFI and CFI, values > .95 constitute good fit and values > .90 acceptable fit (Medsker, Williams, & Holahan, 1994; Marsh & Hau, 1996); and finally, the Akaike's information criterion (AIC; computed as $\chi^2 - 2df$; Akaike, 1987), when smaller the value, better the fit.

For testing the second hypothesis, the equivalence of the Spanish and Chilean sample, we evaluated four nested models that increased progressively the number of restrictions by constraining free parameters. Once the baseline model was established so to test if CFA model fit both samples well we conducted the following sequence of increasingly more restrictive tests of invariance across both samples: fist, unconstrained, without any restrictions across any parameters; second, we fixed factor pattern coefficients; third, factor variances and covariances; and fourth, the equality of the error variances. At each step, to restrict parameters of the previous model are freed degrees of freedom and chi-square increases. When $\Delta\chi^2$ value is statistically significant, is rejects the null hypothesis that the models are equivalent. However, as $\Delta\chi^2$ also is oversensitivity to sample size, Cheung and Rensvold (2002) propose to examine the invariance of nested models via the ΔCFI. On the basis of extensive simulations with 20 different adjustment indexes, Cheung and Rensvold (2002, p. 251) proposed that an absolute ΔCFI value higher than .01 (i.e., $|\Delta CFI| > .01$) was indicative of a meaningful fall in fit.

Results

Preliminary item analyses

We confirmed the expected pattern of univariate skewness, except in item 8 of the Spanish sample, and, 8 and 23 of the Chilean sample. The skew was negative (except for negative worded items, which was positive), to be expected, most individuals reported having relatively high self-concept with a minority reporting lower self-concept (García et al., 2006; Tomás & Oliver, 2004). Univariate kurtosis is met in items 11, 16, 22 and 26 of the Spanish sample and 30 of the Chilean sample. Univariate normality (DeCarlo, 1997: *D'agostino-Pearson* K^2 *omnibus test*; *Jarque-Bera LM test*) was not confirmed for any item of the two samples. The four tests were by setting $\alpha = .05$ (DeCarlo, 1997, p. 304).

Factor structure

Fit indexes for the models applied on the responses scaled from 1 to 99 were consistently worse, but proportional to *d* (applied on the transformed dichotomous scales), as was expected due to the high skewness of the original response scale (Table 1). As well, as expected, the chi-square value did not discriminate between the alternative models, since in all of cases was statistically significant.

In the analysis for the first hypothesis, with the Spanish sample, (Table 1) the results indicated that when constrained the data to be consistent with the single one-factor model, statistics failed to meet the conventional standards completely, fit indexes were the worse (RMSEA = .16; GFI = .44; AIC = 19619). The orthogonal five-factor model, even with the same number of parameters than one-factor model, improved clearly the fit respect to the previous model (RMSEA = .07; GFI = .86; AIC = 3386). The theoretical model oblique, to free the restriction of orthogonality, also improved the fit indexes (RMSEA = .06; GFI = .89; AIC = 2527), indicating that the covariance between factors improves the fit to the data. Finally, freeing the orthogonality in oblique theoretical model for the five pairs of errors more correlated (16-26, 2-17, 3-13, 4-14, and 10-25), was obtained a reasonably good fit to the data. In addition, in the four models is possible observe as the 90% confidence intervals for RMSEA non-overlapping in any case. The analyses of RMSEA confidence intervals also confirm that the progressive changes between four alternative models have been related to a progressive improved of the fit to the data.

Table 1

Confirmatory Factor Analysis for the Spanish Sample and Multi-sample Analysis for the Invariance between the Spanish and Chilean samples

Model	χ^2	*df*	χ^2/df	$\Delta\chi^2$	Δgl	RMSEA (CI 90%)*	GFI	AGFI	CFI	ΔCFI	AIC
Tr. Theoretical+r_{error}#	2332.7	390	5.98	–984.4	–5	.051 (.049 - .053)	.92	.91	.937	.032	1553
T. Theoretical: 5 Fact. obliq.	3317.1	395	8.40	–878.8	–10	.062 (.060 - .064)	.89	.87	.905	.028	2527
O. 5 Factors orthogonal	4195.9	405	10.36	–16233.4	0	.070 (.068 - .072)	.86	.84	.877	.526	3386
U. One-dimensional	20429.3	405	50.44			.161 (.159 - .162)	.44	.35	.351		19619
dTr. Theoretical+r_{error}#	996.3	390	2.55	–317.7	–5	.028 (.026 - .031)	—	—	.960	.021	216
dT. Theoretical: 5 Fact. obliq.	1314.0	395	3.33	–299.6	–10	.035 (.033 - .037)	—	—	.939	.019	524
dO. 5 Factors orthogonal	1613.6	405	3.98	–5603.3	0	.039 (.037 - .041)	—	—	.919	.373	804
dU. One-dimensional	7216.9	405	17.82			.094 (.092 - .096)	—	—	.546		6407
mTr_1. Negatively worded	2176.7	380	5.73	–156.0	–10	.050 (.048 - .052)	.93	.91	.942	.005	1417
mTr_2. Positively worded	1986.7	370	5.37	–346.0	–20	.048 (.046 - .050)	.93	.92	.948	.011	1247
mTr_3. Negatively and positively	1784.2	360	4.96	–548.5	–30	.045 (.043 - .048)	.94	.92	.954	.017	1064
mT_1. Negatively worded	3079.5	385	8.00	–237.6	–10	.060 (.058 - .062)	.90	.88	.913	.008	2310
mT_2. Positively worded	2535.8	375	6.76	–781.3	–20	.055 (.053 - .057)	.91	.89	.930	.025	1786
mT_3. Negatively and positively	2303.3	365	6.31	–1013.8	–30	.053 (.051 - .055)	.92	.90	.937	.032	1573
Tr_0. Theo.+r_{error}# multisamples	4971.5	782	6.36			.035 (.034 - .036)	.93	.91	.935		3408
Tr_1. Equal loading in the factors	5174.0	807	6.41	–202.5	–25	.035 (.034 - .036)	.92	.91	.933	–.002	3560
Tr_2. Equal var./cov. factors	5534.9	822	6.73	–360.9	–15	.036 (.035 - .037)	.92	.91	.927	–.006	3891
Tr_3. Equal variance of errors	6074.7	852	7.13	–539.8	–30	.037 (.037 - .038)	.91	.90	.921	–.006	4371

Note: Chi-square tests statistically significant ($p < .01$). χ^2 = chi-squared; *df* = degrees of freedom; RMSEA = root mean squared error of approximation; GFI = goodness of fit index; AGFI = adjusted goodness of fit index; CFI = comparative fit index; AIC = Akaike information criterion (calculated as: $\chi^2 - 2gl$).

* CI: the 90% confidence interval (CI) for RMSEA.

\# Models *d* have been contrasted with a dichotomous scale obtained from the median and testing used the Satorra-Bentler chi-square statistic. Goodness of fit index (GFI) and adjusted goodness-of-fit index (AGFI) are not available in EQS output.

Model *Tr* is the same that *T*, except that in the *Tr* is has freed the restriction of independence for errors in pairs: Spain, 16-26, 2-17, 3-13, 4-14 and 10-25; and Chile, 16-21, 2-17 and 10-25.

Furthermore, models for method effect of positively and negatively worded items have few increased in fit indexes of *Tr* model to data; with the 10 negatively worded items (RMSEA = .05; GFI = .93; AIC = 1417), with the 20 positively worded items (RMSEA = .05; GFI = .92; AIC = 1247), and with positively and negatively worded items (RMSEA = .05; GFI = .94; AIC = 1064). Nor are major improvements respects to model *T*, with the 10 negatively worded items (RMSEA = .06; GFI = .90; AIC = 2310), with the 20 positively worded items (RMSEA = .05; GFI = .92; AIC = 1247), and with positively and negatively worded items (RMSEA = .05; GFI = .94; AIC = 1064). Should be noted that model for method effects with positive and negative items represents a major increase of 30 degrees of freedom on their basic theoretical models (*Tr* y *T*, Table 1), but only an increase in CFI by one degree of freedom (i.e., CFI/*df*) of .0006 and .0011, respectively.

Equivalence with the Chilean sample

The second hypothesis was tested with the multi-sample confirmatory factor analysis (Table 1). The baseline model was the model Tr_0, the model *Tr* but applied on the two samples. The model Tr_1 constrained the pattern loadings across the two samples, to testing whether an element has a different relative importance in either of the two samples. As the CFI decreased lower than .01 (ΔCFI = –.002; RMSEA = .035, overlaps with the previous, .034 –.036) was maintained the null hypothesis of the equivalence between the two models, suggesting that factor loadings were invariant across the two samples. In the following model, Tr_2, constrained too structural variances and covariances across the two samples, yielded non-significant changes in fit (ΔCFI = –.006; RMSEA = .036, overlaps with the previous model, .034 - .036), suggesting no difference in structural variances and covariances across the two samples. Finally, the model Tr_3, constrained also the error variances of items resulted in no changes in goodness-of-fit (ΔCFI = –.006; RMSEA = .037, overlaps with the previous model, .035 - .037), ensuring too this third assumption of invariance. As noted, is possible observe as the 90% confidence intervals for RMSEA have overlapped in all cases, indicating that the new restrictions for the parameters in each new step did not change significantly the fit of the new model respect to the previous model. Table 2 gives an overview of the parameters and standard errors of the final model.

Reliability

Alpha reliability coefficients for the total scale were .88 in the Spanish sample and .87 in Chilean sample; for academic/professional, .91 and .86; for social, .89 and .87; for emotional, .82 and .84; for family, .89 and .85; and, for physical, .86 and .87.

Table 2
Summary of Parameter Estimates (and Standard Errors) for Multi-Sample Confirmatory Factor Analysis Model between the Spanish and Chilean samples

	F1	F2	F3	F4	F5	
Item	**Standardized regression coefficients**					**Error**
1	.71					.70
6	.83					.55
11	.78					.62
16	.64					.77
21	.87					.50
26	.77					.64
2		.87				.49
7		.76				.65
12		.80				.60
17		.65				.76
22		.71				.71
27		.75				.66
3			.57			.82
8			.75			.67
13			.66			.75
18			.67			.75
23			.63			.78
28			.73			.69
4				.69		.73
9				.74		.67
14				.67		.74
19				.67		.74
24				.77		.64
29				.81		.58
5					.66	.75
10					.65	.76
15					.65	.76
20					.77	.64
25					.70	.72
30					.79	.61

Factor	**Factor Variances, Covariances, and [Correlations]**				
Academic / Professional: *F1*	146.9 (5.6)	[.31]	[.04]	[.34]	[.39]
Social: *F2*	69.0 (4.0)	344.2 (9.8)	[.21]	[.24]	[.40]
Emotional: *F3*	7.3 (3.4)	62.2 (5.4)	257.5 (13.5)	[.06]	[.18]
Family: *F4*	65.6 (3.7)	71.7 (5.3)	15.3 (4.5)	254.0 (10.2)	[.26]
Physical: *F5*	79.5 (4.2)	126.5 (6.3)	49.7 (5.2)	71.7 (5.2)	290.7 (12.7)

Country	**Freed error covariances**				
	E16-26	**E02-17**	**E03-13**	**E04-14**	**E10-25**
Spain	[.22]	[–.27]	[.28]	[.25]	[.61]
	E16-21	**E02-17**			**E10-25**
Chile	[–.30]	[–.27]			[.67]

Note: All estimated parameters were statistically significant for α = .001, except the covariance between *F3* - *F4* (*p* = .001) and *F1* - *F3* (*p* = .036). Negatively worded items (3, 4, 8, 12, 13, 14, 18, 22, 23, and 28) were inverted.

Discussion

In this paper we proposed a dual purpose: first, tested the adjustment of theoretical five-factor model of the AF5 to a Spanish sample of young adults and, second, determine whether this theoretical structure is invariant in another Chilean sample of young adults. The results of the confirmatory factor analysis from the Spanish sample of young adults, have confirmed that the oblique theoretical five-factor model of the AF5 –academic/professional, social, emotional, physical and family– provided a superior fit to the data as compared to two competitive one-dimensional and five-orthogonal-dimensional models. Thus, confirming the first hypothesis of work, results provided support for the AF5 multidimensional model proposed.

Moreover, the results using multi-sample confirmatory factor analysis shows that the AF5 multidimensional model was invariant across two Spanish and Chilean samples of young adults. Concretely, in the Chilean sample the AF5 items underline the same dimensions that in the Spanish sample, i.e., each item in the factor that is assigned theoretically has the same relatively importance to the two samples. Additionally, the five factors have an equivalent structure of variances and keep an equivalent relational pattern of covariances. Also, results met the assumption of equal error variances between the two samples for all items of the questionnaire. Thus, confirming the second hypothesis of work, AF5 multidimensional model was invariant across Spanish and Chilean samples for the three invariant assumptions that we have tested. Is important to note that the assumption of invariance between the residuals has been considerate as a very stringent test for self-concept measures (Byrne & Shavelson, 1996; García et al., 2006; Yin & Fan, 2003), indeed, in other previous studies does not meet this test of invariance (García et al., 2006). As well, the reliability for all items and dimensions for both samples was also good and similar to previous studies (e.g., García et al., 2006; Tomás & Oliver, 2004).

Therefore, the results of this study with the confirmatory factor analysis support the previous findings of other studies with very heterogeneous samples of age; specifically, confirming that the AF5 multidimensional model fit to data from young adults from Spain and Chile (García et al., 2006, Tomás & Oliver, 2004). It is interesting to note that this structure has also been replicated in several studies with the exploratory factorial method (Busso, 2003; García & Musitu, 1999), which reinforces the validity of the multidimensional structure of self-concept as is conceptualized and measured by the AF5. We believe that these results should be viewed positively keeping in mind the difficulty of confirming empirically the multivariate structures of self-concept instruments, as the widely used Tennessee Self Concept Scale (Fitts, 1965), that never has been empirically replicated (see Alfaro-Garcia & Santiago-Negrón, 2002; Bracken, 1996).

Another notable contribution of this study is that showed no method effects associated with negatively worded items, extending previous results of another study that was limited to test only four of the ten negatively worded items of the questionnaire (Tomás & Oliver, 2004). However we agree with Tomás and Oliver (2004, p. 289) that "the method effect is very minor in this scale" since the 4 degrees of freedom of their model for testing the method effects associated with negatively worded items was an increase in CFI by degree of freedom of .0025 (CFI/*df* = .01/4 = .0025). In the present study the largest increase detected, with the 20 positively worded items, was .0013 (Model *T*, .025/20 = .0013), different so more considerable ratios obtained with Rosenberg scale (Rosenberg, 1965) in which has detected increase of CFI by one degree of freedom (theorist single-factor [*df* = 35] vs. two-factor model of positive and negative self-esteem [*df* = 34]) of .19 (Spanish university students: Martín-Albo et al., 2007), .13 (Californians university students: Greenberger, Chen, Dmitrieva, & Farruggia, 2003), .10 (German general population: Roth et al., 2008), and .07 (Spanish adolescents: Tomas & Oliver, 1999). The results of this study showed no method effects associated with negatively worded items, so that affects the meaning of statements and their relationship to the theoretical dimensions; other widely applied instruments (Rosenberg, 1965), depending on whether the statements are made positively or negatively can measure two dimensions (positive or negative self-esteem), assumption subject of debate for substantial theoretical and conceptual questions (Martín-Albo et al., 2007; Tomás & Oliver, 2004; Verkuyten, 2003).

In sum, the results obtained in this study support that AF5 questionnaire is a valid instrument to be used in Spanish-speaking samples, and therefore, advisable for the assessment of self-concept from a multidimensional framework, is possible evaluate main different aspects of self with only one measuring instrument, and measuring them as was suggested by Shavelson et al. (1976). Thus, measures can be obtained sensitively, specifically and adjusted for each of these five major areas of self, being more related to specific behaviour than the global component of one-dimensional model (Rosenberg, 1965; Wylie, 1979).

Finally, this study, like all scientific work, is not without at least three limitations. First, the sampling, although the sample size is large, also is largely resulting from a set of samples of convenience; in subsequent studies we suggest use of random sampling. Second, the questionnaire was applied collectively, which can generate a considerable rate of random responding; in subsequent studies we suggest use infrequent response and social desirability scales to detect such anomalies responses. Thirdly, also there is a limitation inherent in the implementation of any measure of self, which refers particularly to the difficulty of some subjects to self-report about their own behaviors, cognitions and affects (Fonseca-Pedrero et al., 2009).

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