An Alternative Substantive Factor Structure of the Emotional Autonomy Scale

Wim Beyers[1], Luc Goossens[2], Ben Van Calster[2], and Bart Duriez[2]

[1]Ghent University, Belgium, [2]Catholic University Leuven, Belgium

Abstract. This study reexamined the factor structure of the Emotional Autonomy Scale (EAS; Steinberg & Silverberg, 1986) by means of confirmatory factor analysis on a large adolescent sample ($N = 5065$) from the Dutch-speaking part of Belgium. By looking for homogeneous subsets of items within the EAS, the meaning of the measure was clarified. None of the factor structures of the EAS suggested in the literature was supported, because of lack of fit and/or lack of construct validity. Instead, a model with seven first-order factors (Deidealization, Nondependency, Nonimitation, Privacy, Perceived ignorance, Distrust, and Perceived alienation) and two second-order factors (Separation and Detachment) was retained that proved invariant and equal across gender and grade. These findings call for continuing work on the conceptual foundations of the EAS and have important implications for the ongoing debate on the developmental outcomes of emotional autonomy (the detachment debate).

Keywords: Emotional Autonomy Scale, confirmatory factor analysis

Becoming emotionally autonomous involves becoming aware that self and parents are separate individuals (Blos, 1979). To accomplish this, adolescents need to overcome the childish representations of their parents and become aware that their parents are ordinary people with their own mistakes and problems (Levy-Warren, 1999). Blos (1979) has stressed the emotional character of this deidealization process. Similarly, Zimmer-Gembeck and Collins (2003) have stressed that becoming emotionally autonomous is primarily an emotional experience. In an attempt to incorporate Blos' (1979) theory of emotional autonomy into empirical research on adolescents, Steinberg and Silverberg (1986) developed the Emotional Autonomy Scale (EAS). This scale tries to tap four supposedly central components of emotional autonomy: Deidealizing the parents *(deidealization)*, taking responsibility for one's own behavior *(nondependency),* understanding that parents have roles outside of their parental status *(parents as people),* and establishing a sense of oneself as a separate individual *(individuation).* Items included in these subscales are displayed in Table 1, along with their original item number and scoring direction. Distinguishing these components might be important in the debate about the meaning of the EAS.

Although emotional autonomy, as indicated by high total EAS scores, was supposed to be related to better adjustment (Steinberg & Silverberg, 1986), it turns out to be related to indicators of adolescent *mal*adjustment, including susceptibility to peer pressure (Steinberg & Silverberg, 1986), feelings of insecurity with parents and of not being lovable (Ryan & Lynch, 1989), substance use and aggressive behavior (Turner, Irwin, Tschann, & Millstein, 1993), feelings of depression and anxiety (Papini & Roggman, 1992), and distress and internalizing problems (Beyers & Goossens, 1999; Garber & Little, 2001). Because of these negative correlates, some authors (e.g., Ryan & Lynch, 1989) have claimed that the EAS measures detachment (i.e., an extreme form of separation) rather than emotional autonomy. One possible solution for this detachment debate (Silverberg & Gondoli, 1996) lies in the item content of the EAS. Specifically, it can be hypothesized that, whereas some EAS components measure detachment, others do not. Several studies indicate that distinctive EAS parts assess different aspects of adolescent autonomy, and Ryan and Lynch (1989) have shown that these different parts are differently related to measures of separation-individuation and psychosocial adjustment. Individuation has a much stronger negative relation with perceived security in the relationship with parents and peers than other EAS subscales, and Chen and Dornbusch (1998) showed that individuation has a negative impact on adjustment. These results call for an in-depth analysis of the EAS items and their supposed factorial structure, prior to examining the relation with psychosocial adjustment or other theoretically important variables.

European Journal of Psychological Assessment 2005; Vol. 21(3):147–155
DOI 10.1027/1015-5759.21.3.147

Table 1. Emotional Autonomy Scale items by subscale (including direction of scoring).

Deidealization	
1	My parents and I agree on everything (–)
4	Even when my parents and I disagree, my parents are always right (–)
11	I try to have the same opinions as my parents (–)
14	When I become a parent, I'm going to treat my children in exactly the same way that my parents have treated me (–)
18	My parents hardly ever make mistakes (–)
Nondependency	
2	I go to my parents for help before trying to solve a problem myself (–)
5	It's better for kids to go to their best friend than to their parents for advice on some things (+)
6	When I've done something wrong, I depend on my parents to straighten things out for me (–)
13	If I was having a problem with one of my friends, I would discuss it with my mother or father before deciding what to do about it (–)
Parents as people	
3	I have often wondered how my parents act when I'm not around (+)
8	My parents act differently when they are with their own parents from the way they do at home (+)
10	I might be surprised to see how my parents act at a party (+)
12	When they are at work, my parents act pretty much the same way as they do when they are at home (–)
16	My parents probably talk about different things when I am around from what they talk about when I'm not (+)
20	My parents act pretty much the same way when they are with their friends as they do when they are at home with me (–)
Individuation	
7	There are some things about me that my parents don't know (+)
9	My parents know everything there is to know about me (–)
14	My parents would be surprised to know what I'm like when I'm not with them (+)
17	There are things that I will do differently from my mother and father when I become a parent (+)
19	I wish my parents would understand who I really am (+)

Confirmatory factor analysis (CFA) allows for such an internal structure analysis. CFA tests whether a theoretical model, consisting of an a priori specified number of factors and an a priori specified pattern of loadings, fits the data of a given sample. Moreover, CFA allows comparison of competing models and selection of the best fitting model. In this way, it can be tested whether the original four-factor structure of the EAS is the most valid structure, or whether there is a more valid representation. Once the best fitting model is selected, through multi-group analyses, CFA allows checking whether this structure holds across subgroups of respondents. Invariance of a CFA solution across gender and age, for instance, means that the EAS has the same structure both for girls and boys and at different age periods. Although factorial invariance is rarely tested, it is a necessary condition that must be met before one can compare mean scores or examine associations with external variables. Recently, Schmitz and Baer (2001) presented such a CFA of the EAS items. Although they rejected the original four-factor structure, they argued that negatively worded items might be differently interpreted by younger respondents (who often have weaker reading skills), and presented a methodological extension of the original model, consisting of six factors: The four original factors and two method factors (one for the positively worded and one for the negatively worded items). Technically speaking, each item had a loading on one of the four factors as well as on one of the method factors. Fit indices indicated that this model provided a significantly better fit to the data than the original model. However, ethnic differences were found in this model. The obtained factor structure turned out to differ substantially for European-American, Mexican-American, and African-American adolescents. Other types of invariance (e.g., across age and gender) were not systematically examined. Therefore, a range of questions remains unanswered.

First, Schmitz and Baer (2001) failed to compare their model with other models suggested in the literature. Following Steinberg and Silverberg (1986), the EAS was frequently used as a unidimensional measure (e.g., Garber & Little, 2001; Papini & Roggman, 1992; Ryan & Lynch, 1989; Turner et al., 1993). However, none of these studies provided psychometric information beyond Cronbach's α, and hence, no empirical test of unidimensionality was ever undertaken. In addition, several authors used a shortened 14-item EAS, leaving out the items of the Parents as people subscale (e.g., Chen & Dornbusch, 1998; Lamborn & Steinberg, 1993). This subscale had already been found to show low correlations with the other subscales, and unlike other subscales, the average score on this subscale did not increase with age (Steinberg & Silverberg, 1986), leading Smollar and Youniss (1989) to suggest that the ability to perceive parents as people probably does not develop before early adulthood. In short, these studies suggest a two-dimensional structure, with one factor comprising the 6 Parents as people items, and the other factor the remaining 14 items. Again, such a model was never actually tested. So, a complete investigation of the factor structure of the EAS entails at least a comparison of a one-factor, a two- actor, the original four-factor, and the four-plus-two factor models (Schmitz & Baer, 2001).

Second, Schmitz and Baer (2001) failed to divide the EAS items into homogeneous subsets. The ideal situation, in which all items in this four-plus-two factor model have a high loading (Λ) on one of the four factors and a substantial loading on one of the method factors, was not

Table 2. Sample information.

	Gender			Age			
Grade	Girls	Boys	Total	Minimum	Maximum	*M*	*SD*
Grade 7	210	155	365	11 y 10 m	13 y 2 m	12 y 4 m	3.8 m
Grade 8	344	202	546	12 y 10 m	14 y 2 m	13 y 5 m	3.7 m
Grade 9	199	122	321	13 y 10 m	15 y 2 m	14 y 5 m	3.7 m
Grade 10	613	481	1094	14 y 10 m	16 y 2 m	15 y 7 m	3.6 m
Grade 11	606	337	943	15 y 11 m	17 y 2 m	16 y 7 m	3.7 m
Grade 12	542	273	815	16 y 10 m	18 y 2 m	17 y 7 m	3.8 m
College	742	239	981	17 y 10 m	19 y 2 m	18 y 6 m	3.7 m
Total sample	3256	1809	5065	11 y 10 m	19 y 2 m	16 y 1 m	22.8 m

met. Item loadings on each of the four factors differed considerably. Typical illustrations were the high loadings of Items 7 and 9 on the Individuation factor. The other three items in that subscale had much weaker and occasionally even nonsignificant or negative loadings on this factor. Averaged across all age and ethnic groups, about half of the standardized factor loadings on the four substantive factors did not reach their self-imposed criterion of $\Lambda > .40$. This problem applied to items pertaining to Parents as people and Individuation in particular, and resulted in acceptable fit indices for the four-plus-two factor model that were largely due to overfitting (i.e., estimating a large number of nonsignificant parameters). Therefore, Schmitz and Baer (2001) stressed the poor construct validity of the four substantive factors in their conclusion, which might indicate that more than four factors must be differentiated within the EAS. In fact, Schmitz and Baer (2001) point out that the two method factors might also indicate that additional substantive factors have to be considered. Indeed, the fact that 9 items load on the positive and 11 items load on the negative factor does not imply that these factors actually refer to differences in wording or another method effect. It is equally possible that items with loadings on the same method factor are similar in more substantive ways, which were not captured by the original four factors. However, the lack of balance between oppositely worded items within the EAS subscales (see Table 1) makes it hard to distinguish both interpretations of the two extra factors (i.e., method factors vs. substantive factors).

The purpose of the present study is to clarify the internal structure of the EAS by looking for substantive homogeneous subsets of items and to compare this solution with the other factor models that were suggested in the literature. As already mentioned, this solution will probably contain more than four substantive factors. However, introducing higher-order factors might lead to a more parsimonious solution. Recent research (Beyers & Goossens, 2003; Beyers, Goossens, Vansant, & Moors, 2003), in which EAS subscales as well as other measures were used as indicators of two different aspects of parent-adolescent relationships, suggests such a solution. Specifically, these higher-order factors distinguish healthy separation (Beyers et al., 2003) or true independence from parents (Beyers & Goossens, 2003) from more conflictual and radical detachment from parents.

Method

Participants

The sample comprised 5065 adolescents ranging in age from 11 years and 10 months to 19 years and 2 months, of which 4755 adolescents provided complete data. The data of 310 participants were partially missing (resulting in 0.42% missing data) and were estimated using direct maximum likelihood through the expectation-maximization (EM) algorithm as available in Prelis 2.54® (Du Toit & Du Toit, 2001). Adolescents in Grades 7 to 12 were secondary school students from the Dutch-speaking part of Belgium. College students were first-year psychology students from a large university in the same area. The large majority of the participants were Caucasian with a middle class background, living in intact two-parent families, making comparison across ethnic groups impossible. Table 2 shows the distribution by gender and grade and reveals that girls and late adolescents were slightly overrepresented. This distribution allows for tests of invariance and equality of the factor structure across gender and grade.

Measure

Participants completed the Dutch Emotional Autonomy Scale (EAS; Beyers & Goossens, 1999, 2003; Finkenauer, Engels, & Meeus, 2002). Students in Grades 7 to 12 participated voluntarily in small group sessions of 15 to 50 students during normal school time. College students participated in large group sessions and received course credit for their participation. All items

Table 3. Fit indices for the various factor models.

Model description	Fit indices *df*	SBS-χ^2	RMSEA	SRMR	CFI	CAIC
Models in the odd subsample						
1. One-factor model	170	4182.42	.097	.086	.85	4535.90
2. Two-factor model	169	2991.76	.081	.073	.88	3354.09
3. Four-factor model	164	2155.47	.069	.070	.91	2561.98
4. Four-plus-two-factor model	143	941.83	.047	.044	.95	1533.92
5. Seven-factor model	149	884.92	.044	.049	.96	1423.99
6. Seven-factor model +	151	881.89	.044	.049	.96	1403.28
Models in the even subsample						
5. Seven-factor model	149	774.31	.041	.044	.97	1313.36
6. Seven-factor model +	151	770.17	.040	.045	.97	1291.54
Model 6 across gender						
6.1. Boys	151	614.10	.041	.046	.96	
6.2. Girls	151	995.30	.041	.045	.97	
6.3. Invariant pattern	302	1609.40	.041	.046	.96	
6.4. Equal structure	361	1647.69	.038	.050	.96	
Model 6 across grade						
6.1. Grade 7	151	240.24	.040	.057	.94	
6.2. Grade 8	151	327.85	.046	.055	.93	
6.3. Grade 9	151	316.45	.059	.067	.91	
6.4. Grade 10	151	370.54	.037	.043	.96	
6.5. Grade 11	151	401.65	.042	.049	.96	
6.6. Grade 12	151	504.61	.054	.060	.94	
6.7. College	151	405.37	.042	.050	.96	
6.8. Invariant pattern	1057	2567.07	.045	.052	.95	
6.9. Equal structure	1411	2813.16	.037	.062	.94	

were rated on the 4-point Likert scale that is typically used for the EAS, ranging from *don't agree at all* to *completely agree*.

Results

The sample was split in two equal parts, preserving the relative distributions of grade and gender in Table 2. Data of the odd subsample ($N = 2533$) were used for testing several a priori factor models that were found in the literature as well as for the development and testing of a new model. Data of the even subsample ($N = 2532$) were used for cross-validation of the newly developed factor model.

Comparative Analysis of the a priori Models

CFA was performed using Lisrel 8.54® (Jöreskog & Sörbom, 1996) and the maximum likelihood estimation method. Given the ordinal character and the nonnormality of the data, as evidenced by tests of univariate and multivariate normality, all tests used a matrix of polychoric correlations and the Satorra-Bentler Scaled chi-square (SBS-χ^2; Satorra & Bentler, 1994). Because χ^2 is known to be highly sensitive to small deviations from the hypothesized model, particularly when sample size is large (e.g., Loehlin, 1998), fit indices that are less influenced by sample size are also used. The root mean square error of approximation (RMSEA), which assesses model discrepancy per degree of freedom (penalizing model complexity), is particularly sensitive to factor loading misspecification (Hu & Bentler, 1998). The standardized root mean square residual (SRMR), or the average difference between the predicted and observed (co)variances in the model based on standardized residuals, is very sensitive to misspecified latent structures (Hu & Bentler, 1998). Lower RMSEA and SRMR values indicate better fit. Finally, the comparative fit index (CFI) reflects relative model improvement compared to the null model and also penalizes model complexity (Bollen, 1989). Greater CFI values indicate better fit. Hu and Bentler (1999) suggested combined cutoff values of .06 for RMSEA, .08 for SRMR, and .95 for CFI. Because the different models (Table 3, Models 1 to 4) were hierarchically nested, at least statistically, χ^2-difference tests using the Satorra and Bentler (2001) correction were suited for testing fit differences. An additional model comparison tool, which can also be used when models are not hierarchically nested, is the Consistent Akaike Information Criterion (CAIC). The CAIC adjusts the model χ^2 in order to penalize model complexity and to adjust for large sample sizes. Lower CAIC values indicate better fit.

The first four rows of Table 3 show the fit indices of the a priori models in the odd subsample ($N = 2533$), in order of decreasing parsimony (i.e., decreasing *df*). The one-factor model (Model 1) provided very poor fit to the data, as indicated by the extremely high SBS-χ^2 value, values of RMSEA and SRMR that exceed the previously mentioned cutoff values, and the low CFI estimate. This indicates that both the estimated factor loadings and the latent factor structure should be rejected. Models 2 to 4 improved the various fit indices substantially. Pairwise comparisons between models by means of the SBS-χ^2-difference test indicated that all four models differed significantly in terms of fit. Hence, Model 4 provided the best fit to the data. Moreover, in comparison with Models 1 to 3, Model 4 showed the lowest value of RMSEA, SRMR, and CAIC, and the highest value of CFI. In spite of this, 7 out of 20 (35%) standardized factor loadings on the substantive factors in Model 4 were very weak ($\Lambda <$ .40), some even close to zero. This seriously undermines the construct validity of the model (Bollen, 1989). The same lack of validity was partially evident for Models 1, 2, and 3, and particularly applied to items of the Parents as people and Individuation subscales in Models 3 and 4. Also, item loadings on the positive and negative method factor were not as expected. Some items, particularly those with very weak factor loadings on the substantive factors (e.g., Items 3, 10, 14, and 16), had very high loadings on the positive method factor. Other items had low or even negative loadings on the method factors. (The standardized factor loadings of all models can be obtained on request.) Moreover, the two method factors in Model 4 were very highly correlated ($r = .99$; $p < .0001$). Fixing the correlation between these factors to 1 did not increase SBS-χ^2 significantly. Randomly assigning items to the method factors (10 different random assignments were tested) did not increase the overall SBS-χ^2 significantly. These results indicate that the method factors do not represent the positive or negative wording of the items.

Development of a New Factor Model

Based on the previous findings, a new model was developed. Two comments made earlier guided this development. First, the low construct validity particularly applies to the items of the Parents as people and Individuation subscales. The two other subscales (Deidealization and Nondependency) are relatively homogeneous. Second, the rejection of Model 4, which includes two so-called method factors, leads to the conclusion that additional substantive factors are likely to exist. Both the modification indices (MI) associated with Models 3 and 4 and careful inspection of the item content (Table 1) confirmed these ideas. The MI statistic in Lisrel output estimates how much χ^2 is expected to decrease if a constrained parameter is set free and the model is reestimated (Sörbom, 1989).

First, a very high MI was found for the error covariance of Items 15 and 17 (estimated decrease in SBS-χ^2 = 486.04 and 500.50 in Models 3 and 4, respectively), suggesting that these share additional content not accounted for by the relation between their respective latent factors (Deidealization and Individuation). The item content confirms this idea. Both items mention the hypothetical situation of the adolescent being a parent in later life and then not simply copying the opinions and behaviors of his own parents. Because of this content similarity, both items were modeled as a separate latent factor, called *Nonimitation*. Following this reassignment, four items measuring a perception of parents as fallible persons (Items 1, 4, 11, and 18) remain in the original *Deidealization* subscale.

Second, a very high MI was found for the error covariance of Items 7 and 9 (estimated decrease in SBS-χ^2 = 122.17 and 180.06 in Models 3 and 4, respectively), suggesting that these share content not accounted for by their common latent factor (Individuation). Adolescents with high scores on these items indicate that they keep secrets from their parents. Therefore, both items were modeled as a separate factor, called *Privacy*. Following this modification, only two items remain in the Individuation subscale (Items 14 and 19). Adolescents with high scores on these items complain that their parents do not know them. Therefore, the factor that summarizes these two items was labeled *Perceived ignorance*.

Finally, both modification indices and item content suggested splitting up the items pertaining to Parents as people in two separate factors of three items each. First, Items 3, 10, and 16 cluster together to create the factor *Distrust* and describe adolescents who both do not know how their parents behave when they are not around and distrust their parents in such situations. Second, the remaining items of Parents as people indicate that adolescents suspect that their parents act completely different when not at home. This factor was labeled *Perceived alienation*.

Seven factors emerged after these modifications were made: Deidealization (Items 1, 4, 11 and 18; $\alpha = .64$), Non-dependency (Items 2, 5, 6 and 13; $\alpha = .52$), Nonim-

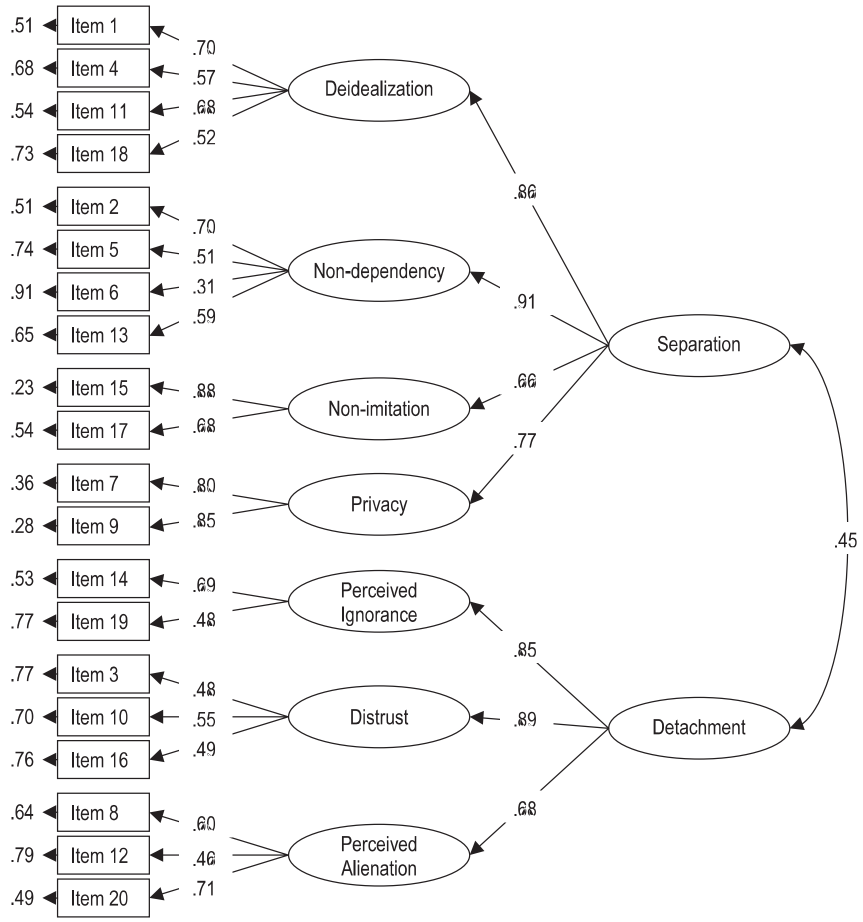

Figure 1. Final model without residual correlations among first-order factors.

itation (Items 15 and 17; α = .65), Privacy (Items 7 and 9; α = .73), Perceived ignorance (Items 14 and 19; α = .44), Distrust (Items 3, 10 and 16; α = .46), and Perceived alienation (item 8, 12 and 20; α = .55). The two-indicator rule (Bollen, 1989) was applied to this model: Latent factors are measured by at least two indicators, no cross-loadings or correlated errors are allowed, and latent factors may be correlated. Hence, the model met the conditions for identification (i.e., enough information is available to appropriately estimate the parameters).

Empirical Test of this New Model

Using the same estimation method and data matrices, this new model (Model 5) fit the data in the odd subsample well (Table 3). Values of RMSEA, SRMR, and CFI met the cutoff criteria. Moreover, this model was more parsimonious (as indicated by extra *df*) than Model 4. In line with the lower CAIC value, virtually all standardized factor loadings were above .40 (mean Λ = .62; *SD* = .15), indicating that Model 5 had better construct validity than the other models. A retest of Model 5 in the even subsample led to high factor loadings (mean Λ = .61; *SD* = .14) and excellent model fit (Table 3).

A preliminary conclusion might be that emotional autonomy, as measured by the EAS, is a many-splintered thing. However, a clear pattern of correlations was observed among the seven latent factors in Model 5. All disattenuated correlations were significantly positive ($p < .01$), but the highest correlations were found among the first four factors, with *r*s varying between .53 (Nonimitation vs. Privacy) and .80 (Deidealization vs. Nondependency). Substantial correlations were also found among the last three factors, with *r*s varying between .59 (Perceived alienation vs. Perceived ignorance) and .74 (Distrust vs. Perceived ignorance). The remaining correlations were moderate to low ($r < .50$; less than 25% of common variance). This pattern of correlations points to a more parsimonious factor structure with two higher-order factors. The first higher-order factor was hypothesized to capture the positive and normative aspects of emotional separation from parents (Deidealization, Nondependency, Nonimitation, and Privacy). Perceived ignorance, Distrust, and Perceived alienation were hypothesized to indicate the second higher-order factor. Together, both factors explained 81% of the correlations among the first-order factors. To account for the remaining 19%, residual correlations rather than additional higher-order factors were allowed, because the latter would lead to under-identification of the higher-order structure. This seven-factor model with two higher-order factors (Table 3, Model 6) fitted the data better than a model with only one higher-order factor (SBS-$\chi^2\Delta$ with 1 df = 85.66; p < .0001). Moreover, comparison of Model 6 to Model 5, favored Model 6. Residual correlations among the first-order factors in the latter model were significant, but relatively low (mean r = .12; SD = .14).

Cross-validation of this model in the even subsample resulted in an equally good model fit (Table 3), and a comparison of the estimated parameters in the odd and even subsample revealed strong similarity. Constraining *all* parameters to be equal across the two subsamples by means of a fully constrained multigroup model resulted in an excellent model fit (SBS-$\chi^2\Delta$ with 361 df = 1659.62; RMSEA = .038; SRMR = .047; CFI = .96) that did not differ significantly from a model in which all parameters were freely estimated within each group (SBS-$\chi^2\Delta$ with 59 df = 54.69; *ns*). The standardized solution is shown in Figure 1. Item loadings on the first-order factors were similar to those in Model 5 and all loadings of first- on second-order factors were very high (Λ > .66). The first higher-order factor was labeled *Separation* because it represents items that refer to emotionally separating oneself from one's parents. Separation takes place through deidealization of the parents, not depending exclusively on parents for help, not simply copying parental opinions and behaviors, and not sharing all secrets with parents. Cronbach's α based on the correlations of the 12 items that indirectly load on this factor was .80. The second higher-order factor was labeled *Detachment* because its items either describe thoughts of alienation and distrust in the relationship with parents or express complaints that their parents do not really know them. Cronbach's α of the 8 items constituting this factor was .65. Both higher-order factors were moderately positive correlated (r = .45; p < .001).

Invariance Across Gender and Grade

Testing whether Model 6 held for boys as well as for girls was done in two steps. First, Model 6 was tested for boys and girls separately. For both genders, all fit indices indicated that Model 6 fit the data well (Table 3, Model 6.1 and 6.2). Because χ^2 and its degrees of freedom are additive measures, the sum of the χ^2 values for boys and girls reflects how well the underlying factor structure fits the data across these groups. Specifically, this measure evaluates model fit when the number of factors and the pattern of loadings is held invariant, but is not constrained to be equal. The fit of the model across boys and girls was good (Table 3, Model 6.3), at least when evaluated using fit indices that are less influenced by sample size. Second, a model was tested in which both the number of factors and the pattern of loadings were held invariant, and in which both the factor loadings and the correlations between the latent factors were constrained to be equal across boys and girls (Table 3, Model 6.4). This model fitted the data well, and when compared to Model 6.3, what was gained in freedom (59 *df*) did not result in a significant increase in misfit. Therefore, the hypotheses of an invariant factor structure and an equal pattern of factor loadings across gender were confirmed.

The same two-step procedure was used to test whether Model 6 held across different grades. For each grade, fit indices (Table 3, Models 6.1 to 6.7) indicated that Model 6 fitted the data well. Moreover, results indicate that the high SBS-χ^2/df ratios found in the odd and even subsamples and in the subsamples of boys and girls were in line with the large sample sizes in earlier analyses. The fit of Model 6 *across* different grades was very good (Table 3, Model 6.8). Finally, the model hypothesizing the factor structure to be equal across grades fitted the data well (Table 3, Model 6.9), and when compared to Model 6.8, what was gained in freedom (354 *df*) did not result in a significant increase in misfit (SBS-$\chi^2\Delta$ = 317.69; *ns*). Therefore, the hypotheses of an invariant factor structure and equal pattern of factor loadings across different grades was confirmed.

Discussion

The seven-factor structure obtained in this study, which refines the original EAS structure (Steinberg & Silverberg, 1986), has better fit and superior construct validity than earlier models. In line with previous research (cf. Schmitz & Baer, 2001), whereas the original subscales Deidealization and Nondependency were largely replicated, the subscales Individuation and Parents as people proved less homogeneous. Most modifications made to the original structure were in line with earlier findings regarding the EAS. First, Chen and Dornbusch (1998) reported a high correlation between the items forming our Nonimitation factor (Items 15 and 17). Second, Schmitz and Baer (2001) reported that the two items forming our Privacy factor (Items 7 and 9) had high standardized factor loadings on their Individuation factor when compared to the other items in this factor. The correlation between the full EAS score and a measure of secrecy toward parents found in earlier research (Finkenauer et al., 2002) can probably be explained by this Privacy factor.

As suggested by recent research (Beyers & Goossens, 2003; Beyers et al., 2003) and based on the correlations among the seven first-order factors, a model with two higher-order factors fitted the data well. The first higher-order factor comprised Deidealization, Nondependency, Nonimitation, and Privacy, and was labeled *Separation*. Label choice was inspired by Bray, Adams, Getz, and Baer (2001), who used a shortened EAS including 9 out of the 12 Separation items, to measure separation. High scores on this factor reflect the separation process, which involves a move away from the childhood representations of the parents toward a representation of self and parents as separate individuals (Levy-Warren, 1999). This separation experience (Zimmer-Gembeck & Collins, 2003) is not necessarily accompanied by negative feelings toward parents, as captured by the second higher-order factor, which was labeled *Detachment* and is comprised of Distrust, Perceived alienation, and Perceived ignorance. All items of this factor have a pejorative and in some cases somewhat paranoid tone suggesting alienation and distrust (cf. Frank, Pirsch, & Wright, 1990). The positive Separation-Detachment correlation could be expected because both factors refer to the parent-adolescent relation and the experience of distance in this relation. The proposed factor structure proved invariant and equal across gender and grade, allowing the comparison of average scores across gender and grade. With regard to invariance across grade, however, longitudinal confirmation is needed, although the probability of finding such invariance is relatively high, because in that type of study, the same people are compared with each other at different points in time.

At first sight, the results of this study may seem discouraging. Emotional autonomy, as measured by the EAS, is a many-splintered thing, at least at the level of first-order factors. However, the multitude of first-order factors could be reduced to two higher-order factors that are clearly distinct, despite their positive correlation. Comparison of the models with two vs. one higher-order factor supports this. Therefore, we recommend no longer using the EAS as a unidimensional emotional autonomy measure. Based on the higher-order structure (see Figure 1), a good emotional separation measure may be formed by the 12 items of the first higher-order factor. The 8 items of the second higher-order factor show low internal consistency but nevertheless might serve as a starting point for the development of a detachment measure. These findings have important implications for the detachment debate (Silverberg & Gondoli, 1996). Any future examination of predictors and outcomes of the EAS should take its multidimensionality into account. Specifically, the finding that the original Individuation subscale is most consistently associated with negative developmental outcomes (Chen & Dornbusch, 1998; Ryan & Lynch, 1989) must be reexamined, and research should determine which part of this subscale is responsible for these outcomes. Based on our analyses, Perceived ignorance seems the most likely candidate because it shows a high loading on the Detachment factor. Generally speaking, one can expect that the Separation factor will be associated with positive or at least less negative outcomes than the Detachment factor. In a recent study, Beyers et al. (2003) found that detachment was negatively related to adolescents' agency or self-governance, while separation from parents showed a modest but significant positive correlation with agency.

It is important to realize that this study has several limitations. First, the factorial invariance across grade should not be generalized beyond the age range examined. A specific limitation is that we did not explore the EAS factor structure with upper elementary school students, a population in which Schmitz and Baer (2001) noted difficulties with the phrasing of the EAS items and with the negatively worded items in particular. Second, model invariance across ethnic groups could not be tested due to the ethnic homogeneity of our sample. Finally, it is unclear whether the findings, which were obtained on European adolescents, can be generalized to North American adolescents (the population on which most research with the EAS has been conducted). Despite these limitations, an in-depth structural analysis of the EAS in a large sample of European adolescents shows that it is possible to give a substantive interpretation to the EAS structure, rather than a solution dominated by method variance (Schmitz & Baer, 2001). Moreover, this substantive factor solution has better construct validity than the method-oriented solution. Both solutions, however, converge on the conclusion that additional work on the conceptual foundations of the EAS, and by extension, of the detachment debate, is in order.

Acknowledgments

Portions of the findings that are presented in this paper were presented at the Seventh Biennial Conference of the European Association for Research on Adolescence (EARA), Jena, Germany, June 2000. This research was supported by a grant ("Krediet aan Navorsers" – 1.5.150.99) from the Fund for Scientific Research – Flanders (Belgium) that was awarded to the second author. The authors are greatly indebted to Alfons Marcoen for his help with translations, and to all SEMNET colleagues for their valuable suggestions.

References

Beyers, W., & Goossens, L. (1999). Emotional autonomy, psychosocial adjustment, and parenting: Interactions, moderating, and mediating effects. *Journal of Adolescence, 22,* 753–769.

Beyers, W., & Goossens, L. (2003). Psychological separation and adjustment to university: Moderating effects of gender, age, and perceived parenting style. *Journal of Adolescent Research, 18*, 363–382.

Beyers, W., Goossens, L., Vansant, I., & Moors, E. (2003). A structural model of autonomy in middle and late adolescence: Connectedness, separation, detachment, and agency. *Journal of Youth and Adolescence, 32*, 351–365.

Blos, P. (1979). *The adolescent passage.* Madison, CT: International Universities Press.

Bollen, K. A. (1989). *Structural equations with latent variables.* New York: Wiley.

Bray, J.H., Adams, G.J., Getz, J.G., & Baer, P.E. (2001). Developmental, family, and ethnic influences on adolescent alcohol use: A growth curve approach. *Journal of Family Psychology, 15,* 301–314.

Chen, Z., & Dornbusch, S.M. (1998). Relating aspects of adolescent emotional autonomy to academic achievement and deviant behavior. *Journal of Adolescent Research, 13,* 293–319.

Du Toit, M., & Du Toit, S. (2001). *Interactive LISREL: User's guide.* Chicago: Scientific Software International.

Finkenauer, C., Engels, R.C.M.E., & Meeus, W. (2002). Keeping secrets from parents: Advantages and disadvantages of secrecy in adolescence. *Journal of Youth and Adolescence, 31,* 123–136.

Frank, S.J., Pirsch, L.A., & Wright, V.C. (1990). Late adolescents' perceptions of their relationships with their parents: Relationships among deidealization, autonomy, relatedness, and insecurity and implications for adolescent adjustment and ego identity status. *Journal of Youth and Adolescence, 19,* 571–588.

Garber, J., & Little, S.A. (2001). Emotional autonomy and adolescent adjustment. *Journal of Adolescent Research, 16,* 355–371.

Hu, L., & Bentler, P.M. (1998). Fit indices in covariance structure modeling: Sensitivity to underparameterized model misspecification. *Psychological Methods, 3,* 424–453.

Hu, L., & Bentler, P.M. (1999). Cutoff criteria for fit indices in covariance structure analysis: Conventional criteria versus new alternatives. *Structural Equation Modeling, 6,* 1–55.

Jöreskog, K.G., & Sörbom, D. (1996). *LISREL 8: Structural equation modeling with the SIMPLIS command language.* Chicago: Scientific Software International.

Lamborn, S.D., & Steinberg, L. (1993). Emotional autonomy redux: Revisiting Ryan and Lynch. *Child Development, 64,* 483–499.

Levy-Warren, M.H. (1999). I am, you are, and so are we: A current perspective on adolescent separation-individuation theory. In A.H. Esman, L.T. Flaherty, & H.A. Horowitz (Eds.), *Adolescent psychiatry: Developmental and clinical studies* (Vol. 24, pp. 3–24). Hillsdale, NJ: Analytic Press.

Loehlin, J.C. (1998). *Latent variable models: An introduction to factor, path, and structural analysis* (3rd ed.). Hillsdale, NJ: Erlbaum.

Papini, D.R., & Roggman, L.A. (1992). Adolescent perceived attachment to parents in relation to competence, depression, and anxiety: A longitudinal study. *Journal of Early Adolescence, 12,* 420–440.

Ryan, R.M., & Lynch, J.H. (1989). Emotional autonomy versus detachment: Revisiting the vicissitudes of adolescence and young adulthood. *Child Development, 60,* 340–356.

Satorra, A., & Bentler, P.M. (1994). Corrections to test statistics and standard errors in covariance structure analysis. In A. von Eye & C.C. Clogg (Eds.), *Latent variable analysis: Applications in developmental research* (pp. 399–419). Newbury Park, CA: Sage.

Satorra, A., & Bentler, P.M. (2001). A scaled difference χ^2 test statistic for moment structure analysis. *Psychometrika, 66*, 507–514.

Schmitz, M.F., & Baer, J.C. (2001). The vicissitudes of measurement: A confirmatory factor analysis of the Emotional Autonomy Scale. *Child Development, 72,* 207–219.

Silverberg, S.B., & Gondoli, D.M. (1996). Autonomy in adolescence: A contextualized perspective. In G.R. Adams, R. Montemayor, & T.P. Gullotta (Eds.), *Psychosocial development during adolescence* (pp. 12–61). Thousand Oaks, CA: Sage.

Sörbom, D. (1989). Model modification. *Psychometrika, 54,* 371–384.

Smollar, J., & Youniss, J. (1989). Transformations in adolescents' perceptions of parents. *International Journal of Behavioral Development, 12,* 71–84.

Steinberg, L., & Silverberg, S.B. (1986). The vicissitudes of autonomy in early adolescence. *Child Development, 57,* 841–851.

Turner, R.A., Irwin, C.E., Tschann, J.M., & Millstein, S.G. (1993). Autonomy, relatedness, and the initiation of health risk behaviors in early adolescence. *Health Psychology, 12,* 200–208.

Zimmer-Gembeck, M.J., & Collins, W.A. (2003). Autonomy development during adolescence. In G.R. Adams & M. Berzonsky (Eds.), *Blackwell handbook of adolescence* (pp. 175–204). Oxford: Blackwell.

Address for correspondence

Wim Beyers
Ghent University
Department of Developmental, Personality, and Social Psychology
Henri Dunantlaan 2
B-3000 Ghent
Belgium
E-mail Wim.Beyers@ugent.be

