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CONFIRMATORY FACTOR ANALYSIS OF THE SCALE MEASURING TEACHER ATTITUDES TOWARDS INCLUSIVE EDUCATION (TAIS)

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Abstract:
All published scales constructed to measure teacher attitudes towards inclusion have been shown to contain several factors. This study explored the factor structure of the Teacher Attitudes towards Inclusive Education Scale (TAIS) in a population of Finnish basic school teachers (n = 1,764) using confirmatory factor analysis. The TAIS scale was shown to be one-dimensional in this population. However, the result does not automatically generalize to other countries where the school system differs from that of Finland.

Keywords: inclusive education, Finland, measurement of attitudes, teachers

1. Introduction

Inclusive education refers to a situation in which students with special educational needs (SEN) are educated in the same classrooms as their non-disabled peers instead of separate special schools or special education classrooms (UNESCO, 1994). An important precondition for successful inclusion is the positive attitude of teachers towards such placements (UNESCO, 2009). Accordingly, teacher attitudes have been an object of extensive study. Several scales have also been developed to measure teacher attitudes towards inclusive education (Avramidis & Norwich, 2002; de Boer, Pijl, & Minnaert, 2011). At least 11 attitude scales have been published, which have been used more often than in a single study and for which sufficient psychometric data is available (Saloviita, 2015). The number of items in these scales has varied between 12 and 30, and their

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reliabilities, as measured by Cronbach's alpha, have shown good or excellent levels. However, their factor structure has been scattered, having 3–5 factors (Saloviita, 2015).

The multifactorial structure of the attitude scales measuring teachers' views on inclusion may reflect problems relating to scale construction rather than the multifactorial nature of the construct itself. A unidimensional scale may be, for several reasons, worth pursuing. At least one such scale has been documented, the Teacher Attitudes towards Inclusive Education Scale (TAIS) (Saloviita, 2015). Despite its unidimensionality, the items of the scale encompass a wide array of contents, such as inclusion as a value, expected outcomes, rights of the child, and workload of the teacher, thus adding to the construct validity of the instrument. The scale consists of ten items measured by a 5-point Likert scale ranging from 'strongly disagree' to 'strongly agree' with a neutral mid-point. To calculate the sum total, the scoring of six items is reversed (Table 1). The reliability of the scale has varied between α = .81–.90 in various samples (Saloviita, 2015). In exploratory factor analyses, the TAIS scale has been shown to be one-dimensional in four Finnish samples of in-service teachers or final-year pre-service teachers (Saloviita, 2015). However, in the population of German teachers, the assessment produced three factors in a principal-axis factor analysis: 'inclusion as value', 'outcomes of inclusion' and 'workload concerns' (Saloviita & Schaffus, 2016). This indicates that the scale is sensitive to the changes in the work environments and conditions of teachers.

The claim of the unidimensionality of the TAIS scale is based on exploratory factor analyses. However, a confirmatory factor analysis is needed to validate this claim. Therefore, a confirmatory factor analysis was performed with a sample of Finnish basic school in-service teachers. We also used reliability analysis to investigate whether a short form of the scale would be conceivable.

2. Methods

2.1 Participants

The sample consisted of 1,764 Finnish basic school teachers, including 783 classroom teachers, 539 subject teachers and 345 special education teachers. Their mean age was 47 years, and 21% were men and 79% women.

2.2 Data Collection

The data was collected by 33 university students who contacted teachers via an e-mail survey. The addresses were obtained from the official websites of the schools, which represented 137 randomly selected municipalities from a total of 317. The survey was returned by 26% of the teachers approached. The questionnaire contained some demographic variables and several other measures, including the TAIS scale.

Table 1: Full texts of the items in the TAIS scale, reversed items (R) and item/total correlations

Item		R	*r*
Inclusion as a value			
2.	The children with emotional and behavioural problems should be educated in mainstream classrooms, with the provision of adequate support.		.541
4.	Children with attention deficit/hyperactive disorder (ADHD) should be admitted in mainstream classrooms with adequate support.		.581
7.	The students with special educational needs should be educated in mainstream classrooms as much as possible.		.724
Expected outcomes			
1.	Children with special educational needs learn best in their own special education classes where they have specially trained teachers.	R	.720
6.	The best result is achieved if each child with special educational needs is placed in a special education classroom that best suits him/her.	R	.733
10.	The learning of children with special educational needs can be effectively supported in mainstream classrooms as well		.699
Rights of the child			
3.	It is the right of a child with special educational needs to be placed in a special education classroom.	R	.625
9.	A child with special educational needs should be transferred to a special education classroom in order not to violate his/her rights	R	.718
Workload of the teacher			
5.	Teachers' workload should not be increased by compelling them to accept children with special educational needs in their classrooms.	R	.664
8.	Integrated children with special educational needs create extra work for teachers in mainstream classrooms.	R	.582

2.3 Data Analysis

The data were analysed by using the IBM SPSS Statistics version 24 and Mplus v.7.3 statistical program. The confirmatory factor analysis was estimated using the full information maximum likelihood method with the Mplus v.7.3 statistical program (Muthén & Muthén, 2002). Missing values (0%–1.5%) were supposed to be missing at random (MAR). Using the MLR estimator in Mplus, the chi-square is scale-corrected and the standard error estimates are robust against non-normal distribution. The theoretical model was modified adding residual correlations with the help of modification indices. The model fit was evaluated using a chi-square test, root mean square error of approximation (RMSEA), the Tucker-Lewis index (TLI), the comparative fit index (CFI) and standardized root mean square error (SRMR). The model fit to the data well if the chi-square test is not statistically significant, if RMSEA is lower than .06, if TLI and CFI are greater than .95 and if SRMR is lower than .08 (Hu & Bentler, 1999; Muthén & Muthén, 2002). The measurement structure invariances were tested across sex and age for competitive factor models. These two theoretical models are nested and can therefore be compared using a scale-corrected chi-square difference test (Satorra & Bentler, 2001). The chi-square difference test was also used to test the measurement invariances. The models were modified according to the modification indices.

3. Results

The reliability of the TAIS scale was α = .90. When only three items correlating highest with the sum total were selected (items 1, 6 and 7), the reliability of this short form was α = .81, and its correlation with the original 10-item scale was r = .93. Thus, it explained 86% of the variance of the larger scale. This three-item form contained items on values and outcomes of inclusion.

When performing the confirmatory factor analysis, the theoretical one-factor model was first tested using multigroup method (11: women age < = 41; N=469, 21: women age > 41 and age < = 51; N=127, 31: women age > 51;N=455, 12: men age < = 41; N=110, 22: men age > 41 and age < = 51; N=431, 32: men age > 51; N=124) and freely estimated all parameters. The model fit was poor, $\chi^2(210) = 798.05, P < .001$, RMSEA = .099, CFI = .91, TLI = .88, SRMR = .05. There were four residual correlations which were freely estimated according to large modification indices. After this modification model fit was acceptable, $\chi^2(186) = 415.25, P < .001$, RMSEA=.066, CFI=.97, TLI = .95, SRMR = .04. When factor loadings were fixed to be equal across groups, the model fit decrease was not statistically significant, $\chi^2(45) = 48.08\ p = .349$. Additionally, when intercepts fixed equal across groups, the model fit worsened statistically significantly, $\chi^2(45) = 104.08\ p < .001$. When freeing 5 intercepts, the model fit compared to the freely estimated intercept was not statistically significant, $\chi^2(40) = 43.97\ p = .307$. In the final step, the residual variances were fixed to be equal across groups. The model fit compared to the model in which factor loadings and intercepts were partially fixed to equal was not significant, $\chi^2(50) = 55.87\ p = .264$. The model fit for these partially invariant model was good, $\chi^2(321) = 569.29, P < .001$, RMSEA = .052, CFI = .96, TLI = .97, SRMR = .06.

Factor means differ between groups, $\chi^2(5) = 20.04, p = .001$. Pairwise comparison reveal that in the p < .05 level, the factor means in group 22, 31 and 32 are lower than in group 11, while those in group 32 are lower than in group 21. Factor variances do not differ between groups, $\chi^2(5) = 2.78, p = .735$.

When estimating the two-factor theoretical model, the model fit was poor, $\chi^2(204) = 662.74, P < .001$, RMSEA = .089, CFI = .93, TLI = .91, SRMR = .05. There were three residual correlations which were freely estimated according to large modification indices. After this modification model fit was acceptable, $\chi^2(186) = 415.25, P < .001$, RMSEA = .066, CFI = .97, TLI = .95, SRMR = .04. The fit was exactly the same as in the one-factor model because the models ended up being equivalent models. Even if the freely estimated model is equivalent the measurement invariance test can differ. When factor loadings were fixed to be equal across groups, the model fit decrease was not statistically significant, $\chi^2(40) = 43.71\ p = .317$. When intercepts were fixed equal across groups, the model fit worsened statistically significantly, $\chi^2(40) = 93.13\ p < .001$. When freeing 4 intercepts, the model fit compared to the freely estimated intercept was not statistically significant, $\chi^2(36) = 43.53\ p = .539$. In the final step, the residual variances were fixed to be equal across groups. The model fit compared to the model in which factor loadings and intercepts were partially fixed to equal was not significant, $\chi^2(50) = 63.61\ p = .094$. The model fit for this partially invariant model was good, $\chi^2(312) = 562.57, P < .001$, RMSEA = .053, CFI = .96, TLI = .97, SRMR = .06. The correlation of two factors varied between .82–.88 across groups.

To be able to compare the partially invariant one- and two-factor models, the freely estimated intercepts must be the same. Therefore, freely estimated intercepts were added to the one-factor model and the two-factor model based on the final solutions. In the two-factor model there was only one additional intercept to be freely estimated. The models are then nested and the chi-square difference test was not statistically significant, $\chi^2(10) = 13.82\ p = .181$. The standardized factor loadings vary between .52 and .82 for the partially invariant one-factor model (Table 2).

4. Discussion

The reliability analysis of the TAIS scale confirmed its high reliability in terms of Cronbach's alpha. Moreover, the short three-item version of the scale containing items 1, 6 and 7 showed that it was also conceivable for further use explaining a high amount of the variance (86%) of the full scale.

Regarding the confirmatory factor analysis, the competitive one- and two-factor models were first analysed separately. The analysis began without constraints in the between-group parameters. These freely estimated models were modified according to modification indices. Both the one- and two-factor models required the freeing of some residual correlations. Upon modification, the competitive model fits were exact due to the equivalence between the models. Even if the models are equivalent, further analyses of invariance can result in different models. These modified models were therefore used as a basic model to test the invariances in the factor loadings, the intercepts of the observed variables and the residual variances.

The invariance test showed that the factor loadings were equal across the age and sex groups in the one-factor and two-factor models. In testing the invariances of the intercepts, five intercepts in the one-factor model and four intercepts in the two-factor model should estimate freely, resulting in partially invariant models. Four freely estimated intercepts were the same in both competitive models. The invariance of the residual variances was then proven in both competitive models, resulting in a final accepted partially invariant model.

Table 2: Standardized factor loading partially invariant measurement for the one-factor model

Variable	Group11 N = 335	Group21 N = 481	Group31 N = 557	Group12 N = 97	Group22 N = 117	Group32 N = 154
TAIS_1R	.75	.75	.76	.76	.74	.77
TAIS_2	.53	.53	.54	.54	.52	.56
TAIS_3R	.64	.64	.66	.65	.64	.67
TAIS_4	.57	.58	.59	.59	.57	.61
TAIS_5R	.68	.68	.70	.69	.67	.71
TAIS_6R	.80	.80	.81	.81	.79	.82
TAIS_7	.74	.74	.75	.75	.73	.76
TAIS_8R	.59	.59	.60	.60	.58	.62
TAIS_9R	.77	.77	.79	.78	.77	.80
TAIS_10	.72	.72	.74	.73	.72	.75

Note: Scoring of items marked with R is reversed when counting the sum total.

In order to compare partially invariant competitive models, they should be nested. Therefore, one additional intercept parameter was allowed to freely estimate in the final two-factor model. Following this, freeing the comparison of the competitive one- and two-factor models demonstrated that there was no difference in the model fits. This result favoured the simpler one-factor model, and therefore, the unidimensionality of the TAIS scale was supported by the study.

The TAIS scale can be recommended as a choice when unidimensional measurements are sought in assessments of teacher attitudes towards inclusion. However, as shown in the case of Brandenburg (Saloviita & Schaffus, 2016) the factorial structure of the scale may not be the same in every country because the working conditions of the teachers differ from each other. The short form of the scale containing only three items can be also recommended for further use.

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