AMERICAN PSYCHOLOGICAL ASSOCIATION

Psychological Methods

© 2020 American Psychological Association
ISSN: 1082-989X

2022, Vol. 27, No. 3, 281–306
http://dx.doi.org/10.1037/met0000355

Mixture Multigroup Factor Analysis for Unraveling Factor Loading Noninvariance Across Many Groups

Kim De Roover
Tilburg University and KU Leuven

Jeroen K. Vermunt
Tilburg University

Eva Ceulemans
KU Leuven

Abstract

Psychological research often builds on between-group comparisons of (measurements of) latent variables; for instance, to evaluate cross-cultural differences in neuroticism or mindfulness. A critical assumption in such comparative research is that the same latent variable(s) are measured in exactly the same way across all groups (i.e., measurement invariance). Otherwise, one would be comparing apples and oranges. Nowadays, measurement invariance is often tested across a large number of groups by means of multigroup factor analysis. When the assumption is untenable, one may compare group-specific measurement models to pinpoint sources of noninvariance, but the number of pairwise comparisons exponentially increases with the number of groups. This makes it hard to unravel invariances from noninvariances and for which groups they apply, and it elevates the chances of falsely detecting noninvariance. An intuitive solution is clustering the groups into a few clusters based on the measurement model parameters. Therefore, we present mixture multigroup factor analysis (MMG-FA) which clusters the groups according to a specific level of measurement invariance. Specifically, in this article, clusters of groups with metric invariance (i.e., equal factor loadings) are obtained by making the loadings cluster-specific, whereas other parameters (i.e., intercepts, factor (co)variances, residual variances) are still allowed to differ between groups within a cluster. MMG-FA was found to perform well in an extensive simulation study, but a larger sample size within groups is required for recovering more subtle loading differences. Its empirical value is illustrated for data on the social value of emotions and data on emotional acculturation.

Translational Abstract

Psychological research often builds on between-group comparisons of (measurements of) latent variables; for instance, to evaluate how mindfulness affects anxiety in several groups. A critical assumption in such comparative research is that the same latent variable(s) are measured in exactly the same way across all groups (i.e., measurement invariance). Otherwise, one would be comparing apples and oranges. Nowadays, measurement invariance is often tested across a large number of groups by means of multigroup factor analysis. When the assumption is untenable, one may inspect group-specific measurement models to pinpoint why this is the case, but the number of pairwise comparisons exponentially increases with the number of groups. This makes it hard to unravel why (and which) groups are incomparable. As an intuitive solution, we present mixture multigroup factor analysis (MMG-FA) for clustering the groups into a few clusters based on their comparability. Specifically, in this article, clusters of groups with so-called "metric invariance" are obtained. As a result, valid comparisons of latent structures (e.g., how one latent variable affects another) can be made for groups that belong to the same cluster. In an extensive simulation study, MMG-FA performed well in terms of recovering clusters of comparable groups and their measurement models. Its empirical value is illustrated for data on the social value of emotions and data on emotional acculturation.

Keywords: measurement invariance, multigroup factor analysis, metric invariance, factor loading invariance, mixture modeling

This document is copyrighted by the American Psychological Association or one of its allied publishers.
This article is intended solely for the personal use of the individual user and is not to be disseminated broadly.

This article was published Online First December 3, 2020.

Kim De Roover, Department of Methodology and Statistics, Tilburg School of Social and Behavioral Sciences, Tilburg University, and Research Group of Quantitative Psychology and Individual Differences, KU Leuven; Jeroen K. Vermunt, Department of Methodology and Statistics, Tilburg School of Social and Behavioral Sciences, Tilburg University; Eva Ceulemans, Research Group of Quantitative Psychology and Individual Differences, KU Leuven.

The research leading to the results reported in this article was funded by the Netherlands Organization for Scientific Research (NWO) [Veni Grant 451-16-004]. The computational resources and services used in this work were provided by the VSC (Flemish Supercomputer Center), funded by the Research Foundation–Flanders (FWO) and the Flemish Government–department EWI. We thank Batja Mesquita (KU Leuven) for allowing us to reanalyze the emotional acculturation data. A preliminary version of the research results reported in this article were previously presented at IMPS 2018 and a preprint of the article was posted on researchgate.net and psyarxiv.com (De Roover, Vermunt, & Ceulemans, 2019). Jeroen K. Vermunt is a co-developer and co-owner of the Latent GOLD computer program (together with Jay Magidson from Statistical Innovations).

Correspondence concerning this article should be addressed to Kim De Roover, Department of Methodology and Statistics, Tilburg School of Social and Behavioral Sciences, Tilburg University, P.O. Box 90153, 5000 LE Tilburg, the Netherlands. E-mail: k.deroover@uvt.nl

In psychological research, one often measures latent variables (e.g., personality traits, attitudes) for several groups in order to evaluate between-group differences therein. A few examples are gender differences in neuroticism (Lynn & Martin, 1997), or cross-cultural differences in mindfulness (Christopher, Charoensuk, Gilbert, Neary, & Pearce, 2009). A critical assumption in such comparative research is that the same latent variable(s) are measured in exactly the same way across all groups. Otherwise, comparing the latent variables across groups would be like comparing apples and oranges (Chen, 2008; Greiff, & Scherer, 2018). This assumption is referred to as "measurement invariance" (MI) or "measurement equivalence" (Meredith, 1993). Specifically, how the latent variables are measured by, for instance, questionnaire items is expressed by the so-called "measurement model" (MM), indicating which items measure which latent variables, and this MM needs to be invariant across groups.

The MM is traditionally evaluated with item response theory (IRT; De Ayala, 2013) in case of dichotomous or ordinal items, and with factor analysis (Lawley & Maxwell, 1962) when the items are considered to be continuous. In this article, we focus on factor analysis, where the so-called "factors" ideally correspond to the latent variables of interest. The extent to which an item relates to a factor is quantified by a "factor loading." When one wants to impose a priori assumptions about which items are measuring which factors (by fixing certain loadings to zero) and evaluate the fit of this MM for the data at hand, confirmatory factor analysis (CFA) is used. In contrast, when one wants to explore whether and how the intended latent variables are measured by the items, exploratory factor analysis (EFA) is used. Regardless of the MM being evaluated with CFA or EFA, measurement invariance pertains to the equality (i.e., invariance) of certain parameters of the factor model across all groups. The tenability of this invariance is tested by means of multigroup factor analysis (MG-FA; Dolan, Oort, Stoel, & Wicherts, 2009; Jöreskog, 1971; Sörbom, 1974) with a sequence of progressively more restricted models (see Method section for more details). Specifically, in multigroup CFA, one starts by inspecting model fit in order to evaluate "configural invariance", that is, whether the number of factors and the imposed pattern of zero loadings holds across the groups (Meredith, 1993). In multigroup EFA, no specific zero loadings are imposed. Next, in both approaches, the tenability of "weak" or "metric invariance" is evaluated by restricting the factor loadings to be equal across groups. When metric invariance holds, latent structures (e.g., how neuroticism affects another latent variable) are comparable across groups. Subsequently, "strong" or "scalar invariance" is tested by also restricting the item intercepts to be equal across groups. The finding of scalar invariance is a prerequisite for the between-group comparability of latent means (e.g., the mean level of neuroticism). Finally, "strict invariance" or invariance of uniquenesses pertains to the equality of the residual or "unique" variances of the items across groups. When combined with equal factor variances, this is a test of the equivalence of item reliability across groups (Vandenberg & Lance, 2000). Each level of invariance is tested by inspecting whether model fit drops significantly when the relevant MM parameters are restricted to be equal across groups (Cheung & Rensvold, 2002).

When a certain level of MI is rejected across groups, one may resort to pairwise comparisons of group-specific MM parameters in an attempt to pinpoint sources of noninvariance—that is, which parameters are noninvariant for which groups—and figure out how to move forward. However, the number of pairwise comparisons of group-specific parameters exponentially increases as the number of groups increases and, nowadays, the number of groups involved is on the rise (Kim, Cao, Wang, & Nguyen, 2017; Rutkowski & Svetina, 2014). The growing abundance of large-scale cross-national surveys such as the World Values Survey, European Social Survey, and International Social Survey Programme exemplify this trend. This poses two important problems (Byrne & van de Vijver, 2010; Rutkowski & Svetina, 2014): First, the multitude of comparisons makes it hard to disentangle invariant and noninvariant parameters and for which groups they differ. Second, it elevates the chances of falsely detecting noninvariance with hypothesis testing. Therefore, after the hard work of collecting data from many groups, researchers often cannot proceed with the comparisons of interest, at least not without risking invalid results.

Though, theoretically, each group may have its own MM, realistically, some groups are likely to have the same measurement parameters. Therefore, a few clusters of groups may emerge with respect to these parameters. To capture these clusters, we present a new method called "mixture multigroup factor analysis" (MMG-FA), which is an extension of multigroup factor analysis that performs a mixture clustering (McLachlan & Peel, 2000) of the groups based on (a specific subset of) the MM parameters, whereas other parameters remain group-specific. Specifically, to tackle metric (non)invariance, the current article focuses on a variant of MMG-FA that clusters the groups purely on their factor loadings, whereas parameters irrelevant for metric invariance are estimated per group. Thus, irrespective of other parameter differences, groups with (near-)identical factor loadings end up in the same mixture cluster and are modeled with one set of cluster-specific factor loadings. Clustering groups based on their MM parameters—that is, factor loadings in this case—not only confines the number of comparisons needed to identify sources of noninvariance, the clustering of the countries is an interesting result in itself. First, it indicates for which groups metric invariance holds. Second, the clustering may indicate substantively interesting between-group differences, for instance, cross-cultural differences in the functioning of a questionnaire item or in the latent variables measured by the items. Obviously, the mixture clustering of groups introduces an important model selection problem, that is, the user needs to determine the most appropriate number of clusters for a given data set. A solution for this model selection problem is discussed and evaluated in this article.

In the literature, several methods have been proposed to evaluate measurement (non)invariance for many groups, but MMG-FA differs from them in two important respects. On the one hand, the existing methods are predominantly CFA-based—for an overview, see Kim et al. (2017)—whereas EFA has some important advantages when it comes to evaluating MI (Marsh, Morin, Parker, & Kaur, 2014): First, assumed MMs often do not hold or not for all groups (i.e., configural invariance fails). In that case, respecifying CFA models in an exploratory way capitalizes on chance (Browne, 2001; MacCallum, Roznowski, & Necowitz, 1992) and using EFA right from the start is the better strategy (Gerbing & Hamilton, 1996). Second, even when the MM holds, fixed zero loadings are often too restrictive (Asparouhov & Muthén, 2009; Muthén & Asparouhov, 2012). For instance, for the well-known Big five

This document is copyrighted by the American Psychological Association or one of its allied publishers.
This article is intended solely for the personal use of the individual user and is not to be disseminated broadly.

model of personality, it was shown that zero loadings are untenable (McCrae, Zonderman, Costa, Bond, & Paunonen, 1996). Third, model misspecifications can severely bias the estimates of other MM parameters, such as the other loadings (Anderson & Gerbing, 1982; Bollen, Kirby, Curran, Paxton, & Chen, 2007), and may differ across groups (e.g., Byrne & van de Vijver, 2010; Christopher et al., 2009). For all these reasons and to prevent the clustering from being affected by model misspecifications, MMG-FA applies EFA for estimating the cluster-specific factor loadings. As a result, MMG-FA simultaneously models differences in the pattern of (near-)zero and nonzero loadings as well as differences in the strength of the nonzero loadings (i.e., both configural and metric noninvariances).

On the other hand, some existing CFA-based (Kim et al., 2017) and EFA-based (De Roover, Vermunt, Timmerman, & Ceulemans, 2017) methods apply a mixture approach similar to MMG-FA, but neither of them clusters the groups exclusively on specific subsets of the MM parameters. The latter is an important step forward as MI is traditionally evaluated in a stepwise manner, where different levels of (non)invariance have different implications in terms of which comparisons are (in)valid (Meredith, 1993). To allow for substantive researchers to focus on the level of invariance they need for a particular research question or to scrutinize noninvariances in a stepwise manner, MMG-FA clusters groups based on their MM parameters in a level-specific way, where metric invariance is the focus of this article. Metric invariance is sufficient for studies where the comparability of latent structures is of interest rather than comparing latent means (e.g., Byrne, Baron, & Balev, 1998; Byrne & Shavelson, 1986; Cooke, Kosson, & Michie, 2001; Marsh, Hau, Artelt, Baumert, & Peschar, 2006). Suggestions on how to continue toward higher levels of MI are given in the Discussion section. Note that clustering the groups based on all MM parameters at the same time (i.e., also on intercepts and unique variances) would imply the rather stringent assumption that one clustering is underlying all MM parameters, whereas some parameter differences may be explained by another clustering—possibly with a higher number of clusters—or they may be group-specific. When this assumption does not hold, the obtained clustering may even fail to capture the underlying factor loading differences. For the same reason, MMG-FA also sets aside so-called "structural" parameters that are irrelevant to the MI question—such as differences in factor (co)variances. Surely, when clustering groups in terms of how the items of a questionnaire measured, for instance, neuroticism and extraversion, it is irrelevant how the groups differ with respect to the (co)variance of neuroticism and extraversion. Clustering the groups based on a specific subset of MM parameters also limits the number of parameter comparisons needed to untangle what is different between which clusters, which adds to the insightfulness and efficiency of the method and again lowers the risk of false positives when performing hypothesis tests for parameter differences.

The remainder of this article is organized as follows: The Method section recaps MG-FA and discusses its extension into MMG-FA, covering details about model specification, model estimation and model selection. The Simulation Studies section describes extensive simulation studies to evaluate the performance of MMG-FA in terms of model estimation and model selection. The Applications section illustrates the added value of MMG-FA for cross-cultural data sets on the social value of emotions and on emotional acculturation. The Discussion section concludes with some points of discussion and directions for future research.

This document is copyrighted by the American Psychological Association or one of its allied publishers.
This article is intended solely for the personal use of the individual user and is not to be disseminated broadly.

Method

Multigroup Factor Analysis

Multigroup factor analysis (MG-FA; Jöreskog, 1971; Sörbom, 1974) operates on data from multiple groups (e.g., patient groups, countries). The groups are indicated by $g = 1, \ldots, G$ and the subjects by $n_g = 1, \ldots, N_g$. The scores for subject n_g on the J items are denoted by the vector $\mathbf{x}_{n_g}$ and, per group g, they are gathered in an $N_g \times J$ matrix $\mathbf{X}_g$. The factor model for $\mathbf{x}_{n_g}$ is written as:

$$\mathbf{x}_{n_g} = \boldsymbol{\tau}_g + \boldsymbol{\Lambda}_g \boldsymbol{\eta}_{n_g} + \boldsymbol{\varepsilon}_{n_g} \quad (1)$$

where $\boldsymbol{\tau}_g$ indicates a J-dimensional group-specific intercept vector, $\boldsymbol{\Lambda}_g$ denotes a $J \times Q$ matrix of group-specific factor loadings, $\boldsymbol{\eta}_{n_g}$ is a Q-dimensional vector of scores on the Q factors and $\boldsymbol{\varepsilon}_{n_g}$ is a J-dimensional vector of residuals. The factor loadings indicate the linear item-factor associations. The factor scores indicate how subject n_g scores on the latent variables and are assumed to be identically and independently distributed (i.i.d.) as $MVN(\boldsymbol{\alpha}_g, \boldsymbol{\Phi}_g)$, independently of $\boldsymbol{\varepsilon}_{n_g}$, which are i.i.d. as $MVN(\mathbf{0}, \boldsymbol{\Psi}_g)$. The factor means of group g are denoted by $\boldsymbol{\alpha}_g$, whereas $\boldsymbol{\Phi}_g$ pertains to the factor (co)variances and $\boldsymbol{\Psi}_g$ to a diagonal matrix containing the residual or unique variances of the items in group g. The model-implied covariance matrix for group g is $\boldsymbol{\Sigma}_g = \boldsymbol{\Lambda}_g \boldsymbol{\Phi}_g \boldsymbol{\Lambda}_g' + \boldsymbol{\Psi}_g$. In multigroup EFA (MG-EFA; Dolan et al., 2009), the group-specific factors have rotational freedom which is dealt with by a rotation criterion (De Roover & Vermunt, 2019).

Estimating Equation 1 per group corresponds to the baseline model for MI testing. To partially identify the model, the factor means $\boldsymbol{\alpha}_g$ are fixed to zero and the factor covariance matrix $\boldsymbol{\Phi}_g$ to identity (i.e., orthonormal factors: uncorrelated with variances equal to one) per group g. That fact that MG-EFA does not impose specific zero loadings on $\boldsymbol{\Lambda}_g$ makes it more flexible than multigroup CFA (MG-CFA; Meredith, & Teresi, 2006; Sörbom, 1974) in terms of the factor loading differences that can be found (De Roover & Vermunt, 2019). MI is tested by the following sequence of progressively more restricted models (Cheung & Rensvold, 2002; Dolan et al., 2009). Weak or metric invariance is evaluated by comparing the fit of the baseline model and the model with invariant loadings, that is, $\boldsymbol{\Lambda}_g = \boldsymbol{\Lambda}$ for $g = 1, \ldots, G$. For the latter model, orthonormality of the factors is no longer imposed per group but, for example, for the mean factor (co)variances across groups; $\frac{1}{N}\sum_{g=1}^{G} N_g \boldsymbol{\Phi}_g = \mathbf{I}$ where $\mathbf{I}$ refers to a $Q \times Q$ identity matrix. Strong or scalar invariance is tested by also restricting the intercepts $\boldsymbol{\tau}_g$ to be equal across groups, while freely estimating factor means $\boldsymbol{\alpha}_g$ for all groups but one. Strict invariance is assessed by restricting the unique variances, that is, the diagonal of $\boldsymbol{\Psi}_g$, to be the same across groups. Several criteria are available to evaluate whether a drop in fit when moving toward a more restricted model is statistically or practically significant. Because χ^2- difference tests for nested models are strongly affected by sample size, we focus on other fit indices such as the CFI and RMSEA. Lack of invariance is indicated when the decrease in CFI (ΔCFI) is larger than .01 and the increase in RMSEA (ΔRMSEA) exceeds .01

when imposing invariant MM parameters (Chen, 2007; Cheung & Rensvold, 2002). However, for detecting metric noninvariance across many groups, more liberal cut-off values should be used, that is, ΔCFI $< -.02$ and ΔRMSEA $> .03$ (Rutkowski & Svetina, 2014).

In case of metric noninvariance—the focus of this article—one can return to the baseline model and compare group-specific loadings to locate noninvariances (e.g., De Roover & Vermunt, 2019), but this becomes infeasible and problematic when more than a few groups are involved (see Introduction). For instance, comparing factor loadings for five groups implies only 10 pairwise comparisons, but 10 groups require 45 comparisons and 47 groups (as in the Applications section) result in 1,081 comparisons. To tie down the number of comparisons needed to identify noninvariances, we present mixture multigroup factor analysis.

Mixture Multigroup Factor Analysis

Model specification. Mixture multigroup factor analysis (MMG-FA) aims to gather groups into a few clusters according to the equivalence of their MM parameters; specifically, their factor loadings. To this end, the observations $\mathbf{x}_{n_g}$ are assumed to be sampled from a mixture of K multivariate normal distributions where all observations of a group are assumed to be sampled from the same normal distribution. Thus, the mixture clustering operates at the group level, which is an important difference from the well-known factor mixture modeling (Lubke & Muthén, 2005). In the remainder of the article, the K mixture components will be referred to as "clusters." Formally, the MMG-FA model for group g is written as follows:

$$f(\mathbf{X}_g; \boldsymbol{\theta}) = \sum_{k=1}^{K} \pi_k f_{gk}(\mathbf{X}_g; \boldsymbol{\theta}_{gk}) = \sum_{k=1}^{K} \pi_k \prod_{n_g=1}^{N_g} MVN(\mathbf{x}_{n_g}; \boldsymbol{\mu}_g, \boldsymbol{\Sigma}_{gk}) \quad with \quad \boldsymbol{\Sigma}_{gk} = \boldsymbol{\Lambda}_k \boldsymbol{\Phi}_{gk} \boldsymbol{\Lambda}_k' + \boldsymbol{\Psi}_g \tag{2}$$

where f is the total population density function, and $\boldsymbol{\theta}$ refers to the total set of parameters. The mixing proportions (i.e., prior probabilities of a group belonging to each of the clusters) are indicated by π_k, with $\sum_{k=1}^{K} \pi_k = 1$, whereas f_{gk} refers to the kth cluster-specific density function for group g and $\boldsymbol{\theta}_{gk}$ to the corresponding set of parameters. It is important to note that the means are group-specific and the covariance matrices are both group- and cluster-specific. A combination of group- and cluster-specific parameters is applied such that the clustering of the groups is driven exclusively by the parameters relevant to metric invariance, that is, the factor loadings. How to deal with higher levels of MI is described in the Discussion. Specifically, the covariance matrices are modeled by means of *cluster-specific* factor loadings $\boldsymbol{\Lambda}_k$, *group- and cluster-specific* factor (co)variances $\boldsymbol{\Phi}_{gk}$, and *group-specific* unique variances on the diagonal of $\boldsymbol{\Psi}_g$. The fact that $\boldsymbol{\Phi}_{gk}$ is not only group-specific but also varies across clusters within groups needs some additional explanation. Because the latent factors have a different meaning across clusters, and moreover have rotational freedom per cluster, it is too restrictive to assume the factor (co)variances of a group to be the same in all clusters. As shown in Appendix A, the factor (co)variances of a group can be estimated for every cluster despite the fact that the mixture model itself assumes that each group belongs to only one cluster. This holds even when group g is assigned to cluster k with a probability of zero. The resulting $\boldsymbol{\Phi}_{gk}$ should be interpreted as the factor (co)variances conditional on group g belonging to cluster k. For each group g, the factor (co)variances for the clusters the group does *not* belong to may be regarded as nuisance parameters.

Thus, in MMG-FA, the (exploratory) factor model is conditional on the cluster membership of group g, indicated by z_{gk}, as follows:

$$[\mathbf{x}_{n_g} | z_{gk} = 1] = \boldsymbol{\tau}_g + \boldsymbol{\Lambda}_k \boldsymbol{\eta}_{n_g k} + \boldsymbol{\varepsilon}_{n_g} \tag{3}$$

where $\boldsymbol{\eta}_{n_g k} \sim MVN(\mathbf{0}, \boldsymbol{\Phi}_{gk})$ and $\boldsymbol{\varepsilon}_{n_g} \sim MVN(\mathbf{0}, \boldsymbol{\Psi}_g)$. Note that, because the factor means are equal to zero per group within each cluster, the item intercepts $\boldsymbol{\tau}_g$ are equal to the means $\boldsymbol{\mu}_g$ in Equation 2. To set the scale of the cluster-specific factors, the mean factor variances are fixed to one over all groups within a cluster k, that is, $\boldsymbol{\Phi}_k = \frac{1}{N_k} \sum_{g=1}^{G} N_g \hat{z}_{gk} \boldsymbol{\Phi}_{gk} = \mathbf{I}$, where $N_k = \sum_{g=1}^{G} N_g \hat{z}_{gk}$. Note that this restriction also fixes the factor covariances to zero over all groups within a cluster, which implies that the initial rotation is orthogonal for each cluster. Afterward, the cluster-specific factors can be (orthogonally or obliquely) rotated to facilitate interpretation and comparability.

Note that the existing method that is most similar to MMG-FA, as specified above, is mixture simultaneous factor analysis (MSFA; De Roover et al., 2017). Like MMG-FA—and unlike multilevel factor mixture modeling (Kim et al., 2017)—MSFA sets apart the means as group-specific parameters and uses EFA within the clusters. This implies that its mixture clustering is also unaffected by intercept or factor mean differences and that it is equally flexible in the factor loading differences it can capture (i.e., both configural and metric noninvariances). However, MSFA differs from MMG-FA in that the covariance matrix depends entirely on the cluster, that is, $\boldsymbol{\Sigma}_k = \boldsymbol{\Lambda}_k \boldsymbol{\Lambda}_k' + \boldsymbol{\Psi}_k$. This implies that it assumes factor (co)variances and unique variances to be the same for groups within a cluster, which is too restrictive when looking for clusters of groups wherein metric invariance holds. Thus, the MSFA clustering also captures between-group differences in factor (co)variances and unique variances, rendering this method less focused on loading differences than MMG-FA.

Model estimation. The unknown parameters $\boldsymbol{\theta}$ of the MMG-FA model are estimated by means of maximum likelihood (ML) estimation. This involves maximizing the logarithm of the likelihood of the data:

$$\log L = \log\left(\prod_{g=1}^{G}\sum_{k=1}^{K}\pi_k\prod_{n_g=1}^{N_g}\frac{1}{(2\pi)^{J/2}|\boldsymbol{\Sigma}_{gk}|^{1/2}}\exp\left(-\tfrac{1}{2}(\mathbf{x}_{n_g}-\boldsymbol{\mu}_g)'\boldsymbol{\Sigma}_{gk}^{-1}(\mathbf{x}_{n_g}-\boldsymbol{\mu}_g)\right)\right) = \sum_{g=1}^{G}\log\left(\sum_{k=1}^{K}\pi_k\prod_{n_g=1}^{N_g}\frac{1}{(2\pi)^{J/2}|\boldsymbol{\Sigma}_{gk}|^{1/2}}\exp\left(-\tfrac{1}{2}(\mathbf{x}_{n_g}-\boldsymbol{\mu}_g)'\boldsymbol{\Sigma}_{gk}^{-1}(\mathbf{x}_{n_g}-\boldsymbol{\mu}_g)\right)\right), \tag{4}$$

where $\boldsymbol{\Sigma}_{gk}$ is decomposed as specified in Equation 2. Note that obtaining the parameter estimates $\hat{\boldsymbol{\theta}}$ by means of Newton-Raphson, Fisher scoring or Quasi-Newton optimization methods—that is, methods that are used in commercial software such as Latent GOLD (Vermunt & Magidson, 2013, 2016) and Mplus (Muthén & Muthén, 2005)—is very slow due to the large number of parameters and very sensitive to starting values. To find the parameter estimates in a time-efficient and stable manner, we developed an

This document is copyrighted by the American Psychological Association or one of its allied publishers.
This article is intended solely for the personal use of the individual user and is not to be disseminated broadly.

expectation-conditional maximization (ECM) algorithm (see Appendix A) and implemented it in Matlab R2017a, R (see https://github.com/KimDeRoover/MixtureMG_FA), and Latent GOLD 6.0 (see Appendix B). An R-package for MMG-FA will be developed in the near future. Because the algorithm may end up in a local maximum, a multistart procedure (based on several random partitions of the groups or several sets of random initial values for the parameters) is applied to increase the probability of obtaining the global maximum (see Appendices A and B). As an indication of computation time, the estimation of MMG-FA with three clusters and two factors for the emotion values data set (see Applications section) took 43 s with the Matlab algorithm, 188 s with the algorithm in R, and 160 s in Latent GOLD 6.0 (where the latter includes the computation of standard errors), when using 25 random starts (preselected from a set of 250 starts, see Appendices A and B). Note that repeating the same analysis in Latent GOLD without the new ECM algorithm—thus, with Fisher scoring to estimate the factor parameters—took more than 7 hrs.

Model selection. In this article, we focus on the case where the number of factors is assumed to be known and equal for all groups, and thus for all clusters. Thus, the model selection problem is confined to selecting the most appropriate number of clusters K for a given data set. For enumerating the number of clusters in related mixture models, minimizing the Bayesian information criterion (BIC; Schwarz, 1978) is often the recommended method (Nylund, Asparouhov, & Muthén, 2007; Tay, Diener, Drasgow, & Vermunt, 2011; Tein, Coxe, & Cham, 2013). The BIC takes into account model complexity in addition to the log$L$ and penalizes a model with more parameters and larger sample size as follows:

$$\text{BIC} = -2\log L + fp\log(N) \tag{5}$$

where fp refers to the number of free parameters and N refers to the total sample size $\sum_{g=1}^{G} N_g$. For MMG-FA, fp is equal to the sum of the number of mixing proportions (minus one restriction), cluster-specific factor loadings (corrected for rotational freedom), factor (co)variances for each group (minus identification restrictions), and group-specific intercepts and unique variances: $fp = K - 1 + K(JQ - Q(Q-1)/2) + (G-K)Q(Q+1)/2 + 2GJ$. Note that only one set of factor (co)variances is counted for each group, because the remaining factor (co)variances (i.e., for the clusters they do not belong to) are nuisance parameters that do not contribute to the model fit.

Several authors (Kim, Joo, Lee, Wang, & Stark, 2016; Lukočienė, Varriale, & Vermunt, 2010) suggested that, for group-level clusters, it is better to use the number of groups G for the sample size in the computation of BIC instead of the number of subjects N. In case of small sample size and low cluster separation in multilevel mixture modeling, it was found that the Akaike information criterion (AIC; Akaike, 1973) outperformed the BIC (Kim et al., 2017; Lukočienė et al., 2010). For growth mixture models, Bauer (2007) and McNeish and Harring (2017) indicated that in less ideal—but empirically more realistic—conditions (e.g., non-normality), BIC (and AIC) may overselect the number of clusters.

Finally, Bulteel, Wilderjans, Tuerlinckx, and Ceulemans (2013) showed that the convex hull procedure (CHull) is a valuable alternative to BIC and AIC in the context of mixtures of factor analyzers. The CHull (Ceulemans & Kiers, 2006; Ceulemans & Van Mechelen, 2005) is a generalization of the scree test (Cattell, 1966). Specifically, the CHull procedure balances fit and complexity by comparing the log$L$ and fp of the obtained solutions and selecting the one with the highest scree ratio. Note that, like a scree test, CHull cannot select the least complex model and thus always selects at least two clusters. But because we are focusing on cases where factor loading invariance was rejected, and loading differences are thus expected to be present, we do not regard this to be a problem. Furthermore, visual inspection of the CHull plot may still lead to the conclusion that no clear elbow is present and thus that an underlying clustering is unlikely. How these methods perform in terms of selecting the correct number of clusters for MMG-FA is evaluated in the next section.

Simulation Studies

In this section, we first present a large simulation study to evaluate the performance of MMG-FA, both in terms of model estimation and model selection, when clusters of groups with metric invariance are underlying the data. Then, a smaller simulation study is presented to evaluate MMG-FA in terms of model selection, when no such clusters are underlying the data and metric invariance holds across all groups (i.e., the number of clusters equals one).

Simulation Study 1

Problem

The goal of Simulation Study 1 is, on the one hand, to evaluate the performance of MMG-FA with respect to the recovery of the clustering of the groups and of the cluster-specific factor loadings when the correct number of clusters is known and to compare this performance to that of MSFA. On the other hand, it is evaluated to what extent the model selection procedures described in the Method section select the correct number of clusters for MMG-FA. We manipulated six factors that were expected to affect the cluster separation and/or the stability of parameter estimates, and thus the performance of MMG-FA and its model selection: (a) the number of groups, (b) the group sizes, (c) the number of clusters, (d) the cluster sizes, (e) the number of factors, and (f) the type and size of the loading differences.

Specifically, in terms of their effect, we hypothesize the following: The number of groups (a) determines how many groups end up within each cluster. Because more groups within a cluster implies more information on that cluster-specific MM (i.e., a higher within-cluster sample size), we hypothesize the performance to improve with a higher number of groups. A higher number of observations per group (b) increases the within-cluster sample size and thus the performance. It also implies more information on each of the cluster memberships and thus a higher cluster separation (Lukočienė et al., 2010). Relatedly, a higher number of clusters (c) lowers the within-cluster sample size (for a given number of groups) and is thus expected to lower the performance. It also increases the number of cluster memberships (posterior classification probabilities; see Appendix A) to be determined for each group and thus makes their recovery more intricate. The cluster sizes (d), corresponding to the mixing proportions, pertain to the groups being equally or unequally divided across the clusters. In the unequal case, larger cluster(s) will compete with

This document is copyrighted by the American Psychological Association or one of its allied publishers.
This article is intended solely for the personal use of the individual user and is not to be disseminated broadly.

smaller cluster(s) and the smaller ones will be much harder to recover both in terms of cluster memberships and factor loading estimates. With respect to the number of factors (e), a higher number of factors—given the same number of variables—implies a lower factor overdetermination and thus probably a lower performance. Finally, the type and size of loading differences (f) greatly determines the extent to which the cluster-specific MMs differ from one another (i.e., cluster separation) and thus affects the recovery of the cluster memberships. For instance, a primary loading that shifts to another factor is a large difference that would be easier to recover than a small difference in the size of a primary loading or crossloading.

Design

These factors were systematically varied in a complete factorial design:

1. the *number of groups G* at two levels: 12, 60;
2. the *group sizes* N_g (i.e., number of observations per group) at five levels: 30, 50, 100, 300, 500;
3. the *number of clusters K* at two levels: 2, 4;
4. the *cluster sizes* at two levels: equal, unequal;
5. the *number of factors Q* at two levels: 2, 4;
6. the *type and size of loading differences* at five levels: primary loading shift, crossloading of .40, crossloading of .20, primary loading decrease of .40, primary loading decrease of .20.

The number of variables J was fixed at 20 and the cluster-specific factor loading matrices were generated by inducing changes to the same simple structure loading matrix. In this "base loading matrix," the variables are equally distributed over the factors, that is, each factor gets 10 nonzero loadings when $Q = 2$ (see Table 1) and five nonzero loadings when $Q = 4$ (see Table 2). Given that the unique variances vary around .40 (see below), the nonzero loadings are equal to $\sqrt{.60}$ to obtain total variances that vary around one. From the common base, K different cluster-specific loading matrices are derived by altering the loadings for a different pair of variables for each cluster (see Tables 1 and 2). Specifically, depending on the type and size of loading differences, the loadings of two variables were altered as follows: In case of a primary loading shift, when $Q = 2$, the loadings $[\sqrt{.6}\ \ 0]$ of the base matrix are replaced by $[0\ \ \sqrt{.6}]$or vice versa (see Table 1). When $Q = 4$, primary loadings are shifted similarly between Factors 1 and 2, leaving Factors 3 and 4 unaffected; for example, $[\sqrt{.6}\ \ 0\ \ 0\ \ 0]$ becomes $[0\ \ \sqrt{.6}\ \ 0\ \ 0]$. In case of the crossloading differences, the loadings $[\sqrt{.6}\ \ 0\ \ (0)\ \ (0)]$ become $[\sqrt{.6}\ \ .4\ \ (0)\ \ (0)]$ or $[\sqrt{.6}\ \ .2\ \ (0)\ \ (0)]$ depending on the size of the crossloadings (see Table 2). Note that a crossloading of .20 may be considered "ignorable," whereas one of .40 is not (Stevens, 1992). To manipulate a primary loading decrease, the loadings $[\sqrt{.6}\ \ 0\ \ (0)\ \ (0)]$ are replaced by $[\sqrt{.6}-.4\ \ 0\ \ (0)\ \ (0)]$ or $[\sqrt{.6}-.2\ \ 0\ \ (0)\ \ (0)]$ depending on the size of the decrease (see Table 3). A primary loading decrease of .40 is considered a large

Table 1
Base Loading Matrix and the First Two Derived Cluster-Specific Loading Matrices, in Case of Two Factors and Primary Loading Shifts

	Base Loading Matrix		Cluster-Specific Loading Matrix 1		Cluster-Specific Loading Matrix 2	
Variable	F1	F2	F1	F2	F1	F2
V1	$\sqrt{.6}$	0	**0**	$\sqrt{\mathbf{.6}}$	$\sqrt{.6}$	0
V2	$\sqrt{.6}$	0	$\sqrt{.6}$	0	**0**	$\sqrt{\mathbf{.6}}$
V3	$\sqrt{.6}$	0	$\sqrt{.6}$	0	$\sqrt{.6}$	0
V4	$\sqrt{.6}$	0	$\sqrt{.6}$	0	$\sqrt{.6}$	0
V5	$\sqrt{.6}$	0	$\sqrt{.6}$	0	$\sqrt{.6}$	0
V6	$\sqrt{.6}$	0	$\sqrt{.6}$	0	$\sqrt{.6}$	0
V7	$\sqrt{.6}$	0	$\sqrt{.6}$	0	$\sqrt{.6}$	0
V8	$\sqrt{.6}$	0	$\sqrt{.6}$	0	$\sqrt{.6}$	0
V9	$\sqrt{.6}$	0	$\sqrt{.6}$	0	$\sqrt{.6}$	0
V10	$\sqrt{.6}$	0	$\sqrt{.6}$	0	$\sqrt{.6}$	0
V11	0	$\sqrt{.6}$	$\sqrt{\mathbf{.6}}$	**0**	0	$\sqrt{.6}$
V12	0	$\sqrt{.6}$	0	$\sqrt{.6}$	$\sqrt{\mathbf{.6}}$	**0**
V13	0	$\sqrt{.6}$	0	$\sqrt{.6}$	0	$\sqrt{.6}$
V14	0	$\sqrt{.6}$	0	$\sqrt{.6}$	0	$\sqrt{.6}$
V15	0	$\sqrt{.6}$	0	$\sqrt{.6}$	0	$\sqrt{.6}$
V16	0	$\sqrt{.6}$	0	$\sqrt{.6}$	0	$\sqrt{.6}$
V17	0	$\sqrt{.6}$	0	$\sqrt{.6}$	0	$\sqrt{.6}$
V18	0	$\sqrt{.6}$	0	$\sqrt{.6}$	0	$\sqrt{.6}$
V19	0	$\sqrt{.6}$	0	$\sqrt{.6}$	0	$\sqrt{.6}$
V20	0	$\sqrt{.6}$	0	$\sqrt{.6}$	0	$\sqrt{.6}$

Note. Differences are indicated in bold face. When $K = 4$, the third and fourth cluster-specific loading matrices are created by shifting the primary loadings of Items 3 and 13 and Items 4 and 14, respectively.

noninvariance (Stark, Chernyshenko, & Drasgow, 2006) that can lead to incorrect statistical inference and biased parameter estimates (Hancock, Lawrence, & Nevitt, 2000). Please observe the following: First, a primary loading shift maintains the item's communality whereas a crossloading increases it and a primary loading decrease lowers it. Second, primary loading shifts and crossloadings are violations of configural invariance and thus differences that are very hard to trace by any of the existing CFA-based methods (Kim et al., 2017).

Note that the number of groups of 12 and 60 nicely correspond to the range of group numbers that are generally encountered in large-scale surveys (Rutkowski & Svetina, 2014). In case of equal cluster sizes, the groups are equally divided across the clusters, that is, each cluster contains 50% of the groups in case of two clusters or 25% in case of four clusters. In the unequal cluster size conditions, the groups are divided over the clusters such that one cluster contains 75% of the groups, whereas the remaining groups are equally divided over the other clusters. Thus, in case of two clusters, 75% of the groups are in one cluster and 25% in the other one. In case of four clusters, each of the three smaller clusters contains 8.33% of the groups. Note that the latter correspond to singleton clusters (i.e., including only one group) in case of 12 groups, whereas in case of 60 groups they hold five groups each. The cluster memberships were generated by randomly assigning the correct number of groups to each cluster, according to these cluster sizes.

This document is copyrighted by the American Psychological Association or one of its allied publishers.
This article is intended solely for the personal use of the individual user and is not to be disseminated broadly.

Table 2
Base Loading Matrix and the First Two Derived Cluster-Specific Loading Matrices, in Case of Four Factors and Crossloading Differences

	Base Loading Matrix				Cluster-Specific Loading Matrix 1				Cluster-Specific Loading Matrix 2			
Variable	F1	F2	F3	F4	F1	F2	F3	F4	F1	F2	F3	F4
V1	$\sqrt{.6}$	0	0	0	$\sqrt{.6}$	**CL**	0	0	$\sqrt{.6}$	0	0	0
V2	$\sqrt{.6}$	0	0	0	$\sqrt{.6}$	0	0	0	$\sqrt{.6}$	**CL**	0	0
V3	$\sqrt{.6}$	0	0	0	$\sqrt{.6}$	0	0	0	$\sqrt{.6}$	0	0	0
V4	$\sqrt{.6}$	0	0	0	$\sqrt{.6}$	0	0	0	$\sqrt{.6}$	0	0	0
V5	$\sqrt{.6}$	0	0	0	$\sqrt{.6}$	0	0	0	$\sqrt{.6}$	0	0	0
V6	0	$\sqrt{.6}$	0	0	**CL**	$\sqrt{.6}$	0	0	0	$\sqrt{.6}$	0	0
V7	0	$\sqrt{.6}$	0	0	0	$\sqrt{.6}$	0	0	**CL**	$\sqrt{.6}$	0	0
V8	0	$\sqrt{.6}$	0	0	0	$\sqrt{.6}$	0	0	0	$\sqrt{.6}$	0	0
V9	0	$\sqrt{.6}$	0	0	0	$\sqrt{.6}$	0	0	0	$\sqrt{.6}$	0	0
V10	0	$\sqrt{.6}$	0	0	0	$\sqrt{.6}$	0	0	0	$\sqrt{.6}$	0	0
V11	0	0	$\sqrt{.6}$	0	0	0	$\sqrt{.6}$	0	0	0	$\sqrt{.6}$	0
V12	0	0	$\sqrt{.6}$	0	0	0	$\sqrt{.6}$	0	0	0	$\sqrt{.6}$	0
V13	0	0	$\sqrt{.6}$	0	0	0	$\sqrt{.6}$	0	0	0	$\sqrt{.6}$	0
V14	0	0	$\sqrt{.6}$	0	0	0	$\sqrt{.6}$	0	0	0	$\sqrt{.6}$	0
V15	0	0	$\sqrt{.6}$	0	0	0	$\sqrt{.6}$	0	0	0	$\sqrt{.6}$	0
V16	0	0	0	$\sqrt{.6}$	0	0	0	$\sqrt{.6}$	0	0	0	$\sqrt{.6}$
V17	0	0	0	$\sqrt{.6}$	0	0	0	$\sqrt{.6}$	0	0	0	$\sqrt{.6}$
V18	0	0	0	$\sqrt{.6}$	0	0	0	$\sqrt{.6}$	0	0	0	$\sqrt{.6}$
V19	0	0	0	$\sqrt{.6}$	0	0	0	$\sqrt{.6}$	0	0	0	$\sqrt{.6}$
V20	0	0	0	$\sqrt{.6}$	0	0	0	$\sqrt{.6}$	0	0	0	$\sqrt{.6}$

Note. The crossloadings (CL) are either equal to .40 or .20. Differences are indicated in bold face. When $K = 4$, the third and fourth cluster-specific loading matrices are created by adding CL for Items 3 and 8 and Items 4 and 9, respectively.

The group- and cluster-specific factor correlations are randomly sampled from a uniform distribution between −.50 and .50, that is, $U(-.50, .50)$, and factor variances from $U(.50, 1.50)$. Whenever a resulting $\mathbf{\Phi}_{gk}$ is not positive definite, the sampling is repeated. Group-specific unique variances (i.e., diagonal of $\mathbf{\Psi}_g$) are sampled from $U(.20, .60)$. Factor scores are sampled from $MVN(\mathbf{0}, \mathbf{\Phi}_{gk})$ and residuals from $MVN(\mathbf{0}, \mathbf{\Psi}_g)$, according to the specified group sizes. The group size of 100 corresponds to the absolute minimal sample size for obtaining accurate factor loading estimates (Gorsuch, 1983), whereas higher sample sizes are recommended in case of lower factor overdetermination and/or item communalities (Fabrigar, Wegener, MacCallum, & Strahan, 1999; MacCallum, Widaman, Zhang, & Hong, 1999). Note that, for MMG-FA, the accuracy of the factor loadings will be determined by the sample size of a cluster of groups, rather than of a single group, whereas the accuracy of the factor (co)variances of groups within a cluster will depend on the group sizes. To evaluate the extent to which MMG-FA can find loading differences among really small groups, we included the group sizes of 30 and 50. Finally, the simulated data are created according to Equation 3. The intercepts $\boldsymbol{\tau}_g$ are zero, because the focus is on loading differences.

According to this procedure, 50 data sets were generated per cell of the design, using Matlab R2017a. Thus, 2 (number of groups) × 5 (group sizes) × 2 (number of clusters) × 2 (cluster sizes) × 2 (number of factors) × 5 (type/size of loading differences) × 50 (replications) = 20,000 data sets were generated. The data were analyzed by the ECM algorithm for MMG-FA detailed in Appendix A, using the correct number of factors Q and using 25 starts. On the one hand, the correct number of clusters K was specified to evaluate the performance of the algorithm itself. On the other hand, for the first five replications of each cell of the design (i.e., for 2,000 data sets), MMG-FA analyses were performed with one to six clusters to evaluate the performance of the model selection procedures described in the Method section. No convergence problems were encountered in this simulation study. The analyses were performed on a supercomputer consisting of Xeon E5-2680 v2 processors with a clock frequency of 2.8 GHz and with 64 GB RAM. The average CPU time for MMG-FA with the correct number of clusters K was 80 s and, for the model selection, estimating the six models with an increasing number of clusters took about 9 min.[1] To compare the performance of MMG-FA to that of MSFA, the data sets were also analyzed by MSFA with the correct K and Q and 25 starts (for details, see De Roover, Timmerman, & Ceulemans, 2017).

Results

Sensitivity to local maxima. To evaluate the frequency of local maximum solutions, we should compare the log$L$ value of the best solution obtained by the multistart procedure (i.e., starting from 25 random partitions; see Appendix A) with the global ML solution for each simulated data set. Because of sampling fluctuations, the global maximum is unknown, however. Therefore, we made use of a "proxy" of the global ML solution; that is, the solution that is obtained when the algorithm starts from the true clustering of the groups. The final solution from the multistart

[1] These are the average CPU times for the conditions with group sizes of 100, 300, and 500. The group sizes of 30 and 50 were added for the revision and, on an i7 processor with 8GB RAM, the average CPU time for these conditions was 44 s for estimating the model with the correct number of clusters and 8 min for estimating models with one to six clusters.

This document is copyrighted by the American Psychological Association or one of its allied publishers.
This article is intended solely for the personal use of the individual user and is not to be disseminated broadly.

Table 3
Base Loading Matrix and the First Two Derived Cluster-Specific Loading Matrices, in Case of Two Factors and Primary Loading Decreases

	Base Loading Matrix		Cluster-Specific Loading Matrix 1		Cluster-Specific Loading Matrix 2	
Variable	F1	F2	F1	F2	F1	F2
V1	$\sqrt{.6}$	0	$\sqrt{.6}$ − **PLD**	0	$\sqrt{.6}$	0
V2	$\sqrt{.6}$	0	$\sqrt{.6}$	0	$\sqrt{.6}$ − **PLD**	0
V3	$\sqrt{.6}$	0	$\sqrt{.6}$	0	$\sqrt{.6}$	0
V4	$\sqrt{.6}$	0	$\sqrt{.6}$	0	$\sqrt{.6}$	0
V5	$\sqrt{.6}$	0	$\sqrt{.6}$	0	$\sqrt{.6}$	0
V6	$\sqrt{.6}$	0	$\sqrt{.6}$	0	$\sqrt{.6}$	0
V7	$\sqrt{.6}$	0	$\sqrt{.6}$	0	$\sqrt{.6}$	0
V8	$\sqrt{.6}$	0	$\sqrt{.6}$	0	$\sqrt{.6}$	0
V9	$\sqrt{.6}$	0	$\sqrt{.6}$	0	$\sqrt{.6}$	0
V10	$\sqrt{.6}$	0	$\sqrt{.6}$	0	$\sqrt{.6}$	0
V11	0	$\sqrt{.6}$	0	$\sqrt{.6}$ − **PLD**	0	$\sqrt{.6}$
V12	0	$\sqrt{.6}$	0	$\sqrt{.6}$	0	$\sqrt{.6}$ − **PLD**
V13	0	$\sqrt{.6}$	0	$\sqrt{.6}$	0	$\sqrt{.6}$
V14	0	$\sqrt{.6}$	0	$\sqrt{.6}$	0	$\sqrt{.6}$
V15	0	$\sqrt{.6}$	0	$\sqrt{.6}$	0	$\sqrt{.6}$
V16	0	$\sqrt{.6}$	0	$\sqrt{.6}$	0	$\sqrt{.6}$
V17	0	$\sqrt{.6}$	0	$\sqrt{.6}$	0	$\sqrt{.6}$
V18	0	$\sqrt{.6}$	0	$\sqrt{.6}$	0	$\sqrt{.6}$
V19	0	$\sqrt{.6}$	0	$\sqrt{.6}$	0	$\sqrt{.6}$
V20	0	$\sqrt{.6}$	0	$\sqrt{.6}$	0	$\sqrt{.6}$

Note. The primary loading decreases (PLD) are either equal to .40 or .20. Differences are indicated in bold face. When $K = 4$, the third and fourth cluster-specific loading matrices are created by decreasing the primary loading for Items 3 and 13 and Items 4 and 14, respectively.

procedure was considered to be a local maximum when its log$L$ value is smaller than the one from the proxy. To exclude mere calculation precision differences, we only considered such differences with an absolute value higher than .0001 as a local maximum. By this definition, 3,002 (15.0%) local maxima were detected over all 20,000 simulated data sets. Most of these occurred in case of four clusters; that is, 2,844 of the 3,002 local maxima are found when $K = 4$. Not surprisingly, the sensitivity to local maxima also depends strongly on the group sizes. Specifically, the percentage of local maxima equals 19.7%, 19.4%, 18.8%, 9.8%, and 7.5% for groups of 30, 50, 100, 300, and 500 subjects, respectively. Note that, for 1,176 out of these 3,002 data sets, rerunning the analysis with 50 starts (i.e., starting from 50 random partitions) was sufficient to avoid the local maxima, reducing the percentage of local maxima to 9.1% across all data sets. Per group size, it reduced to 13.0%, 12.6%, 11.7%, 5.3%, and 3.1% for groups of 30, 50, 100, 300, and 500, respectively.

Goodness of cluster recovery. To examine the goodness of recovery of the cluster memberships of the groups, we compared the modal clustering (i.e., assigning each group to the cluster for which the posterior classification probability is the highest) to the true clustering by means of the *Adjusted Rand Index* (*ARI*; Hubert & Arabie, 1985). The *ARI* equals 1 if the two partitions are identical, and equals 0 when the overlap between the two partitions is at chance level. According to Steinley (2004), *ARI* values greater than .90 indicate excellent recovery, whereas values greater than .80 indicate good recovery and values greater than .65 are considered moderate recovery. The mean *ARI* over all data sets amounted to .82 ($SD = .31$), which indicates a good recovery. The *ARI* was equal to 1 for 67.5% of the data sets. Table 4 presents the mean *ARI* values (for the analyses with 25 starts) in function of the simulated conditions. Clearly, and not surprisingly, the recovery of the clustering depends very strongly on the group sizes and on the size of the between-cluster loading differences. On the one hand, the mean *ARI* is .95 and .97 for groups of 300 and 500 subjects, but only .86 for group sizes of 100 and smaller than .80 for group sizes of 30 and 50. On the other hand, the *ARI* is .95 for the primary loading shift differences, around .89 for loading differences (cross-loadings or primary loading decreases) of .40 and around .67 for the differences of .20. Note that the latter still indicates a moderate recovery according to the guidelines in Steinley (2004) for loading differences so small that they may be considered ignorable (Stevens, 1992). On top of that, the *ARI* results were affected by the local maxima detected in the previous paragraph. Out of the 6,508 solutions with at least one incorrect cluster assignment, 2,994 were in fact a local maximum. After replacing the 3,002 local maxima (obtained with 25 starts) by the solutions obtained with 50 starts, the number of data sets with incorrect assignments reduced from 6,508 to 5,751 and the overall mean of the *ARI* amounted to .84 ($SD = .30$).

To have a more detailed look at how the recovery of the clustering is affected by the above mentioned aspects, Table 5 presents the mean *ARI* values for all combinations of the group sizes on the one hand and the other simulated conditions on the other hand. In addition to the results for MMG-FA with 25 random starts, it also includes the *ARI* results after replacing the 3,002 local maxima by the solutions obtained with 50 starts. When inspecting the latter results, we find that, for each level of the group sizes, the cluster recovery depends most on the size of the loading differences. It is interesting to see how this recovery improves with

Table 4
Mean Adjusted Rand Index (ARI) of the Estimated Clustering of the Groups in Comparison to the True Clustering, in Function of the Simulated Conditions, for MMG-FA and MSFA

Simulated conditions	*ARI* MMG-FA	*ARI* MSFA
$G = 12$	.78	.49
$G = 60$	.86	.65
$N_g = 30$	.59	.44
$N_g = 50$	.73	.51
$N_g = 100$	.86	.59
$N_g = 300$	.95	.66
$N_g = 500$	.97	.67
Equal clusters	.87	.66
Unequal clusters	.77	.49
$K = 2$	.92	.70
$K = 4$	.72	.45
$Q = 2$	.86	.64
$Q = 4$	.78	.50
Primary loading shift	.95	.95
Crossloading .40	.89	.74
Crossloading .20	.68	.22
Primary loading decrease .40	.90	.73
Primary loading decrease .20	.67	.22
Overall	.82	.57

This document is copyrighted by the American Psychological Association or one of its allied publishers.
This article is intended solely for the personal use of the individual user and is not to be disseminated broadly.

Table 5
Mean Adjusted Rand Index (ARI) of the Estimated Clustering of the Groups in Comparison to the True Clustering, for All Combinations of the Group Sizes N_g (Columns) and the Other Simulated Conditions (Rows), for MMG-FA With 25 Random Starts (Left) and With 25 or 50 Random Starts (i.e., 50 Starts When 25 Starts Resulted in a Local Maximum; Right)

	ARI 25 starts					*ARI* 25 or 50 starts				
Simulated conditions	30	50	100	300	500	30	50	100	300	500
$G = 12$	.51	.64	.80	.95	.98	.52	.67	.83	.96	.99
$G = 60$	.67	.82	.92	.96	.96	.68	.84	.95	.99	.99
Equal clusters	.65	.80	.92	.99	1.00	.66	.81	.93	.99	1.00
Unequal clusters	.52	.66	.80	.92	.93	.54	.70	.85	.96	.97
$K = 2$	.73	.88	.97	1.00	1.00	.74	.88	.98	1.00	1.00
$K = 4$	.44	.58	.75	.91	.93	.47	.62	.80	.95	.97
$Q = 2$	.65	.79	.90	.97	.97	.67	.81	.93	.99	.99
$Q = 4$	.52	.67	.82	.94	.96	.54	.69	.85	.96	.98
PLS	.90	.94	.96	.98	.98	.93	.96	.98	1.00	.99
CL .40	.74	.85	.92	.98	.98	.77	.88	.95	.99	.99
CL .20	.29	.50	.75	.91	.95	.29	.51	.78	.94	.97
PLD .40	.75	.86	.93	.98	.99	.77	.88	.96	1.00	1.00
PLD .20	.25	.50	.74	.92	.94	.25	.51	.77	.94	.97
Overall	.59	.73	.86	.95	.97	.60	.75	.89	.98	.99

Note. PLS = primary loading shifts; CL = crossloadings; PLD = primary loading decreases.

increasing group sizes. For loading differences of .20, the recovery is very bad for group sizes of 30 but becomes moderate (mean *ARI* of .77 or .78) when the groups contain 100 subjects and excellent (mean *ARI* of .94 to .97) when the groups contain 300 or 500 subjects. For loading differences of .40, the recovery is moderate (mean *ARI* of .77) for the smallest group sizes and it quickly improves for larger group sizes, with a good recovery (mean *ARI* of .88) when groups contain 50 subjects each and excellent to perfect recovery when groups contain at least 100 subjects. For the primary loading shift differences, the mean *ARI* exceeds .90—indicating excellent recovery—for all group sizes. Other important factors are the number of groups, the number of clusters and whether the clusters are of equal size or not. Specifically, in addition to larger groups, having more groups, less clusters and/or clusters of equal size increases the amount of information that is available on the cluster-specific MMs (i.e., the within-cluster sample size) and thus improves the cluster recovery. On top of that, the cluster recovery is better in case of two rather than four factors. To scrutinize this further, Table 6 shows the *ARI* values separately for the conditions with two and four factors and for the two smallest group sizes (i.e., 30 and 50), crossed with the type and size of loading differences. For comparison, we simulated 1,600 additional data sets with one factor only and group sizes of 30 or 50 and added the resulting *ARI* values to Table 6. Note that, in case of one factor, primary loading decreases are the only relevant loading differences. Table 6 clearly shows that, in case of two factors, even group sizes of 30 are sufficient to achieve an excellent cluster recovery for primary loading shift differences (mean *ARI* of .98) and a good recovery for crossloadings or primary loading decreases of .40 (mean *ARI* of .83 and .85, respectively). The recovery for the differences of .40 becomes excellent (mean *ARI* of .94 or .95) with group sizes of 50. In case of one factor, the mean *ARI* for primary loading decreases of .40 equals .90 even with group sizes of 30.

To examine the occurrence of classification uncertainty, we computed the minimum posterior probability with which a group was assigned to a cluster (according to the modal cluster assignments), that is, the minimum "classification certainty" or "CC_{min}," for each data set. For the data sets with a perfect cluster recovery (i.e., *ARI* = 1), CC_{min} varied between .50 and 1.00, with a mean of .9979 (*SD* = .02). For the data sets with at least one misclassification, CC_{min} varied between .32 and 1.00 with a mean of .93 (*SD* = .12). Thus, for the simulated conditions, classification uncertainty is quite infrequent and hardly related to misclassification.

Finally, we compared the cluster recovery of MMG-FA to that of MSFA. Over all data sets, the mean *ARI* of MSFA amounted to .57 (*SD* = .43) and the *ARI* was equal to 1 for only 43% of the data sets. Thus, the performance of MSFA is clearly inferior to that of MMG-FA when it comes to recovering the clustering that is underlying the between-group loading differences. Table 4 includes the mean *ARI* values for MSFA in function of the simulated conditions. It is obvious that MSFA performs almost as well as MMG-FA when it comes to picking up the largest loading differ-

Table 6
Mean Adjusted Rand Index (ARI) of the Estimated Clustering of the Groups in Comparison to the True Clustering, for All Combinations of the Two Smallest Group Sizes N_g, Number of Factors Q, and the Type and Size of Loading Differences

	$N_g = 30$			$N_g = 50$		
Loading differences	$Q = 1$	$Q = 2$	$Q = 4$	$Q = 1$	$Q = 2$	$Q = 4$
PLS		.98	.88		.98	.95
CL .40		.83	.70		.94	.83
CL .20		.36	.22		.59	.44
PLD .40	.90	.85	.69	.95	.95	.82
PLD .20	.48	.32	.19	.68	.60	.42

Note. PLS = primary loading shifts; CL = crossloadings; PLD = primary loading decreases.

This document is copyrighted by the American Psychological Association or one of its allied publishers.
This article is intended solely for the personal use of the individual user and is not to be disseminated broadly.

ences (i.e., primary loading shifts) and that its performance becomes inferior when the loading differences are smaller, probably because the clustering then focuses on the differences in factor (co)variances and/or unique variances.

Goodness of loading recovery. To evaluate the recovery of the cluster-specific loading matrices, we obtained a *goodness-of-cluster-loading-recovery statistic* (*GOCL*; De Roover, Ceulemans, Timmerman, Vansteelandt, Stouten, & Onghena, 2012) by computing congruence coefficients φ (Tucker, 1951) between the loadings of the true and estimated factors and averaging across factors and clusters as follows:

$$GOCL = \frac{\sum_{k=1}^{K}\sum_{q=1}^{Q}\varphi(\boldsymbol{\lambda}_{kq}, \hat{\boldsymbol{\lambda}}_{kq})}{KQ} \quad (6)$$

where $\boldsymbol{\lambda}_{kq}$ and $\hat{\boldsymbol{\lambda}}_{kq}$ indicate the true and estimated loading vector of the q-th factor for cluster k, respectively. The rotational freedom of the factors per cluster was dealt with by an oblique procrustes rotation of the estimated toward the true loading matrices. To account for the permutational freedom[2] of the cluster labels (also referred to as "label switching"; Tueller, Drotar, & Lubke, 2011), the estimated clusters were matched to the true clusters such that the *GOCL* value was maximized. The *GOCL* statistic takes values between 0 (*no recovery at all*) and 1 (*perfect recovery*). For the simulation, the average *GOCL* is .9940 (*SD* = .01), which corresponds to an excellent recovery that is hardly affected by the manipulated conditions—probably because misclassifications of groups occur mostly when between-cluster loading differences are small.

Model selection. Because the cluster recovery was not even moderate with group sizes of 30 (i.e., mean *ARI* < .65), we only report model selection results for group sizes of 50 or larger. Per data set, the number of clusters with the optimal balance between log*L* and the number of free parameters *fp* was determined according to four model selection procedures: BIC using the number of subjects *N* as the sample size (BIC_N), BIC using the number of groups *G* as the sample size (BIC_G), AIC and CHull. For BIC_N, the percentage of data sets for which the correct number of clusters was chosen is 55.6%. Specifically, BIC_N has a tendency to select one cluster. BIC_G selects the correct number of clusters for 76.3% of the data sets, whereas AIC and CHull do so for 79.9% and 81.4% of the data sets, respectively. For these three criteria, most of the model selection mistakes pertain to the number of clusters being underestimated.

The main effects of the simulated conditions on the performance of the four criteria are given in Table 7. Because BIC_N performed clearly inferior, we focus on the other three criteria in what follows. When looking at the effect of the type and size of the loading differences, it becomes clear that BIC_G, AIC, and CHull show comparable performances in case of the more pronounced primary loading shift differences, with percentages correct of 92.5%, 91.3%, and 91.6%, respectively. The performance drops when loading differences become more subtle. For crossloading differences or primary loading decreases of .40, all three criteria still selected the number of clusters with an accuracy of about 87%. When differences are as subtle as .20, the accuracy of BIC_G and AIC drops to about 56% and 66%, respectively, whereas the accuracy of CHull is 68.4 to 72.2%. It is also interesting to note that, for all three criteria, the performance depends strongly on the group sizes and that CHull outperformed BIC_G and AIC when groups consist of only 50 subjects.

Table 7
Percentage of Data Sets for Which the Model Selection Procedures (Specifically, BIC_N, BIC_G, AIC, and CHull) Select the Correct Number of Clusters K, in Function of the Simulated Conditions

Simulated conditions	BIC_N	BIC_G	AIC	CHull
$G = 12$	40.8	71.1	75.1	76.5
$G = 60$	70.4	81.4	84.8	86.4
$N_g = 50$	29.3	51.3	59.8	66.3
$N_g = 100$	45.3	73.5	77.3	79.8
$N_g = 300$	71.3	88.5	90.5	89.0
$N_g = 500$	76.5	91.8	92.3	90.8
Equal clusters	65.4	86.1	90.4	91.9
Unequal clusters	45.8	66.4	69.5	71.0
$K = 2$	72.4	92.0	94.1	97.8
$K = 4$	38.8	60.5	65.8	65.1
$Q = 2$	61.9	80.8	82.9	84.1
$Q = 4$	49.3	71.8	77.0	78.8
Primary loading shift	87.8	92.5	91.3	91.6
Crossloading .40	65.0	86.9	88.1	87.5
Crossloading .20	31.3	58.4	66.3	72.2
Primary loading decrease .40	64.4	87.2	86.6	87.5
Primary loading decrease .20	29.4	56.3	67.5	68.4
Overall	55.6	76.3	79.9	81.4

Of course, one may argue that the comparison of CHull to the other two criteria is not entirely fair, because AIC and BIC_G may select just one cluster in case of very subtle differences, whereas CHull always selects at least two clusters. Therefore, we checked the performance of AIC and BIC_G when only considering solutions with two or more clusters. By doing so, the overall accuracy of BIC_G and AIC increased to 80.3% and 82.1%, respectively, which is comparable to the 81.4% of CHull.

Simulation Study 2

The goal of Simulation Study 2 is to investigate the performance of the MMG-FA model selection criteria when metric invariance exists among all groups and, thus, the number of clusters is one. To this aim, we simulated data as detailed in Simulation Study 1, with five replications per cell of the design and retaining the following manipulated factors: (a) the number of groups; (b) the group sizes (excluding the smallest group sizes of 30); and (c) the number of factors. Additionally, we manipulated: (d) exact or approximate metric invariance across groups. To achieve exact metric invariance across all groups, the loadings of each group were made equal to the so-called "base loading matrix" from Simulation Study 1 (i.e., no loading differences were induced). Approximate metric invariance (Muthén & Asparouhov, 2013) was achieved by inducing small differences across groups for each loading, sampled from a normal distribution with a mean of zero and a variance of .0009. In this way, the ±2 *SD* difference in the loadings between groups was between −.064 and .064, which is considered to be negligible

[2] Permutational freedom refers to the fact that different combinations of the estimated and true clusters are possible.

This document is copyrighted by the American Psychological Association or one of its allied publishers.
This article is intended solely for the personal use of the individual user and is not to be disseminated broadly.

(Kim et al., 2017). The resulting 160 data sets were analyzed with MMG-FA with one to six clusters.

Out of the 80 data sets with exact invariance, BIC_G selected one cluster for all of them and AIC selected one cluster for 78 data sets. Thus, AIC overestimated the number of clusters for two data sets; specifically, two clusters were selected for both of them. For the 80 data sets with approximate invariance, BIC_G and AIC selected one cluster for 52 and 22 data sets, respectively. For the other data sets, BIC_G mostly selected two or three clusters, whereas AIC selected up to six clusters. Furthermore, BIC_G only selected more than one cluster in case of large group sizes (i.e., 300 or 500), whereas AIC does so for all simulated conditions.

Because no scree ratio can be computed for the least complex solution, CHull automatically selects at least two clusters. Specifically, across the 160 data sets, CHull selected two clusters for 55 data sets, three clusters for 47 data sets, four clusters for 31 data sets, and five clusters for 27 data sets. However, upon visual inspection of the CHull plot, one may still conclude that the elbow for the selected number of clusters is barely visible and thus that an underlying clustering is unlikely. Because visual inspection for all simulated data sets is infeasible, we examined the scree ratios for the selected number of clusters (i.e., the highest scree ratio) for all data sets. An elbow that is hardly visible would correspond to a scree ratio that is close to one. On average, the selected scree ratio was equal to 1.37 (SD = .20) in case of exact invariance and 1.42 (SD = .17) in case of approximate invariance. Boxplots of the scree ratios are given in Figure 1. For comparison, the selected scree ratio was on average equal to 40.37 (SD = 118.85) for the data sets from Simulation Study 1 and the corresponding boxplots are included in Figure 1, per number of clusters K and type and size of loading differences. Note that the smallest selected scree ratios occurred for the crossloadings and primary loading decreases of .20: On average, they were 9.91 for $K = 2$ and 4.33 for $K = 4$. In general, we conclude that the scree ratio of the selected solution is (a lot) larger—and, thus, that the elbow is more outspoken—when $K > 1$ than when $K = 1$.

Conclusion

Regarding the sensitivity to local maxima, we conclude that the multistart procedure of MMG-FA with 25 starts is sufficient to largely avoid local maxima, but that it is certainly advisable to increase the number of starts to at least 50 when the number of clusters is higher or when group sizes are small. The recovery of the cluster memberships of the groups was excellent for large loading differences (i.e., primary loading shifts) but depended strongly on the simulated conditions for loading differences of .40 or .20. Especially when detecting loading differences of .20 and/or when the number of factors is higher, a larger within-cluster sample size is essential (i.e., more groups, larger groups, and/or less clusters). However, it is important to note that these differences are so small that they are not harmful (Stevens, 1992) and that the insufficient group sizes of 30 and 50 would be even more problematic for a standard MG-FA (i.e., because groups are not clustered together). Anyhow, MMG-FA clearly outperformed MSFA in terms of cluster recovery, especially for smaller loading differences. The recovery of the cluster-specific factor loadings was excellent overall.

For selecting the most appropriate number of clusters, BIC_G, AIC, and CHull were found to perform quite similarly, at least for

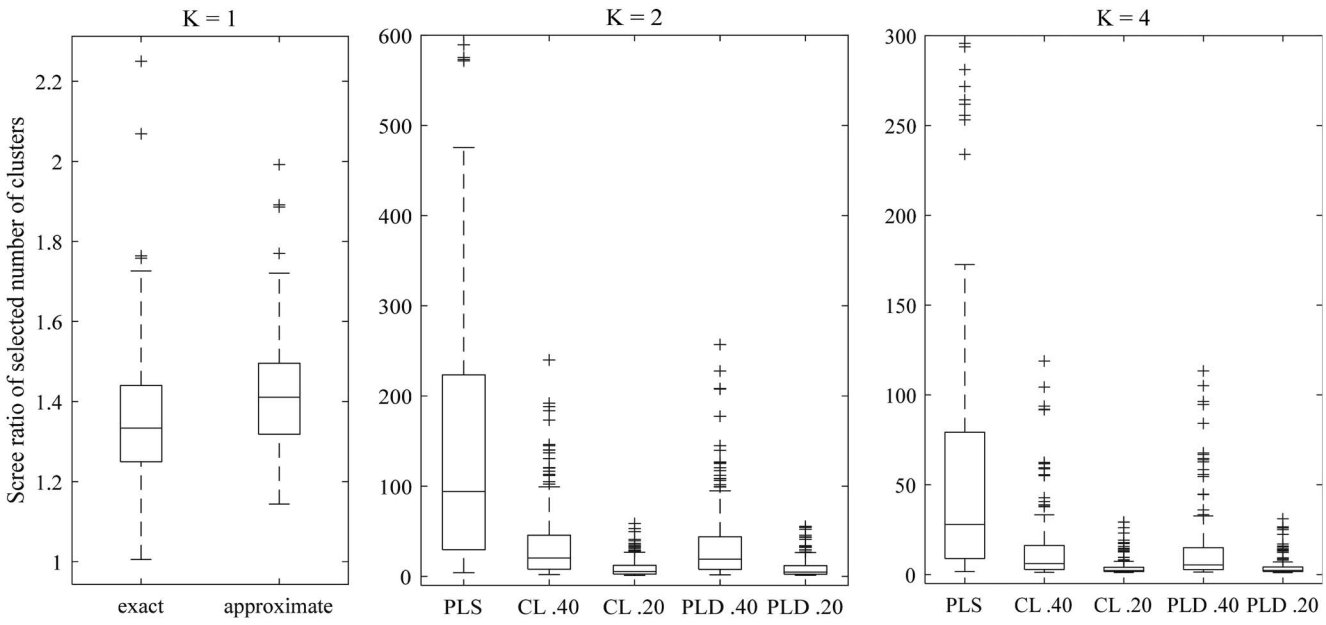


Figure 1. Boxplots of the scree ratios of the selected number of clusters for the data sets from Simulation Study 2 (true number of clusters K equal to 1), for exact and approximate metric invariance across all groups, and for the data sets from Simulation Study 1 ($K = 2$ and $K = 4$), for each level of the type and size of loading differences. "PLS" refers to primary loading shifts, "CL" to crossloadings, and "PLD" to primary loading decreases. Note that the scale of the y-axis differs across the subplots.

This document is copyrighted by the American Psychological Association or one of its allied publishers.
This article is intended solely for the personal use of the individual user and is not to be disseminated broadly.

the simulated conditions in Simulation Study 1. BIC_G and AIC have the added value that they can automatically distinguish between one cluster (i.e., metric invariance across all groups) and more clusters (i.e., metric noninvariance across all groups), but CHull makes no distributional assumptions and, thus, may perform better for empirical data. From Simulation Study 2, we conclude that, when K equals one but the metric invariance across groups is approximate rather than exact, AIC usually overestimated the number of clusters, whereas BIC_G did so in case of large groups.[3] For both exact and approximate metric invariance across groups, the selected CHull scree ratio was close to one, which indicates that the selected elbow would hardly be visible in the CHull plot. Therefore, our recommendation is to use BIC_G and AIC in combination with CHull whenever possible. For performing CHull, one can use the free software presented by Wilderjans, Ceulemans, and Meers (2013) or the R-package that can be downloaded from https://cran.r-project.org/package=multichull.

Applications

Social Value of Emotions

To illustrate the empirical value of MMG-FA, we applied it to cross-cultural data on how much experiencing certain emotions is socially valued. The data was collected as part of the International College Survey, 2001 (Diener et al., 2001; Kuppens, Ceulemans, Timmerman, Diener, & Kim-Prieto, 2006), which included 10,018 participants out of 48 different nations. Each of them rated, among other things, how much each emotion of a given set is appropriate, valued and approved in their society, using a 9-point Likert scale (1 = *people do not approve it at all* to 9 = *people approve it very much*). This data set was used by Bastian, Kuppens, De Roover, and Diener (2014) to show that living in a country that places more social value on positive emotions is related to a higher life satisfaction, even when allowing for an interaction with the frequency of experiencing positive and negative emotions. Such an association was not found for the social value of negative emotions. Specifically, they included the following positive emotions: happy, love, cheerful, pride, and gratitude. The negative emotions were: sad, jealousy, worry stress, anger, guilt and shame. Bastian et al. (2014) excluded Egypt from the analysis (for an unspecified reason) and, therefore, we also excluded it from the analyses reported below. For the remaining countries, participants with missing data were omitted, so that 8,773 participants were retained in the data set. The countries included in the analyses (with their retained sample size between brackets) are: Australia (171), Austria (119), Bangladesh (89), Belgium (112), Brazil (234), Bulgaria (122), Cameroon (85), Canada (98), Chile (339), China (328), Colombia (331), Croatia (135), Cyprus (90), Georgia (98), Germany (138), Ghana (135), Greece (211), Hong Kong (174), Hungary (514), India (106), Indonesia (236), Iran (171), Italy (280), Japan (164), South Korea (177), Kuwait (64), Malaysia (351), Mexico (298), Nepal (91), Netherlands (37), Nigeria (264), Philippines (187), Poland (527), Portugal (221), Russia (104), Singapore (89), Slovakia (100), Slovenia (270), South Africa (26), Spain (311), Switzerland (138), Thailand (182), Turkey (115), Uganda (106), United States (340), Venezuela (196), Zimbabwe (99).

Before computing social value indices for positive emotions and for negative emotions per country, measurement invariance testing is necessary to avoid drawing invalid conclusions. Bastian et al. (2014) performed invariance tests for positive and negative emotions separately and found the factor loading of "pride" to be noninvariant across the countries, which was thus excluded from their index of social value of positive emotions. By testing for factor loading invariance per factor, that is, for the positive and negative emotions separately, the possibility of crossloadings between the two factors—and the misfit that may result from restricting them to zero—was disregarded. To properly evaluate the tenability of the measurement model with the two factors—that is, "social value of positive emotions" (POS) and "social value of negative emotions" (NEG)—and potentially gain more insight in the noninvariance of the loadings for "pride," we performed a multigroup CFA by means of the R-packages lavaan 0.6–5 and semTools 0.5–2 (Rosseel, 2012). We specified one factor with nonzero loadings of happy, love, cheerful, pride, and gratitude, and a second factor with nonzero loadings of sad, jealousy, worry stress, anger, guilt, and shame. Even though the emotions are rated on a Likert scale, Dolan (1994) showed that ordinal data can be considered continuous and the ML estimator is quite robust when the number of response categories is at least five and when the data is not severely non-normal. Because the ratings have nine response categories and none of the variables have a skewness or kurtosis outside of the acceptable range (i.e., skewness < 2 and kurtosis < 7; George & Mallery, 2010), we expect ML parameter estimates to be robust. However, to stay clear of biased standard errors and χ^2-based fit indices, we report Satorra-Bentler corrected fit indices (Satorra & Bentler, 1994). Note that, due to the very large sample size across the 47 countries, the χ^2-tests of exact model fit as well as the χ^2-difference tests for comparing nested models (Satorra & Bentler, 2001) were significant for all analyses reported below. Therefore, we evaluate model fit based on the CFI and RMSEA indices and consider CFI $\geq$ .90 and RMSEA $\leq$ .08 as indicators of acceptable model fit whereas CFI $\geq$ .95 and RMSEA $\leq$ .06 would indicate very good model fit (Hu & Bentler, 1999). When imposing invariant loadings, we make use of the guidelines reported by Rutkowski and Svetina (2014) to evaluate metric invariance and check whether ΔCFI $\geq$ –.02 and ΔRMSEA $\leq$ .03.

The fit for the configural invariance model—with nonzero loadings as specified above—was bad (CFI = .819, RMSEA = .106), indicating that the a priori assumed measurement model does not hold or not for all countries. Note that negative unique variances were found for "happy" in Russia and for "cheerful" in Uganda. From an inspection of the modification indices, we concluded that a crossloading for "pride" on the NEG factor would improve the model fit for many groups. With this modification, we obtained a slightly improved model fit; that is, CFI = .844 and RMSEA = .10. For some groups, other crossloadings and residual covariances were suggested by the modification indices, but to avoid capitalizing on chance we refrained from further model modification

[3] Of course, strictly speaking, we cannot speak of overselection of clusters in case of approximate metric invariance, because the "true" number of clusters is larger than one due to the small loading differences, which are not captured within the clusters of MMG-FA. However, for the sake of parsimony, we are optimistic about the fact that often only one cluster was selected. In the Discussion section, we discuss how the MMG-FA solution can be used to evaluate whether exact or approximate invariance holds within the clusters.

This document is copyrighted by the American Psychological Association or one of its allied publishers.
This article is intended solely for the personal use of the individual user and is not to be disseminated broadly.

(Browne, 2001; MacCallum et al., 1992; Silvia, & MacCallum, 1988).

To evaluate crossloadings, and differences therein, as well as primary loading differences across countries, without making the 1,081 pairwise comparisons across the 47 country-specific EFA models, we switch to MMG-FA. To select the most appropriate number of clusters, we performed MMG-FA with one up to eight clusters. BIC_G and AIC select eight clusters (see Table 8), which may be an overselection. According to CHull, the best number of clusters is two (with a scree ratio of 2.06) and the second best is three (with a scree ratio of 1.61). From the CHull plot given in Figure 2, it becomes clear that the fit still improves considerably by adding a third cluster whereas the plot clearly levels off after three clusters. Thus, we looked at both the two- and three-cluster solution and found that the two-cluster solution is mainly about differences in the loadings of "pride," whereas in the three-cluster solution we also see some interesting differences for "guilt" and "shame." Therefore, we chose to report the three-cluster solution.

The clustering of the selected model is given in Table 9. Most countries are assigned to one of the clusters with a posterior probability of 1, whereas a small amount of classification uncertainty is found for Belgium, Switzerland, and India. The first thing to note is that Cluster 1 mainly contains Western countries, whereas Cluster 3 mainly gathers non-Western countries. Cluster 2 contains an interesting mix of Western and non-Western countries. To see which loading differences resulted in this clustering, we inspect the cluster-specific loadings in Table 10. For each cluster, these loadings are rotated toward a target where the positive emotions make out the first factor (with target loadings of 1) and the negative ones the second factor. In each cluster, the distinction between positive and negative emotions is found, at least to some extent, such that we can label the factors as "social value of positive emotions" (POS) and "social value of negative emotions" (NEG) in all clusters. Some differences clearly stand out, however, the most important ones pertaining to the self-conscious or self-reflective emotions pride, shame, and guilt.

In Cluster 1, consisting of mainly Western countries, pride has a strong loading on POS and a near-zero loading on NEG, which corresponds to the fact that "pride" was considered to be a positive emotion by Bastian et al. (2014). However, in Cluster 3, pride has a strong crossloading on NEG and, in Cluster 2, it even loads primarily on NEG (but still has a strong crossloading on POS). This implies that, in Cluster 2, the value of pride is affected more by the value placed on negative emotions than by the value of positive emotions. In Cluster 3, the value rating for pride is affected primarily by the POS factor, but also to a large extent by the NEG factor. Thus, in contrast to its positive status in many Western countries, pride (also) belongs to the negative emotions—in terms of its social value—in the countries of Cluster 2 and 3.

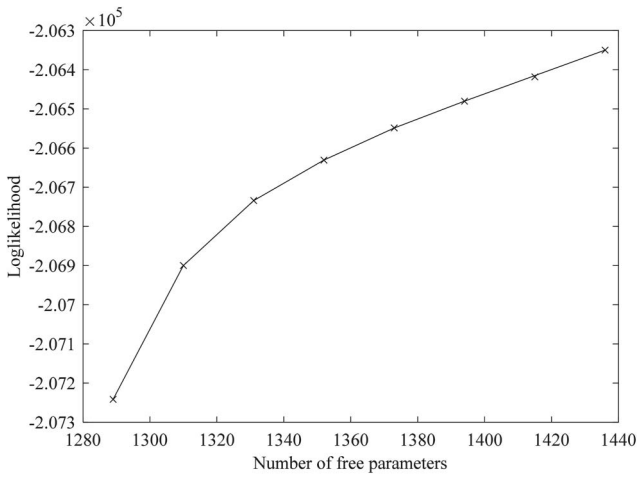


Figure 2. Convex Hull (CHull) plot of the loglikelihood in function of the number of free parameters for MMG-FA models with one to eight clusters for the social value of emotions data set.

For guilt and shame, important differences are found as well. Even though they both load primarily on NEG in all three clusters, the sizes of these loadings as well as the crossloadings on POS differ across clusters. In Cluster 1, guilt and shame have small crossloadings on POS and, in Cluster 2, larger—but still subtle—crossloadings are found and the primary loadings on NEG are somewhat weaker. In Cluster 3, however, the crossloadings are very salient and almost as strong as the primary loadings, since the latter are a lot lower than in Clusters 1 and 2. Thus, for the (mainly non-Western) countries in Cluster 3, the ratings for guilt and shame are affected both by the value of negative emotions and the value of positive emotions, indicating that they are not unambiguously part of the negative emotions in terms of their social value.

Cross-cultural differences in the appropriateness of the self-conscious emotions have been studied extensively. Specifically,

Table 8
Loglikelihood (logL), Number of Free Parameters (fp), BIC_G, AIC, and CHull Scree Ratio for MMG-FA Models With One to Eight Clusters for the Social Value of Emotions Data Set

Number of clusters	log*L*	fp	BIC_G	AIC	CHull scree ratio
$K = 1$	−207241.6	1289	419446.0	417061.1	/
$K = 2$	−206899.9	1310	418843.4	416419.7	**2.06**
$K = 3$	−206734.0	1331	418592.6	416130.1	**1.61**
$K = 4$	−206630.8	1352	418467.0	415965.6	1.26
$K = 5$	−206548.6	1373	418383.5	415843.3	1.20
$K = 6$	−206480.3	1394	418327.8	415748.7	1.05
$K = 7$	−206415.2	1415	**418278.4**	**415660.4**	—
$K = 8$	−206349.9	1436	**418228.6**	**415571.8**	/

Note. The model with seven clusters was not on the convex hull. For each criterion, the values for the two best models are in bold face.

This document is copyrighted by the American Psychological Association or one of its allied publishers.
This article is intended solely for the personal use of the individual user and is not to be disseminated broadly.

Table 9
Clustering of the Countries of the MMG-FA Model With Three Clusters and Two Factors per Cluster for the Social Value of Emotions Data Set

Cluster 1	Belgium ($\hat{z}_{g1}$ = .98), Canada, Chile, Colombia, Croatia, Germany, Kuwait, Mexico, Netherlands, Slovenia, Spain, Switzerland ($\hat{z}_{g1}$ = .96), Venezuela, Zimbabwe
Cluster 2	Australia, Austria, Brazil, Cameroon, China, Cyprus, Ghana, Greece, Hong Kong, Iran, Italy, Malaysia, Nepal, Nigeria, Philippines, Poland, Portugal, Russia, Singapore, South Africa, Turkey, United States
Cluster 3	Bangladesh, Bulgaria, Georgia, Hungary, India ($\hat{z}_{g3}$ = .92), Indonesia, Japan, Slovakia, South Korea, Thailand, Uganda

Note. All countries are assigned to the clusters with a posterior probability $\hat{z}_{gk}$ of 1, except for Belgium, India, and Switzerland. The probabilities for the latter countries are given between parentheses.

guilt and shame were found to be more desirable in countries with collectivistic values (Bedford, 2004; Cole, Bruschi, & Tamang, 2002; Eid & Diener, 2001; Moore, Romney, Hsia, & Rusch, 1999; Mosquera, Manstead, & Fischer, 2000), which is explained by the fact that they result from violating social norms and failure to fulfill social obligations (Eid & Diener, 2001). In contrast, pride is evoked when personal goals are achieved and is thus highly valued in individualistic cultures and much less so in collectivistic ones (Eid & Diener, 2001).

In Figure 3, boxplots of four collectivism measures (which were available for 35 out of the 47 countries) are given per cluster of the MMG-FA solution. Specifically, institutional and in-group collectivism were measured by the Societal Cultural Practices Scale on the one hand and the Societal Cultural Values Scale on the other hand (House, Hanges, Javidan, Dorfman, & Gupta, 2004). In-group collectivism pertains to family and friend groups, whereas institutional collectivism has more to do with the work environment and society. Cultural practices are perceptions of how people behave in a culture (how it is) and cultural values are ideals of a culture (how it should be; Frese, 2015). The correlation between cultural practices and values is often insignificant or even negative (House et al., 2004). With regard to cultural practices, Cluster 1 seems to contain the least collectivistic countries, whereas Cluster 3 contains the most collectivistic countries and Cluster 2 lies in between. The three clusters overlap to a large degree in terms of institutional and in-group collectivism, however. This indicates that collectivism is probably not the only dimension that explains cross-cultural differences in value of emotions. For instance, the tightness by which cultural norms are enforced on the members of a society may also have an impact (Triandis, 1989). Additionally, linguistic factors could play a role in the construction of emotional realities (Wierzbicka, 1999). Another interesting thing to note is that experiencing pride is clearly the least desirable in the countries of Cluster 2, whereas these countries are *not* the ones with the largest crossloadings for guilt and shame. Thus, it seems that the value of pride varies independently of the value of guilt and shame. Together with the fact that collectivism is not sufficient to explain these differences, this is not only an interesting finding in itself but also an incentive for future research.

To conclude, MMG-FA pointed out some fascinating cross-cultural differences in the underlying structure of emotion values, but what would have been our advice to Bastian et al. (2014) regarding their study? First and most importantly, they were successful in detecting and excluding the most important noninvariant item from their index of social value of positive emotions: that is, pride. Second, they failed to detect important factor loading noninvariances with respect to guilt and shame. Other noninvariances that are observed when comparing the cluster-specific loadings in

Table 10
Target Rotated Loadings of the MMG-FA Model With Three Clusters and Two Factors per Cluster for the Social Value of Emotions Data Set

	Cluster 1		Cluster 2		Cluster 3	
Emotion	Social Value of Positive Emotions (POS)	Social Value of Negative Emotions (NEG)	Social Value of Positive Emotions (POS)	Social Value of Negative Emotions (NEG)	Social Value of Positive Emotions (POS)	Social Value of Negative Emotions (NEG)
Happy	**1.29**	−0.03	**1.40**	−0.20	**1.22**	−0.14
Love	**1.23**	0.05	**1.40**	−0.15	**1.25**	−0.03
Sad	0.16	**1.29**	0.09	**1.28**	0.29	**1.36**
Jealousy	0.02	**1.37**	0.17	**1.29**	0.14	**1.18**
Cheerful	**1.18**	0.02	**1.10**	−0.16	**1.07**	−0.01
Worry	0.30	**1.48**	0.15	**1.49**	0.06	**1.80**
Stress	−0.05	**1.77**	0.09	**1.73**	−0.26	**1.93**
Anger	−0.01	**1.60**	−0.07	**1.69**	−0.16	**1.70**
Pride	**0.93**	0.02	**0.40**	**0.94**	**0.72**	**0.42**
Guilt	0.14	**1.40**	0.24	**1.18**	**0.58**	**0.69**
Shame	0.07	**1.33**	0.28	**1.08**	**0.54**	**0.71**
Gratitude	**1.02**	−0.08	**0.89**	−0.19	**0.98**	−0.03

Note. For each cluster, the loadings are obliquely procrustes rotated toward a target structure representing the a priori assumed distinction between positive and negative emotions (i.e., with a "1" for positive emotions on the first factor and for negative emotions on the second factor, whereas other entries are equal to zero). Loadings with an absolute value greater than .40 are indicated in bold face.

This document is copyrighted by the American Psychological Association or one of its allied publishers.
This article is intended solely for the personal use of the individual user and is not to be disseminated broadly.

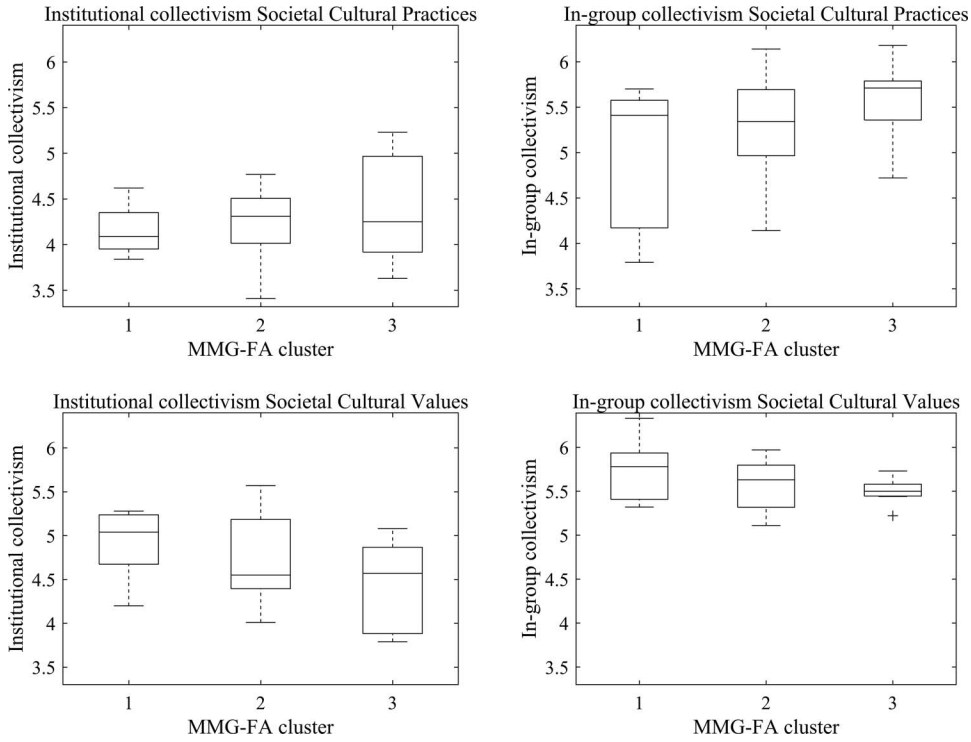


Figure 3. Boxplots of institutional and in-group collectivism (House et al., 2004), as measured by the Societal Cultural Practices Scale (above) and by the Societal Cultural Values Scale (below) per cluster of the MMG-FA solution for the social value of emotions data set.

Table 10 are a lot more subtle. Because guilt and shame are clearly noninvariant in the extent to which they measure the value of negative emotions, they should be excluded from the social value index of negative emotions to avoid invalid conclusions about its effect on life satisfaction. In fact, by removing pride, guilt, and shame and assuming that the remaining positive emotions have nonzero loadings on POS and that the remaining negative ones load on NEG, an acceptable fit is obtained for the configural invariance model, that is, CFI = .933 and RMSEA = .078. When imposing metric invariance, CFI becomes .917 and RMSEA amounts to .077. Thus, ΔCFI equals −.016 and ΔRMSEA is .001, which indicate that metric invariance holds. Because Bastian et al. (2014) used country-level social value indices—rather than individual indices—to predict the life satisfaction of its inhabitants, intercept or strong invariance was also required. Suggestions on how to move forward to further measurement invariance testing and between-group comparisons are given in the Discussion section.

Emotional Acculturation

As a second empirical illustration, we reanalyzed data from a study on emotional acculturation of immigrants (De Leersnyder, Mesquita, & Kim, 2011). The data pertain to samples from two host cultures (i.e., U.S. and Belgium), immigrant groups in the host cultures, and two heritage cultures (i.e., cultures of origin for some of the immigrants), yielding 13 groups in total (see Table 11). The participants reported on one to four situations that differed in valence (positive, negative), social engagement (engaged, disengaged), and social context (with friends, at home/with family, at school/work). The situations were chosen according to three designs. In Design 1, participants rated three situations of the same type (e.g., positive disengaging situation) for different social contexts. In Design 2, participants rated four different types of situations (i.e., positive disengaging, positive engaging, negative disengaging, or negative engaging situation) for the same social context. Design 3 was similar to Design 2, but participants only reported on two types of situations for the same context. The design was fixed within each group (see Table 11), which implies that differences between groups may partly be due to design differences. Participants rated on a 7-point Likert scale to what extent each situation elicited each of 17 emotions (see Table 13). Skewness and kurtosis values are inside the acceptable range (i.e., skewness < 2 and kurtosis < 7), except for the emotion "relying" (i.e., skewness = 2.4 and kurtosis = 37.6). Subject-situation combinations with missing values were discarded.

Again, we assume two factors, pertaining to positive and negative emotions, respectively. Interestingly, De Roover, Timmerman, De Leersnyder, Mesquita, and Ceulemans (2014) already performed measurement invariance testing on these data, confirming that metric invariance fails across groups and that configural invariance holds for almost all groups. In search for sources of noninvariance, they used Clusterwise SCA-P (De Roover, Ceulemans, Timmerman, & Onghena, 2013). This approach yields a heuristic clustering of groups and applies simultaneous component analysis within each cluster of groups. Although it is similar to MMG-FA in many ways, Clusterwise SCA-P does not include

This document is copyrighted by the American Psychological Association or one of its allied publishers.
This article is intended solely for the personal use of the individual user and is not to be disseminated broadly.

Table 11
The 13 Cultural Groups of the Emotional Acculturation Data Set and the Associated Host Country, Design, Sample Size and MMG-FA Cluster

Cultural group	Host country	Design	N_g after removing observations with missing values	N_g after sampling 1 observation per subject	MMG-FA cluster
European Americans 1	U.S.	1	120	42	1
Korean immigrants	U.S.	1	126	46	1
Mexican immigrants	U.S.	1	188	67	1
East-Asian immigrants	U.S.	2	159	37	1
Latino immigrants	U.S.	2	142	40	1
European Americans 2	U.S.	2	122	32	1
Koreans	Korea	2	298	79	1
Flemish students 1	Belgium	3	183	183	2
Flemish students 2	Belgium	3	516	264	2
Belgian community	Belgium	3	166	90	2
Turkish second generation immigrants	Belgium	3	157	83	2
Turkish first generation immigrants	Belgium	3	143	79	3
Turkish students	Turkey	3	699	375	3

Note. Sample size per group N_g is indicated (note: each subject-situation combination counts as one observation) after removing observations with missing values and after randomly selecting one situation per subject. The last column indicates to which cluster the cultural group is assigned in the MMG-FA model with three clusters and two factors per cluster. All countries are assigned to the clusters with a posterior probability $\hat{z}_{gk}$ of 1.

This document is copyrighted by the American Psychological Association or one of its allied publishers.
This article is intended solely for the personal use of the individual user and is not to be disseminated broadly.

unique variances, which implies that it does not account for differences in unique variances between groups within clusters. To see whether MMG-FA finds the same clustering, we performed analyses with one up to eight clusters and two factors per cluster. Note that we analyzed the raw data, whereas De Roover, Timmerman, De Leersnyder, Mesquita, and Ceulemans (2014) rescaled each variable to have a variance of one across all groups.[4] BIC_G and AIC select seven and eight, respectively, as the best number of clusters (see Table 12), whereas CHull selects three clusters (with a scree ratio of 1.94). The CHull plot (see Figure 4) clearly shows that the improvement in fit levels off after three clusters, which suggests that BIC_G and AIC may be overestimating the number of clusters. The clustering of the groups into three clusters is shown in Table 11 and agrees perfectly with the one reported by De Roover et al. (2014). The groups living in the U.S. are in Cluster 1, together with the Koreans. Cluster 2 contains the indigenous Belgian groups as well as the second generation Turkish immigrants. Cluster 3 contains the group living in Turkey and the first generation Turkish immigrants in Belgium. The fact that the second generation Turkish immigrants were assigned to the Belgian cluster suggests that they acculturated with respect to their emotions.

The fact that each subject rated up to four situations may imply that the conditional independence assumption of MMG-FA is violated. Retaining only one situation per subject drastically reduces the sample size per group, but allows to investigate whether MMG-FA still finds the same clustering, despite small sample sizes. To this end, we randomly sampled one situation per subject, retaining less than 100 subjects for 10 out of the 13 groups and less than 50 for five groups (see Table 11). We repeated the MMG-FA analyses on this subset of the data. Again, CHull clearly indicated three clusters (see Table 12 and Figure 4), containing exactly the same groups. Also, the loadings were highly similar, so we only discuss the loadings for the total data set (see Table 13).

The loadings of the three clusters were target rotated toward a positive (POS) factor (i.e., loadings of 1 for the first eight emotions) and a negative (NEG) factor (i.e., loadings of 1 for the remaining nine emotions). When inspecting the rotated loadings, some remarkable between-cluster differences become apparent. For instance, "proud about myself" has a strong positive loading on POS in the "U.S. and Koreans" and "Turkish" clusters, whereas, in the "Belgian" cluster, this emotion has a weak loading on POS and a stronger negative loading on NEG. Thus, when Belgians experience negative emotions, they feel less proud about themselves than people belonging to the other cultural groups. A similar but less outspoken difference is observed for "strong". As another example, "relying" has a moderate positive loading on NEG in the U.S. and Koreans cluster, where this loading is lower for the other clusters. Thus, in this cluster, relying on someone else may have a negative connotation. Finally, "resigned" has a lower loading on NEG in the Belgian and Turkish clusters in comparison to the U.S. and Koreans cluster, while it loads primarily on POS in the Turkish cluster, which indicates that resignation is regarded as more positive by Turkish people. For a more elaborate interpretation, see De Roover et al. (2014).

To summarize, we found important factor loading differences, indicating that some emotions covary differently with the other emotions or are even valued differently in some cultural groups. On the one hand, these configural and metric noninvariances preclude comparisons in terms of the latent variables POS and NEG between groups of different clusters. On the other hand, these cross-cultural differences are very interesting in itself, because the data were collected to study differences and similarities in emotional covariation. On top of that, the clustering indicated which cultural groups are highly similar, and, thus, which immigrant groups emotionally acculturated.

[4] They also centered the data per group, which corresponds to our group-specific treatment of the means.

Table 12
Loglikelihood (logL), Number of Free Parameters (fp), BIC_G, AIC, and CHull Scree Ratio for MMG-FA Models With One to Eight Clusters for the Emotional Acculturation Data Set: For the Total Data Set (Above) and for the Subset (Below)

Number of clusters	log*L*	fp	BIC_G	AIC	CHull scree ratio
$K = 1$	−102484.3	511	206279.3	205990.6	/
$K = 2$	−102107.3	542	205604.7	205298.5	1.72
$K = 3$	−101888.0	573	205245.7	204921.9	**1.94**
$K = 4$	−101774.8	604	205098.8	204757.6	**1.74**
$K = 5$	−101709.8	635	205048.3	204689.5	1.36
$K = 6$	−101661.8	666	205031.9	204655.6	1.06
$K = 7$	−101616.7	697	**205021.1**	**204627.3**	1.17
$K = 8$	−101578.2	728	**205023.6**	**204612.4**	/
$K = 1$	−47937.8	511	97186.3	96897.6	/
$K = 2$	−47793.3	542	96976.7	96670.5	1.05
$K = 3$	−47655.1	573	96780.0	96456.3	**2.61**
$K = 4$	−47602.3	604	**96753.8**	96412.5	**1.47**
$K = 5$	−47566.4	635	**96761.5**	**96402.7**	1.12
$K = 6$	−47534.3	666	96777.0	**96400.7**	1.11
$K = 7$	−47505.6	697	96798.9	96405.1	1.15
$K = 8$	−47480.5	728	96828.4	96417.1	/

Note. For each criterion, the values for the two best models are in bold face.

Discussion

MMG-FA is a promising new method that identifies clusters of groups sharing the same factor loadings and allows the user to gain insight in the cluster-specific measurement models that are underlying the data. It is especially useful for examining how and which factor loadings differ between many groups, because it ties down the number of loading matrices to compare, making it easier to identify items and/or groups causing factor loading invariance to fail. For instance, for the data on the social value of emotions in 47 countries, the comparison of only three cluster-specific loading matrices made it obvious that the self-conscious emotions were the main reason why configural invariance was failing. On top of that, the obtained clusters of groups are often empirically interesting in itself. For instance, in the empirical example on emotional acculturation, the clustering of the host cultures, heritage cultures, and immigrant groups indicated which immigrant groups had acculturated to their host cultures (i.e., were assigned to the same cluster as their host cultures).

Table 13
Target Rotated Loadings of the MMG-FA Model With Three Clusters and Two Factors per Cluster for the Emotional Acculturation Data Set

	Cluster 1 (U.S. & Koreans)		Cluster 2 (Belgian)		Cluster 3 (Turkish)	
Emotion	Positive (POS)	Negative (NEG)	Positive (POS)	Negative (NEG)	Positive (POS)	Negative (NEG)
Respect	**1.53**	−0.30	**1.53**	−0.07	**1.74**	−0.26
Interested	**1.47**	−0.33	**1.19**	**−0.46**	**1.39**	−0.02
Helpful	**1.55**	−0.06	**1.22**	−0.06	**1.31**	−0.12
Close	**1.31**	**−0.47**	**1.70**	0.28	**1.56**	**−0.43**
Strong	**1.41**	−0.23	**0.92**	**−0.82**	**1.62**	**−0.41**
Proud about myself	**1.45**	**−0.49**	**0.97**	**−1.26**	**1.68**	**−0.53**
Relying	**0.96**	**0.64**	**1.55**	0.31	**1.35**	−0.11
Surprised	**0.55**	**0.78**	**0.54**	−0.18	**1.32**	−0.16
Ill feelings	**−0.52**	**1.10**	**−0.51**	**1.08**	**−0.60**	**1.13**
Upset	**−0.78**	**1.42**	**−0.57**	**1.34**	**−0.70**	**1.40**
Irritated	**−0.43**	**1.42**	**−0.95**	**1.03**	−0.09	**1.46**
Embarrassed	0.20	**1.72**	−0.22	**1.09**	0.27	**1.77**
Ashamed	0.22	**1.72**	−0.23	**1.38**	0.17	**1.47**
Guilty	−0.05	**1.55**	−0.13	**1.63**	−0.05	**1.53**
Bored	0.10	**0.66**	−0.30	**0.41**	**−0.61**	**1.10**
Indebted	**0.87**	**0.80**	**0.55**	**1.42**	**0.61**	**0.94**
Resigned	−0.23	**1.08**	0.14	**0.52**	**0.47**	0.25

Note. For each cluster, the loadings are obliquely procrustes rotated toward a target structure representing the distinction between positive and negative emotions (i.e., with a "1" for the first eight emotions on Factor 1 and for the remaining emotions on Factor 2, whereas other entries are equal to zero). Loadings with an absolute value greater than .40 are indicated in bold face.

This document is copyrighted by the American Psychological Association or one of its allied publishers.
This article is intended solely for the personal use of the individual user and is not to be disseminated broadly.

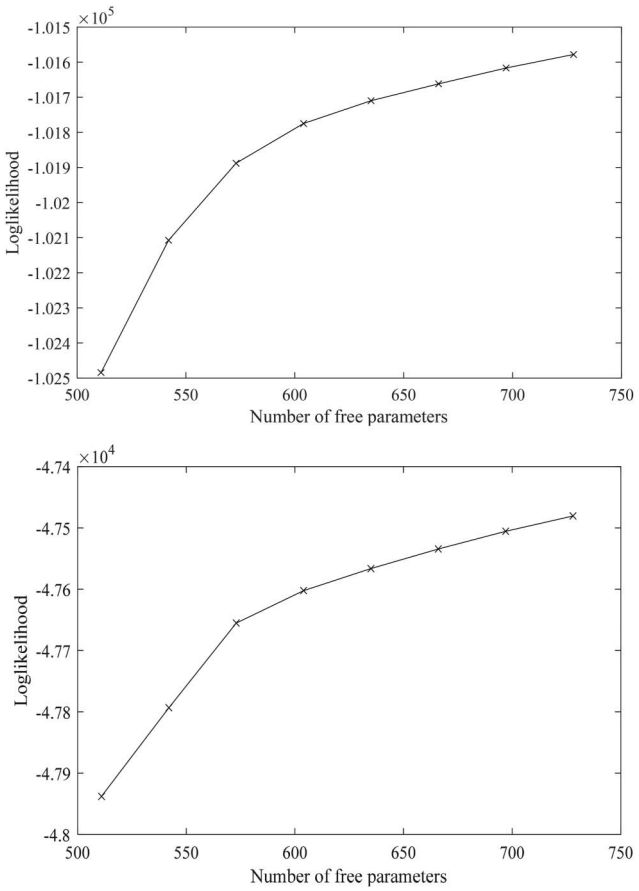


Figure 4. Convex Hull (CHull) plot of the loglikelihood in function of the number of free parameters for MMG-FA models with one to eight clusters for the emotional acculturation data set: for the total data set (above) and for the subset (below).

When comparing the cluster-specific factor loadings, it would be interesting to perform hypothesis testing to determine which between-cluster differences in factor loadings are significant and which loadings are significantly different from zero in each cluster. To this end, for each cluster, the rotational freedom of the factors should be resolved in the estimation procedure such that a fully identified model is obtained with optimally rotated estimates and proper standard errors. The recently proposed multigroup factor rotation (De Roover & Vermunt, 2019) solves this problem for the standard MG-EFA with maximum likelihood estimation and optimizes simple structure as well as agreement. It would thus be very useful to extend this rotation approach to the more complicated MMG-FA.

Currently, MMG-FA applies maximum likelihood estimation, which assumes continuous items and multivariate normality whereas many questionnaire items are Likert scale items and often some deviation from normality is present. Having five or more response categories allows for the items to be analyzed as continuous and a large sample size alleviates the effects of non-normality (Dolan, 1994; Lei & Lomax, 2005). At least, we expect the within-cluster sample sizes to be (mostly) large enough when multiple groups are gathered within each cluster. However, to properly deal with the non-normal and ordinal nature of items, we will consider extensions with robust estimators (Mîndrilă, 2010; Muthén, 1993) on the one hand and IRT-like multinomial logit or probit specifications for item responses (Agresti, 2013) on the other hand.

When one wants to continue toward further measurement invariance testing for the current data set and, potentially, between-group comparisons with respect to the latent variables, the MMG-FA solution provides the user with a few possibilities to do so. On the one hand, the comparison of the cluster-specific factor loadings may indicate that one or a few items are causing the noninvariance and, thus, that excluding these items—or making their loadings noninvariant to continue with partial invariance—makes it possible to continue with measurement invariance testing for all groups in the data set. Of course, one should make sure that the latent concepts of interest are still sufficiently reflected in the remaining set of items. On the other hand, the clustering may identify a few groups with deviating factor loadings and, in that case, excluding these groups is an option. A combination of noninvariant items and groups may be found as well, in which case one should consider which (combination of) items or groups to remove based on substantive considerations and maximizing the amount of retained data. Finally, when the noninvariant items and/or groups are too extensive, one can look into further invariance testing per cluster of groups. When one wants to improve the measurement instrument toward future studies, MMG-FA provides some useful clues as well. For instance, one could consider to exclude or rephrase items that are found to be noninvariant, potentially only for certain groups (e.g., for certain cultures or languages) as indicated by the clustering.

For investigating intercept invariance across many groups, the MMG-FA framework will be extended to include a variant that addresses intercept noninvariance in the same way; that is, by clustering the groups on intercept differences only. To this end, in addition to making the intercepts cluster-specific and imposing invariance of the factor loadings across all groups, group-specific factor means will be added whereas the factor (co)variances and unique variances are also group-specific. For the time being, the existing method that most closely approximates this specification is multilevel factor mixture modeling (MLFMM; Kim et al., 2017; Varriale & Vermunt, 2012), with the difference that neither factor means nor any of the other parameters are group-specific and that—for model identification—the clustering of the groups is based either on the intercepts (restricting the factor means to be zero per group) or on the factor means (restricting the intercepts to be zero per group). This implies that, when using MLFMM for clustering based on the intercepts, between-group differences in factor means will also be captured by the clustering and the cluster-specific intercepts.

To explore covariates of the noninvariances captured by the clustering, MMG-FA can easily be extended to allow for the inclusion of covariates; specifically, to model the relation between covariates and cluster membership (Lubke & Muthén, 2005). Alternatively, covariates can be added after estimating the MMG-FA model by means of the three-step approach (Vermunt, 2010). This would give researchers the opportunity to use group-level data, such as geographical region or economic, political or cultural indicators, to explain how the measurement model differs across groups. Furthermore, this could even improve the accuracy

This document is copyrighted by the American Psychological Association or one of its allied publishers.
This article is intended solely for the personal use of the individual user and is not to be disseminated broadly.

of the group clustering and parameter recovery (Lubke & Muthén, 2007). In the empirical example on the social value of emotions, it would have been interesting and potentially advantageous to include collectivism, and/or other potential determinants of emotion values, as a covariate.

Using EFA within the clusters may imply some capitalization on chance, but we believe this to be limited and to distort the clustering (and the parameter estimates) less than using CFA with potentially misspecified or overly restrictive zero loadings. Also, in the simulation study, the recovery of the clusters was shown to be excellent for larger loading differences and good for small ones in case of sufficient within-group sample sizes. Of course, for theory building, it would be interesting to return to a CFA approach at some point. On the one hand, based on the insights gained by MMG-FA, one can evaluate (partial) invariance of the established measurement models by MG-CFA for a subset of items and/or groups, where this subset is determined as described above. Preferably, this is performed for a new data set or, in case of a large enough sample size, by using split-sample crossvalidation for the current data set (Gerbing & Hamilton, 1996). On the other hand, a CFA-based variant of mixture multigroup factor analysis (MMG-CFA) could be developed, allowing to (cross)validate both the clustering and the cluster-specific model specifications. To this end, an extensive evaluation of MMG-CFA's robustness against cluster-specific model misspecifications is required. When configural invariance holds, MMG-CFA would also allow to investigate violations of metric invariance without including all the crossloadings and without the hassle of rotational freedom.

Furthermore, exact and full invariance of measurement parameters within a cluster may be too restrictive and unrealistic, especially in case of many groups. It could be that the groups within a cluster are identical with respect to most—but not all—MM parameters (i.e., partial invariance) or that approximate rather than exact invariance exists within clusters (Muthén & Asparouhov, 2013). Note that both partial and approximate invariance do not preclude further invariance tests or latent variable comparisons (Byrne, Shavelson, & Muthén, 1989; Muthén & Asparouhov, 2013). One can investigate partial and approximate invariance by means of existing approaches—such as modification indices (Sörbom, 1989), item-deletion strategies (Byrne & van de Vijver, 2010; De Roover et al., 2017; De Roover et al., 2014; Gvaladze, De Roover, Tuerlinckx, & Ceulemans, 2020), and multigroup Bayesian structural equation modeling (Muthén & Asparouhov, 2013)—within each cluster of the MMG-FA solution.

Finally, MMG-FA with cluster-specific numbers of factors fell beyond the scope of this article, but it is certainly an interesting extension to consider in the future. For instance, in cross-cultural research, differences in the number of factors are likely to occur when a latent variable is more differentiated in one culture than in another (Chen, 2008). For example, the concept of individuation is two-dimensional in China, whereas it is unidimensional in the United States (Kwan, Bond, Boucher, Maslach, & Gan, 2002). Also, for some groups, an additional factor may occur due to response styles (Billiet & McClendon, 2000; Watson, 1992). It is not a straightforward extension, however, mainly because it requires tackling a much more extensive model selection problem since the best number of factors needs to be selected for each cluster (see, e.g., De Roover, Ceulemans, Timmerman, Nezlek, & Onghena, 2013).

To conclude, the proposed MMG-FA approach is an important step in building a very promising framework of mixture-based methods for unraveling measurement noninvariances across many groups. Its added value lies in the fact that between-group differences and similarities in MM parameters are captured by means of a clustering of the groups and for a specific level of measurement invariance. The latter allows the user to examine MM differences that are relevant to the required level of invariance for the research question at hand. In combination with existing methods, MMG-FA opens up a realm of possibilities to identify noninvariances and figure out which parts of the data (i.e., items, groups) are comparable with respect to the latent variables of interest.

References

Agresti, A. (2013). *Categorical data analysis*. Hoboken, NJ: Wiley.

Akaike, H. (1973). Information theory and an extension of the maximum likelihood principle. In B. N. Petrov & F. Caski (Eds.), *Proceedings of the Second International Symposium on Information Theory* (pp. 267–281). Budapest, Hungary: Akademiai Kiado.

Anderson, J. C., & Gerbing, D. W. (1982). Some methods for respecifying measurement models to obtain unidimensional construct measurement. *Journal of Marketing Research, 19,* 453–460. http://dx.doi.org/10.1177/002224378201900407

Asparouhov, T., & Muthén, B. (2009). Exploratory structural equation modeling. *Structural Equation Modeling, 16,* 397–438. http://dx.doi.org/10.1080/10705510903008204

Bastian, B., Kuppens, P., De Roover, K., & Diener, E. (2014). Is valuing positive emotion associated with life satisfaction? *Emotion, 14,* 639–645. http://dx.doi.org/10.1037/a0036466

Bauer, D. J. (2007). Observations on the use of growth mixture models in psychological research. *Multivariate Behavioral Research, 42,* 757–786. http://dx.doi.org/10.1080/00273170701710338

Bedford, O. A. (2004). The individual experience of guilt and shame in Chinese culture. *Culture and Psychology, 10,* 29–52. http://dx.doi.org/10.1177/1354067X04040929

Billiet, J. B., & McClendon, M. J. (2000). Modeling acquiescence in measurement models for two balanced sets of items. *Structural Equation Modeling, 7,* 608–628. http://dx.doi.org/10.1207/S15328007SEM0704_5

Bollen, K. A., Kirby, J. B., Curran, P. J., Paxton, P. M., & Chen, F. (2007). Latent variable models under misspecification: Two-stage least squares (2SLS) and maximum likelihood (ML) estimators. *Sociological Methods & Research, 36,* 48–86. http://dx.doi.org/10.1177/0049124107301947

Browne, M. W. (2001). An overview of analytic rotation in exploratory factor analysis. *Multivariate Behavioral Research, 36,* 111–150. http://dx.doi.org/10.1207/S15327906MBR3601_05

Bulteel, K., Wilderjans, T. F., Tuerlinckx, F., & Ceulemans, E. (2013). CHull as an alternative to AIC and BIC in the context of mixtures of factor analyzers. *Behavior Research Methods, 45,* 782–791. http://dx.doi.org/10.3758/s13428-012-0293-y

Byrne, B. M., Baron, P., & Balev, J. (1998). The Beck Depression Inventory: A cross-validated test of second-order factorial structure for Bulgarian adolescents. *Educational and Psychological Measurement, 58,* 241–251. http://dx.doi.org/10.1177/0013164498058002007

Byrne, B. M., & Shavelson, R. J. (1986). On the structure of adolescent self-concept. *Journal of Educational Psychology, 78,* 474–481. http://dx.doi.org/10.1037/0022-0663.78.6.474

Byrne, B. M., Shavelson, R. J., & Muthén, B. (1989). Testing for the equivalence of factor covariance and mean structures: The issue of partial measurement equivalence. *Psychological Bulletin, 105,* 456–466. http://dx.doi.org/10.1037/0033-2909.105.3.456

Byrne, B. M., & Van de Vijver, F. J. (2010). Testing for measurement and structural equivalence in large-scale cross-cultural studies: Addressing

This document is copyrighted by the American Psychological Association or one of its allied publishers.
This article is intended solely for the personal use of the individual user and is not to be disseminated broadly.

the issue of nonequivalence. *International Journal of Testing, 10,* 107–132. http://dx.doi.org/10.1080/15305051003637306

Cattell, R. B. (1966). The scree test for the number of factors. *Multivariate Behavioral Research, 1,* 245–276. http://dx.doi.org/10.1207/s15327906mbr0102_10

Ceulemans, E., & Kiers, H. A. (2006). Selecting among three-mode principal component models of different types and complexities: A numerical convex hull based method. *British Journal of Mathematical & Statistical Psychology, 59,* 133–150. http://dx.doi.org/10.1348/000711005X64817

Ceulemans, E., & Van Mechelen, I. (2005). Hierarchical classes models for three-way three-mode binary data: Interrelations and model selection. *Psychometrika, 70,* 461–480. http://dx.doi.org/10.1007/s11336-003-1067-3

Chen, F. F. (2007). Sensitivity of goodness of fit indexes to lack of measurement invariance. *Structural Equation Modeling, 14,* 464–504. http://dx.doi.org/10.1080/10705510701301834

Chen, F. F. (2008). What happens if we compare chopsticks with forks? The impact of making inappropriate comparisons in cross-cultural research. *Journal of Personality and Social Psychology, 95,* 1005–1018. http://dx.doi.org/10.1037/a0013193

Cheung, G. W., & Rensvold, R. B. (2002). Evaluating goodness-of-fit indexes for testing measurement invariance. *Structural Equation Modeling, 9,* 233–255. http://dx.doi.org/10.1207/S15328007SEM0902_5

Christopher, M. S., Charoensuk, S., Gilbert, B. D., Neary, T. J., & Pearce, K. L. (2009). Mindfulness in Thailand and the United States: A case of apples versus oranges? *Journal of Clinical Psychology, 65,* 590–612. http://dx.doi.org/10.1002/jclp.20580

Cole, P. M., Bruschi, C. J., & Tamang, B. L. (2002). Cultural differences in children's emotional reactions to difficult situations. *Child Development, 73,* 983–996. http://dx.doi.org/10.1111/1467-8624.00451

Cooke, D. J., Kosson, D. S., & Michie, C. (2001). Psychopathy and ethnicity: Structural, item, and test generalizability of the Psychopathy Checklist—Revised (PCL-R) in Caucasian and African American Participants. *Psychological Assessment, 13,* 531–542. http://dx.doi.org/10.1037/1040-3590.13.4.531

De Ayala, R. J. (2013). *The theory and practice of item response theory*. New York, NY: Guilford Press Publications.

De Leersnyder, J., Mesquita, B., & Kim, H. S. (2011). Where do my emotions belong? A study of immigrants' emotional acculturation. *Personality and Social Psychology Bulletin, 37,* 451–463. http://dx.doi.org/10.1177/0146167211399103

De Roover, K., Ceulemans, E., Timmerman, M. E., Nezlek, J. B., & Onghena, P. (2013). Modeling differences in the dimensionality of multiblock data by means of clusterwise simultaneous component analysis. *Psychometrika, 78,* 648–668. http://dx.doi.org/10.1007/s11336-013-9318-4

De Roover, K., Ceulemans, E., Timmerman, M. E., & Onghena, P. (2013). A clusterwise simultaneous component method for capturing within-cluster differences in component variances and correlations. *British Journal of Mathematical & Statistical Psychology, 66,* 81–102. http://dx.doi.org/10.1111/j.2044-8317.2012.02040.x

De Roover, K., Ceulemans, E., Timmerman, M. E., Vansteelandt, K., Stouten, J., & Onghena, P. (2012). Clusterwise simultaneous component analysis for analyzing structural differences in multivariate multiblock data. *Psychological Methods, 17,* 100–119. http://dx.doi.org/10.1037/a0025385

De Roover, K., Timmerman, M. E., & Ceulemans, E. (2017). How to detect which variables are causing differences in component structure among different groups. *Behavior Research Methods, 49,* 216–229. http://dx.doi.org/10.3758/s13428-015-0687-8

De Roover, K., Timmerman, M. E., De Leersnyder, J., Mesquita, B., & Ceulemans, E. (2014). What's hampering measurement invariance: Detecting non-invariant items using clusterwise simultaneous component analysis. *Frontiers in Psychology, 5,* 604. http://dx.doi.org/10.3389/fpsyg.2014.00604

De Roover, K., & Vermunt, J. K. (2019). On the exploratory road to unraveling factor loading non-invariance: A new multigroup rotation approach. *Structural Equation Modeling, 26,* 1–19. http://dx.doi.org/10.1080/10705511.2019.1590778

De Roover, K., Vermunt, J., & Ceulemans, E. (2019, December 19). *Mixture multigroup factor analysis for unraveling factor loading non-invariance across many groups*. http://dx.doi.org/10.31234/osf.io/7fdwv

De Roover, K., Vermunt, J. K., Timmerman, M. E., & Ceulemans, E. (2017). Mixture simultaneous factor analysis for capturing differences in latent variables between higher level units of multilevel data. *Structural Equation Modeling, 24,* 506–523. http://dx.doi.org/10.1080/10705511.2017.1278604

Diener, E., Kim-Prieto, C., Scollon, C., and Associates. (2001). [International College Survey 2001]. Unpublished raw data.

Dolan, C. V. (1994). Factor analysis of variables with 2, 3, 5 and 7 response categories: A comparison of categorical variable estimators using simulated data. *British Journal of Mathematical & Statistical Psychology, 47,* 309–326. http://dx.doi.org/10.1111/j.2044-8317.1994.tb01039.x

Dolan, C. V., Oort, F. J., Stoel, R. D., & Wicherts, J. M. (2009). Testing measurement invariance in the target rotated multigroup exploratory factor model. *Structural Equation Modeling, 16,* 295–314. http://dx.doi.org/10.1080/10705510902751416

Eid, M., & Diener, E. (2001). Comparing norms for affect across cultures: Inter-and intranational differences. *Journal of Personality and Social Psychology, 81,* 869–885. http://dx.doi.org/10.1037/0022-3514.81.5.869

Fabrigar, L. R., Wegener, D. T., MacCallum, R. C., & Strahan, E. J. (1999). Evaluating the use of exploratory factor analysis in psychological research. *Psychological Methods, 4,* 272–299. http://dx.doi.org/10.1037/1082-989X.4.3.272

Frese, M. (2015). Cultural practices, norms, and values. *Journal of Cross-Cultural Psychology, 46,* 1327–1330. http://dx.doi.org/10.1177/0022022115600267

George, D., & Mallery, M. (2010). SPSS for Windows step by step: A simple guide and reference, 17.0 update (10a ed.). Boston, MA: Pearson.

Gerbing, D. W., & Hamilton, J. G. (1996). Viability of exploratory factor analysis as a precursor to confirmatory factor analysis. *Structural Equation Modeling, 3,* 62–72. http://dx.doi.org/10.1080/10705519609540030

Gorsuch, R. L. (1983). *Factor analysis* (2nd ed.). Hillside, NJ: Lawrence Erlbaum Associates.

Greiff, S., & Scherer, R. (2018). Still comparing apples with oranges? Some thoughts on the principles and practices of measurement invariance testing. *European Journal of Psychological Assessment, 34,* 141–144. http://dx.doi.org/10.1027/1015-5759/a000487

Gvaladze, S., De Roover, K., Tuerlinckx, F., & Ceulemans, E. (2020). Detecting which variables alter component interpretation across multiple groups: A resampling-based method. *Behavior Research Methods, 52,* 236–263.

Hancock, G. R., Lawrence, F. R., & Nevitt, J. (2000). Type I error and power of latent mean methods and MANOVA in factorially invariant and noninvariant latent variable systems. *Structural Equation Modeling, 7,* 534–556. http://dx.doi.org/10.1207/S15328007SEM0704_2

House, R. J., Hanges, P. J., Javidan, M., Dorfman, P., & Gupta, V. (2004). *Culture, leadership, and organizations: The GLOBE Study of 62 societies*. Thousand Oaks, CA: Sage.

Hu, L., & Bentler, P. M. (1999). Cutoff criteria for fit indexes in covariance structure analysis: Conventional criteria versus new alternatives. *Structural Equation Modeling, 6,* 1–55. http://dx.doi.org/10.1080/10705519909540118

Hubert, L., & Arabie, P. (1985). Comparing partitions. *Journal of Classification, 2,* 193–218. http://dx.doi.org/10.1007/BF01908075

This document is copyrighted by the American Psychological Association or one of its allied publishers.
This article is intended solely for the personal use of the individual user and is not to be disseminated broadly.

Jöreskog, K. G. (1969). Efficient estimation in image factor analysis. *Psychometrika, 34,* 51–75. http://dx.doi.org/10.1007/BF02290173

Jöreskog, K. G. (1971). Simultaneous factor analysis in several populations. *Psychometrika, 36,* 409–426. http://dx.doi.org/10.1007/BF02291366

Kim, E. S., Cao, C., Wang, Y., & Nguyen, D. T. (2017). Measurement invariance testing with many groups: A comparison of five approaches. *Structural Equation Modeling, 24,* 524–544. http://dx.doi.org/10.1080/10705511.2017.1304822

Kim, E. S., Joo, S. H., Lee, P., Wang, Y., & Stark, S. (2016). Measurement invariance testing across between-level latent classes using multilevel factor mixture modeling. *Structural Equation Modeling, 23,* 870–887. http://dx.doi.org/10.1080/10705511.2016.1196108

Kuppens, P., Ceulemans, E., Timmerman, M. E., Diener, E., & Kim-Prieto, C. H. U. (2006). Universal intracultural and intercultural dimensions of the recalled frequency of emotional experience. *Journal of Cross-Cultural Psychology, 37,* 491–515. http://dx.doi.org/10.1177/0022022106290474

Kwan, V. S., Bond, M. H., Boucher, H. C., Maslach, C., & Gan, Y. (2002). The construct of individuation: More complex in collectivist than in individualist cultures. *Personality and Social Psychology Bulletin, 28,* 300–310. http://dx.doi.org/10.1177/0146167202286002

Lawley, D. N., & Maxwell, A. E. (1962). Factor analysis as a statistical method. *The Statistician, 12,* 209–229. http://dx.doi.org/10.2307/2986915

Lei, M., & Lomax, R. G. (2005). The effect of varying degrees of nonnormality in structural equation modeling. *Structural Equation Modeling, 12,* 1–27. http://dx.doi.org/10.1207/s15328007sem1201_1

Lubke, G. H., & Muthén, B. (2005). Investigating population heterogeneity with factor mixture models. *Psychological Methods, 10,* 21–39. http://dx.doi.org/10.1037/1082-989X.10.1.21

Lubke, G., & Muthén, B. O. (2007). Performance of factor mixture models as a function of model size, covariate effects, and class-specific parameters. *Structural Equation Modeling, 14,* 26–47. http://dx.doi.org/10.1080/10705510709336735

Lukočienė, O., Varriale, R., & Vermunt, J. K. (2010). The simultaneous decision(s) about the number of lower-and higher-level classes in multilevel latent class analysis. *Sociological Methodology, 40,* 247–283. http://dx.doi.org/10.1111/j.1467-9531.2010.01231.x

Lynn, R., & Martin, T. (1997). Gender differences in extraversion, neuroticism, and psychoticism in 37 nations. *The Journal of Social Psychology, 137,* 369–373. http://dx.doi.org/10.1080/00224549709595447

MacCallum, R. C., Roznowski, M., & Necowitz, L. B. (1992). Model modifications in covariance structure analysis: The problem of capitalization on chance. *Psychological Bulletin, 111,* 490–504. http://dx.doi.org/10.1037/0033-2909.111.3.490

MacCallum, R. C., Widaman, K. F., Zhang, S., & Hong, S. (1999). Sample size in factor analysis. *Psychological Methods, 4,* 84–99. http://dx.doi.org/10.1037/1082-989X.4.1.84

McLachlan, G., & Krishnan, T. (2007). *The EM algorithm and extensions* (Vol. 382). New York, NY: Wiley.

McLachlan, G., & Peel, D. (2000). *Finite mixture models*. New York, NY: Wiley. http://dx.doi.org/10.1002/0471721182

Marsh, H. W., Hau, K. T., Artelt, C., Baumert, J., & Peschar, J. L. (2006). OECD's brief self-report measure of educational psychology's most useful affective constructs: Cross-cultural, psychometric comparisons across 25 countries. *International Journal of Testing, 6,* 311–360. http://dx.doi.org/10.1207/s15327574ijt0604_1

Marsh, H. W., Morin, A. J., Parker, P. D., & Kaur, G. (2014). Exploratory structural equation modeling: An integration of the best features of exploratory and confirmatory factor analysis. *Annual Review of Clinical Psychology, 10,* 85–110. http://dx.doi.org/10.1146/annurev-clinpsy-032813-153700

McCrae, R. R., Zonderman, A. B., Costa, P. T., Jr., Bond, M. H., & Paunonen, S. V. (1996). Evaluating replicability of factors in the Revised NEO Personality Inventory: Confirmatory factor analysis versus Procrustes rotation. *Journal of Personality and Social Psychology, 70,* 552–566. http://dx.doi.org/10.1037/0022-3514.70.3.552

McNeish, D., & Harring, J. R. (2017). The effect of model misspecification on growth mixture model class enumeration. *Journal of Classification, 34,* 223–248. http://dx.doi.org/10.1007/s00357-017-9233-y

Meredith, W. (1993). Measurement invariance, factor analysis and factorial invariance. *Psychometrika, 58,* 525–543. http://dx.doi.org/10.1007/BF02294825

Meredith, W., & Teresi, J. A. (2006). An essay on measurement and factorial invariance. *Medical Care, 44,* S69–S77. http://dx.doi.org/10.1097/01.mlr.0000245438.73837.89

Mîndrilă, D. (2010). Maximum likelihood (ML) and diagonally weighted least squares (DWLS) estimation procedures: A comparison of estimation bias with ordinal and multivariate non-normal data. *International Journal of Digital Society, 1,* 60–66.

Moore, C. C., Romney, A. K., Hsia, T. L., & Rusch, C. D. (1999). The universality of the semantic structure of emotion terms: Methods for the study of inter-and intra-cultural variability. *American Anthropologist, 101,* 529–546. http://dx.doi.org/10.1525/aa.1999.101.3.529

Mosquera, P. M. R., Manstead, A. S., & Fischer, A. H. (2000). The role of honor-related values in the elicitation, experience, and communication of pride, shame, and anger: Spain and the Netherlands compared. *Personality and Social Psychology Bulletin, 26,* 833–844. http://dx.doi.org/10.1177/0146167200269008

Muthén, B. O. (1993). Goodness of fit with categorical and other nonnormal variables. In K. A. Bollen & J. S. Long (Eds.), *Testing structural equation models* (pp. 205–243). Newbury Park, CA: Sage.

Muthén, B., & Asparouhov, T. (2012). Bayesian structural equation modeling: A more flexible representation of substantive theory. *Psychological Methods, 17,* 313–335. http://dx.doi.org/10.1037/a0026802

Muthén, B. O., & Asparouhov, T. (2013). BSEM measurement invariance analysis. *Mplus web notes, 17,* 1–48.

Muthén, L. K., & Muthén, B. O. (2005). *Mplus: Statistical analysis with latent variables: User's guide*. Los Angeles, CA: Author.

Nylund, K. L., Asparouhov, T., & Muthén, B. O. (2007). Deciding on the number of classes in latent class analysis and growth mixture modeling: A Monte Carlo simulation study. *Structural Equation Modeling, 14,* 535–569. http://dx.doi.org/10.1080/10705510701575396

Rosseel, Y. (2012). Lavaan: An R package for structural equation modeling. *Journal of Statistical Software, 48,* 1–36. http://dx.doi.org/10.18637/jss.v048.i02

Rutkowski, L., & Svetina, D. (2014). Assessing the hypothesis of measurement invariance in the context of large-scale international surveys. *Educational and Psychological Measurement, 74,* 31–57. http://dx.doi.org/10.1177/0013164413498257

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 variables analysis: Applications for developmental research* (pp. 399–419). Thousand Oaks, CA: Sage.

Satorra, A., & Bentler, P. M. (2001). A scaled difference chi-square test statistic for moment structure analysis. *Psychometrika, 66,* 507–514. http://dx.doi.org/10.1007/BF02296192

Schwarz, G. (1978). Estimating the dimension of a model. *Annals of Statistics, 6,* 461–464. http://dx.doi.org/10.1214/aos/1176344136

Silvia, E. S. M., & MacCallum, R. C. (1988). Some factors affecting the success of specification searches in covariance structure modeling. *Multivariate Behavioral Research, 23,* 297–326. http://dx.doi.org/10.1207/s15327906mbr2303_2

Sörbom, D. (1974). A general method for studying differences in factor means and factor structure between groups. *British Journal of Mathe-*

This document is copyrighted by the American Psychological Association or one of its allied publishers.
This article is intended solely for the personal use of the individual user and is not to be disseminated broadly.

matical & Statistical Psychology, 27, 229–239. http://dx.doi.org/10.1111/j.2044-8317.1974.tb00543.x

Sörbom, D. (1989). Model modification. *Psychometrika, 54,* 371–384. http://dx.doi.org/10.1007/BF02294623

Stark, S., Chernyshenko, O. S., & Drasgow, F. (2006). Detecting differential item functioning with confirmatory factor analysis and item response theory: Toward a unified strategy. *Journal of Applied Psychology, 91,* 1292–1306. http://dx.doi.org/10.1037/0021-9010.91.6.1292

Steinley, D. (2004). Properties of the Hubert-Arabie adjusted Rand index. *Psychological Methods, 9,* 386–396. http://dx.doi.org/10.1037/1082-989X.9.3.386

Stevens, J. (1992). *Applied multivariate statistics for the social sciences*. Hillsdale, NJ: Erlbaum.

Tay, L., Diener, E., Drasgow, F., & Vermunt, J. K. (2011). Multilevel mixed-measurement IRT analysis: An explication and application to self-reported emotions across the world. *Organizational Research Methods, 14,* 177–207. http://dx.doi.org/10.1177/1094428110372674

Tein, J. Y., Coxe, S., & Cham, H. (2013). Statistical power to detect the correct number of classes in latent profile analysis. *Structural Equation Modeling, 20,* 640–657. http://dx.doi.org/10.1080/10705511.2013.824781

Tipping, M. E., & Bishop, C. M. (1999). Probabilistic principal components. *Journal of the Royal Statistical Society Series B. Methodological, 61,* 611–622. http://dx.doi.org/10.1111/1467-9868.00196

Triandis, H. C. (1989). The self and social behavior in differing cultural contexts. *Psychological Review, 96,* 506–520. http://dx.doi.org/10.1037/0033-295X.96.3.506

Tucker, L. R. (1951). *A method for synthesis of factor analysis studies* (Personnel Research Section Report No. 984). Washington, DC: Department of the Army.

Tueller, S. J., Drotar, S., & Lubke, G. H. (2011). Addressing the problem of switched class labels in latent variable mixture model simulation studies. *Structural Equation Modeling, 18,* 110–131. http://dx.doi.org/10.1080/10705511.2011.534695

Vandenberg, R. J., & Lance, C. E. (2000). A review and synthesis of the measurement invariance literature: Suggestions, practices, and recommendations for organizational research. *Organizational Research Methods, 3,* 4–70. http://dx.doi.org/10.1177/109442810031002

Varriale, R., & Vermunt, J. K. (2012). Multilevel mixture factor models. *Multivariate Behavioral Research, 47,* 247–275. http://dx.doi.org/10.1080/00273171.2012.658337

Vermunt, J. K. (2010). Latent class modeling with covariates: Two improved three-step approaches. *Political Analysis, 18,* 450–469. http://dx.doi.org/10.1093/pan/mpq025

Vermunt, J. K., & Magidson, J. (2013). *Technical Guide for Latent GOLD 5.0: Basic, advanced, and syntax*. Belmont, MA: Statistical Innovations Inc.

Vermunt, J. K., & Magidson, J. (2016). *Upgrade manual for Latent GOLD 5.1*. Belmont, MA: Statistical Innovations Inc.

Watson, D. (1992). Correcting for acquiescent response bias in the absence of a balanced scale. *Sociological Methods & Research, 21,* 52–88. http://dx.doi.org/10.1177/0049124192021001003

Wierzbicka, A. (1999). *Emotions across languages and cultures: Diversity and universals*. New York, NY: Cambridge University Press. http://dx.doi.org/10.1017/CBO9780511521256

Wilderjans, T. F., Ceulemans, E., & Meers, K. (2013). CHull: A generic convex-hull-based model selection method. *Behavior Research Methods, 45,* 1–15. http://dx.doi.org/10.3758/s13428-012-0238-5

Zhou, H., & Lange, K. L. (2010). On the bumpy road to the dominant mode. *Scandinavian Journal of Statistics, 37,* 612–631. http://dx.doi.org/10.1111/j.1467-9469.2009.00681.x

This document is copyrighted by the American Psychological Association or one of its allied publishers.
This article is intended solely for the personal use of the individual user and is not to be disseminated broadly.

Appendix A

ECM Algorithm and Multistart Procedure MMG-FA

As in all mixture models, log$L$ (Equation 4)—also referred to as the "observed-data loglikelihood"—is complicated by the latent clustering of the groups, making it hard to maximize log$L$ directly. Therefore, the ECM algorithm makes use of the so-called "complete-data loglikelihood," that is, the likelihood when the cluster memberships z_{gk} of all groups as well as the factor scores $\boldsymbol{\eta}_{n_g k}$ are assumed to be known (i.e., the joint distribution of the observed and latent data):

$$
\begin{aligned}
\log L_c &= \prod_{k=1}^{K}\prod_{g=1}^{G}\prod_{n_g=1}^{N_g}\left(f(z_{gk})f(\boldsymbol{\eta}_{n_g k}\,|\,z_{gk})f(\mathbf{x}_{n_g}\,|\,\boldsymbol{\eta}_{n_g k}, z_{gk})\right)\\
&= \log\left[\prod_{k=1}^{K}\prod_{g=1}^{G}\left(\pi_k\prod_{n_g=1}^{N_g}\left(MVN(\boldsymbol{\eta}_{n_g k};\,\mathbf{0},\boldsymbol{\Phi}_{gk})MVN(\mathbf{x}_{n_g};\,\boldsymbol{\mu}_g+\boldsymbol{\Lambda}_k\boldsymbol{\eta}_{n_g k},\boldsymbol{\Psi}_g)\right)\right)^{z_{gk}}\right]\\
&= \sum_{g=1}^{G}\sum_{k=1}^{K} z_{gk}\left(\begin{array}{l}\log(\pi_k)+\sum_{n_g=1}^{N_g}\log\left(\frac{1}{(2\pi)^{Q/2}\,|\,\boldsymbol{\Phi}_{gk}\,|^{1/2}}\exp\left(-\frac{1}{2}\boldsymbol{\eta}'_{n_g k}\boldsymbol{\Phi}_{gk}^{-1}\boldsymbol{\eta}_{n_g k}\right)\right)\\ +\sum_{n_g=1}^{N_g}\log\left(\frac{1}{(2\pi)^{J/2}\,|\,\boldsymbol{\Psi}_g\,|^{1/2}}\exp\left(-\frac{1}{2}(\tilde{\mathbf{x}}_{n_g}-\boldsymbol{\Lambda}_k\boldsymbol{\eta}_{n_g k})'\boldsymbol{\Psi}_g^{-1}(\tilde{\mathbf{x}}_{n_g}-\boldsymbol{\Lambda}_k\boldsymbol{\eta}_{n_g k})\right)\right)\end{array}\right)\\
&= \sum_{g=1}^{G}\sum_{k=1}^{K} z_{gk}\log(\pi_k) - \frac{N(J+Q)}{2}\log(2\pi) - \frac{1}{2}\sum_{g=1}^{G}\sum_{k=1}^{K} z_{gk}N_g\log(\,|\,\boldsymbol{\Phi}_{gk}\,|\,)\\
&\quad -\frac{1}{2}\sum_{g=1}^{G} N_g\log(\,|\,\boldsymbol{\Psi}_g\,|\,) - \frac{1}{2}\sum_{g=1}^{G}\sum_{k=1}^{K} z_{gk}\left(\sum_{n_g=1}^{N_g}\left(\begin{array}{l}tr\left(\boldsymbol{\eta}_{n_g k}\boldsymbol{\eta}'_{n_g k}\boldsymbol{\Phi}_{gk}^{-1}\right)+tr\left(\tilde{\mathbf{x}}_{n_g}\tilde{\mathbf{x}}'_{n_g}\boldsymbol{\Psi}_g^{-1}\right)\\ -2\tilde{\mathbf{x}}'_{n_g}\boldsymbol{\Psi}_g^{-1}\boldsymbol{\Lambda}_k\boldsymbol{\eta}_{n_g k}+tr\left(\boldsymbol{\Lambda}'_k\boldsymbol{\Psi}_g^{-1}\boldsymbol{\Lambda}_k\boldsymbol{\eta}_{n_g k}\boldsymbol{\eta}'_{n_g k}\right)\end{array}\right)\right)
\end{aligned}
\tag{7}
$$

where $\tilde{\mathbf{x}}_{n_g} = \mathbf{x}_{n_g} - \boldsymbol{\mu}_g$. Because the latent data are unknown, the following expected values of z_{gk}, $\boldsymbol{\eta}_{n_g k}$ and $\boldsymbol{\eta}_{n_g k}\boldsymbol{\eta}'_{n_g k}$ (McLachlan & Krishnan, 2007) are inserted in Equation 7:

$$
E[z_{gk}=1\,|\,\mathbf{X}_g] = \frac{\pi_k f_{gk}(\mathbf{X}_g;\,\boldsymbol{\theta}_{gk})}{\sum_{k'=1}^{K}\pi_{k'}f_{gk'}(\mathbf{X}_g;\,\boldsymbol{\theta}_{gk'})} \tag{8}
$$

$$
E[\boldsymbol{\eta}_{n_g k}\,|\,\mathbf{x}_{n_g}, z_{gk}=1] = \boldsymbol{\beta}_{gk}\tilde{\mathbf{x}}_{n_g} \quad with \quad \boldsymbol{\beta}_{gk} = \boldsymbol{\Phi}_{gk}\boldsymbol{\Lambda}'_k(\boldsymbol{\Lambda}_k\boldsymbol{\Phi}_{gk}\boldsymbol{\Lambda}'_k+\boldsymbol{\Psi}_g)^{-1} \tag{9}
$$

$$
E[\boldsymbol{\eta}_{n_g k}\boldsymbol{\eta}'_{n_g k}\,|\,\mathbf{x}_{n_g}, z_{gk}=1] = \boldsymbol{\Phi}_{gk} - \boldsymbol{\beta}_{gk}\boldsymbol{\Lambda}_k\boldsymbol{\Phi}_{gk} + \boldsymbol{\beta}_{gk}\tilde{\mathbf{x}}_{n_g}\tilde{\mathbf{x}}'_{n_g}\boldsymbol{\beta}'_{gk} \tag{10}
$$

where Equation 8 corresponds to the posterior classification probability $\hat{z}_{gk}$. In this way, the following expected value of logL_c or E[logL_c] is obtained:

$$
\begin{aligned}
E[\log L_c] &= \sum_{k=1}^{K} G_k\log(\pi_k) - \frac{N(J+Q)}{2}\log(2\pi) - \frac{1}{2}\sum_{g=1}^{G}\sum_{k=1}^{K} N_{gk}\log(\,|\,\boldsymbol{\Phi}_{gk}\,|\,)\\
&\quad -\frac{1}{2}\sum_{g=1}^{G} N_g\log(\,|\,\boldsymbol{\Psi}_g\,|\,) - \frac{1}{2}\sum_{k=1}^{K}\sum_{g=1}^{G} N_{gk}tr(\boldsymbol{\Theta}_{gk}\boldsymbol{\Phi}_{gk}^{-1}) - \frac{1}{2}\sum_{g=1}^{G} N_g tr(\mathbf{S}_g\boldsymbol{\Psi}_g^{-1})\\
&\quad +\sum_{k=1}^{K}\sum_{g=1}^{G} N_{gk}tr(\boldsymbol{\Psi}_g^{-1}\boldsymbol{\Lambda}_k\boldsymbol{\beta}_{gk}\mathbf{S}_g) - \frac{1}{2}\sum_{k=1}^{K}\sum_{g=1}^{G} N_{gk}tr(\boldsymbol{\Lambda}'_k\boldsymbol{\Psi}_g^{-1}\boldsymbol{\Lambda}_k\boldsymbol{\Theta}_{gk})
\end{aligned}
\tag{11}
$$

(*Appendices continue*)

This document is copyrighted by the American Psychological Association or one of its allied publishers.
This article is intended solely for the personal use of the individual user and is not to be disseminated broadly.

where $G_k = \sum_{g=1}^{G} \hat{z}_{gk}$, $N_{gk} = \hat{z}_{gk} N_g$, $\mathbf{S}_g = \frac{1}{N_g} \sum_{n_g=1}^{N_g} \tilde{\mathbf{x}}_{n_g} \tilde{\mathbf{x}}'_{n_g}$, and $\mathbf{\Theta}_{gk} = \mathbf{\Phi}_{gk} - \boldsymbol{\beta}_{gk} \mathbf{\Lambda}_k \mathbf{\Phi}_{gk} + \boldsymbol{\beta}_{gk} \mathbf{S}_g \boldsymbol{\beta}'_{gk}$. By taking the derivative of $E[\log L_c]$ and equating it to zero—using the Lagrange multiplier method for imposing $\sum_{k=1}^{K} \hat{\pi}_k = 1$—we obtain the following parameter updates:

$$\hat{\pi}_k = \frac{1}{G} \sum_{g=1}^{G} \hat{z}_{gk} \tag{12}$$

$$\hat{\boldsymbol{\lambda}}_{kj} = \left(\sum_{g=1}^{G} \frac{N_{gk}}{\hat{\psi}_{gj}} \left(\mathbf{S}_g \hat{\boldsymbol{\beta}}'_{gk} \right)_{j.} \right) \left(\sum_{g=1}^{G} \frac{N_{gk}}{\hat{\psi}_{gj}} \hat{\mathbf{\Theta}}_{gk} \right)^{-1} \tag{13}$$

$$\hat{\mathbf{\Phi}}_{gk} = \hat{\mathbf{\Theta}}_{gk} \tag{14}$$

$$diag\left(\hat{\mathbf{\Psi}}_g \right) = diag\left(\mathbf{S}_g - \frac{1}{N_g} \left(\sum_{k=1}^{K} N_{gk} (2 \hat{\mathbf{\Lambda}}_k \hat{\boldsymbol{\beta}}_{gk} \mathbf{S}_g - \hat{\mathbf{\Lambda}}_k \hat{\mathbf{\Theta}}_{gk} \hat{\mathbf{\Lambda}}'_k) \right) \right) \tag{15}$$

Note that the factor loadings (Equation 13) are updated per row.

To set the scale of the cluster-specific factors, the mean factor variances are fixed to one over all groups within a cluster k, that is, $\hat{\mathbf{\Phi}}_k = \frac{1}{N_k} \sum_{g=1}^{G} N_g \hat{z}_{gk} \hat{\mathbf{\Phi}}_{gk} = \mathbf{I}$, where $N_k = \sum_{g=1}^{G} N_g \hat{z}_{gk}$ and $\mathbf{I}$ is an identity matrix. Note that this restriction also fixes the factor covariances to zero over all groups within a cluster, which implies that the initial rotation is orthogonal for each cluster. Afterward, one can choose to rotate the cluster-specific factors according to an orthogonal or oblique rotation criterion and counterrotate the group- and cluster-specific factor covariances accordingly. To impose the above-mentioned restriction, the factor (co)variances and loadings are rescaled and rotated as follows:

$$\hat{\mathbf{\Phi}}^*_{gk} = \left(\hat{\mathbf{\Theta}}_k \right)^{-\frac{1}{2}} \hat{\mathbf{\Phi}}_{gk} \left(\hat{\mathbf{\Theta}}_k \right)^{-\frac{1}{2}} \quad with \quad \hat{\mathbf{\Theta}}_k = \left(\sum_{g'=1}^{G} \frac{N_{g'k}}{N_k} \hat{\mathbf{\Theta}}_{g'k} \right) \tag{16}$$

$$\hat{\mathbf{\Lambda}}^*_k = \hat{\mathbf{\Lambda}}_k \left(\hat{\mathbf{\Theta}}_k \right)^{\frac{1}{2}}. \tag{17}$$

Note that this can be done either in the final iteration only (i.e., upon convergence), or in each iteration.

Algorithm

The group-specific means are determined as $\hat{\boldsymbol{\mu}}_g = \frac{1}{N_g} \sum_{n_g=1}^{N_g} \mathbf{x}_{n_g}$ and the algorithm continues with the centered data for each group: $\tilde{\mathbf{x}}_{n_g} = \mathbf{x}_{n_g} - \hat{\boldsymbol{\mu}}_g$.

For a user-specified number of starts, perform the following steps for each start:

1. Start from a preselected random partition (see Multistart Procedure section), that is, with binary values for $\hat{z}_{gk}$.
2. Perform the following initialization of the parameters per cluster k of the random partition, based on probabilistic principal component analysis[5] (Tipping & Bishop, 1999): $\hat{\mathbf{\Lambda}}_k = \mathbf{U}_Q \sqrt{(\mathbf{V}_Q - \hat{\sigma}^2 \mathbf{I}_Q)}$ for $k = 1, \ldots, K$ and $\hat{\mathbf{\Psi}}_g = \hat{\sigma}^2 \mathbf{I}_J$ for $g = 1, \ldots, G$ if $\hat{z}_{gk} = 1$, where the columns of $\mathbf{U}_Q$ correspond to the first Q eigenvectors and the diagonal matrix $\mathbf{V}_Q$ contains the Q largest eigenvalues of the eigenvalue decomposition of $\mathbf{S}_k = \frac{1}{N_k} \sum_{g=1}^{G} \hat{z}_{gk} (\sum_{n_g=1}^{N_g} \tilde{\mathbf{x}}_{n_g} \tilde{\mathbf{x}}'_{n_g})$, $\hat{\sigma}^2$ is the average of the J–Q smallest eigenvalues, and $\mathbf{I}_Q$ and $\mathbf{I}_J$ are $Q \times Q$ and $J \times J$ identity matrices, respectively. The factor (co)variance matrices $\hat{\mathbf{\Phi}}_{gk}$ are initialized as $\mathbf{I}_Q$.
3. Iterate the following steps while δ_1 & $\delta_2 > 1 \times 10^{-4}$ and $v \leq 100$:
   a. Update the iteration number: $v = v + 1$.
   b. Update the posterior classification probabilities $\hat{z}_{gk}$ (Equation 8) for g = 1, . . . , G and $k = 1, \ldots, K$.
   c. Update the mixing proportions $\hat{\pi}_k$ (Equation 12) for $k = 1, \ldots, K$.
   d. Update the factor loadings $\hat{\mathbf{\Lambda}}_k$ for $k = 1, \ldots, K$ (Equation 13) and the factor (co)variance matrices $\hat{\mathbf{\Phi}}_{gk}$ for $g = 1, \ldots, G$ and $k = 1, \ldots, K$ (Equation 14) if $N_k > 0$. Update the unique variances $\hat{\mathbf{\Psi}}_g$ for $g = 1, \ldots, G$ (Equation 15). Rescale $\hat{\mathbf{\Phi}}_{gk}$ and $\hat{\mathbf{\Lambda}}_k$ (Equations 16 and 17) for $g = 1, \ldots, G$ and $k = 1, \ldots, K$. To remedy Heywood cases, fix unique variances to .0001 when they are smaller than this number.
   e. Compute $\log L^v$ value (Equation 4).
   f. Evaluate convergence with respect to $\log L^v$ and the parameter estimates $\hat{\theta}^v_r$: $\delta_1 = \sum_{r=1}^{R} \left| \frac{\hat{\theta}^v_r - \hat{\theta}^{v-1}_r}{\hat{\theta}^{v-1}_r} \right|$ and $\delta_2 = \log L^v - \log L^{v-1}$.
4. After (preliminary) convergence is reached (or 100 iterations), check whether the obtained solution is the best one in terms of $\log L$ (across all starts up to now) and, if so, save the parameter estimates $\hat{\theta}^{best}_r = \hat{\theta}^v_r$ and iteration number $v^{best} = v$.

After performing this procedure for all starts, iterate further until full convergence is reached for the best solution: that is, starting from $\hat{\theta}^v_r = \hat{\theta}^{best}_r$ and $v = v^{best}$, iterate Steps 3a to 3f while δ_1 & $\delta_2 > 1 \times 10^{-6}$ and $v \leq 1{,}000$ (or another user-specified maximal number of iterations).

[5] These starting values are similar to the maximum likelihood estimates of image factor analysis described by Jöreskog (1969).

(*Appendices continue*)

This document is copyrighted by the American Psychological Association or one of its allied publishers.
This article is intended solely for the personal use of the individual user and is not to be disseminated broadly.

Multistart Procedure

Because the ECM algorithm described above is not guaranteed to converge to the global maximum, a multistart procedure is used to increase the probability of finding the global maximum. The multistart procedure applies a tiered testing strategy with respect to several sets of starting values. Specifically, given the user-specified number of starts (e.g., 25), it starts from 10 times as many random partitions of the groups (e.g., 10 × 25 partitions). For each of these partitions, the parameter estimates per cluster k are initialized as described in Step 2 of the algorithm (see Algorithm section). Subsequently, the parameter estimates are updated once by means of Equations 12 to 17 and the log$L$ value (Equation 4) is determined. The 10% most promising partitions (i.e., with the highest log$L$) are selected as the starts for the algorithm described in the previous section.

Appendix B

Latent GOLD 6.0 Syntax for MMG-FA

An example syntax for MMG-FA with three clusters and two factors per cluster for a data set with 12 variables is given and explained below (for more details, see Vermunt & Magidson, 2013):

```
options
    algorithm
        tolerance = 1e-006 emtolerance = 1e-006 emiterations = 1000 nriterations = 0
        emfa;
    startvalues
        seed = 0 sets = 250 tolerance = 1e-004 iterations = 50 PCA annealing;
    bayes
        categorical = 0 variances = 0 latent = 0 poisson = 0;
    missing includeall;
    output
        iterationdetail parameters = first standarderrors probmeans = posterior
        reorderclasses;
    rotation oblimin;
variables
    groupid Country;
    dependent (V1 - V12) continuous;
    independent Country nominal;
    latent
        F1 continuous,
        F2 continuous,
        Cluster nominal Group 3;// 1 - 8 to estimate models with 1 to 8 clusters
equations
// group- and cluster-specific factor (co)variances
    F1 | Country Cluster;
    F2 | Country Cluster;
    F1 <-> F2 | Country Cluster;
// logistic regression model for clusters (only intercept)
    Cluster <- 1;
// regression models for items: group-specific intercepts and
// cluster-specific loadings
    V1 - V12 <- 1 | Country + F1 | Cluster + F2 | Cluster;
// group-specific unique variances
    V1 - V12 | Country;
```

(*Appendices continue*)

This document is copyrighted by the American Psychological Association or one of its allied publishers.
This article is intended solely for the personal use of the individual user and is not to be disseminated broadly.

The Latent GOLD syntax contains three sections, that is, "options," "variables," and "equations." First, the "options" section pertains to specifications regarding the estimation process and to output options. The parameters in the "algorithm" subsection indicate when the algorithm should proceed with Newton-Raphson instead of EM iterations and when convergence is reached. To apply only EM iterations, set "nriterations" to zero, "emiterations" to a high number and "emtolerance" to the same value as "tolerance." The new option "emfa" makes sure that the factor model parameters are estimated by means of the time-efficient EM procedure detailed in Appendix A. The "startvalues" subsection includes the parameters pertaining to the multistart procedure used by Latent GOLD. Specifically, for each set of starting values (the number of sets is specified by "sets"), the model is reestimated for as many iterations as specified by "iterations" or until δ_1 or δ_2 drops below the "tolerance" value. Subsequently, it continues with the 10% (rounded upward) most promising sets (i.e., with the highest log$L$), performing another two times the specified number of iterations (i.e., 2 $\times$ the value of "iterations"). Finally, it continues with the best solution until convergence. In the example syntax above, 250 starts are requested by the user. The thus obtained multistart procedure is actually more elaborate than the one evaluated in the Simulation Studies section, because the clustering and factor parameters are updated 50 times before choosing the 10% most promising sets of starting values (as opposed to one update of the factor parameters only). "PCA" prompts Latent GOLD to use starting values for the factor loadings and unique variances that are based on principal component analysis (PCA; Vermunt & Magidson, 2016). More specifically, PCA is performed on the entire data set and, to get cluster-specific starting values, randomness is added to the PCA solution per cluster k. For more details on the PCA-based starting values and the entire multistart procedure see De Roover et al. (2017). When group sizes are large, the algorithm may be prone to local maxima because the posterior classification probabilities quickly approach one and zero, even for a clustering that is far from the one that is actually underlying the data. This may happen especially when between-cluster loading differences are small. To avoid this, the "annealing" option—referring to "deterministic annealing"—is used which implies that an auxiliary variable is used to keep the posterior classification probabilities more fuzzy for the first few iterations (Zhou & Lange, 2010). It is advised to set the options in the "bayes" subsection to zero when using the "emfa" algorithm. Finally the "rotation" option is used to specify how the MMG-FA parameters should be identified (see below). In the "output" subsection, the desired output can be specified by the user.

Second, the "variables" section specifies the different types of variables included in the model. Because MMG-FA operates on multilevel data, after "groupid," the variable in the data file that indicates the group structure (i.e., the group number for each observation) should be specified, using its label in the data file (e.g., "country"). In the "dependent" subsection, the dependent variables of the model (i.e., the observed variables) are specified, by means of their label in the data file and their measurement scale. Next, the "independent" variables are listed. For MMG-FA, one has to include the grouping variable as an independent variable, since parameters will be allowed to vary across groups. Finally, the "latent" variables of the MMG-FA model are the factors (i.e., "F1" and "F2" in the example syntax) and the mixture model clustering (i.e., "Cluster"). In particular, the former are specified as continuous latent variables, whereas the latter is specified as a nominal latent variable at the group level with a specified number of categories (i.e., the desired number of clusters). For estimating models with, for instance, one to eight clusters, use "1–8."

In the "equations" section, the model equations are listed. First, the factor variances and covariances are specified and they are allowed to differ among groups and clusters by adding "| Country Cluster." Next, a logistic regression model for the categorical latent variable "cluster" is specified, which contains only an intercept term in this case. Then, regression models are defined for the observed variables, that is, which variables are regressed on which factors. Note that all variables are regressed on all factors (i.e., EFA is applied) and that an intercept term is included. To obtain factor loadings that differ between clusters and intercepts that differ between groups, "| Cluster" is added to each regression effect and "| Country" is added to the intercepts. By default, factor means are equal to zero and, since no effect is specified to let them differ between groups or clusters, they are zero for all groups. Finally, unique variances are added, which are allowed to differ across groups. At the end of the syntax, additional restrictions may be specified or starting values for the parameters may be given, either by directly typing them in the syntax or by referring to a text file.

The specified MMG-FA model is not identified without additional constraints, but identification is achieved using the "rotation" option. The scales of the latent factors are fixed by restricting the weighted means of their variances to be equal to 1 across groups within a cluster. The rotational freedom is dealt with by applying the specified rotation method (e.g., oblimin) to the cluster-specific loadings. As described in De Roover and Vermunt (2019), it is possible to use target rotation and to add a weighted agreement term (e.g., "procrustes = 0.5") to the rotation criterion.

Received December 20, 2019
Revision received July 22, 2020
Accepted July 27, 2020 ■

This document is copyrighted by the American Psychological Association or one of its allied publishers.
This article is intended solely for the personal use of the individual user and is not to be disseminated broadly.