Dimensionality and measurement invariance in the Satisfaction with Life Scale in Norway

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Abstract

Purpose Results from previous studies examining the dimensionality and factorial invariance of the Satisfaction with Life Scale (SWLS) are inconsistent and often based on small samples. This study examines the factorial structure and factorial invariance of the SWLS in a Norwegian sample.

Methods Confirmatory factor analysis (AMOS) was conducted to explore dimensionality and test for measurement invariance in factor structure, factor loadings, intercepts, and residual variance across gender and four age groups in a large (N = 4,984), nationally representative sample of Norwegian men and women (15–79 years).

Results The data supported a modified unidimensional structure. Factor loadings could be constrained to equality between the sexes, indicating metric invariance between genders. Further testing indicated invariance also at the strong and strict levels, thus allowing analyses involving group means. The SWLS was shown to be sensitive to age, however, at the strong and strict levels of invariance testing.

Conclusion In conclusion, the results in this Norwegian study seem to confirm that a unidimensional structure is acceptable, but that a modified single-factor model with correlations between error terms of items 4 and 5 is preferred. Additionally, comparisons may be made between the genders. Caution must be exerted when comparing age groups.

Keywords Satisfaction with life · Mental health · Measurement invariance · Gender · Age · Dimensionality · Partial measurement invariance

Abbreviations

AMOS Analysis of moment structures
CFA Confirmatory factor analysis
CFI Comparative fit index
CI Confidence interval
EM Expectation-Maximization
ML Maximum likelihood
PGFI Parsimony goodness of fit index
PRATIO Parsimony ratio
PNFI Parsimony normed fit index
RMSEA Root mean square error of approximation
SE Standard error
SPSS Statistical Package for Social Sciences
SWLS Satisfaction with life score
F1 Simple one-factor model
F2 Two-factor model
F1cov One-factor model with covariance between residuals of items 4 and 5
ADF Asymptotically distribution free

Introduction

Satisfaction with life is one of several aspects of positive mental health. It is not a direct, verifiable experience, nor a known personal fact, but a cognitive product that involves a comparative process between the individual’s current life situation and internalized standards, allowing respondents...
Table 1 Overview of the five items of the Satisfaction with Life Scale—percent response for each item (Norwegian health interview survey 2005; \( N = 4,984 \))

| Item | Question | Strongly disagree | Disagree | Disagree slightly | Neither agree nor disagree | Agree slightly | Agree | Strongly agree |
|------|----------|-------------------|----------|-------------------|---------------------------|----------------|-------|---------------|
| 1    | In most ways my life is close to ideal | 2.1 | 6.2 | 6.6 | 12.6 | 24.1 | 37.4 | 11.0 |
| 2    | The conditions of my life are excellent | 1.4 | 3.2 | 4.1 | 7.8 | 16.7 | 46.5 | 20.2 |
| 3    | I am satisfied with my life | 1.1 | 2.9 | 4.8 | 5.9 | 14.9 | 48.1 | 22.5 |
| 4    | So far, I have gotten the important things I want in life | 1.6 | 4.2 | 6.1 | 9.5 | 22.6 | 40.2 | 15.8 |
| 5    | If I could live my life over, I would change nothing | 4.8 | 10.2 | 12.1 | 12.6 | 21.7 | 28.4 | 10.2 |

to use the information they subjectively deem relevant when evaluating their own lives [1].

The Satisfaction with Life Scale (SWLS) [2, 3] is perhaps the most commonly used measure of life satisfaction worldwide. The scale consists of five statements (Table 1) and was originally developed to circumvent problems inherent in previous scales based on single items, or scales based toward domain or culture-specific items. As people derive their life satisfaction from different sources and vary considerably in their ideas about what constitutes a good life, the SWLS measures people’s perception of their life as a whole, using items that are supposedly free from the varying criteria people use when evaluating their lives. The scale thus reflects a global evaluative judgment, partly determined by the respondent’s current mood and immediate context, and partly by stable personality factors [4, 5] and genetic influences [6].

Although the SWLS is extensively studied and shows good psychometric properties including validity, internal consistency, and test–retest reliability [2, 3, 7, 8], there are still important issues that need to be addressed.

One issue concerns the dimensionality of the scale. Many studies have supported a unidimensional model, attesting a single latent factor accounting for a majority of the variance in life satisfaction scores [2, 9–12]. Some of these studies were based on traditional factor analysis, however, and when there are well founded hypotheses about dimensionality, confirmatory factor analysis (CFA) is a preferred analytical method. Some studies report the fifth item to be more weakly associated with the latent life satisfaction construct than the remaining four items (Table 1). Other studies claim essential, but not strict unidimensionality, as item 5 shows a weaker association with the latent variable than the remaining four items [13–18]. Yet other studies support a modified unidimensional structure [19]. Some studies even suggest that a two-factor structure consisting of strongly correlated “present” (i.e., items 1–3 measure the status at the moment) and “past” (i.e., items 4 and 5 measure the individual to reflect the status over the life sequence) factors should be considered [14, 20]. Most studies involved small, non-random samples, however.

Another issue concerns the invariance of the scale. Measurement invariance indicates that the same underlying construct is measured across the relevant comparison groups. This ensures that group differences can be interpreted in terms of group differences in the underlying construct. Should the assumption of invariance not hold, comparisons across groups may not be valid, the subsequent interpretations may not be meaningful, and the conclusions incorrect.

Findings concerning the invariance of the SWLS are somewhat inconsistent. Some studies have reported the SWLS to be invariant (factor loadings, unique variances, factor variance) across gender [21] and age groups [22–24], whereas other studies have reported sensitivity to either sex [25] or age [26]. These inconsistencies may partly be explained by inadequate sample sizes and/or composition of samples. To explore invariance sufficiently well, respondents should represent the entire adult life span and both genders. Most studies, however, are based on small to moderately sized, e.g., [21, 26, 27] or highly homogenous samples such as Spanish junior high school students [25], Taiwanese [18] and British [21] university students and Swedish student teachers [14] and consequently exhibit both a restricted age range, biased sex ratio, and limited socio-demographic profiles.

This study explores the dimensionality and measurement invariance of the SWLS across gender and age in a large (\( N = 4,984 \)), nationally representative subsample of persons aged 15–79, thus including both male and female participants from emerging to older adulthood. The respondents are Norwegian and along with the other Scandinavian countries, Norwegian SWLS scores generally rank among the highest in the world, perhaps due to the distribution of welfare benefits in these countries. Scandinavian studies may therefore provide insights into differences in SWLS that may relate to benefits associated with the welfare state that attempt to equalize income and social/health benefits over the entire age span.
Method

Sample

The data are from the 2005 wave of a regularly repeated (every 3 years) health investigation in Norway. The cross-sectional investigation is based on a nationally representative subsample of 10,000 persons living at home. The data are selected to be representative based on a stratified selection by municipality of residence. Information was collected through a postal questionnaire (one reminder) that each individual completed and returned through the postal services. Of the 9,187 that received the questionnaire, 5,212 responded (57%). Individuals with 3 or more missing values on SWLS, or missing gender or age were removed prior to analysis, leaving altogether 4,984 respondents. The final sample consisted of 2,369 men (mean age 46.2 years) and 2,615 women (mean age 44.1 years). Sample size by age group can be found in Table 3.

The study was approved by the Regional Committees for Medical Health and Research Ethics, and each participant gave informed consent.

Measures

Satisfaction with life was measured using the five-item Satisfaction with Life Scale (SWLS) [2, 3]. Responses were rated on a Likert scale ranging from 1 (strongly disagree) to 7 (strongly agree) (Table 1). This battery includes the following five questions:

1. In most ways my life is close to ideal
2. The conditions of my life are excellent
3. I am satisfied with my life
4. So far, I have gotten the important things I want in life
5. If I could live my life over, I would change almost nothing"

Statistical methods

All preliminary analyses were performed by the Statistical Package for Social Sciences (SPSS) version 17.0. Factor analysis operations were then conducted using Maximum Likelihood (ML) estimation by means of Analysis of Moment Structures (AMOS 17) [28].

There were 26, 33, 33, 67, and 15 cases with missing values for questions 1–5, respectively. For respondents with two or less missing items, the Expectation Maximization (EM) option in SPSS was used to impute missing values for each SWLS item using the remaining SWLS items. The EM procedure is a process of regression imputation based on the observed relationship between variables. Missing values are replaced iteratively until successful iterations are sufficiently similar, and yield a complete set of data.

The data were handled as continuous data based on observations that 7-point Likert scales are best handled using continuous methodology [29]. To test the validity of handling the data as continuous, the analyses were repeated using Bayesian methodology, which is the preferred method for ordinal data.

To evaluate the dimensional structure, we performed confirmatory factor analyses (CFA) [30] using responses both from the entire sample and from each of the different subgroups. The analyses were run by means of ML estimation. The use of ML estimation can cause problems when using non-normal data, but is considered to be robust when used with moderately non-normal data from large samples [31].

The data were tested for normality and found to be univariate normal (highest kurtosis value was 2.42) [32], but not multivariate normally distributed (multivariate kurtosis was equal to 25.4; Marida’s normalized estimate 73.9) [33]. The analyses were therefore repeated using asymptotic free distribution (ADF) estimation. In addition, the data were normalized using Tukey’s formula, and ML estimation repeated on normalized data. Finally, results of analyses using ML were tested with Bootstrapping, using 2,000 samples, 95% CI, and significance tested with bias corrected confidence intervals.

Due to inconsistencies in the previous literature regarding the factorial structure (dimensionality) of the scale (Table 2), two alternative baseline models were specified. Altogether four models were tested in this study:

1. A simple one-factor model
2. A two-factor model including “past” (last two items) and “present” (first three items)
3. A modified one-factor model allowing the residual terms of items 4 and 5 to be correlated. This model is nested under model 1 and the modification based on modification indices (Fig. 1)
4. A model testing for inter-item correlation when the items were presented consecutively and successively rather than scattered throughout the questionnaire [34]

As the chi² has been shown to be problematic for assessing model fit in large samples [33, 35], model fit was primarily assessed using the root mean square error of approximation (RMSEA) with values of 0.08, 0.05, and 0, and the comparative fit index (CFI), with values 0.90, 0.95, and 1.0 demonstrating reasonable, close, and exact fit, respectively. It is strongly recommended to include
measures of parsimony that control for degrees of freedom, especially when testing complex models [33, 35]. Parsimony was evaluated here using the parsimony ratio (PRATIO), the parsimony goodness of fit index (PGFI) and the parsimony normed fit index (PNFI).

Testing of measurement invariance was conducted by multigroup CFAs using ML estimation in AMOS 17. This method employs successive analyses where constraints to the models are added consecutively. The baseline model is an unconstrained model, with one-factor loading constrained to unity. The weak (metric) model, nested under the baseline model, constrains the factor loadings to be equal across groups. Non-significance at this level allows comparing relationships. The strong (scalar) model also constrains factor loadings and intercepts to equality across comparison groups, thus allowing comparing means, and the strict model additionally constrains the residuals. Invariance at the strict level is very seldom achieved.

The chi^2 alone was not deemed useable in this large sample, but the ΔChi^2 was measured and reported when comparing the model fit in different subgroups. The ΔCFI was considered a more appropriate test, however, and a cut-off ≤0.01 has been suggested when testing for significant differences between subgroups [33, 35].

Partial measurement invariance was examined with a successive removal of constraints at each level of invariance testing based on examination of modification indices [36]. Level 1 removed constraints on factor loadings one by one. Thereafter constraints were removed for the intercepts keeping the factor loading structure achieved in the partial

| Study reference | Sample characteristics | Sample size | Gender | Age |
|-----------------|------------------------|-------------|--------|-----|
| Anaby et al. [9] | Israeli adults         | 487         | 190    | 27–60 |
| Arrindel et al. [10] | Dutch young adults   | 2,800       | 888    | 18–30 |
| Atienza et al. [25] | Spanish junior high students | 2,080 | 1,023  | 1,057|
| Balatsky and Dieri [11] | Soviet students    | 116         |        | 18.9 |
| Blais et al. [22] | French-Canadian students | 871       |        |       |
| Durak et al. [23] | Turkish univ students, correctional officers and elderly adults (3 groups) | 547, 166 and 123 | 20.7, 37.2, 68.2 |
| Lewis et al. [12] | Czech university students | 109       | 38     | 23.0 |
| Oishi [15]      | Chinese and American students | 556 chinese; 442 American |        |       |
| Pons et al. [26] | Spanish junior high students | 266       | 65     | 11–15 |
| Shevlin et al. [21] | Undergraduates        | 258         | 173    | 18–57 |
| Swami and Chamorro-Premuzic [41] | Malay community sample | 816       |        |       |
| Vauhtier et al. [34] | Group 1 Successive item presentation | 494       | 233    | 47.7 |
| Gouveia et al. [13] | Five groups, high school students, teachers undergraduate students, physicians, general population | 2,180 (306–797) | 2,369 | 16–79 |
| Hultell and Gustavsson [14] | Swedish student teachers | 2,900       | 453    | 28.9 |
| Sachs [19]      | Hong Kong University students | 123       | 43     | 32   |
| Slocum-Gori et al. [17] | Canadian (BC) adults  | 410         | 239    | 18–90 |
| Wu and Yao [18] | University students   | 476         | 207    | 269  |

Table 2 Overview of the literature examining dimensionality of the Satisfaction with Life Scale
variance testing at level 1. This method was then repeated at the strict level with progressive removal of constraints on the variances [37]. Tables 5 and 6 indicate which parameters are constrained at each level.

Results

Descriptives

The average response category endorsed by respondents for items 1 through 5 of the SWLS were 5.1, 5.6, 5.6, 5.3, and 4.6, respectively, on a scale of 1–7. Cronbach’s alpha was estimated to be 0.91. This is consistent with values found elsewhere in the literature [16, 38].

Common factoring with principal axis extraction and varimax rotation resulted in 74% of variance explained by a single factor.

Dimensionality

CFAs were then used to compare a one-factor (F1) to a two-factor solution (F2). The two-factor model including “past” (last two items) and “present” (first three items) factors yielded better fit than the unconstrained one-factor model (CFI = 0.995 vs. 0.986; RMSEA = 0.065 vs. 0.094). However, the correlation between the factors was close to unity (r = 0.93), indicating that the two factors could not be easily differentiated. In addition, parsimony was slightly better with the unidimensional model (PNFI: F1 = 0.493, F2 = 0.398).1

A modified unidimensional model allowing the residual variance for items 4 and 5 to correlate (F1cov) showed improved fit relative to the baseline model (CFI: F1 = 0.986, F1cov = 0.995; RMSEA: F1 = 0.094, F1cov = 0.065) and identical fit to the two-factor solution (Fig. 1), since the two models are equivalent [39]. This latter model reflects, however, the time dependency in items 4 and 5 more specifically.

A fourth variant suggested by Vautier [34] tested for inter-item correlation when the items were presented to the participants consecutively and successively rather than scattered throughout the questionnaire. The fit of this model was very high (CFI = 0.999; RMSEA = 0.043), but consideration of parsimony indicated that this model should be rejected (PNFI = 0.100, P6FI = 0.067).

Since the correlation between the two factors in the two-factor model was very high (0.93) and minor secondary factors are inherent in most psychological measures [17], the modified single-factor model (Fig. 1) was retained for the subsequent analyses.

1 Detailed results of the analyses of these two models and the model for inter-item correlation is available from the corresponding author.
Table 3 and Fig. 1 shows the factor loadings and fit measures for the modified single-factor model—for the total sample and for the different subgroups. The factor loadings showed basically the same pattern across subgroups and were generally high (>0.70). For the youngest age group, however, factor loadings for items 4 and 5 were estimated to be <0.70 and relatively lower for item 2 (0.72) than observed in the remaining age groups (0.82–0.90). In the oldest age group, the factor loading for item 5 was also <0.70.

Since the responses are based on a 7-point Likert scale, we assumed continuous variables. The analyses were repeated using Bayesian techniques, however, which are recommended for ordinal data. The results from the two estimation techniques were identical principally to the third decimal (Table 4).

To further examine the effect of the non-normality of the data on results obtained using ML estimation, we also reran the analysis (main model, Fig. 1) using asymptotically distribution free (ADF) testing on the original data set (Table 4). Comparing results from ML estimation with ADF testing resulted in very similar estimates of factor loadings and variance, and all parameters were significant using both methods. Model fit using ADF testing was slightly worse when measured by CFI and PNFI and slightly better when measured by RMSEA.

To further examine the effect of non-normality of the data, the ML testing were repeated on data normalized using Tukey’s formula. The estimated factor loadings, intercepts, and variance differed as expected from those based on ML estimation. The factor loadings remained, however, significant, but the intercepts were no longer found to be significant. Additionally, tests of model fit resulted in worse fit as measured by CFI and parsimony (Table 4).

Finally, bootstrapping, the recommended analysis technique for non-normal data, confirmed the results obtained with standard ML estimation and indicated significance for factor loadings, intercepts, and variance (Table 4). A total assessment based on Table 4, thus seems to indicate that ML yields satisfactory results even when accounting for the non-normality and ordinal nature of the data.

Measurement invariance

Gender

The results of the tests for multigroup invariance between genders are given in Table 5. No significant differences in Δχ² were found, indicating weak (metric) invariance between the sexes. Both the strong and the strict invariance tests indicated significant differences across men and women. Due to the large sample sizes, tests involving χ² can be misleading, however. We therefore used the ΔCFI test which is more appropriate for large sample sizes [33, 40]. The ΔCFI results (Table 5) indicate measurement invariance at the weak, strong, and strict levels between genders. Partial invariance techniques indicated invariance at the strict level.

Model fit as measured by RMSEA improved as more constraints were imposed to the model while fit measured by CFI remained consistently high.

Age groups

Results of the tests for multigroup invariance between age groups are shown in Table 6. All three tests (weak, strong,
strict) indicated non-invariance across age as measured by significant differences in $\Delta \text{CFI}$. The $\Delta \text{CFI}$ test, however, indicated invariance at the weak level, but not at the strong or strict levels of invariance testing. Partial invariance testing, however, did indicate invariance at the weak level when removing constraints on factor loadings for item 1 (data not shown), as well as better model fit. Further testing of partial invariance at the strong and strict level did not support measurement invariance across age. In conclusion, the finding of invariance at the weak level assures that comparisons can be made as to the relationships between the factors (factor coefficients) across age groups. The results indicate, however, that caution should be used in analyses involving comparison of means between groups.

**Discussion**

The Satisfaction with Life Scale [3] is perhaps the most widely used measure of well-being worldwide. The dimensionality of SWLS has been widely discussed, but most studies have been based on specialized sample groups limited in size and biased with respect to gender, age, and relevant socio-demographic parameters. This study aimed to examine the dimensionality of the SWLS in a large and representative sample from Norway (including nearly 5,000 respondents), and to study the robustness of the scale in different subpopulations. This was done by exploring the (1) dimensional structure and (2) measurement invariance across gender and age. No other study has studied dimensionality or subgroup invariance across a continuous age distribution in a comparatively large community sample. This study also examined the comparability of results from different estimation techniques, including standard ML estimation using raw scores, Bayesian estimation, ADF estimation, and ML estimation using normalized data. The results were consistent regardless of estimation technique, indicating that use of standard ML estimation is satisfactory when studying dimensionality of SWLS when scored on a 7-point Likerts scale.

| ML (SE) | Bayesian analysis (SE) | ADF (SE) | Normalized data | Bootstrappingc ML (SE) |
|---------|------------------------|----------|-----------------|------------------------|
| $\lambda_{11}^a$ | 1.000 | 1.000 | 1.000 | 1.000 |
| $\lambda_{21}$ | 0.859 (0.011)* | 0.859 (0.011) | 0.843 (0.014)* | 0.921 (0.013)* | 0.859 (0.015)* |
| $\lambda_{31}$ | 0.881 (0.011)* | 0.881 (0.011) | 0.871 (0.014)* | 0.957 (0.012)* | 0.881 (0.014)* |
| $\lambda_{41}$ | 0.850 (0.013)* | 0.850 (0.013) | 0.849 (0.016)* | 0.882 (0.013)* | 0.850 (0.016)* |
| $\lambda_{51}$ | 0.948 (0.016)* | 0.969 (0.016) | 0.951 (0.017)* | 0.844 (0.014)* | 0.949 (0.016)* |
| $\tau_1$ | 2.935 (0.02)* | 2.934 (0.020) | – | $-0.017$ (0.013) | 2.936 (0.02)* |
| $\tau_2$ | 2.443 (0.019)* | 2.443 (0.018) | – | $-0.025$ (0.013)* | 2.443 (0.018)* |
| $\tau_3$ | 2.355 (0.018)* | 2.354 (0.018) | – | $-0.026$ (0.013)* | 2.355 (0.018)* |
| $\tau_4$ | 2.688 (0.02)* | 2.687 (0.019) | – | $-0.021$ (0.013) | 2.689 (0.019)* |
| $\tau_5$ | 3.377 (0.024)* | 3.376 (0.024) | – | $-0.011$ (0.013) | 3.378 (0.024)* |
| $\theta_1$ | 0.486 (0.014)* | 0.486 (0.015) | 0.450 (0.019)* | 0.204 (0.006)* | 0.486 (0.021)* |
| $\theta_2$ | 0.530 (0.014)* | 0.530 (0.014) | 0.533 (0.021)* | 0.255 (0.007)* | 0.530 (0.022)* |
| $\theta_3$ | 0.392 (0.011)* | 0.393 (0.012) | 0.386 (0.018)* | 0.201 (0.006)* | 0.391 (0.019)* |
| $\theta_4$ | 0.759 (0.018)* | 0.760 (0.018) | 0.753 (0.028)* | 0.333 (0.008)* | 0.757 (0.028)* |
| $\theta_5$ | 1.396 (0.031)* | 1.400 (0.032) | 1.362 (0.042)* | 0.404 (0.009)* | 1.395 (0.043)* |
| cov45 | 0.195 (0.018)* | 0.195 (0.017) | 0.174 (0.023)* | 0.071 (0.006)* | 0.195 (0.023)* |
| $\tau$ | 1.608 (0.042)* | 1.610 (0.042) | 1.610 (0.045)* | 0.650 (0.017)* | 1.607 (0.045)* |
| Chi$^2$ ($df$)$^b$ | 89.4 (4)* | – | 44.4 (4)* | 64.3 (4)* | 97.5 (0.611) ($df = 4$)$^d$ |
| CFI | 0.995 | – | 0.978 | 0.996 |
| RMSEA | 0.065 | – | 0.048 | 0.055 |
| PRatio | 0.400 | – | 0.400 | 0.267 |
| PNFI | 0.398 | – | 0.390 | 0.266 |

$\lambda = \text{item factor loading between latent variable (1) and items 1 through 5}$; $\tau = \text{item intercepts for items 1 through 5}$; $\theta = \text{item residuals for items 1 through 5}$; $\text{cov}$ = covariance between item 4 and 5; $\tau$ = latent mean; $^a$fixed at 1.0; $^b$in 1,000 s; $^c$Bootstrapping with 2,000 samples, 95% CI, and significance tested with bias corrected CI; $^d$mean chi$^2$; $* P < 0.01$. Model as described in Fig. 1
Table 5 Non-standardized parameter estimates and fit indices for measurement invariance models for men and women: baseline (unconstrained), weak (measurement weights), strong (measurement intercept), and strict (measurement residual) (Norwegian health interview survey 2005; N = 4,984)

| Parameter | Baseline | Weak | Strong | Strict |
|-----------|----------|------|--------|--------|
|           | M F      | M F  | M F    | M F    |
| λ11       | 1.00 1.00a | 1.00 1.00a | 1.00 1.00a | 1.00 1.00a |
| λ21       | 0.87 0.85  | 0.86 0.86b | 0.86 0.86b | 0.86 0.86b |
| λ31       | 0.90 0.87  | 0.88 0.88b | 0.88 0.88b | 0.88 0.88b |
| λ41       | 0.88 0.83  | 0.85 0.85b | 0.85 0.85b | 0.85 0.85b |
| λ51       | 0.97 0.93  | 0.95 0.95b | 0.95 0.95b | 0.96 0.96b |
| a         | 2.91 2.96  | 2.91 2.96 | 2.93 2.93b | 2.93 2.93b |
| b         | 2.43 2.46  | 2.43 2.46 | 2.44 2.44b | 2.44 2.44b |
| c         | 2.34 2.37  | 2.34 2.37 | 2.35 2.35b | 2.35 2.35b |
| d         | 2.72 2.66  | 2.72 2.66 | 2.69 2.69b | 2.69 2.69b |
| e         | 3.43 3.33  | 3.43 3.33 | 3.37 3.37b | 3.38 3.38b |
| f         | 0.45 0.51  | 0.44 0.52 | 0.44 0.52 | 0.49 0.49b |
| g         | 0.53 0.53  | 0.53 0.53 | 0.53 0.53 | 0.53 0.53b |
| h         | 0.38 0.40  | 0.38 0.40 | 0.38 0.40 | 0.39 0.39b |
| i         | 0.72 0.79  | 0.72 0.78 | 0.73 0.79 | 0.76 0.76b |
| j         | 1.47 1.32  | 1.47 1.32 | 1.47 1.33 | 1.40 1.40b |
| k         | 0.20 0.19  | 0.20 0.18 | 0.21 0.18 | 0.20 0.19 |
| l         | 1.48 1.73  | 1.51 1.69 | 1.51 1.69 | 1.51 1.70 |
| RMSEA     | 0.047 0.039 | 0.036 0.034 | 0.036 0.034 |

λ = item factor loading between latent variable (1) and items 1 through 5; τ = item intercepts for items 1 through 5; θ = item residuals for items 1 through 5; cov = covariance between item 4 and 5; z = latent mean. M male; F female; a = fixed at 1.0; b = constrained to equality with first group in the same model. * P < 0.01.

Dimensions

Our data essentially support a single-factor solution for the SWLS with 74% of variance explained by this single factor. The loadings are on the high side compared to previous studies, and there is a tendency for the last two items to load on a second, less important factor reflecting past accomplishments. This finding is in accordance with several previous reports, but the finding has been interpreted differently across the studies. The correlation between the two factors estimated in this study was very high (r = 0.93), however, and similar to previously reported estimates [14, 18], indicating that the two factors could not be easily differentiated. A post-hoc modification test on the data showed gained fit for the single-factor model when allowing the residual variances for items 4 and 5 to be correlated. This modified single-factor model improved the fit relative to the baseline model and produced fit measures identical to the two-factor model. The single-factor model also agrees with the theoretical development of the scale and measurement processes have been shown to elicit minor secondary factors for psychological measures [17].

Invariance

Between genders

No gender differences were observed at the level of factor loadings, indicating metric or “weak” invariance across gender in the total sample. This attests that the latent variable is related to the items in the same way for men and women. Further constraints equating the intercepts (strong invariance) and the residuals (strict invariance) resulted in significantly reduced fit in terms of the chi² test. Analyses based on large samples may result in high chi² values, however, and increased risk for rejecting good models. In the current study, additional fit indices were either improved (RMSEA) or only slightly decreased (CFI) when adding further constraints (equating intercepts and residuals) to the baseline model. This suggests that the intercepts and residuals may be fixed to equality in men and women, thereby supporting the assumption of strict invariance across gender. This implies that group means on the latent variable as well as analyses involving correlations with the latent variable are comparable across gender. This finding corroborates a number of previously described findings [14, 18, 21, 41], although Atienz et al. [25] in Spanish junior high students did not agree.

Between age groups

People differ in what they require for a satisfying life, and different dimensions of well-being seem to be meaningful.
Table 6 Non-standardized parameter estimates and fit indices for measurement invariance models for the subgroups of age: baseline (unconstrained), weak (measurement weights), strong (measurement intercept), and strict (measurement residual) (Norwegian health interview survey 2005; \( N = 4,984 \))

| Parameter | Baseline | Weak | Strong | Strict |
|-----------|----------|------|--------|--------|
| \( \lambda_{11} \) | 1.00 | 1.00a | 1.00a | 1.00a |
| \( \lambda_{21} \) | 0.65 | 0.85 | 0.91 | 0.95 |
| \( \lambda_{31} \) | 0.87 | 0.88 | 0.91 | 0.94 |
| \( \lambda_{41} \) | 0.81 | 0.87 | 0.85 | 0.85 |
| \( \lambda_{51} \) | 0.88 | 0.97 | 0.94 | 0.95 |
| \( \tau_{1} \) | 3.07 | 2.90 | 2.99 | 2.75 |
| \( \tau_{2} \) | 2.33 | 2.41 | 2.54 | 2.37 |
| \( \tau_{3} \) | 2.26 | 2.34 | 2.45 | 2.24 |
| \( \tau_{4} \) | 3.12 | 2.68 | 2.64 | 2.44 |
| \( \tau_{5} \) | 3.46 | 3.36 | 3.45 | 3.08 |
| \( \theta_{1} \) | 0.46 | 0.46 | 0.52 | 0.48 |
| \( \theta_{2} \) | 0.66 | 0.53 | 0.53 | 0.29 |
| \( \theta_{3} \) | 0.45 | 0.37 | 0.39 | 0.36 |
| \( \theta_{4} \) | 1.29 | 0.71 | 0.60 | 0.60 |
| \( \theta_{5} \) | 1.67 | 1.40 | 1.31 | 1.46 |
| \( \text{cov}_{45} \) | 0.16 | 0.15 | 0.20 | 0.31 |
| \( k \) | 1.64 | 1.58 | 1.69 | 1.39 |

\( k = \) item factor loading between latent variable (1) and items 1 through 5; \( s = \) item intercepts for items 1 through 5; \( h = \) item residuals for items 1 through 5; \( \text{cov} = \) covariance between item 4 and 5; \( \alpha = \) latent mean.

\( \chi^2 (df) \) = 163.7 (18); \( \Delta \chi^2 (\Delta df) \) = 80.3 (12)*; \( \text{CFI} \) = 0.991; \( \text{RMSEA} \) = 0.040; \( \Delta \chi^2 (\Delta df) \) = 209.3 (15)*; \( \text{CFI} \) = 0.976; \( \text{RMSEA} \) = 0.043; \( \Delta \chi^2 (\Delta df) \) = 291.3 (14)*; \( \text{CFI} \) = 0.959; \( \text{RMSEA} \) = 0.048

\( \chi^2 (df) \) = 244.0 (30); \( \Delta \chi^2 (\Delta df) \) = 453.3 (45); \( \text{CFI} \) = 0.987; \( \text{RMSEA} \) = 0.038

\( \chi^2 (df) \) = 744.7 (59); \( \Delta \chi^2 (\Delta df) \) = 291.3 (14)*; \( \text{CFI} \) = 0.959; \( \text{RMSEA} \) = 0.048

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entire country—all levels of society and a large age span. The shortcomings are related to a moderate response rate, perhaps leading to a less representative sample. When compared to population statistics, women and the age group from 45 to 64 years are overrepresented in this study. The eldest population group (>65 years) consists of fewer individuals, and only includes those living at home, and not in institutions. Likewise, immigrants with a non-Western ethnic background are clearly underrepresented in this material. In addition, using the AMOS analytical package did not allow robust ML testing (Satorra-Bentler scaled statistic) that would have strengthened the analysis when using ordinal non-normal data.

Conclusions

The overall results indicate that the one-factor latent structure of the SWLS is valid in the Norwegian data and that comparing men and women is feasible whereas some caution should be exerted when comparing age groups.

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References

1. Cummins, R. A., & Nistico, H. (2002). Maintaining life satisfaction: The role of positive cognitive bias. *Journal of Happiness Studies, 3*, 37–69.
2. Diener, E., Emmons, R. A., Larsen, R. J., & Griffin, S. (1985). The satisfaction with life scale. *Journal of Personality Assessment, 49*, 71–75.
3. Pavot, W., Diener, E., Colvin, C. R., & Sandvik, E. (1991). Further validation of the satisfaction with life scale: Evidence for the cross-method convergence of well-being measures. *Journal of Personality Assessment, 57*, 149–161.
4. Fujita, F., Diener, E., & Sandvik, E. (1991). Gender differences in negative affect and well-being: The case for emotional intensity. *Journal of Personality and Social Psychology*, 61, 427–434.
5. Lucas, R. E., & Fujita, F. (2000). Factors influencing the relation between extraversion and pleasant affect. *Journal of Personality and Social Psychology, 79*, 1039–1056.
6. Stubbe, J. H., Posthuma, D., Boomsma, D. I., & De Geus, E. J. (2005). Heritability of life satisfaction in adults: A twin-family study. *Psychological Medicine, 35*(11), 1581–1588.
7. Lucas, R. E., Diener, E., & Suh, E. (1996). Discriminant validity of well-being measures. *Journal of Personality and Social Psychology, 71*, 616–628.
8. Pavot, W., & Diener, E. (1993). Review of the Satisfaction with Life Scale. *Psychological Assessment, 5*, 164–172.
9. Anaby, D., Jarus, T., & Zumbo, B. (2010). Psychometric evaluation of the Hebrew language version of the Satisfaction with Life Scale. *Social Indicators Research, 96*, 267–274.
10. Arrindell, W. A., Heesink, J., & Feij, J. A. (1999). The Satisfaction with Life Scale (SWLS): Appraisal with 1700 healthy young adults in The Netherlands. *Personality and Individual Differences, 26*, 815–826.
11. Balatsky, G., & Diener, E. (1993). A comparison of the well-being of Soviet and American students. *Social Indicators Research, 28*, 225–243.
12. Lewis, C. A., Shevlin, M. E., Smekal, V., & Dorahy, M. J. (1999). Factor structure and reliability of a Czech translation of the Satisfaction with Life Scale among Czech university students. *Studia Psychologica, 41*, 239–244.
13. Gouveia, V., Milfont, T., da Fonseca, P., & Coelho, J. (2009). Life satisfaction in Brazil: Testing the psychometric properties of the Satisfaction with Life Scale (SWLS) in five Brazilian samples. *Social Indicators Research, 90*, 267–277.
14. Hultell, D., & Gustavsson, J. P. (2008). A psychometric evaluation of the Satisfaction with Life Scale in a Swedish nationwide sample of university students. *Personality and Individual Differences, 44*, 1070–1079.
15. Oishi, S. (2006). The concept of life satisfaction across cultures: An IRT analysis. *Journal of Research in Personality, 40*, 411–423.
16. Pavot, W., & Diener, E. (2008). The Satisfaction with Life Scale and the emerging construct of life satisfaction. *The Journal of Positive Psychology: Dedicated to furthering research and promoting good practice, 3*, 137–152.
17. Slocum-Gori, S., Zumbo, B., Michalos, A., & Diener, E. (2009). A note on the dimensionality of quality of life scales: An illustration with the Satisfaction with Life Scale (SWLS). *Social Indicators Research, 92*, 489–496.
18. Wu, Ch., & Yao, G. (2006). Analysis of factorial invariance across gender in the Taiwan version of the Satisfaction with Life Scale. *Personality and Individual Differences, 40*, 1259–1268.
19. Sachs, J. (2003). Validation of the Satisfaction with Life Scale in a sample of Hong Kong university students. *Psychologia: An International Journal of Psychology in the Orient, 46*, 225–234.
20. McDonald, R. P. (1999). *Test theory: A unified treatment*. Mahwah, N.J: Lawrence Erlbaum Associates.
21. Shevlin, M., Brunsden, V., & Miles, J. N. V. (1998). Satisfaction With Life Scale: analysis of factorial invariance, mean structures and reliability. *Personality and Individual Differences, 25*, 911–916.
22. Blais, M. R., Vallerand, R. J., Pelletier, L. G., & Briere, N. M. (1989). L’echelle de satisfaction de vie: Validation Canadienne-Francaise du “Satisfaction With Life Scate” [French-Canadian validation of the Satisfaction With Life Scale]. *Canadian Journal of Behavioral Science, 21*, 210–223.
23. Durak, M., Senol-Durak, E., & Genco, T. (2010). Psychometric properties of the Satisfaction with Life Scale among Turkish university students, Correctional officers, and elderly adults. *Social Indicators Research, doi:10.1007/s11205-010-9589-4*.
24. Siedlecki, K. L., Tucker-Drob, E. M., Oishi, S., & Salthouse, T. A. (2008). Life satisfaction across adulthood: Different determinants at different ages? *The Journal of Positive Psychology, 3*, 153–164.
25. Atienza, F. L., Balaguera, I., & Garcia-Merita, M. L. (2003). Satisfaction with Life Scale: Analysis of factorial invariance across sexes. *Personality and Individual Differences, 35*, 1255–1260.
26. Pons, D., Atienza, F. L., Balaguera, I., & Garcia-Merita, M. L. (2000). Satisfaction with Life Scale: Analysis of factorial invariance for adolescents and elderly persons. *Perceptual and Motor Skills, 91*, 62–68.
27. Tucker, K., Ozer, D., Lyubomirsky, S., & Boehm, J. (2006). Testing for measurement invariance in the Satisfaction with Life Scale: A comparison of Russians and North Americans. *Social Indicators Research, 78*, 341–360.

28. Arbuckle, J. L., & Wothke, W. (2008). *AMOS 17.0 User’s Guide*. Chicago: SPSSInc.

29. Rhemtulla, M., Brosseau-Liard, P., & Savalei, V. (2010). How many categories is enough to treat data as continuous? A comparison of robust continuous and categorical SEM estimation methods under a range of non-ideal situations. Manuscript under review, pp. 1–51.

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

31. Byrne, B. M. (2010). *Structural equation modeling with AMOS—basic concepts, applications and programming*. New York, London: Routledge, Taylor and Francis group, LLC.

32. West, S. G., Finch, J. F., & Curran, P. J. (1995). Structural equation models with non-normal variables: Problems and remedies. In R. H. Hoyle (Ed.), *Structural equation modeling: Concepts issues and applications* (pp. 56–75). Thousand Oaks: Sage.

33. Hooper, D., Coughlan, J., & Mullen, M. R. (2008). Structural equation modelling: Guidelines for determining model fit. *The Electronic Journal of Business Research Methods, 6*, 53–60.