

1 **Title:** Testing the Factor Structure of the Warwick-Edinburgh Mental Well-Being Scale in
2 Adolescents: a Bi-Factor Modelling Methodology.

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14 *Conflict of interest*

15 All authors have no conflict of interest to declare.

16 *Ethical approval*

17 Ethical approval was granted by Ulster University's ethical committee (approval number=
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19 *Informed consent*

20 Informed consent was provided by all participants included in the study.

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1 **Abstract**

2 **Aim:** Despite extensive use in mental health research and practice, limited evidence exists for
3 the hypothesised unidimensional model of the Warwick-Edinburgh Mental Well-Being Scale
4 in adolescents. Few studies have assessed competing Confirmatory Factor Analysis (CFA)
5 models, and the instrument has yet to be assessed in younger adolescents in Northern Ireland,
6 a jurisdiction characterised by high rates of mental illness.

7 **Subject and Methods:** School pupils ($n=1,673$) aged 13-18 years ($M = 14.87$, $SD = 1.16$),
8 including 1,036 females, 997 urban children, and 312 from lower socio-economic status,
9 completed psychometric tests. Seven CFA models based on extant research were tested,
10 including unidimensional, bi-factor, higher-order and clustered.

11 **Results:** Several models, including the default unidimensional model, did not achieve
12 recommended CFA fit thresholds. Model 6 comprising one strong 'general well-being' factor
13 and three residual factors (i.e., figuratively labelled: 'Affective', 'Psychological Functioning'
14 and 'Social Relationships') was confirmed as the superior model. Most item variance was
15 explained by the general factor, relative to residual factors.

16 **Conclusions:** Adolescents predominantly conceptualise well-being as a unitary construct that
17 coexists with relatively weak affective, psychological and social relationship domains.
18 Researchers and practitioners should foremost calculate a composite score of well-being, and
19 explore sub-domains to supplement understanding of adolescent well-being.

20

21 **Keywords:** Well-being; mental health; psychology; confirmatory factor analysis; validity;

22

23 **Abbreviations**

24 CFA= confirmatory factor analysis;

Mental Health Measurement in Adolescents

- 1 CFI= Comparative Fit Index
- 2 RMSEA = root mean square error of approximation
- 3 SDQ=Strengths & Difficulties Questionnaire;
- 4 STROBE =Strengthening the Reporting of Observational Studies in Epidemiology;
- 5 TLI = Tucker-Lewis Index
- 6 χ^2 = Chi-Square Value
- 7 WEMWBS = Warwick-Edinburgh Mental Well-Being Scale;
- 8 WRMSR = Weighted Root Mean Square Residual
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1 **Introduction**

2 Well-being is defined as a state of optimal functioning (Ryan & Deci, 2017) and represents
3 one dimension of a two continua model of mental health, wherein mental ill-being and well-
4 being coexist as two distinct, but correlated factors (Keyes, 2005). Within the well-being
5 construct, Keyes' (2002) theory outlines three factors of hedonic (i.e., positive affective
6 states), eudemonic (e.g., psychological functioning, sense of purpose), and social (i.e.,
7 relationships, integration) well-being. Individual well-being is considered a measurable
8 marker of societal health beyond economic indicators such as Gross Domestic Product (GDP)
9 per-capita (Huppert, 2009), with educational success, living in a safe neighbourhood, family
10 support, and economic prosperity all correlates of self-reported well-being (United Nations,
11 2015). Further, when high in well-being, one is at reduced likelihood of mental illness, and
12 more likely to flourish in society (Keyes, 2007; Ryan & Deci, 2017). Most mental health
13 disorders have their onset during early adolescence (i.e., age 11-14) (Jones, 2013), and
14 relatedly, the importance of young people's well-being development is acknowledged as a
15 human right in Article 12 of the United Nations Convention of the Rights of the Child
16 (UNICEF, 1989). As such, the United Kingdom (UK) and Irish Governments aim is to
17 continually assess and improve young people's well-being (Coulter, 2017), delineating the
18 need for valid and reliable measurement instruments.

19 Adolescents conceptualisation of well-being may be diverse, wherein cultural
20 contexts, socio-economic circumstances, and age-related developmental stages (e.g., puberty)
21 likely exerting a role (Witten, Savahl & Adams, 2019). As such, competing psychometric
22 measurement models require assessment among diverse population groups (Park, Han &
23 Cho, 2011), and the accurate reporting of psychometric measures of well-being affects public
24 health to the extent that an instrument's validity evidence informs clinical practice, research,
25 economic and policy decisions (Doran, Wallace & Woods, 2014). Fried (2017) outlined that

1 when observed items on a questionnaire correlate with a latent variable(s) (e.g., general well-
2 being factor), it is more likely that a psychometric instrument is measuring its underlying
3 ‘true’ or ‘natural’ construct. Confirmatory Factor Analysis (CFA) is a robust analytic method
4 that can explain and establish construct validity among observed and latent variables (Hagger
5 & Chatzisarantis, 2009). Specifically, researchers can assess competing CFA models through
6 inspection of fit indices, helping determine whether items, structural pathways and/or factors
7 are acceptable or require modifications (e.g., removal of items/constructs) (Schreiber et al.,
8 2006). A number of potential CFA models can be specified, including; unidimensional (i.e.,
9 one underlying construct), higher order (i.e., general dimension as well as specific sub-
10 dimensions), and clustered (i.e., multiple correlated dimensions with no general dimension)
11 (Jackson et al., 2009). Recent bi-factor approaches to CFA also permit items to correlate with
12 a strong general factor alongside residual sub-factors (Reise, 2012), wherein sub-factors and
13 general factors are orthogonal, and compete for explaining item covariance (Beaujean, 2015).

14 Beyond CFA, a sound measure should display nomological validity, which refers to
15 how the proposed construct correlates with constructs derived from the same theory (Hagger
16 & Chatzisarantis, 2009). Lastly, to be confident in an instrument’s responsiveness to change
17 over time (e.g., in response to intervention), Terwee et al. (2007) indicate that a measure
18 should not display floor or ceiling effects, wherein 15% of a sample score at the extreme ends
19 of the scale.

20 *The Warwick-Edinburgh Mental Well-Being Scale*

21 The Warwick-Edinburgh Mental Well-Being Scale (WEMWBS) is a 14-item questionnaire,
22 developed in the UK to standardise the measurement of the population’s well-being (Tennant
23 et al., 2007) and captures the positive well-being dimension of mental health. Respondents
24 self-report their experiences using a 5-point Likert scale, and items are theorised to asses

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1 hedonic or eudemonic well-being components (Tennant et al., 2007). The instrument
2 development phase concluded a hypothesised unidimensional general well-being factor
3 (Tennant et al., 2007; Stewart-Brown et al., 2009a). Psychometric support for the original
4 hypothesised unitary WEMWBS measurement model has been found among diverse UK
5 population groups, although relatively fewer studies exist among adolescents compared to
6 adults (Stewart-Brown et al., 2009a; Stewart-Brown et al., 2009b; Maheswaran, Weich,
7 Powell & Stewart-Brown, 2012; Melendez-Torres et al., 2019). Further evidence has shown
8 that the unidimensional factor structure has not been robustly supported by recommended
9 CFA fit thresholds (e.g., Kim et al., 2014; Lopez et al., 2015), including a study among
10 adolescents showing an unacceptable fit (Hunter, Houghton & Wood, 2015).

11 Some authors have re-specified the model with item residual error covariances (e.g.,
12 Hunter, Houghton & Wood, 2015; Haver et al., 2015; Clarke et al., 2011), which may be
13 unintentionally omitting theoretically meaningful variable(s) (Hermida, 2015). For example,
14 in a community-based sample of adults in Austria, Lang and Bachinger (2017) found that the
15 unidimensional model was not supported, but a bi-factor model comprised one strong general
16 well-being factor alongside three residual domain factors (i.e., positive affect, psychological
17 functioning and social relationships) was. That said, the retained model included item
18 residual error covariances across different residual factors, which given such limited semantic
19 overlap (e.g., 'I've been feeling optimistic about the future' with 'I've been feeling useful'),
20 provides insufficient theoretical explanation (Reise, 2012). Furthermore, given bi-factor
21 modelling restricts correlations between residual domain factors, the research may have
22 benefitted from testing a higher order model, in which well-being may hypothetically
23 subsume the shared variance between these domain factors (Beaujean, 2015). In another
24 study with Argentine elders (Azcurra, 2015) a lack of fit was found for the unidimensional
25 structure, but a clustered model comprising two covaried factors labelled as emotional and

1 psychological well-being factors was retained. However, the authors did not test competing
2 bi-factor or higher order models in their research.

3 Most notable for the rationale for the present study, in the few adolescent studies that
4 have been conducted on the WEMWBS, there has yet to be a comprehensive assessment of
5 competing CFA measurement models, with most adopting the traditional default
6 unidimensional model (Reise et al., 2013). Such incomplete and contrasting statistical
7 evidence is significant, as Fried (2017) has suggested that model misfit issues should be
8 identified for improved construct clarity. Hence, iterations of Tennant et al.'s (2007), Lang
9 and Bachinger (2017) and Azcurra's (2015) models require further assessment among
10 adolescents. In addition to the theoretical underpinning of the structure of the instrument, the
11 demography of samples included in WEMWBS psychometric validation studies in Northern
12 Ireland (NI) (i.e., Lloyd & Devine, 2012; McKay & Andretta, 2017) was restricted to
13 individuals aged ≥ 16 . NI has been identified as having the poorest adult mental health rates
14 across the UK and Republic of Ireland (Leavey, Galway, Rondon, & Logan, 2009). Given the
15 aforesaid point that most mental health disorders have their onset in early adolescence (i.e.,
16 aged 11-14) (Jones, 2013), it is important to include a broader spectrum of adolescents (i.e.,
17 aged 12-18) for improved model accuracy in research (Wang et al., 2007; Park, Han & Cho,
18 2011).

19 *Aims and Objectives*

20 Given the extensive application of the WEMWBS in mental health research and practice, the
21 current study responded to the contrasting statistical evidence on the WEMWBS factor
22 structure in adolescents. Hence, the objective was to assess competing measurement models
23 of the WEMWBS, and its psychometric properties (i.e., nomological validity, floor and

1 ceiling effects) across a large, understudied, sample of adolescents who took part in the
2 Northern Ireland Study of Adolescent Well-Being (NISAW; Corry & Leavey, 2017).

3 **Materials and Methods**

4 ***Study Design, recruitment and participants***

5 This study adhered to the Strengthening the Reporting of Observational Studies in
6 Epidemiology (STROBE) Statement. This was a cross-sectional study of adolescent school
7 pupils in NI aged 13-18 years, collected as part of the NISAW (Corry & Leavey, 2017;
8 Leavey et al., 2019; Leavey et al., 2020). Inclusion criteria was based on being a registered
9 adolescent (aged 11-18) school pupil in NI. Individuals were excluded if they were under or
10 above adolescent years, and resided outside the jurisdiction of NI. To achieve a representative
11 sample, we randomly selected eight post-primary schools from Education and Library Board
12 databases, which were stratified by markers of urbanicity, school type (i.e.,
13 Grammar/Secondary Modern), and socio-economic index (e.g., average income, lone
14 parenthood). The research was ethically approved by Ulster University Research Ethics
15 Committee. Further methodological information on the data collection procedures can be
16 found in Leavey et al. (2019).

17 ***2.2 Outcome measures***

18 *Well-being*

19 The WEMWBS (Tennant et al., 2007) was used to assess well-being. The instrument
20 comprises 14-items designed to measure hedonic (e.g., happiness and life satisfaction), and
21 eudemonic (i.e., relationships, self-actualisation) well-being components (Stewart-Brown et
22 al., 2011). Each item is positively worded and scored on a 5-point Likert scale ranging from
23 'none of the time' (1), to 'all of the time' (5). The measure has undergone robust
24 psychometrics (i.e., test-retest, convergent validity, discriminant validity) validated across

1 [adolescents in the UK, with a proposed unitary factor structure \(Tennant et al., 2007; Stewart-](#)
2 [Brown et al., 2009a\)](#). Total scores can range from 14 through to 70, with higher scores
3 indicating higher well-being (Stranges et al., 2014; Fat et al., 2017).

4 *Psychological difficulties*

5 The Strengths & Difficulties Questionnaire (SDQ) (Goodman, 1997) was used to assess
6 psychological difficulties such as emotional problems, conduct problems, hyperactivity and
7 peer problems. The SDQ has been widely validated in youth samples (Theunissen, de Wolff,
8 & Reijneveld, 2019). Due to construct overlap with the SDQ prosocial items (e.g., ‘Other
9 people my age generally like me’) and social questions within the WEMWBS (e.g., ‘I have
10 been feeling close to other people’), we only computed the SDQ difficulties scale (McCrory
11 & Layte, 2012). The SDQ difficulties scale showed an acceptable Cronbach’s alpha value
12 within the sample ($\alpha=.75$). All items were scored on a 3-point scale (i.e., 0 =Not True, 1=
13 Somewhat True, 2=Certainly True). Total difficulty scores can range between 0 to 40, with
14 higher scores indicating more psychological difficulties.

15 *Ill-being*

16 A 3-item ill-being scale used in the NISAW study (Corry & Leavey, 2017) was used to
17 measure common negative emotions reflective of psychological ill-being (Keyes, 2002). The
18 items were scored on a 5-point Likert scale ranging from 1 (all to time) to 5 (none of the
19 time), and included: “How often do you feel stressed?”, “How often do you feel depressed?”,
20 and “How often do you feel like your emotions have become too much for you?”. This scale
21 demonstrated good internal consistency within the sample ($\alpha = .86$) and higher scores reflect
22 the presence of fewer negative emotions.

23 *Home life*

1 The stability and support in one's home environment was measured using a questionnaire
2 from the Understanding Society UK Household Longitudinal Study (McFall & Garrington,
3 2011). The 12-item questionnaire measures aspects of home life including: sibling
4 relationships; family activities and meals; and perceived support during times of difficulty.
5 Items are scored on Likert scales ranging from 2 to 6-points, and negatively worded items
6 (e.g., "In the past month, how many times have you stayed out after 9.00pm at night without
7 your parents knowing where you are?") were reverse coded for analysis. Higher scores
8 indicate a stable, supportive home environment (McFall & Garrington, 2011).

9 *Parental relationships*

10 An eight-item questionnaire from the Determinants of Adolescent Social Well-Being and
11 Health study (DASH; Maynard & Harding, 2010) was used to measure the quality of parental
12 relationships. This scale was internally consistent within the sample (Cronbach's $\alpha = .81$),
13 and used a 4-point Likert scale (i.e., 1 = "Never" to 4 = "Always"). Items are scored so that
14 higher scores are indicative of a more positive, healthy relationship with parents (Maynard &
15 Harding, 2010), and therefore negatively worded items (e.g., "They treat me like a baby")
16 were reverse coded for analysis.

17 *Data Analysis*

18 Detailed information on the data input, cleaning and preparation process can be found in
19 Corry & Leavey (2017). Seven competing measurement models based on prior research were
20 tested in Mplus (version 7) using Weighted Least Squares estimation. CFA was firstly
21 conducted on the default hypothesised unidimensional structure (Model 1) (Tennant et al.,
22 2007). Three model iterations based on the Azcurra (2015) study were then tested, including:
23 a clustered model with two correlated factors (Model 2); a bi-factor model comprising a
24 strong general factor and two orthogonal residual domain factors (Model 3), and; a higher

1 order model comprising two correlated domain factors loading onto general well-being
2 (Model 4). Furthermore, three iterations of Lang and Bachinger's (2017) model were tested,
3 including: a clustered model (Model 5) with three correlated domain factors; the previously
4 retained bi-factor model with one strong general factor, and three orthogonal residual factors
5 of social relationships, affect and psychological functioning (Model 6), and; a higher order
6 model with three correlated domain factors loading onto a general well-being factor (Model
7 7). Parameter estimates were fixed to 1, and cross-loadings on unintended factors were
8 estimated at 0.

9 The adequacy of the competing models were assessed through inspection and
10 comparison of multiple recommended goodness-of-fit indices (Hu & Bentler, 1999), with
11 5000 Bollen-Stine bootstraps to improve the accuracy of model parameters (Byrne, 2001).
12 The Chi-Square [χ^2] goodness-of-fit index was reported, however given that large sample
13 sizes are sensitive to χ^2 values (Bentler, 1990), we approached this value with caution. The
14 Tucker-Lewis Index (TLI) and Comparative Fit Index (CFI) were reported with values of
15 $>.90$ or $>.95$ considered as acceptable or good model fit, respectively; the root mean square
16 error of approximation (RMSEA) with values of 0.08 or below considered acceptable, and;
17 the Weighted Root Mean Square Residual (WRMSR) with values ~ 1 considered acceptable.
18 We adopted the recommendations of Comrey and Lee (1992) for determining the perceived
19 strength of factor loadings (i.e., $<.30$ = poor; $>.45$ = fair; $>.55$ =good; $>.63$ = very good, and;
20 $>.71$ = excellent).

21 After choosing a bifactor model as the superior model (see results below), factor
22 weightings were entered in the Bi-factor Indices Calculator (Dueber, 2017). For items loading
23 onto the general factor internal reliability was assessed through McDonald's omega
24 hierarchical coefficient (ω_H); McDonald's omega specific (ω_S) was calculated to determine
25 whether items loading onto their residual domains could reliability explain residual variances.

1 Further, total model variance was calculated through the explained common variance (ECV)
2 value, that determines the ratio of variance explained by the general factor, divided by the
3 variance explained by the general factor plus the residual domain factors. For interpretation,
4 ω^2 values of $> .80$ and higher ECV values attributable to the general factor suggest the
5 presence of a unitary construct (Reise, Schienens, Widaman & Haviland, 2013). A table
6 comprising the fit statistics above was produced, and a second table for the retained model
7 describing the items and corresponding factor loadings. A visual figure describing the
8 retained model was also produced.

9 Ceiling and floor effects were calculated and determined present if more than 15%
10 scored at the highest (i.e., 70) or lowest (i.e., 14) end of the WEMWBS respectively (Terwee
11 et al., 2007). Nomological validity assessments were based on Keyes (2002) theory of mental
12 health, which as previously discussed, encompasses two distinct but correlated mental well-
13 being and mental ill-being continuums, represented by psychological, social and emotional
14 factors (Keyes, 2002). Specifically the WEMWBS composite score was correlated with
15 psychological difficulties (SDQ; Goodman, 1997) and the ill-being scale (NISAW study;
16 Corry & Leavey, 2017) to reflect indicators of emotional, social and psychological ill-being
17 (Keyes, 2002). Positive home environment (McFall & Garrington, 2011), and parental
18 relationships (Maynard & Harding, 2010) were correlated as proxy measures of social well-
19 being. Pearson's Product-Moment Correlation (r) was used with alpha significance set at $p <$
20 0.05 . A correlation matrix was calculated and inputted into a table, considering values of $0.0 -$
21 0.3 as weak, $0.31 - 0.70$ as moderate, and 0.71 and above as strong (Field, 2013). Based on
22 Keyes (2002) theory we hypothesised that the mental ill-being (i.e., ill-being and
23 psychological difficulties) scales would be strongly inversely correlated with well-being,
24 whereas the social factors (i.e., home life, parental relationships) would be moderately
25 correlated.

1 Results

2

3 Data was collected from 1,673 participants comprising 1,036 females (61.9%) and 623 males
4 (37.2%), with 14 participants (0.8%) not disclosing gender. Participant ages ranged between
5 13-18 ($M = 14.87$, $SD = 1.16$), with 997 (59.80%) living in an urban and 670 (40.20%) rural
6 environment. Ethnicity was collected from 1,649 participants (98.6%), with 1,562
7 participants (93.4%) reporting 'White European' and 87 (5.2%) as 'minority ethnic groups
8 (e.g., 'Indian' =0.7%; 'Irish Traveller' = 0.5%; 'Black African' = 0.4%, and; 'Mixed
9 Ethnicity'=1.6%). In the total sample, 312 participants (20.9%) claimed free school meals
10 which is a proxy measure of lower socio-economic status for youths in NI (Corry & Leavey,
11 2017). Table 1 details a description of the mean scores for the total sample and demographic
12 groupings for each study outcome.

13 *Please insert Table 1*

14 Fit indices for the competing CFA models are presented in Table 2. Given the χ^2
15 value sensitivities with sample size, all values were significant and thus did not lead to
16 rejection of the models. All factor loadings in Model 1 were statistically significant ($p <$
17 0.05) ranging from 0.36 (item 4) to 0.80 (item 8) as reflected in the original unidimensional
18 WEMWBS structure. However, the RMSEA value was above the recommended threshold of
19 equal or close to 0.08 (i.e., 0.100). However, other fit statistics were within the acceptable-to-
20 good ranges.

21 Similarly, all factor loadings were statistically significant in iterations of Azcurra's
22 (2015) Models 2 and 4 ($p < 0.05$), but both RMSEA values were unacceptable. Further, the
23 covariation pathway was .97 in Model 2, suggesting little distinct variation between the two
24 constructs beyond that of a general factor. The bi-factor version of Azcurra's (2015) model
25 (Model 3) yielded a substantially better fit, and the RMSEA value was close to acceptance at

1 0.087. With that said, none of the items significantly loaded onto their residual factors ($p >$
2 0.05), suggesting very little contribution of the residual factors to the explained variance
3 beyond that of the unitary well-being construct.

4 Lang and Bachinger's (2017) Models 5 and 7 iterations displayed significant ($p <$
5 0.05) item factor loadings on their latent variables, but again were above the recommended
6 RMSEA threshold. However, the bi-factor Model 6 yielded a substantially better fit, and was
7 subsequently confirmed as the superior model among the competing CFA models.
8 Specifically, all values were within the recommended confirmatory fit thresholds, including
9 an excellent CFI value of 0.97. Applying Lang and Bachinger's (2017) figurative labels to the
10 factors, Model 6 comprised a strong 'general well-being' factor, and three comparatively
11 weak residual factors labelled: 'Affect'(n.b., items 1, 3, 5, 8, 10, 14), 'Social Relationships'
12 (n.b., items 2, 4, 9, 12) and 'Psychological Functioning' (n.b., items 6, 7, 11, 13). As visually
13 depicted in Figure 1, paths between the items and GF symbol refer to loadings on to the
14 general well-being unitary construct, whereas the item loadings onto SR (Social
15 Relationships), AF (Affect), and PF (Psychological Functioning) represent the sub-domain
16 factors.

17 As further outlined in Table 3, all factor loadings onto the general factor were
18 significant ($p < 0.05$), and ranged from 0.44 - 0.83, thus deemed in the 'fair' to 'excellent'
19 ranges. All but two item loadings (n.b., 5 and 9) on residual factors were statistically
20 significant ($p < 0.05$), and ranged from 0.07 – 0.59 (n.b., item 9). By including the residual
21 factor loadings alongside the general factor loadings, up to 93% of item variance was
22 explained (item 7), to as low as 26% (item 4). However, inspection of the construct paths on
23 each side of the Figure 1's items demonstrated that general factor item loadings were
24 consistently larger than residual factor item loadings. Notably, one item on each residual
25 factor loaded negatively, albeit 'poorly' (i.e., items 12 and 13 = -0.12 and -0.13, respectively)

1 or 'fairly' (i.e., item 1 = -0.32). In these cases, the items appear to have an opposite
2 contribution to that factor compared to the other items loading onto that factor. Therefore,
3 when calculating residual factor subscale scores, items 1, 12 and 13 would require subtraction
4 from the remaining items.

5 Further model reliability assessment through the Bi-Factor Indices Calculator
6 (Dueber, 2017), revealed a high (i.e., 0.92) omega hierarchical coefficient (ω_H) for items
7 loading onto the general factor, indicating that the composite score reliability assesses a
8 unitary dimension of well-being. However, Omega scores for specific residual factors (ω_S)
9 suggested that the subscales have a sound ability to explain residual variances when assessed
10 in isolation, i.e., 0.83 (Affect), 0.87 (Psychological Functioning), and 0.66 (Social
11 Relationships). Yet, and aligned with the larger item loadings onto the general factor relative
12 to the residual factor item loadings, ECV values showed that the general factor accounted for
13 the substantial proportion of model variance (86%). However, residual factors did explain a
14 significant proportion of item variance, and contributed to the model's overall parsimony;
15 i.e., 6% for Psychological Functioning, 6% for Affect, and 2% for Social Relationships.
16 Taken the results collectively, adolescents perceive well-being as a largely unidimensional
17 construct, with comparatively weak residual subdomains.

18 *Please insert Table 2, 3 and Figure 1 here*

19 In the retained model, no floor or ceiling effects were present, such that 0.05%
20 reported the highest possible 70, and 0.01% reported the lowest possible of 14. The
21 correlation matrix for the retained Model 6 factors and study outcomes is detailed in Table 4.
22 All correlations were statistically significant at $p < .001$, and moderate, with r ranging from
23 0.36 to 0.70. As hypothesised, the composite score representing the strong general well-being
24 factor showed larger correlations with 'Ill-being' and 'Psychological Difficulties' than that of

1 'Parental Relationships' and 'Home Life'. Within the WEMWBS residual factors,
2 correlations were moderate-to-strong.

3 *Please insert Table 4 here: Correlation matrix for the retained Model 6 and study outcomes*

4 **Discussion**

5
6 This study tested the factor structure and psychometric properties of the WEMWBS in a large
7 and diverse sample of NI adolescents and extends existing theoretical and methodological
8 assessments of the WEMWBS. CFA conducted on seven competing CFA models found that,
9 consistent with some prior research (Kim et al., 2014; Lopez et al., 2015; Santos et al., 2015),
10 the original unidimensional model was not supported by recommended fit thresholds in
11 adolescents. However, and for the first time in adolescents, an adapted model proposed by
12 Lang & Bachinger (2017) comprising one strong general well-being factor, and three
13 relatively weak residual factors was tested and confirmed as the superior model. Inspection of
14 fit indices, factor loadings and model reliability confirmed the adequacy of the retained
15 model, which did not display floor or ceiling effects. Moderate nomological validity
16 correlations were present with several mental health outcomes. Overall, findings provide a
17 theoretical and methodological basis for an adapted WEMWBS measurement model in
18 mental health research and practice in adolescents.

19 CFA conducted on the originally theorised unidimensional WEMWBS model
20 (Tennant et al., 2007; Stewart-Brown et al., 2009a) displayed an unacceptable RMSEA fit
21 threshold, despite the factor loadings all displaying statistical significance. Findings provide
22 further evidence of suboptimal fit for the unidimensional model (e.g., Kim et al., 2014; Lopez
23 et al., 2015; Hunter, Houghton & Wood, 2015; Azcurra, 2015; Santos et al., 2015; Lang &
24 Bachinger, 2017). While some studies among adolescents (e.g., Clarke et al., 2011) have
25 found good-to-excellent fit indices for the unidimensional model among urban adolescents,

1 Hunter, Houghton and Wood's (2015) study among rural adolescents did not. Having
2 included a close to 60/40% split of urban to rural children, our study supports the view that
3 adolescents may indeed have a broader conceptualisation of well-being (Santos et al., 2015).
4 Therefore, it was pertinent to test competing alternative models in order to improve construct
5 clarity (Fried, 2017).

6 Clustered and higher order versions of Azcurra's (2015) adapted WEMBWS model
7 comprising 'positive emotions' and 'positive functioning' factors did not display any
8 substantial changes in model fit compared to the unidimensional model. While initial
9 inspections of a bi-factor version of Azcurra's (2015) model showed promise, namely
10 positive changes across all fit indices, item loadings onto residual factors were not
11 significant. Within bi-factor models, residual sub-factors should account for, at least in some
12 part, item variance (Reise, 2012). Therefore, despite the improved fit indices there is limited
13 contribution of the addition of subfactors in this case (Beaujean, 2015).

14 Including three well-being subfactors in clustered and higher order adaptations of
15 Lang & Bachinger's (2017) model showed improvements in model fit indices compared to
16 the unidimensional model, specifically through WRMR and RMSEA values. However, the
17 bi-factor Model 6 was chosen as the superior model as findings (see Table 2) revealed the
18 most positive fit indices. Indeed, all fit values exceeded or were within acceptable ranges (Hu
19 & Bentler, 1999). Inspection of the three residual factors and their corresponding significant
20 item loadings provided a further level of confidence in the retained model (see Table 3).
21 Corresponding figurative labels by Lang & Bachinger (2017) were thus applied: (i) an
22 'Affective' hedonic domain wherein 'feelings' are emphasised; (ii) a 'Social Relationships'
23 domain that is dependent on closeness of others, and; (iii) a 'Psychological Functioning'
24 domain emphasising higher-order cognitive processes (e.g., problem-solving, interests).
25 Importantly, we did not allow item residual errors to correlate with each other within or

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1 across domain factors, as we did not see a semantic or theoretical justification for doing so
2 (Hermida, 2015).

3 Model 6 arguably taps somewhat into Keyes' (2002) definition of positive mental
4 health that comprises hedonic, eudemonic and social well-being domains. However, further
5 analysis of Keyes' (2002) model by Jovanovic' (2015) have, likewise to the present study,
6 found more convincing evidence that one should foremost calculate a composite score of
7 general well-being, and approach residual domain factors with caution. Specifically, and as
8 outlined in Table 2, item loadings on the general factor were consistently larger than their
9 respective residual factor loadings. A high (i.e., 0.92) omega hierarchical coefficient (ω_H) for
10 the unitary well-being construct was also found, and ECV values showed that common item
11 variation was largely explained by the general factor (86%), with relatively weak contribution
12 of the residual factors ranging from 2% to 6%.

13 Despite these findings, some residual factor item loadings were as high as 0.59 (i.e.,
14 item 9: 'I've been thinking clearly'), wherein the collective contribution of the general and
15 residual factor resulted in 93% of item variance explained. Therefore, residual factors did
16 contribute in small part to the model's overall parsimony, and subscales may be useful
17 alongside the composite score in clinical practice, research, or public health provision
18 (Doran, Wallace & Woods, 2014). For example, researchers and practitioners should
19 foremost calculate a composite score of adolescent well-being, whilst supplementing their
20 understanding of adolescent mental health through using specific residual subscales for
21 further interpretation (Hajian-Tilaki, 2013). For example, the effect of school-based social
22 support interventions can be assessed from the perspective of total and domain factors of
23 adolescent well-being. However, given items 1, 12 and 13 loaded negatively, they appear to
24 have an opposite contribution to that factor compared to the other items (DiStefano, Zhu &
25 Mindrila, 2009), and would require subtraction from their remaining items.

1 An additional feature of our results was that the WEMWBS did not display ceiling or
2 floor effects, meaning the WEMWBS will likely have sound responsiveness to change over
3 time in intervention studies (Terwee et al., 2007). Moreover, moderate correlations were
4 found between the WEMWBS and several additional mental health outcomes, including
5 fewer psychological difficulties as assessed by the SDQ (Goodman, 1997), and reduced
6 symptoms of ill-being. Consistent with Keyes' (2002) model, findings suggest that well-
7 being and ill-being are linked, but likely distinct dimensions of the mental health construct.

8 Notwithstanding the study contributions, a limitation of our research was the cross-
9 sectional design implemented. A longitudinal design would have provided the opportunity for
10 test-retest reliability assessments, and if performed over a longer period (i.e., transition from
11 adolescence to adulthood) researchers could examine if individuals' well-being conceptions
12 change over time. Furthermore, an additional measure of well-being such as the mental health
13 continuum short-form (Lamers et al., 2011) and inclusion of a social desirability scale, would
14 have helped provide a more comprehensive validity assessment.

15 **Conclusion**

16 Overall, our study suggest that adolescents' well-being is predominantly represented by an
17 overarching strong general well-being factor, that coexists with relatively weak affective,
18 psychological and social relationship domains. As such, practitioners, researchers and public
19 health providers should foremost calculate a composite score of well-being in the knowledge
20 that it is explaining the vast majority of the WEMWBS item variance. However, given some
21 contribution of the residual factors to item and model variance, and sound subscale internal
22 reliability, further analysis may calculate subdomain scores with the aforesaid
23 recommendations - albeit strictly to supplement the composite score. The advantages of
24 adopting these recommendations is that following analysis of the composite well-being score,
25 domains (e.g., social relationships) relevant to the effect of mental health interventions (e.g.,

1 anti-bullying campaigns) could be explored. Limitations of this cross-sectional study are that
2 the test-retest reliability, and concurrent and divergent validity of the WEMWBS remain
3 unassessed in this population. Further longitudinal research study designs examining the
4 WEMWBS with additional instruments can provide a more comprehensive psychometric
5 assessment.

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14

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