Running head: DRIVE FOR MUSCULARITY SCALE

Factor Structure and Psychometric Properties of a Romanian Translation of the Drive for Muscularity Scale (DMS) in University Men

**Abstract**

We examined the psychometric properties of a Romanian translation of the 15-item Drive for Muscularity Scale (DMS). Male university students from Romania (*N* = 343) completed the DMS, as well as measures of self-esteem, body appreciation, and muscle discrepancy. Exploratory factor analysis indicated that DMS scores reduced to two factors that related to muscularity-oriented attitudes and behaviours, with both first-order factors loading onto a higher-order factor. However, confirmatory factor analysis indicated that a model with two first-order factors and a higher-order factor had poor fit. A two-factor model without a higher-order construct achieved acceptable but mediocre fit. Scores on the two-factor DMS model had adequate internal consistency and demonstrated acceptable convergent validity (significant correlations with self-esteem, body appreciation, and muscle discrepancy). These results provide support for a two-factor model of DMS scores in a Romanian-speaking sample and extends the availability of the DMS to a rarely-examined linguistic group.

**Keywords:** Drive for muscularity; Body image; Psychometrics; Romania; University men

**Introduction**

The ideal male physique in many socioeconomically developed settings is muscular and toned (Karazsia, Murnen, & Tylka, 2017; Leit, Pope, & Gray, 2001; Pope, Phillips, & Olivardia, 2000; Swami & Tovée, 2005) and a large proportion of men in these sites report a discrepancy between their current and desired levels of muscularity (e.g., Edwards, Tod, Morrison, & Molnar, 2014; Frederick et al., 2007; McCreary, 2007; Schneider, Rollitz, Voracek, & Henning-Fast, 2016). Findings such as these have led some scholars to propose that an important component of men’s body image is *drive for muscularity*, or a perception of having an underdeveloped musculature combined with a desire to increase muscle mass (McCreary, 2012; Morrison, Morrison, Hopkins, & Rowan, 2004). Higher drive for muscularity in men is reliably associated with a range of negative outcomes, including lower psychological well-being, higher rates of anabolic steroid and supplement use, and disordered eating (e.g., Lavender, Brown, & Murray, 2017; Parent, 2016; Parent & Bradstreet, 2017).

 A number of scales have been developed to assess the drive for muscularity construct in men (for reviews, see Cafri & Thompson, 2007; Tod, Morrison, & Edwards, 2012), but the most widely-used of these measures is McCreary and Sasse’s (2000) Drive for Muscularity Scale (DMS). The DMS is a self-report measure consisting of 15 items that are rated on a 6-point scale ranging from 1 (*Always*) to 6 (*Never*). Principal components analysis (PCA) with data from Canadian men suggested that DMS scores reduced to two first-order factors, termed Muscularity-Oriented Body Image Attitudes (7 items) and Muscularity-Oriented Behaviours (7 items; McCreary, Sasse, Saucier, & Dorsch, 2004). In this study, one item (#10: “I think about taking anabolic steroids”) was found to have very little variability and was omitted from the subscale calculations. In other studies, however, Item #10 has been found to load onto the Behaviours subscale (McPherson, McCarthy, McCreary, & McMillan, 2010). In addition, some studies have indicated that both subscales loaded onto a single higher-order DMS factor (McCreary et al., 2004; McPherson et al., 2010).

 McCreary and colleagues (2004) also reported that DMS subscale and total scores had adequate internal consistency coefficients (Attitudes α = .88, Behaviours α = .81). Reviews of the measure also highlight good test-retest reliability coefficients, as well as acceptable patterns of concurrent, convergent, and discriminant validity in English-speaking samples (McCreary, 2007; Tod et al., 2012). In addition, the two-factor structure of DMS scores has been supported through confirmatory factor analysis (CFA) in sexual minority men from the United States, with all items loading onto their respective factors (DeBlaere & Brewster, 2017). In contrast, however, recent research has questioned the factorial validity of DMS scores in Asian American men (Keum, Wong, DeBlaere, & Brewster, 2015). More specifically, these authors reported that the 15 DMS items provided poor fit to the data using CFA and that exploratory factor analysis (EFA) suggested the removal of three behavioural items (Items #4, 5, 10). A more recent study of Asian American men indicated through CFA that a fourth item (Item #12) should also be removed for the two-factor model of DMS to achieve good fit (Cheng, McDermott, Wong, & La, 2016).

 The results of studies with Asian American men suggest that the parent factor structure of DMS scores may not present good fit in some social identity groups (Cheng et al., 2016; Keum et al., 2015). However, a complementary body of work has examined the factorial validity of DMS scores in non-English-speaking samples and presents equivocal findings. For example, CFA studies have reported that the original two-factor model had good fit in Argentinian university students (Compte, Sepúlveda, de Pellegrin, & Blanco, 2015), Spanish adolescents (Sepúlveda, Parks, de Pellegrin, Anastasiadou, & Blanco, 2016), German weight-training men (Waldorf, Cordes, Vocks, & McCreary, 2014), and Italian heterosexual and gay men (Nerini, Matera, Baroni, & Stefanile, 2016). With the exception of Nerini and colleagues (2016), who did not include Item #10 in their analyses, all other studies have reported that Item #10 loads onto the Behavioural subscale. Likewise, EFA with principal-axis factoring with Malaysian Malay men indicated that DMS scores reduced to two dimensions mirroring the parent study, with Item #10 again loading onto the Behavioural subscale (Swami, Barron, Lau, & Jaafar, 2016).

 In contrast, a CFA of DMS scores in Brazilian men indicated that the parent two-factor model achieved poor fit (Campana, Gomes, Swami, & da Silva, 2013). Three items (Items #7, 9, 10) had high residuals and the two-factor model was reported to have good fit following the removal of these items. Campana and colleagues (2013) also tested a novel three-factor model consisting of Muscularity Concern, Muscularity Investment, and Ambiguity of Muscularity Investment subscales, but found that it had poorer fit compared to the modified two-factor model. In addition, Escoto and colleagues (2013) examined the factor structure of DMS scores in Mexican university students using EFA and reported that the Attitudes subscale mirrored its parent version, with all 7 items having adequate factor loadings. However, the Behaviour subscale reduced to two dimensions reflecting substance intake and training adherence. A CFA with a sample of Mexican men provided support for this revised three-factor model, although internal consistency coefficients for the Behavioural dimensions were less-than-adequate (Escoto et al., 2013; but see Escoto Ponce de León et al., 2018, who reported adequate internal consistency coefficients for all three factors in Mexican bodybuilders). Escoto and colleagues (2013) also reported, using CFA, that the parent two-factor model of DMS scores had good fit.

These translational studies have also indicated that the two DMS subscales have adequate internal consistencies and good patterns of convergent, concurrent, and discriminant validity (Campana et al., 2013; Compte et al., 2015; Escoto et al., 2013; Nerini et al., 2016; Sepúlveda et al., 2016; Swami et al., 2016; Waldorf et al., 2014). However, one equivocal issue relates to the fit of a higher-order dimension of drive for muscularity scores (that is, the extent to which the two first-order DMS factors adequately load onto a higher-order dimension). Most translational studies have not examined fit of this higher-order dimensionality (Campana et al., 2013; Compte et al., 2015; Escoto et al., 2013; Waldorf et al., 2014). Conversely, Nerini and colleagues (2016) reported that the higher-order dimensionality had good fit in Italian men, while Sepúlveda and colleagues (2016) found that a model that included the higher-order factor had poor fit in Spanish adolescents. Thus, examining the higher-order dimensionality of the DMS scores would be useful direction for future research.

**The Present Study**

 As a contribution to the literature reviewed above, we sought to examine the factor structure and psychometric properties of a Romanian (*limba română* or лимба ромынэ in Moldovan Cyrillic) translation of the DMS. Doing so is important for a number of reasons. First, there remains a dearth of research on body image in the Romanian context (Swami, Tudorel, Goian, Barron, & Vintila, 2017). This is notable because Romanian is spoken by around 24 million people as a first language, mainly in Romania and Moldova (where it has official status), as well as by several million more as a second language (European Commission, 2012). The availability of a translation of the DMS would, therefore, allow for more systematic investigations of the drive for muscularity construct in a population that has traditionally been neglected within the body image literature. Validation of a Romanian version of the DMS would also provide scholars with an appropriate tool to investigate the relationships between drive for muscularity and potential negative outcomes, such as poorer psychological well-being and negative health behaviours, in Romanian-speaking populations.

Second, Romania remains at an early stage of the nutrition transition (Popescu-Spineni, Glavce, David-Rus, Manuc, & Roville-Sausse, 2011), with ongoing changes to diet (e.g., increased intake of foods rich in carbohydrates and saturated fats) and dietary habits that are contributing to increasing rates of obesity (Ulijaszek & Koziel, 2007). Despite, or possibly concomitant to these trends, Romanian boys may desire bodily bigness in order to conform to traditional appearance ideals that celebrate male strength and masculinity (Mocanu, 2013). Indeed, despite recent social and political changes, it has been suggested that cultural norms and traditions have not changed at the same pace (Gavreliuc, 2012), with men expected to appear masculine through bodily expression, primarily in terms of muscularity and self-accomplishment (Mîndruţ, 2006). Recent studies have also suggested that a desire for greater muscularity may have been heightened in Eastern European men (e.g., Babusa, Czeglédi, Túry, Mayville, & Urbán, 2015), possibly as result of a “crisis of masculinity” taking root in a region experiencing social and economic transitions, as well as changing consumerist patterns following the fall of the Iron Curtain (Matlak, 2014).

Because it is difficult to know how these issues might impact on latent dimensionality of DMS scores in Romanian participants, we adopted a two-step strategy (Worthington & Whittaker, 2006) to examine the factor structure of a Romanian translation of the DMS. Based on the framework of classical test theory, we began by using EFA to examine the factor structure of Romanian DMS scores. This allowed us to explore latent dimensionality without any *a priori* limitations in terms of modelling. Next, for the purposes of cross-sample validation, we examined the fit of the model derived from EFA using confirmatory analytic methods. Although we acknowledge that other models have been proposed in the literature (e.g., Campana et al., 2013; Escoto et al., 2013), we also note that these models have either not been supported through CFA or suffer from psychometric limitations (e.g., poor internal consistency coefficients). Thus, we did not examine the fit of these alternative models in the present work. Moreover, the present two-step strategy allows us to determine the most appropriate DMS model for use in the present sample and eliminates the likelihood of spurious model testing (Worthington & Whittaker, 2006). As a preliminary hypothesis based on previous studies (e.g., McCreary et al., 2004), we hypothesised that EFA would show that Romanian DMS scores reduce a two-factor structure, with a single higher-order drive for muscularity factor. In addition, we expected that both a two-factor model with a higher-order drive for muscularity factor, as well as a two-factor model without the higher-order dimension, would show good fit through CFA. Furthermore, based on previous translation studies (Compte et al., 2015; Sepúlveda et al., 2016; Waldorf et al., 2014), we expected that Item #10 would load onto a Behavioural dimension of DMS scores.

 In addition to examining factorial validity, we also conducted an assessment of the convergent validity of DMS scores. More specifically, we examined associations between DMS scores and self-esteem, current-ideal muscle discrepancy, and body appreciation. Selection of these indices of validity was based on the availability of validated measures in Romanian (i.e., self-esteem and body appreciation) or measures that required minimal translation (current-ideal muscle discrepancy). Further, they were also based on theoretical considerations: muscle discrepancy offers a *prima facie* index of muscularity dissatisfaction that should be associated with greater drive for muscularity (e.g., Nerini et al., 2016; Swami et al., 2016), whereas self-esteem has been shown to be negatively associated with drive for muscularity in previous research (e.g., Bergeron & Tylka, 2007). To demonstrate acceptable convergent validity, we expected that DMS scores would be significantly and negatively associated with self-esteem, and body appreciation, and positively associated with current-ideal muscle discrepancy. Finally, we also assessed reliability of derived DMS scores in terms of internal consistency, with the expectation that subscale scores would demonstrate adequate reliability.

**Method**

**Participants**

The participants of this study were 343 male students of Romanian ethnicity from a university in Timișoara, the capital of Timiș County in the west of Romania. Timișoara is the third most populous city in Romania and an important economic and sociocultural hub in the region. Participants ranged in age from 18 to 58 years (*M* = 22.48, *SD* = 6.02) and in self-reported body mass index (BMI) from 16.32 kg/m2 to 35.27 kg/m2 (*M* = 23.66, *SD* = 3.58). Most participants were enrolled on undergraduate programmes (66.2%), with 17.5% enrolled on Masters courses and the remainder on some other programme of study. No further demographic information was collected.

**Measures**

**Drive for muscularity**. Participants were asked to complete a Romanian translation of the 15-item DMS (McCreary & Sasse, 2000). All items were rated on a 6-point scale ranging from 1 (*Always*) to 6 (*Never*) and were reverse-coded prior to analysis so that higher scores reflect greater drive for muscularity.

**Body appreciation.** The survey package included the Body Appreciate Scale-2 (Tylka & Wood-Barcalow, 2015; Romanian translation: Swami et al., 2017), a 10-item measure of positive body image (sample item: “I respect my body”). Items on the BAS-2 were rated on a 5-point scale, ranging from 1 (*Never*) to 5 (*Always*). EFA and CFA indicated that the factor structure of BAS-2 scores was one-dimensional in Romanian men (Swami et al., 2017). Scores on the translation of the scale have also been reported to have adequate internal consistency, adequate test-retest reliability up to three weeks, and acceptable patterns of convergent validity in Romanian men (Swami et al., 2017). In the present study, we computed an overall body appreciation score by taking the mean of all 10 items, such that higher scores reflect greater body appreciation. Cronbach’s α for this measure in the present study was .91 (CI = .89-.92).

**Self-esteem**. Self-esteem was measured using the Rosenberg Self-Esteem Scale (RSES; Rosenberg, 1965; Romanian translation: Schmitt & Allik, 2005), a 10-item measure of global self-evaluations of worth as a human being. All items were rated on a 4-point scale (1 = *Definitely disagree*, 4 = *Definitely agree*). The Romanian version of the RSES has a one-dimensional factor structure (Schmitt & Allik, 2005) and acceptable patterns of convergent validity (e.g., Sava, Maricuțoiu, Rusu, Macsinga, & Vîrgā, 2011). Here, an overall score was computed as the mean of all 10 items, following reverse-coding of 5 items (higher scores on reflect greater self-esteem). In the present study, Cronbach’s α for this scale was .87 (CI = .85-.89).

 **Muscle discrepancy**. To measure current-ideal muscle discrepancy, we used the Muscle Silhouette Measure (MSM; Frederick et al., 2007). The MSM is a figural rating scale that includes 8 line-drawings of the male form that increase linearly in muscularity. Participants were asked to rate the figure that they felt best represented their current body and the figure that best represented their ideal muscularity. All ratings were made on an 8-point scale ranging from 1 (*Least muscular figure*) to 8 *(Most muscular figure*). A measure of muscle discrepancy was computed as the difference between absolute (unsigned) current and ideal ratings1, so that higher scores reflect greater muscle discrepancy. Frederick et al. (2007) reported that the MSM has acceptable construct validity.

**Demographics**. Participants provided their demographic details consisting of age, degree programme, height, and weight. The latter two items were used to compute self-reported BMI as kg/m2 (used for sample descriptive purposes).

**Procedures**

Once ethics approval for the study was obtained from the relevant university ethics committee, we translated the DMS and MSM from English into Romanian using the parallel back-translation procedure (Brislin, 1986). Specifically, a bilingual individual unaffiliated with the study translated the scale from English to Romanian, while a second individual translated this version back into English. Next, the items obtained were assessed by a committee consisting of the individuals who participated in the translation process, the second to fourth authors, and two psychology professors. There were no discrepancies in the translation of the MSM and very minor discrepancies (e.g., word choice differences) in the translation of the DMS were resolved through consensus. The items of the DMS in English and Romanian are presented in Table 1.

Data collection took place between September and December 2017. Participation in the study was solicited through flyers and posters, which advertised a study on men’s health, placed in various campus locations. Inclusion criteria included being of the age of majority and of Romanian ethnicity. Potential participants were provided with an information sheet that contained brief information about the study (e.g., task requirements and estimated duration). Those who agreed to participate provided written informed consent and completed a paper-and-pencil questionnaire in a private cubicle. The questionnaire took between 8-12 minutes to complete. The order of presentation of the scales described above was pre-randomised for each participant. All participation was voluntary and participants did not receive any remuneration. Upon return of completed questionnaires, participants were provided with written debrief information.

**Statistical Analyses**

Missing data accounted for < 0.2% of the total dataset and were missing completely at random (MCAR), as determined by Little’s (1988) MCAR analysis. We, therefore, imputed missing values using pooled estimates from multiple imputations. To examine the factor structure of DMS scores, we employed a two-step analytic strategy consisting of EFA and CFA. To accommodate this strategy, we split the total dataset in two by randomly allocating 150 men to a first subsample and including the remaining 193 men in a second subsample. The two subsamples did not differ significantly in terms of age, *t*(341) = 0.95, *p* = .344, *d* = 0.10, or BMI, *t*(341) = 0.60, *p* = .548, *d* = 0.06. In the first subsample, we assessed the factor structure of DMS scores using EFA with principal-axis factoring in IBM SPSS Statistics v.20. Based on item distribution, average correlation with other items, and item-total correlations (Clark & Watson, 1995), these data were suitable for factor analysis. The subsample size satisfied Tabachnick and Fidell’s (2013) recommendation that sample sizes for EFA meet a 10:1 item-to-participant ratio. We used a promax rotation because we expected latent factors that would be inter-correlated (Fabrigar, Wegener, MacCallum, & Strahan, 1999). To determine the number of factors to be extracted, we used parallel analysis (Hayton, Allen, & Scarpello, 2004). Monte Carlo analyses have shown that, of all the criteria for deciding on factor-extraction, parallel analysis is the most accurate and does not over-factor (e.g., Velicer, Eaton, & Fava, 2000). Parallel analysis works by creating a random dataset with the same number of cases and variables as the actual dataset. Factors in the actual data are only retained if their eigenvalues are greater than the eigenvalues from the random data (Brown, 2006; Fabrigar et al., 1999). Factor loadings were interpreted using Tabachnick and Fidell’s (2013) recommendations, with loadings of .71 and above considered excellent, .63-.70 considered very good, .55-.62 considered good, .33-.54 considered fair, and .32 or lower considered poor.

Data from the second subsample were subjected to CFA using Analysis of Moment Structures (AMOS v.23). Skewness and kurtosis were below critical limits for CFA (skewness < |3|, kurtosis < |10|; Weston & Gore, 2006). The standard maximum likelihood method, which assumes multivariate normality, and the robust estimation method were therefore applied to test for fit. Our sample size met Muthén and Muthén’s (2002) recommendation that, with normally distributed indicator variables, *N* should be at least 150. Hypothesised modelling was based on the results of the EFA in the first subsample as well as the proposed higher-order model from the parent study. To assess the fit of the measurement, absolute and incremental fit indices were selected a priori , namely the normed model chi-square (χ²normed; a goodness-of-fit test that minimises the impact of sample size), the Steiger-Lind root mean square error of approximation (RMSEA; provides a correction for model complexity), the standardised root mean square residual (SRMR; assesses the mean absolute correlation residual), and the comparative fit index (CFI; measures the proportionate improvement in fit by comparing a target model with a more restricted, nested baseline model). Values ≤ 3.00 for χ²normed, and close to .06 for RMSEA, .08 for SRMR, and .95 for CFI indicate good fit of the model to the data (Hu & Bentler, 1999). However, these cut-off values should not be interpreted rigidly (Heene, Hilbert, Draxler, Ziegler, & Bühner, 2011; Perry, Nicholls, Clough, & Crust, 2015) and values between 3.01-5.00 for χ²normed, between .08 to .10 for RMSEA, .09 and .10 for SRMR, and between .87 and .95 for CFI can indicate acceptable but mediocre fit to the data (Hooper, Couglan, & Mullen, 2008; MacCallum, Browne, & Sugawara, 1996). We also report the Parsimony Goodness-of-Fit Index (PGFI), an adjustment to goodness-of-fit that penalises models that are less parsimonious. There is some debate as to whether parsimony adjustments are useful (Marsh & Hau, 1996), but when used in combination with other fit indices, they assist researchers in making decisions about models *vis-á-vis* parsimony, while not penalising models for having more parameters (Mulaik et al., 1989). No thresholds have been recommended for PGFI, but Mulaik and colleagues (1989) suggest that values should be in the region of .50-.90.

To assess reliability, we computed Cronbach’s α in each subsample. Nunnally (1976) recommended that, for basic research tools, Cronbach’s α should be at least .80 to be considered adequate. Finally, we examined convergent validity by computing bivariate correlations between DMS scores and scores on all other included measures using the total sample.

**Results**

**Exploratory Factor Analysis**

Bartlett’s test of sphericity, χ2(105) = 1559.19, *p* < .001, and the Kaiser-Meyer-Olkin (KMO) measure of sampling adequacy, KMO = .88, showed that the items of the DMS had adequate common variance for factor analysis (Tabachnik & Fidell, 2013). The results of the EFA revealed three factors with eigenvalues > 1.0 and inspection of the scree plot suggested that there were two primary factors, with a drop-off to the third factor. The results of parallel analysis also suggested that only two factors should be extracted: the first two factors from the actual data had an eigenvalue greater than the criterion eigenvalue generated from the random data (i.e., 6.72 [actual data] compared to 3.65 [random data] for the first factor, 2.90 [actual data] compared to 1.98 [random data] for the second factor), whereas the third factor had an eigenvalue that was lower than the corresponding criterion eigenvalue generated from the random data (i.e., 1.05 [actual data] compared to 1.39 [random data]. Based on these results, we extracted two factors, which explained 64.1% of the common variance (44.8% and 19.3%, respectively). Additional support for the extraction of two, rather than three, factors comes from the fact that the third factor explained less than 5% of the variance, which Guadagnoli and Velicer (1988) recommended as a cut-off for retaining a factor as interpretable.

 Factor loadings and item-level descriptive statistics are reported in Table 1. The first extracted factor included 7 items that mirror the items included in the Muscularity-Oriented Behaviours factor of the parent scale, plus Item #10. All but one factor loadings on this factor were good-to-excellent and the factor score had adequate internal consistency (Cronbach’s α = .82, CI = .76-.89). The second factor included 7 items that were included in Muscularity-Oriented Body Image Attitudes in the parent study. All items in this factor had good-to-excellent factor loadings and the factor score had adequate internal consistency with all items having very good to excellent loadings. This second factor had adequate internal consistency (Cronbach’s α = .83, CI = .77-.90). To test whether the two factors loaded onto a higher-order drive for muscularity factor, we computed a second EFA with principal-axis factoring using a quartimax rotation (because of the expectation of a single factor). The results indicated that both factors did load onto a higher-order structure (eigenvalue = 1.30, 64.83% of variance explained, factor loadings = .81 and .80).

**Confirmatory Factor Analysis**

We first investigated fit of the model where all items loaded onto two latent variables and fed into a global higher-order latent factor. Fit indices indicated poor fit of the model to the data, χ²(90, *N* = 193) = 730.587, χ²normed = 8.118, CFI = .584, RMSEA = .193 with 90% CI = .180-.206, SRMR = .150, PGFI = .447. Since the fit indices were poor, modification indices were taken into account to improve the model. Schumacker and Lomax (2004) recommended examination of modification indices to locate potential areas of misspecification and model improvement based on substantive meaning. In addition, conservative convention was followed in judging modification indices to have a significant effect on χ² when > 5.00 (Bryne, Shavelson, & Muthén, 1989). To ensure that the fit of the model was not due to chance alone, and to maximise cross-validity of the model (Steenkamp & Baumgartner, 1998), the number of covaried error terms was kept to a maximum of three. Modification indices were consulted to free error covariances on both factors: Items #3 and #4 (137.084), #8 and #10 (60.796) from the Behaviour factor, and Items #13 and #14 (30.848) from the Attitudes factor. However, fit indices values for this revised model were still found to be poor, χ²(87, *N* = 193) = 400.201, χ²normed = 4.600, CFI = .797, RMSEA = 137 with 90% CI = .124-.151, SRMR = .133, PGFI = .531.

Next, we investigated a two-factor solution without the higher-order latent factor. Fit indices values were again found to be poor, χ²(89, *N* = 193) = 367.255, χ²normed = 4.126, CFI = .819, RMSEA = .128 with 90% CI = .114-.141, SRMR = .098, PGFI = .573. Inspection of the modification indices suggested allowing the error variances for Items #8 and #10 (41.438) to correlate; although these items appear conceptually dissimilar, they load onto the same first-order factor. Inspection of the modification indices also suggested allowing error variances for Items #2 and #6 (23.570; both related to weight-training), and Items #13 and #14 (24.723; related to perceiving one’s chest and arms to be insufficiently muscular) to correlate. Following the covariance of these residuals, model fit was improved and reached acceptable but mediocre fit, χ²(86, *N* = 193) = 269.049, χ²normed = 3.128, CFI = .881, RMSEA = .100 with 90% CI = .091-.120, SRMR = .095, PGFI = .595, although some indices (i.e., RMSEA and SRMR) were at, or close to, the limit for acceptable but mediocre fit. The standardised factor loadings were all good-to-excellent, with values ranging from .40 to .93 (see Figure 1). Therefore, from these data, a two-factor structure of DMS scores was found to have acceptable mediocre fit, but without a higher-order factor. The internal consistency coefficients for the DMS subscale scores were adequate: Attitudes α = .80 (CI = .72-.86) and Behavioural α = .84 (CI = .78-.90).

**Convergent Validity**

We computed DMS factor scores in the total sample by computing the means of items associated with each factor. Next, using the total sample, we examined the convergent validity of these factor scores by examining bivariate correlations with self-esteem, body appreciation, and current-ideal muscle discrepancy. As reported in Table 2, the two DMS factors were significantly but weakly correlated with one another. In addition, both subscale scores were significantly correlated with self-esteem, body appreciation, and muscle discrepancy in the hypothesised directions. These significant correlations were weak-to-moderate in strength.

**Discussion**

 In the present study, we examined the psychometric properties of a Romanian translation of the DMS. In broad outline, our results are consistent with both the parent studies (McCreary & Sasse, 2000; McCreary et al., 2004), as well as most previous translational studies (Compte et al., 2015; Nerini et al., 2016; Sepúlveda et al., 2016; Swami et al., 2016; Waldorf et al., 2014), in suggesting that Romanian DMS scores consist of two first-order factors that distinguish between muscularity-oriented body image attitudes and behaviours. Furthermore, across both our EFA and CFA, Item #10 was found to adequately load onto the Behaviours subscale, which is consistent with some previous studies (e.g., Compte et al., 2015; McPherson et al., 2010; Sepúlveda et al., 2016; Swami et al., 2016).

 However, several analytic issues are worth highlighting. First, it should be noted that, to attain acceptable but mediocre fit in our CFA analysis, we had to allow the error variances for three pairs of items to correlate. Second, our findings in relation to the higher-order dimensionality of DMS scores were equivocal. Using EFA, we found that the two first-order DMS factors loaded onto a higher-order drive for muscularity factor; however, our CFA indicated that a model that incorporated the higher-order factor failed to achieve good fit, even following modifications. Third, fit of the two-factor model of Romanian scores was mediocre but not good. It is possible that all of these issues reflect some of the limitations of CFA. Marsh and colleagues (2009, 2013), for example, have suggested that the independent clusters model used in CFA, which requires each indicator to load onto only one factor, may be too restrictive compared with EFA methods, where cross-loadings are freely estimated. Indeed, Marsh, Morin, Parker, and Kaur (2014) highlighted the fact that many psychological measures have well-defined structures based on EFA, yet do not evidence good fit through CFA. The Romanian DMS may be one such measure: here, our EFA quite clearly indicated that scores reduced to two dimensions, whereas the results of our CFA were more equivocal.

Given these issues, our advice for scholars wishing to use the Romanian DMS is that, where possible and theoretically meaningful, they should perhaps use the first-order factor scores. For those wishing to use total DMS scores, we recommend first examining whether subscale scores load onto a higher-order factor in their samples. In fact, this may be good advice for any researchers wishing to use total DMS scores, irrespective of the linguistic or cultural site of the study. Alternatively, one useful way of advancing knowledge in this area would be through the application of exploratory structural equation modelling (ESEM; Asparouhov & Muthén, 2009; Marsh et al., 2009), which integrates EFA, CFA, and structural equation modelling. As Marsh and colleagues (2014) have noted, ESEM estimates of factor correlations are typically more accurate than CFA estimates. Furthermore, although ESEM is primarily a confirmatory tool, it can be used with care as an exploratory analytic method in a manner that has many advantages over EFA and CFA. In short, (re-)examining the factor structure of the Romanian DMS through ESEM may be a useful direction for future research and may help resolve some of the discrepancies between our EFA and CFA results.

 Beyond factorial validity, the two DMS factor scores were found to have adequate internal consistency coefficients by Nunnally’s (1976) standards for basic research tools. Our data also provide preliminary evidence for the convergent validity of DMS subscale scores, insofar as both subscales were significantly correlated with scores of self-esteem, body appreciation, and muscle discrepancy. Although the strength of the relationships was generally small-to-moderate, they were in line with those reported in a previous study that established convergent validity of DMS subscale scores using similar measures (Swami et al., 2016). Future research could improve on the present work by gathering further evidence of the construct validity of Romanian DMS scores. This could be established, for example, by examining relationships with internalisation of an athletic ideal of appearance (Schaefer, Harriger, Heinberg, Soderberg, & Thompson, 2017), personality traits (see Benford & Swami, 2014), social dominance orientation (see Swami et al., 2013), and other measures of drive for muscularity (see Tod et al., 2013).

 Our results should be considered in light of a number of additional limiting issues. First, the present study relied on a sample of university students and the generalisability of our findings to community samples of men is as yet unknown. In future work, it would also be useful to examine the fit of the two-factor model of DMS scores in other Romanian-speaking populations, such as in Moldova. An additional limitation of the present study is the fact that we did not examine test-retest reliability of DMS scores, which will be an important task for future research. Additional support for the construct validity of DMS scores could also be established through examination of latent scores in weight-training or bodybuilding men. Finally, we did not examine the factor structure of Romanian DMS scores in women, although it should be noted that a measure specifically examining muscularity concerns in women has recently been developed (i.e., the Female Muscularity Scale; Rodgers et al., 2018).

 These limitations aside, the present results suggest that scores on a Romanian version of the DMS reduces to two factors tapping muscularity-oriented attitudes and behaviours in Romanian university men. A higher-order drive for muscularity construct was supported through EFA, but not CFA. These results are important in their own right because they provide a useful tool for examining men’s body image in Romanian-speaking samples. More broadly, the availability of the DMS in multiple languages (i.e., English, Romanian, Spanish, Italian, Portuguese, and Malay) offers opportunities for cross-cultural comparisons of drive for muscularity scores, although it will be important to first establish that DMS scores are invariant across linguistic groups. Doing so would help scholars better understand the extent to which drive for muscularity should be considered a public health concern across cultural and national groups, and would help practitioners develop appropriate intervention techniques to reduce drive for muscularity (e.g., Jankowski et al., 2017).

**Footnotes**

1 Only 3.2% of the total sample wanted to be less muscular, whereas 16.3% wanted no change, and the remainder wanted to be more muscular.

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Table 1. Factor loadings for the Malay Drive for Muscularity Scale with values in bold representing items that loaded onto a given factor) and items in italics representing the Romanian translations.

|  |  |  |  |  |
| --- | --- | --- | --- | --- |
| Items | *M* | *SD* | Factor 1 | Factor 2 |
| 3. I use protein or energy supplements / *Folosesc suplimente proteine sau suplimente energizante.* | 2.37 | 1.73 | **.89** | .30 |
| 4. I drink weight gain or protein shakes / *Beau bauturi sportive sau shake-uri proteice care ma ajuta sa cresc in masa musculara.* | 2.34 | 1.73 | **.89** | .31 |
| 8. Other people think I work out with weights too often / *Ceilalti cred despre mine ca ma antrenez cu greutati prea des.* | 2.09 | 1.60 | **.82** | .24 |
| 12. I think that my weight training schedule interferes with other aspects of my life / *Consider ca antrenamentele mele cu greutati interfereaza cu alte aspecte ale vietii mele.* | 4.55 | 1.72 | **.81** | .28 |
| 10. I think about taking anabolic steroids / *Iau in considare consumul de steroizi anabolizanti.* | 1.86 | 1.61 | **.80** | .15 |
| 6. I feel guilty if I miss a weight training session / *Ma simt vinovat daca lipsesc de la un antrenament cu greutati.* | 2.31 | 1.67 | **.72** | .27 |
| 5. I try to consume as many calories as I can in a day / *Incerc sa consum cat mai multe calorii pot intr-o zi.* | 2.59 | 1.50 | **.69** | .17 |
| 2. I lift weights to build up muscle / *Ma antrenez cu greutati pentru a-mi construi masa musculara.* | 2.99 | 1.66 | **.49** | .27 |
| 13. I think that my arms are not muscular enough / *Consider ca musculatura bratelor mele nu este suficient de dezvoltata.* | 3.93 | 1.62 | .22 | **.84** |
| 11. I think that I would feel stronger if I gained a little more muscle mass / *Consider ca m-as simti mai puternic daca mi-as mai dezvolta putin masa musculara.* | 3.72 | 1.62 | .22 | **.82** |
| 9. I think that I would look better if I gained 10 pounds in bulk / *Consider ca as arata mai bine daca as creste 5 kg in masa musculara.*  | 3.26 | 1.80 | .13 | **.82** |
| 7. I think I would feel more confident if I had more muscle mass / *Cred ca m-as simti mai sigur pe mine daca as avea o masa musculara mai dezvoltata.* | 3.23 | 1.59 | .29 | **.82** |
| 14. I think that my chest is not muscular enough / *Consider ca musculatura pieptului meu nu este suficient de dezvoltata.*  | 3.85 | 1.62 | .26 | **.78** |
| 1. I wish that I were more muscular / *Mi-as dori am masa musculara mai mare.* | 3.57 | 1.63 | .05 | **.70** |
| 15. I think that my legs are not muscular enough / *Consider ca musculatura picioarelor mele nu este suficient de dezvoltata.* | 4.15 | 1.73 | .15 | **.64** |

*Note*. Values in bold indicate that an item loaded onto the corresponding factor.

Table 2. *Descriptive Statistics and Inter-Scale Correlations between Drive for Muscularity Scale Factor Scores and All Remaining Variables.*

|  |  |  |  |  |  |
| --- | --- | --- | --- | --- | --- |
|  | (1) | (2) | (3) | (4) | (5) |
| (1) DMS-Attitudes |  | .30\*\* | -.12\* | -.18\* | .32\*\* |
| (2) DMS-Behavioural |  |  | -.18\* | -.22\*\* | .38\*\* |
| (3) Self-esteem |  |  |  | .49\*\* | -.05 |
| (4) Body appreciation |  |  |  |  | -.11\* |
| (5) Muscle discrepancy |  |  |  |  |  |
| *M* | 2.54 | 3.68 | 3.12 | 3.90 | 1.72 |
| *SD* | 0.90 | 0.53 | 0.63 | 0.75 | 1.20 |

*Note.* DMS-Attitudes = Muscularity-Oriented Body Image Attitudes, DMS-Behaviours = Muscularity-Oriented Behaviours. \* *p* < .05, \*\* *p* < .001.



*Figure 1.* Path diagram and estimates for the two-dimensional model of the Drive for Muscularity Scale. The large ovals are the latent constructs, with the rectangles representing measured variables and the small circles with numbers representing the unstandardized residual variables (variances). The path factor loadings are unstandardised, and standardised in parenthesis, with significance levels determined by critical ratios (all *p* < .001).