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Evaluating the psychometric properties of the Social Identity Questionnaire for Sport (SIQS)

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ABSTRACT

Objectives: The strength with which individuals identify with a sport team (i.e., social identity) has important implications for athletes' cognitions, affect, and behavior. Yet there remains ambiguity surrounding the optimal way to conceptualize and thus assess social identity. The purpose of the current study was to examine the psychometric properties of a nine-item version of the Social Identity Questionnaire for Sport (SIQS).
Design: Athletes completed a self-report measure of social identity related to sport team involvement.
Method: In a sample of 869 youth and young adult athletes (M_{age} = 14.84, *SD* = 3.79; male = 375; female = 493), we evaluated the psychometric properties of a nine-item version of the Social Identity Questionnaire for Sport (SIQS) by using bifactor analysis and subsequently testing single-factor and three-factor structures.
Results: Overall, a theoretically conceived three-factor structure is empirically supported, where ingroup ties, cognitive centrality, and ingroup affect are represented as distinct dimensions of social identity. Empirical support was also found for a global factor of social identity, but only when the residuals among the subscales were correlated. There was support for strong measurement invariance across sexes for the unidimensional and three-factor structure models.
Discussion: The findings from the study support the SIQS as a psychometrically sound measure of social identity in sport that can be used to either model social identity along three specific dimensions *or* as a global construct.

1. Introduction

Social identity refers to “that part of an individual's self-concept which derives from his/her knowledge of his/her membership of a social group (or groups) together with the value and emotional significance attached to that membership” (Tajfel, 1981, p. 255). The definition put forward by Tajfel provided a conceptual foundation for theory (Social Identity Theory; Tajfel, 1978; Tajfel & Turner, 1979) and research across several domains including psychology and recently in sport. Recent calls have drawn attention to social identity as a key construct for understanding group behavior and interpersonal relations (e.g., Hornsey, 2008). The strength with which individuals identify with a sport team (i.e., social identity) has important implications for athletes' cognitions, affect, and behavior (Bruner, Dunlop, & Beauchamp, 2014; Rees, Haslam, Coffee, & Lavallee, 2015). For example, stronger perceptions of social identity are positively associated with team outcomes such as team performance (Murrell & Gaertner, 1992) as well as individual outcomes including initiative, self-worth, commitment, perceived effort, and personal and social skills (Bruner, Balish, et al., 2017; Martin, Balderson, Hawkins, Wilson, & Bruner, 2017). This aligns with theoretical accounts that highlight how the identity associated with a sport context can influence self-evaluation, and motivates individual behaviors toward ingroup and outgroup members (Rees et al., 2015). Despite the influential role of social identity, issues concerning the conceptualization and measure of the construct have arisen in a number of domains, including sport. Notably, social identity has been conceptualized and operationalized in several ways (Bruner, Dunlop et al., 2014).

Traditionally, social identity has been conceptualized as one global construct (e.g., Terry, Hogg, & White, 1999). Example studies include investigations highlighting a positive link between team performance (i.e., winning) and team identification (Murrell & Gaertner, 1992), social identity as a mediator of the positive link between athletes' perceptions of coach-related procedural justice and cohesion (De Backer et al., 2011), and athletes' discourses related to their social identities following different performance outcomes in soccer (Zucchermaglio, 2005). A second approach to conceptualizing social identity is a multidimensional perspective. Although the multidimensionality of social

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identity in sport is supported by cogent theoretical and empirical support in other domains (Brown, Condor, Mathews, Wade, & Williams, 1986; Jackson, 2002; see Leach et al., 2008 for a review), only recently has this conceptualization gained attention in relation to sport team involvement. Notably, Bruner, Boardley et al. (2014) adapted a model and measure of social identity in sport based on the three-factor structure developed in social psychology by Cameron (2004). This conceptualization includes (a) ingroup ties – perceptions of similarity, bonding, and belongingness with other group members; (b) cognitive centrality – the importance of being a group member; and (c) ingroup affect – the positive feelings associated with group membership (Cameron, 2004). Based on these definitions, two of the dimensions are cognitive in nature (ingroup ties, cognitive centrality), whereas ingroup affect operates on an affective level. Further support for Cameron's (2004) multidimensional conceptualization has been reported in the sport psychology literature as specific dimensions of social identity differentially relate to team and athlete outcomes. As examples, ingroup ties positively predicted personal skills, social skills, and initiative (Bruner, Balish et al., 2017), ingroup affect linked with cohesion and prosocial teammate behavior (Bruner, Boardley et al., 2014), and cognitive centrality moderated the relationship between group norms and personal behavior (Benson, Bruner, & Eys, 2017).

Given the distinct ways social identity has been conceptualized, it is perhaps not surprising that there is a lack of consensus as to how social identity should be measured. In some instances when researchers perceived social identity as a unidimensional construct, social identity has been evaluated using a range of measures, including a rating scale from 1 to 5 for 30 statements (Murrell & Gaertner, 1992). Alternatively, De Backer et al. (2011) assessed social identity using the three highest loading items from a team identification scale (Boen, Vanbeselaere, Pandelaere, Schutters, & Rowe, 2008). Other social identity sport researchers have adapted multidimensional measures of social identity from the social psychology literature and other group settings (e.g., Amiot, Sansfaçon, & Louis, 2013). Most recently, however, several researchers have adapted Cameron's (2004) three dimensional measure to the domain of sport (Bruner, Boardley et al., 2014; Bruner, Eys, Evans, & Wilson, 2015; Martin et al., 2017). To our knowledge, however, the psychometric integrity of the questionnaire has not been rigorously evaluated. As noted by Carron, Eys, and Martin (2012) to advance the field of sport and exercise psychology, "measurement protocols are at the heart of science; they are fundamental to the advancement of knowledge" (p. 411). Given the strong theoretical basis driving the study of social identity in sport, it is equally important to ensure contemporary measures of social identity are psychometrically sound in the sport setting.

Thus, the current study sought to evaluate the psychometric properties of a questionnaire to assess social identity strength in sport. We first evaluated current approaches to assessing social identity in sport by examining the questionnaire items used in recent work. Using a decision-making framework informed by conceptual and empirical considerations, we then proceeded to evaluate a positively worded version of the Social Identity Questionnaire for Sport (SIQS). To accomplish this aim, we first applied bifactor modeling to assess the psychometric properties of the SIQS. Informed by this procedure, we then tested theoretically informed measurement models based on unidimensional (i.e., global construct of social identity) and three-factor (i.e., ingroup ties, ingroup affect, and cognitive centrality) structures. Finally, we tested the measurement invariance of models determined to be suitable.

2. Method

2.1. Conceptualization of the SIQS

In the current study, we evaluated an adapted version of Cameron's (2004) multidimensional measure of social identity and thus it is important to highlight this formative piece. Briefly, Cameron generated items to reflect three underlying dimensions (i.e., ingroup ties, cognitive centrality, ingroup affect) and conducted five programmatic studies to develop a 12-item measure. This work provides valuable insight into the conceptual basis and utility of a multidimensional view of social identity, while also highlighting the value of assessing social identity in relation to sport team involvement. As seen in the development of other questionnaires in the domain of sport (Benson & Eys, 2017; Eys, Loughead, Bray, & Carron, 2009), it is important to develop a descriptive understanding of a construct prior to attempting to measure the intensity or frequency of its occurrence. For example, in some instances, young athletes have viewed phenomenon in less complex ways (cf. Eys et al., 2009; Martin, Carron, Eys, & Loughead, 2013). Nonetheless, several qualitative studies report that athletes—including youth athletes—are able to differentiate between the various dimensions of social identity when relaying their personal sport-related experiences (Bruner, Boardley, Allen et al., 2017, Bruner, Boardley, Forrest et al., 2017). Overall, these studies lend support to the multidimensional conceptualization of social identity in sport.

2.2. Questionnaire development

The origin of the items evaluated in the current study are derived from Bruner, Boardley et al.'s (2014) adaptation of Cameron's (2004) measure to assess social identity related to sport team involvement. This measure included 12-items assessing three dimensions (i.e., ingroup ties, cognitive centrality, ingroup affect) of social identity with eight positively and four negatively (i.e., reverse-coded) worded items. Although using positively and negatively worded items may help in the detection of response acquiescence, mixed item wording might be problematic in terms of subscale reliability (Eys, Carron, Bray, & Brawley, 2007). Indeed, certain subscales from the 12-item version of the questionnaire have exhibited reliability issues (e.g., cognitive centrality, Bruner, Boardley et al., 2014). In response to the low alpha coefficients, the four negatively worded items were modified to positively worded items in subsequent work (Benson et al., 2017; Bruner, Balish et al., 2017; Bruner, Boardley, Benson et al., 2017). In the current paper, we combined these three datasets to scrutinize the psychometric properties of the SIQS. Although each of these datasets contained 12 positively worded items, we made a decision to retain only nine items due to several considerations. First, we identified two potentially problematic items based on item wording. Specifically, the ingroup ties item "I have a lot in common with other members in this team" appeared to capture similarity with teammates more than perceived connections with teammates. The ingroup affect item "I rarely regret that I am a member of this team" is potentially confusing because it combines two negative terms (rarely, regret). We also excluded a cognitive centrality item ("I often think about the fact that I am a team member") based on its similarity with another subscale item ("The fact that I am a member of this team often enters my mind"). Second, we examined the factorial structure of the 12 items and elected to retain the more concise nine item version.[1] Please see Appendix A for the nine items evaluated in the current study for the brief version of the SIQS as well as the three items removed.

2.3. Participants

Based on a re-analysis of three existing data-sets (Benson et al., 2017; Bruner, Balish, et al., 2017; Bruner, Boardley, Benson et al., 2017), there were 869 ($M_{\text{age}} = 14.84$, $SD = 3.79$) responses to the SIQS from a heterogeneous group of athletes (males = 375, females = 493, unidentified sex = 1). Represented team sports include ice hockey, soccer, basketball, flag football, volleyball, and baseball from a range of recreational and competitive levels.

[1] Please see the supplemental file for the factor loadings and fit indices of the 12 items.

2.4. SIQS

A nine-item, positively worded version of the SIQS was evaluated (see Appendix A). Based on how athletes felt about their current team, participants were asked to indicate the extent to which they agreed with a series of statements, on a scale ranging from 1 (*strongly disagree*) to 7 (*strongly agree*).

2.5. Analyses

As described by Rodriguez, Reise, and Haviland (2016a), "a bifactor measurement model specifies that for a given set of item responses, correlations among items can be accounted for by: (a) a general factor representing shared variance among all the items and (b) a set of group factors where variance over and above the general factor is shared among subsets of items presumed to be highly similar in content" (p. 137). Thus, modeling data with a bifactor structure enables researchers to directly evaluate the extent to which a general factor represents a well-defined construct when the variance attributable to the specific factors is assumed to be orthogonal. Bonifay, Lane, and Reise (2017) identified two distinct approaches to applying bifactor analysis. Consistent with how we applied bifactor analysis in the current research, one approach is using bifactor analysis as an evaluative tool for investigating the psychometric properties of a questionnaire. In this regard, bifactor analysis is an effective investigative tool to inform decisions regarding (a) how well a set of items reflects a latent dimension and (b) whether subscales provide unique information beyond a general factor. The second is using bifactor analysis to represent the structure of a psychological construct. Bonifay et al. (2017) cautioned against this latter approach because of the ambiguity of deciphering what specific factors represent after partitioning out the variance attributed to the global construct. Rodriguez, Reise, and Haviland (2016b) also noted that this approach can be misleading if one relies solely on traditional indices of model fit (e.g., Comparative Fit Index [CFI], Tucker-Lewis Index [TLI], Root Mean Square Error of Approximation [RMSEA], Standardized Root Mean Square Residual [SRMR]) because bifactor models are biased toward producing superior model fit. To be clear, we did not use bifactor analysis to represent the structure of the social identity measure. Rather, we used a multiphase analysis strategy—beginning with bifactor analysis as an *evaluative procedure*, before testing theoretically informed models. Although bifactor modeling affords additional information about a questionnaire's psychometric properties, it was important to follow-up the bifactor analysis by constructing single-factor and three-factor models to represent the construct of social identity.

All analyses were carried out in Mplus 7.3, where all models were estimated using maximum likelihood estimation with standard errors that are robust to non-normality (MLR) (Muthén & Muthén, 2012). In Phase 1, we specified a confirmatory bifactor model, where a general factor of social identity (G-SI) as well as specific factors pertaining to ingroup ties (IGT), cognitive centrality (CC), and ingroup affect (IGA) were modeled to reflect the SIQS items (Fig. 1, Model 1-1). In the confirmatory bifactor model, the general and specific factors are orthogonal to one another. This means that G-SI represents the shared variance among all of the scale items, and three specific factors represent the remaining shared variance among clusters of items with similar content (see Table 1). In Phase 2, we then tested several models based on confirmatory factor analysis (i.e., without fitting the data to a bifactor structure). We first tested a unidimensional factor structure without correlated residuals for the subscales (Fig. 1, Model 2-1) and then tested a unidimensional factor structure with correlated residuals (Fig. 1, Model 2-2). It should be noted that allowing the residual error terms to correlate creates a model that is, in many ways, statistically equivalent to the previously described bifactor model (Model 1-1). The key difference between these models is that the specific factors in Model 1-1 are represented by correlated error terms in Model 2-2. We also evaluated a three-factor structure with independent clusters using confirmatory factor analysis (Model 3-1). In Phase 3, we assessed measurement invariance for the models deemed to be acceptable. In testing invariance according to participant sex, we first examined the factor structure across sexes with no constraints (configural invariance); constraints were then imposed on the factor loadings (factor loading invariance) and finally the factor intercepts (factor loading invariance) (Muthén & Muthén, 2012).

In Phase 1 specifically, following Rodriguez et al. (2016b), we used bifactor analysis to calculate additional statistical indices: (a) the degree of unidimensionality using explained common variance (ECV), (b) construct reproducibility using the *H* value (Hancock & Mueller, 2001), and the (c) reliability of the general factor using coefficient omega hierarchical (ω_H) as well as the reliability of specific factors using coefficient omega hierarchical subscale (ω_{HS}). Even in the presence of good model fit, an ECV score in a bifactor model greater than 0.70 warrants consideration of unidimensionality because most of the variance is attributable to the general factor (Quinn, 2014). We also evaluated the construct reproducibility of each factor to determine the variance explained by each latent variable relative to its unexplained variance—described as the *H* value (Hancock & Mueller, 2001). The internal consistency for the general factor was evaluated with the *omegaH* (ω_H) which refers to variance attributed to the general factor (McDonald, 1999). Subscale reliability was evaluated using the *omegaHS* (ω_{HS}), which refers to the reliability of each specific factor, after statistically controlling for the variance of the general factor (Reise, Bonifay, & Haviland, 2013).[2] In Phase 1–2, the CFI and TLI were inspected as incremental indices of model fit, whereas the RMSEA and SRMR were inspected as absolute indices of model fit (Hu & Bentler, 1999). For guidance, it has been suggested that good fit is achieved when CFI and TLI values are close to or higher than 0.95, the SRMR is less than 0.08, and the RMSEA is less than 0.06 (Hu & Bentler, 1999). In addition, there are an infinite number of solutions when computing factor scores (Gorsuch, 1983). Thus, factor determinacy scores were computed for all models to evaluate whether observed individuals differences on estimated factor scores reflect actual differences on the factor, where values greater than 0.90 indicate that factor scores are determinate (Gorsuch, 1983). In Phase 3, when evaluating measurement invariance, we followed recommendations to inspect changes in model fit in conjunction with χ^2 difference tests (Cheung & Rensvold, 2002). Measurement invariance is supported if additional constraints on a model corresponded to a ΔCFI < -0.010, ΔRSMEA $< +0.015$, and ΔSRMR $< +0.030$ (Chen, 2007). Although concrete guidelines are not available for evaluating ΔTLI, the TLI was inspected to determine whether it stayed within general guidelines for model fit (Hu & Bentler, 1999).

3. Results

3.1. Phase 1: confirmatory bifactor analysis

The confirmatory bifactor model (Model 1-1) with a general factor and three specific factors (Fig. 1, Panel A) demonstrated good model fit based on traditional benchmarks, χ^2 (18) = 43.52, CFI = 0.99, TLI = 0.98, RMSEA = 0.04 90% CI [0.03, 0.06], SRMR = 0.02. As depicted in Table 1, all of the items exhibited high and significant standardized factor loadings on the general factor as well as their respective specific factor. The ECV for the general factor was 0.50. However, whereas the *H* value for the general factor was 0.88, the *H*

[2] For an overview of how these statistical indices are calculated, please refer to Rodriguez et al. (2016a). In addition, Rodriguez et al. provide a thorough overview of why relying solely on Cronbach's alpha in the context of a bifactor model can lead researchers to overestimate the internal consistency of subscale items. In Phase 2, we report the omega coefficient (ω) as a measures of internal consistency for the unidimensional and three dimensional factor structures.

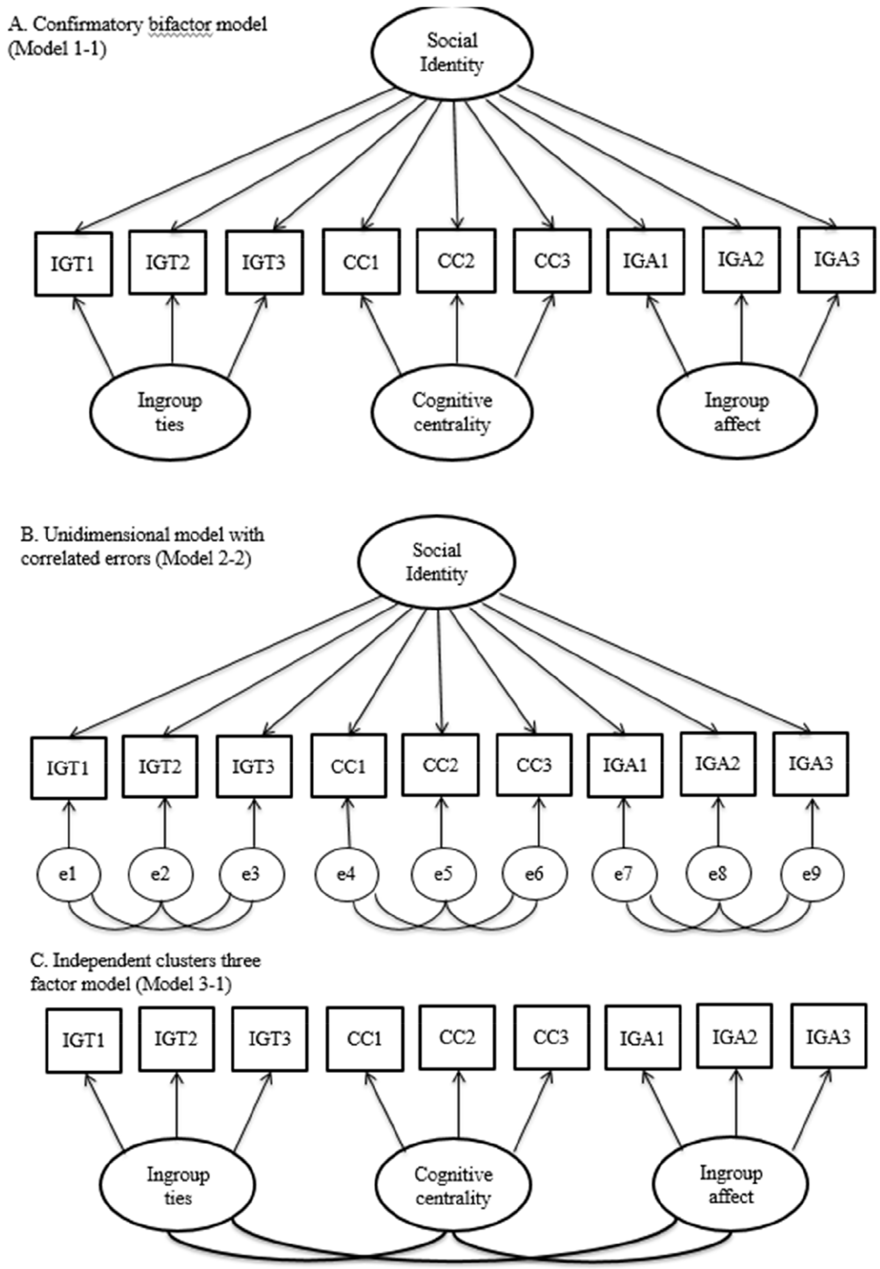

Fig. 1. Conceptual representation of models evaluated for the SIQS. Errors are uncorrelated in Panels A and C, but are not visually depicted.

Table 1
Confirmatory bifactor model: Standardized factor loadings for SIQS (model 1-1).

Item	General factor		Ingroup Ties		Cognitive Centrality		Ingroup Affect		R^2
	ë	*SE*	ë	*SE*	ë	*SE*	ë	*SE*	
IGT1	0.72	0.03	0.43	0.04					0.70
IGT2	0.68	0.04	0.47	0.05					0.68
IGT3	0.72	0.03	0.52	0.05					0.79
CC1	0.60	0.03			0.49	0.05			0.61
CC2	0.51	0.04			0.70	0.04			0.75
CC3	0.57	0.04			0.47	0.05			0.54
IGA1	0.68	0.03					0.41	0.06	0.62
IGA2	0.66	0.03					0.67	0.05	0.89
IGA3	0.72	0.03					0.39	0.05	0.67

Note. IGT = ingroup ties; CC = cognitive centrality; IGA = ingroup affect; λ = standardized factor loading; *SE* = standard error estimate. All factor loadings are significant at $p < .001$.

Table 2
Unidimensional factor structure: Standardized factor loadings for SIQS (models 2-1 and 2-2).

Item	Model 2-1			*Model 2-2*		
	ë	*SE*	R^2	*ë*	*SE*	R^2
SI1	0.78	0.03	0.61	0.72	0.03	0.52
SI2	0.76	0.04	0.58	0.68	0.04	0.46
SI3	0.80	0.03	0.64	0.72	0.03	0.52
SI4	0.61	0.03	0.37	0.60	0.03	0.36
SI5	0.53	0.03	0.28	0.51	0.04	0.26
SI6	0.57	0.03	0.33	0.57	0.04	0.32
SI7	0.71	0.03	0.51	0.68	0.03	0.46
SI8	0.72	0.03	0.52	0.66	0.03	0.44
SI9	0.74	0.03	0.54	0.72	0.03	0.51

Note. SI = social identity; order of items corresponds to Table 1; λ = standardized factor loading; *SE* = standard error estimate. All factor loadings are significant at $p < .001$. Models 1-1 and Models 2-2 exhibit identical factor loadings. However, the r^2 values are lower in Model 2-2 due to the absence of a general factor compared to the bifactor model (Model 1-1).

values for each specific factor were substantially lower (IGT = 0.47, CC = 0.61, IGA = 0.54). These *H* values indicate that the general factor of social identity is a well-defined latent variable, but there is a fair amount of variance unexplained among the specific factors. Moreover, factor determinacy scores indicated potential issues of indeterminacy with the specific factors modeled by the bifactor structure (IGT = 0.70, CC = 0.81, IGA = 0.81), although G-SI was higher (G-SI = 0.89). For the general factor, the ω_H was 0.78. For the specific factors, the ω_{HS} were considerably lower (IGT = 0.27, CC = 0.41, IGA = 0.30). Overall, despite the bifactor model evidencing good fit based on traditional indices, using additional statistical indices illustrates that partitioning out the variance of a general factor renders the specific factors difficult to interpret (i.e., low factor determinacy and low subscale reliability). Notwithstanding the issues related to the specific factors, the general factor exhibited high construct reproducibility, a high degree of internal consistency, and was only marginally below the suggested cutoff for factor determinacy.

3.2. *Phase 2: confirmatory factor analyses*

Confirmatory factor analysis based on a unidimensional factor structure (Model 2-1, not displayed in Fig. 1) exhibited poor model fit, χ^2 (27) = 895.67, CFI = 0.71, TLI = 0.61, RMSEA = 0.19, 90% CI [0.18, 0.20], SRMR = 0.10. In contrast, a unidimensional factor structure with correlated residual error terms (Model 2-2, Fig. 1, Panel B) for each of the purported subscales (nine in total) corresponded to excellent model fit, χ^2 (18) = 43.52, CFI = 0.99, TLI = 0.98, RMSEA = 0.04 90% CI [0.03, 0.06], SRMR = 0.02. As depicted in Table 2, all of the items exhibited high and significant standardized factor loadings on the single factor. The general scales exhibited an acceptable level of internal consistency (SI: ω = 0.89).

Finally, we tested a confirmatory factor model based on a three-factor structure, with the three latent dimensions of IGT, IGA, and CC allowed to correlate (Model 3-1, Fig. 1, Panel C). The model fit well based on traditional benchmarks, χ^2 (24) = 87.10, CFI = 0.98, TLI = 0.97, RMSEA = 0.06, 90% CI [0.04, 0.07], SRMR = 0.03. As depicted in Table 3, all of the items exhibited standardized factor loadings that were high and significant on their respective factor. Moreover, the interfactor correlations between social identity dimensions were moderate and positive (CC with IGT, r = 0.59, IGA with IGT, r = 0.68, IGA with CC, r = 0.55). Factor determinacy scores for all three factors were in an acceptable range (IGT = 0.95, CC = 0.92, IGA = 0.95). The subscales corresponding to each factor exhibited acceptable levels of internal consistency (IGT: ω = 0.89, CC: ω = 0.84, IGA: ω = 0.89).

3.3. *Phase 3: measurement invariance*

Table 4 illustrates the measurement invariance tests across sexes (male and female) based on the models supported through confirmatory factor analyses (Model 2-2, Model 3-1). As it pertains to the unidimensional model, configural invariance was supported by the indices of model fit in Model 4-1. Further, adding constraints to the factor loadings for males and females (Model 4-2) and the factor intercepts for males and females (Model 4-3) did not significantly worsen model fit (ΔCFI = −0.002, ΔRMSEA = 0.004). As it pertains to the three-factor structure, configural invariance was supported by the indices of model fit in Model 5-1. Further, adding constraints to the factor loadings for males and females (Model 5-2) and the factor intercepts for males and females (Model 5-3) did not significantly worsen model fit (ÄCFI = −0.003, ÄRMSEA = 0.001). Across both models, there appears to be support for strong measurement invariance across sexes. Descriptive statistics for social identity including by sex are presented in Table 5.

4. Discussion

In the past decade, there has been a burgeoning interest from sport and exercise psychology researchers in how social identity relates to variables at the individual, interpersonal, and group level (e.g., see Rees et al., 2015; for a review). To evaluate support for a three-dimensional conceptualization of social identity (Cameron, 2004) and present a measure that yields valid scores of these dimensions in a sport setting, we introduced a nine-item version of the Social Identity Questionnaire for Sport (SIQS). We applied a novel approach advocated by Rodriguez et al. (2016a, 2016b) to scrutinize the dimensionality of the SIQS via bifactor analysis. After using bifactor analysis as an investigative tool, we evaluated several theoretically supported models pertinent to the SIQS (i.e., unidimensional and multidimensional factor structures).

Table 3
Confirmatory Factor Analysis based on Three Factor Structure: Standardized Factor Loadings and Item-level Statistics for SIQS (Model 3-1).

Item	Mean (Variance)	Skewness/Kurtosis	Ingroup Ties		Cognitive Centrality		Ingroup Affect		R^2
			ë	*SE*	*ë*	*SE*	*ë*	*SE*	
IGT1	5.56 (1.67)	−0.92/0.57	0.84	0.02					0.71
IGT2	5.83 (1.54)	−1.27/1.66	0.83	0.03					0.68
IGT3	5.64 (1.65)	−1.08, 1.08	0.89	0.02					0.79
CC1	5.09 (2.32)	−0.69, −0.07			0.81	0.02			0.65
CC2	5.06 (2.33)	−0.69, −0.10			0.80	0.02			0.64
CC3	5.00 (2.01)	−0.56, −0.15			0.75	0.02			0.56
IGA1	6.39 (0.98)	−2.10, 5.04					0.81	0.02	0.65
IGA2	6.28 (1.01)	−1.59, 2.41					0.88	0.02	0.77
IGA3	6.06 (1.20)	−1.22, 1.27					0.83	0.02	0.69

Note. IGT = ingroup ties; CC = cognitive centrality; IGA = ingroup affect; λ = standardized factor loading; *SE* = standard error estimate. All factor loadings are significant at $p < .001$.

Table 4
Tests of measurement invariance (male and female).

	χ^2 (*df*)	CFI	TLI	RMSEA [90% CI]	SRMR	CM	$\Delta S\chi^2$ (df)	ΔCFI	ΔTLI	ΔRMSEA
Unidimensional factor structure										
4-1: Configural invariance	62.719 (45)	0.994	0.990	0.030 [0.006, 0.047]	0.028	–	–	–	–	–
4-2: Metric invariance	64.527 (53)	0.996	0.995	0.022 [0.000, 0.040]	0.040	1–1	1.707 (8)	0.002	0.005	−0.008
4-3: Scalar invariance	79.133 (61)	0.994	0.993	0.026 [0.000, 0.041]	0.045	1–2	15.579 (8)*	−0.002	−0.002	0.004
Three-factor structure										
5-1: Configural invariance	100.919 (48)	0.982	0.973	0.050 [0.037, 0.064]	0.034	–	–	–	–	–
5-2: Metric invariance	105.199 (54)	0.983	0.977	0.047 [0.033, 0.060]	0.042	1–1	4.252 (6)	0.001	0.004	−0.003
5-3: Scalar invariance	119.052 (60)	0.980	0.976	0.048 [0.035, 0.060]	0.042	1–2	14.284 (6)*	−0.003	−0.001	0.001

Note. Unidimensional factor structure is based on allowing the residual error terms to correlate between each of the three specific factors used in the bifactor model. CFI = comparative fit index; TLI = Tucker-Lewis index; RMSEA = root mean-square error of approximation; 90% CI = 90% confidence interval of RMSEA; SRMR = standardized root mean square residual. CM = Comparison model; * = $p < .05$; Gender comparisons based on male (n = 375) and female (n = 493).

Table 5
Descriptive statistics.

	Total sample		Males (n = 375)		Females (n = 493)	
	M	*SD*	*M*	*SD*	*M*	*SD*
SI	5.65	0.93	5.61	0.89	5.69	0.96
IGT	5.68	1.14	5.63	1.12	5.70	1.17
CC	5.05	1.28	5.03	1.25	5.07	1.32
IGA	6.24	0.92	6.17	0.90	6.29	0.93

Note. SI = social identity; IGT = ingroup ties; CC = cognitive centrality; IGA = ingroup affect.

Overall, a theoretically conceived three-factor structure is empirically supported, where ingroup ties, cognitive centrality, and ingroup affect are represented as distinct dimensions of social identity. Empirical support was also found for a global factor of social identity, but only when correlated residuals are included among the items of each subscale. This modelling approach recognizes that there is variance shared among the subscale items that is not accounted for by a global factor of social identity.

As noted above, confirmatory factor analysis of the SIQS based on a three-factor structure offered evidence of good fit and all of the items loaded highly onto their respective dimensions. This model also demonstrated acceptable factor determinacy scores (i.e., above 0.90, Gorsuch, 1983), high subscale reliability scores, and showed evidence of strong measurement invariance across sexes. The latter point indicates that the SIQS is assessing the same construct (i.e., three dimensions of social identity) among males and females, which is reassuring for researchers who aim to examine social identity in samples comprised of males and females. Moreover, obtaining support for measurement invariance across sexes is important for researchers who are interested in testing whether a relationship between social identity and a variable of interest differs across males and females (i.e., moderation). Collectively, the results provide further support for the multidimensional nature of social identity (cf. Cameron, 2004) and offer strong empirical support for using the SIQS to evaluate the three dimensions of social identity (ingroup ties, cognitive centrality, ingroup affect) in a sport context.

A key insight from the bifactor model is that although the items in the SIQS can be modeled to reflect social identity as a global construct, there is no empirical justification for attempting to simultaneously model three specific factors (i.e., ingroup ties, cognitive centrality, ingroup affect). As it pertains to the former point, the bifactor model showed that (a) a fair amount of variance is accounted for by the global construct of social identity, (b) social identity as a global construct is a well-defined latent variable (high *H* value), and (c) the items exhibited acceptable levels of reliability (e.g., high omega*H*) and high factor loadings. As it pertains to the former point, partitioning out the variance of a general factor renders the specific dimensions difficult to interpret (e.g., low factor determinacy and low subscale reliability). To be clear, in a bifactor model that includes an orthogonal general factor, the specific factors represented by ingroup ties, cognitive centrality, and ingroup affect performed poorly on a range of statistical indices. As expected, the latent dimensions related to ingroup ties, cognitive centrality, and ingroup ties cannot be satisfactorily assessed once the common variance attributed to social identity as a global construct is removed. Despite evidencing good fit based on traditional indices, we are hesitant to recommend attempting to model social identity as a global construct while simultaneously modeling three specific factors that are orthogonal to social identity.

If the goal is to model social identity as a global construct with the SIQS, researchers can model a unidimensional factor structure where the residuals among the subscales are allowed to correlate (unidimensional approach). Although this approach (Model 2-2) produces a global factor of social identity that is statistically equivalent to the global factor in the bifactor structure (Model 1-1), the inclusion of correlated error terms (rather than specific factors orthogonal to a general factor) shifts the substantive foci to a single latent dimension of social identity. With Model 2-2, researchers are not left with the problem of attempting to interpret what the specific factors represent in substantive terms after removing the variance explained by social identity as a global construct. In addition to Model 2-2 demonstrating good fit, there was also evidence for strong measurement invariance across sexes. Taken together, the results provide a strategy for modelling social identity as a global construct when the substantive research question merits such an approach.

More broadly, the findings of the current paper are consistent with the cautionary note raised by Bonifay et al. (2017) as well as Rodriguez et al. (2016a, 2016b) regarding the increasing popularity of bifactor analysis as a way to model psychological constructs. If we relied solely on indices of model fit, the bifactor model could have been interpreted as a suitable factor structure for the SIQS. Echoing Rodriguez et al. (2016a), sport and exercise psychology researchers who are interested in evaluating a bifactor measurement model should calculate and interpret a range of statistical indices to ensure the general and specific factors represent well-defined latent constructs (e.g., ECV, H values, hierarchical omega, hierarchical subscale omega, factor determinacy scores). Indeed, as it pertains to evaluating measurement models of sport and exercise psychology constructs, there may be instances where a bifactor structure performs well across a range of statistical indices (e.g., ECV, H values, hierarchical omega, hierarchical subscale omega, factor determinacy scores) *and* provides a more accurate approximation of substantive theory (e.g., Myers, Martin, Ntoumanis, Celimli, & Bartholomew, 2014).

Consistent with theoretical conceptualizations of social identity (Cameron, 2004; Hornsey, 2008), our findings suggest that the SIQS can be used to either model social identity along three specific dimensions *or* as a global construct. However, the SIQS does not enable researchers to simultaneously model social identity as a global construct with three specific factors that are unrelated. The decision to use a multidimensional or a global measure of social identity should be driven by theory and the research question. On the one hand, researchers interested in investigating how the cognitive and affective components of ingroup identification differentially relate to psychological outcomes would benefit from a multidimensional approach, and thus by modeling their data based on the three-factor structure. For example, researchers may be interested in examining how the affective component of social identity (i.e., ingroup affect) is associated with personal emotional outcomes related to sport, such as guilt or pride (Tangney, 1999). As another example, researchers may manipulate the salience of group membership (i.e., cognitive centrality) to evaluate potential changes in member behavior toward other ingroup and outgroup members. This proposed research would extend findings indicating the potential salient role of cognitive centrality on antisocial behavior toward team members (Benson et al., 2017). Finally, with growing interest in understanding the structure of relationships in groups (e.g., cliques; Martin, Wilson, Evans, & Spink, 2015), the connections with group members (i.e., ingroup ties) may be of particular interest. A common thread running across these examples is that certain hypotheses generated by a social identity approach are multidimensional in nature.

On the other hand, researchers who are interested in the global construct of social identity (i.e., a single latent dimension) to evaluate the general strength of social identity would benefit from modeling social identity as a single latent dimension where the error terms among the subscale items are allowed to correlate. This enables researchers to assess the global construct of social identity as commonly discussed in theoretical accounts (e.g., Tajfel et al., 1981) and empirical work (e.g., Terry et al., 1999) in instances in which researchers might not have multidimensional hypotheses. A practical example of a global approach may involve comparing youth sport participants high and low on social identity in relation to specific outcomes (e.g., friendship, dropout). Another practical example, may involve a team building intervention to bring together subgroups on a sport team to enhance team social identity. Ultimately, the decision to use a multidimensional versus global approach should remain driven by the research question and theory.

Given that this is the first in-depth examination of the psychometric properties and dimensionality of the SIQS, it is important to consider a number of limitations and future directions. First, it is important acknowledge the three independent samples analyzed in this paper were comprised primarily of team sport athletes and mostly youth participants. A fruitful avenue of research may involve examining the social identities of individual sport groups, some of which train together but compete individually (e.g., cross-country skiing, see Evans, Eys, & Bruner, 2012; for a sport typology). Individual sport athletes might exhibit greater variation in the extent to which they identify with their fellow group members. Although it would be prudent to evaluate the measurement invariance of the SIQS when applied to the study of individual sport populations, this context would allow researchers to test novel research questions. For example, researchers could examine how objective changes in outcome interdependence (e.g., team-based versus individual events) over the course of a season are associated with social identity (Evans & Eys, 2015).

Moving forward, it should be noted that the SIQS offers a shorter assessment of social identity and thus lessens participant burden. Reducing participant burden may facilitate research on social identity across multiple time points. We encourage researchers to consider the antecedents and consequences of social identity strength at the intra-individual level, as this would provide deeper insight into the role of social identity in sport. As an example, it may be fruitful to examine how the four dimensions of identity leadership (identity prototypicality, identity advancement, identity entrepreneurship, identity impresarioship, Steffens et al., 2014) in coaches and athletes relate to the three dimensions of social identity (ingroup ties, cognitive centrality, ingroup affect). Such research would build upon recent work indicating coaches' identity entrepreneurship was uniquely related to global identification (Study 4; Steffens et al., 2014). A related avenue of future research is that less participant burden would also facilitate recommended intervention work in sport settings (Bruner, Boardley et al., 2014, Bruner, Dunlop et al., 2014).

4.1. Conclusion

Over the past 45 years, there has been strong and continued documentation of the influential role of social identity on human behavior. Despite this empirical evidence, the measurement of social identity as a theoretical construct has yet to be closely scrutinized in sport. In the current research, we evaluated the factor structure and measurement invariance of the SIQS as a measure of social identity in sport. The findings from the study support the SIQS as a psychometrically sound measure of social identity (Cameron, 2004) in sport among both male and female athletes.

Appendix A. Supplementary data

Supplementary data related to this article can be found at http://dx.doi.org/10.1016/j.psychsport.2017.12.006.

Appendix A

Social Identity Questionnaire for Sport (SIQS)

The following questions are designed to reflect **how you feel about being a part of your team**. Please CIRCLE a number from 1 (*strongly disagree*) to 7 (*strongly* agree) to indicate your agreement with each of the statements.

1. I feel strong ties to other members of this team.
2. I find it easy to form a bond with other members in this team.
3. I feel a sense of being "connected" with other members in this team.
4. Overall, being a member of this team has a lot to do with how I feel about myself.
5. In general, being a member of this team is an important part of my self-image.
6. The fact that I am a member of this team often enters my mind.
7. In general, I'm glad to be a member of this team.
8. I feel good about being a member of this team.
9. Generally, I feel good when I think about myself as a member of this team.

Ingroup ties: Items 1, 2, and 3.
Cognitive centrality: Items 4, 5, and 6.
Ingroup affect: 7, 8, and 9.
Three items removed:
Ingroup Ties: I have a lot in common with other members in this team.
Cognitive Centrality: I often think about the fact that I am a team member.
Ingroup Affect: I rarely regret that I am a member of this team.

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