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A team fares well with a fair coach: Predictors of social loafing in interactive female sport teams

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The present research aimed to develop and test a theoretical model that links players' perceived justice of the coach to a more optimal motivational climate, which in turn increases players' team identification and cohesion, and results in lower levels of social loafing in female sport teams. Belgian elite female basketball, volleyball, and football players (study 1; $N = 259$; $M_{age} = 22.6$) and Norwegian world-class female handball players (study 2; $N = 110$; $M_{age} = 22.8$) completed questionnaires assessing players' perceived justice (distributive and procedural), motivational climate, team identification, team cohesion (task and social), and social loafing (perceived and self-reported). In both studies, confirmatory and exploratory path analyses indicated that perceived justice was positively related to a mastery climate ($P < 0.05$) and negatively to a performance climate ($P < 0.05$). In turn, a mastery climate was linked to increased levels of team identification ($P < 0.05$) and task cohesion ($P < 0.05$). Consequently, players' perceived and self-reported social loafing decreased ($P < 0.05$). The findings of both independent studies demonstrated the impact of coaches' fairness, and consequently, the motivational climate created by the coach on the optimal functioning of female sport teams.

There seems to be a general consensus in the research literature that coaches play a crucial role in realizing team sport success (Adie et al., 2008). Despite this prevailing belief, sport scientists have paid rather little attention to the specific behavior that makes coaches' leadership more effective. By contrast, research in the business setting has focused frequently on the connection between justice and leadership and concluded that leader's justice is positively associated with leadership effectiveness (Colquitt & Greenberg, 2003; van Knippenberg et al., 2007). Based on this finding, sport psychologists have started to show an increasing interest in the role of organizational justice and have begun to examine the impact of coaches' justice on group dynamic processes in team sports (Jordan et al., 2004).

Organizational justice theory (OJT) attempts to describe and explain the role of fairness in the workplace (Greenberg, 1990). This theory often differentiates justice into distributive and procedural justice. Distributive justice refers to the fairness of outcome distributions or allocations (Greenberg, 1990). A typical outcome in a team sport context is the playing time of an athlete. Procedural justice focuses on the fairness of the procedures used to determine these outcome distributions or allocations (Greenberg, 1990). An example of the application of a fair procedure in sport teams is the use of objective scouting information to select the starting team.

We assume that coach-related justice will be particularly important in team sports, because the coach is in charge of the sportive outcomes (i.e., distributive justice) as well as of the procedures followed to achieve these outcomes (i.e., procedural justice). In support of this assumption, previous research has demonstrated that athletes' perceived justice of the coach positively predicts team identification and team cohesion in elite sport teams (De Backer et al., 2011). Based on these results, we wanted to test whether the relation between coaches' justice and players' team identification is mediated by the motivational climate created by the coach. In addition, we aimed to study the effect of coaches' justice on players' social loafing. Social loafing has been defined as the reduction in motivation and effort when individuals work collectively, compared with when they work individually (Karau & Williams, 1993). Considerable evidence has shown that social loafing is responsible for a substantial portion of the decreased performance of teams (Karau & Williams, 1993).

Liden et al. (2004) already observed a significant correlation between justice and social loafing in a business context. Their model showed that justice was indirectly related to social loafing through its impact on other relevant predictors, such as task visibility and task interdependence. In the current study, we developed and tested a theoretical model that links perceived justice to a more optimal motivational climate, which in turn increases athletes' team identification and cohesion, and finally results in lower levels of social loafing in female interactive sport teams.

Coach-related justice, the motivational climate, and team identification

In order to construct this model, we will first discuss the underlying mechanisms that might explain the previously demonstrated link between coach-related justice and athletes' team identification (De Backer et al., 2011). Team identification is a concept derived from the social identity approach (SIA; Haslam, 2004). SIA makes a distinction between a personal identity (i.e., the self as a unique individual) and a social identity (i.e., the self as an interchangeable team member). Furthermore, this theory assumes that the self-categorization process (i.e., the process of perceiving the self as an interchangeable member of a category) is the cognitive process associated with social identity. Consequently, when an athlete identifies with the team, his social identity predominates, which in turn leads to the internalization of group norms and behavior in favor of the welfare of the group.

In an attempt to explain how leaders can shape team identification, Haslam et al. (2011) stated that leaders actively shape the social environment in which their team operates. This social environment subsequently influences the social identification of group members. In line with the first part of this statement, it has been shown that coaches play an important role in creating a productive team environment that is associated with success (Ntoumanis & Biddle, 1999). The features of this environment are frequently studied through one of the major concepts of the achievement goal theory (AGT; Ames, 1992), namely the motivational climate. Ames (1992) distinguished between two different motivational climates of the social environment. A mastery motivational climate is characterized by an emphasis on effort, improvement, members' contributions to the team's efforts, and helping each other to learn. In contrast, a performance motivational climate emphasizes normative feedback, intra-team member rivalry, inequity, and favoritism to a particular athlete or group (Heuzé et al., 2006).

Based on the definitions of both climates, we expected that the relation between players' perceived fairness of the coach and their team identification runs through the motivational climate created by the coach. In particular, fair coaches distribute outcomes based on objective information about athletes' effort and improvement. As a result, players will probably focus more on their own effort and improvement, which is a key characteristic of an optimal mastery climate. Furthermore, several researchers have observed that fairly treated people were more willing to subordinate their own short-term individual interests to the interest of the group, and were more willing to contribute to the betterment of their fellow employees (Tyler & Degoey, 1995). Both members' willingness to contribute to the team's interest and their support of the learning process of their teammates characterize a mastery climate. Duda and Hall (2001) showed that this mastery climate increased satisfaction with team membership. Consequently, we assume that fair coaches foster a mastery climate in which athletes focus more on the long-term improvement of their own skills and less on the rivalry with their team members, which in turn promotes a sense of team identification.

Contrary to fair behavior, inequity (i.e., distributive injustice) and favoritism to a particular athlete or group within the team (i.e., procedural injustice) are perceived as coach characteristics fostering a performance climate. Such a performance climate promotes intra-team rivalry, which is detrimental for players' team identification. As a result, we expected that fair coaches decrease players' perceptions of a performance climate and in turn increase team identification. Based on this theoretical background, we formulated two hypotheses:

> Hypothesis 1a: Athletes' perceived justice of the coach is positively related to a mastery climate, which in turn is positively related to athletes' team identification.

> Hypothesis 1b: Athletes' perceived justice of the coach is negatively related to a performance climate, which in turn is negatively related to athletes' team identification.

Team identification, team cohesion, and social loafing

The second purpose of the current study was to examine whether this team identification, in turn, predicts the amount of social loafing and to discuss the underlying mechanism that might explain the link between team identification and social loafing in female sport teams. We approached the concept of social loafing from two different perspectives: (a) athletes' perceptions of the social loafing behavior of their teammates (i.e., perceived social loafing); and (b) athletes' own social loafing behavior (i.e., self-reported social loafing). In line with previous research, we expected that both forms of social loafing are distinct, but modestly related concepts (Høigaard & Ommundsen, 2007). Moreover, previous research (Jackson & Harkins, 1985) indicated that people who expect their co-performers to loaf, reduce their own efforts to establish an equitable division of labor.

Haslam (2004) stated that when individuals strongly identify themselves with the team and define themselves as interchangeable group members, social loafing would be eliminated. Because research on the consequences of team identification in sport settings has been limited, the second part of the theoretical model of the current research examined the mechanisms that might lead from athletes' team identification to social loafing. In contrast to team identification, the closely related concept of team cohesion, which can be seen as a logical consequence of team identification (Dutton et al., 1994), has been extensively studied in sport settings and is generally acknowledged as a crucial concept determining group behavior and social loafing (Everett et al., 1992; Carron et al., 2002; Heuzé et al., 2007).

Team cohesion is defined as "a dynamic process that is reflected in the tendency for a group to stick together and remain united in the pursuit of its instrumental objectives and/or for the satisfaction of member affective needs" (Carron et al., 1998, p. 213). Carron et al. (1998) differentiated team cohesion into an individual (i.e., individual attraction to the group) and a group component (i.e., group integration). Both components are further differentiated into a social (i.e., a general orientation toward developing and maintaining social relationships within the group) and a task subcomponent (i.e., a general orientation toward achieving the group's goals and objectives).

Recent research has supported the statement that the perception of shared team identification leads to intragroup cohesion (Dutton et al., 1994; De Backer et al., 2011). As a result, we assume that the relation between team identification and social loafing runs through cohesion. Particularly task cohesion has been showing to play a crucial role in eliminating social loafing in female swimming relay teams, and male soccer teams (Everett et al., 1992; Høigaard et al., 2006). As a result, we expected that the relation between team identification and social loafing runs through task and not through social cohesion. However, sport scientists have paid rather little attention to the relation between social cohesion and social loafing; therefore, we will not exclude social cohesion from our research model (i.e., we will test an additional model with two extra paths from social cohesion to both forms of social loafing). Based on this theoretical background, we formulated two hypotheses:

> Hypothesis 2a: Athletes' team identification is positively related to task cohesion, which in turn is negatively related to athletes' perceived social loafing.

> Hypothesis 2b: Athletes' team identification is positively related to task cohesion, which in turn is negatively related to athletes' self-reported social loafing.

We combined all four hypotheses into one comprehensive research model. This model examines the effects of coach-related justice on perceived and self-reported social loafing through the interrelationships among motivational climate, team identification, and team cohesion. We performed two separate studies to test this model. Study 1 examines the model in Belgian senior female sport teams (i.e., basketball, volleyball, and football). Study 2 examines the same model in Norwegian world-class female handball teams.

Study 1

Method

Participants and procedures

The head coaches of 65 Belgian female senior sport teams were contacted by telephone and informed about the purpose of the study. Forty-five allowed their teams to take part in the study. The coaches that decided not to participate gave as main reason the high workload of their players. The final research sample consisted of 259 Belgian female senior athletes (i.e., a response rate of 49.8%). All participants were engaged at club level: six basketball teams of 12 players ($n = 52$), six soccer teams of 14 players ($n = 65$), and 33 volleyball teams of 11 players ($n = 142$). They were ranked from the national to the regional level of the Belgian competition and trained on average 4.8 h per week [standard deviation (SD) = 3.7]. The mean age of the players was 22.6 years (SD = 5.3) and they had played at senior level from 1 to 25 years (M = 3.8, SD = 3.9).

The players of the participating teams were sent an e-mail, in which they were informed about the objectives of our study. The players who agreed to participate were invited to fill in our web-based questionnaire. The link to the web-based questionnaire was included in the e-mail. Players also had the possibility to receive a paper and pencil questionnaire by post together with an information letter. They had to put their completed questionnaires in a separate envelope, which was enclosed with the questionnaire, and send it to the authors. The American Psychological Association (APA) ethical standards were followed in the conduct of the study. No rewards were given for participation, informed consent was obtained from all participants, and confidentiality was assured.

Measures

Perceived justice (eight items). We assessed distributive and procedural justice with a Dutch version of the perceived justice questionnaire of Colquitt (2001) using a 5-point Likert scale (*strongly disagree* = 1; *strongly agree* = 5). van Prooijen (2009) translated the perceived procedural justice items into Dutch. The scores of these items showed to be valid and reliable within a sample of Dutch junior-level management students and a sample of

Dutch employees of automobile factories. However, the perceived distributive justice items were not yet used in a Dutch study; therefore, these items were back-translated by two native-speaking Dutch researchers. Finally, we slightly adapted the items to fit the specific team sport context (e.g., play minutes replaced salary). Four items assessed the perception of distributive justice (e.g., “The minutes I play per game are a true reflection of my qualities and amount of effort on the field”) and four assessed the perception of procedural justice (e.g., “The decisions of my coach are based on objective information”).

Motivational climate (20 items). The perceived motivational climate was assessed with a Dutch version of the validated 20-item Perceived Motivational Climate in Sport Questionnaire (PMCSQ; Seifriz et al., 1992) using a 5-point Likert scale (*strongly disagree* = 1; *strongly agree* = 5). The scores of the Dutch PMCSQ showed acceptable reliability and validity with a sample of Dutch high-level female athletes (de Bruin et al., 2009). The 20-item PMCSQ consists of 11 items measuring the performance climate (e.g., “On this team, out-playing teammates is important”) and nine items measuring the mastery climate (e.g., “On this team, each player’s improvement is important”).

Identification with the team (six items). Team identification was measured using six Dutch items. Three items were already used in previous research (De Backer et al., 2011), the other three items were adapted from the fan identification scale constructed by Boen et al. (2008). The six items used a 7-point Likert scale (*strongly disagree* = 1; *strongly agree* = 5). An example item is “The successes of my team feel like my own successes.”

Cohesion (18 items). Cohesion was questioned with the Group Environment Questionnaire (GEQ; Carron et al., 1998) using a 9-point Likert scale (*strongly disagree* = 1; *strongly agree* = 9). The items were back-translated by two native-speaking Dutch group dynamics experts. Four items assessed the individual attraction to the group-social subcomponent (e.g., “Some of my best friends are on this team”), four items assessed the individual attraction to the group-task subcomponent (e.g., “I am unhappy with the team’s level of desire to win”), five items assessed the group involvement-social subcomponent (e.g., “Our team would like to spend time together in the off-season”), and five items assessed the group involvement-task subcomponent (e.g., “Our team is united in trying to reach its performance goals”).

Social loafing (nine items). Social loafing was measured by two questionnaires: the Perceived Social Loafing Questionnaire (PSLQ; Høigaard et al., 2006; Høigaard & Ommundsen, 2007; Høigaard et al., 2010) and the Self-Reported Social Loafing Questionnaire (SRSLQ; Høigaard et al., 2010). Scores of both scales were reliable and valid within a sample of top-level Norwegian athletes (Høigaard et al., 2010). These questionnaires were back-translated by two native-speaking Dutch experts. The PSLQ consists of five items (e.g., “Members of my team contribute less than I expected”). The SRSLQ consists of four items (e.g., “I contribute less than I should”). Both scales use a 5-point Likert scale (*strongly disagree* = 1; *strongly agree* = 5).

Data analysis

First, we tested the factorial structure of all scales using a confirmatory factor analysis (CFA). More specifically, we performed multiple group measurement invariance tests to evaluate the factor structure of the instruments across the samples of study 1 (Belgian athletes) and study 2 (Norwegian athletes). Second, the internal consistency was calculated using Cronbach’s alpha, and correlations were examined. Third, to determine the effect of nesting on the results, and to decide whether multilevel analyses were necessary, design effects (DEs) were computed. Finally, confirmatory and exploratory path analyses were performed with the Amos maximum likelihood estimation method to investigate our research model.

Results

Descriptive statistics, correlations, and scale reliabilities

Scales, means, SDs, and correlations for the variables of Study 1 are provided in Table 1. Scale reliabilities (Cronbach’s alphas) are provided on the diagonal.

CFA and multiple group measurement invariance testing

First, we performed CFAs to test the factorial structure of each scale. The findings showed that all scales of study 1 (except the GEQ) exhibited acceptable to good fit indices with significant factor loadings. The results of the separate CFAs for the Belgian sample are provided in Appendix 1.

Second, because the establishment of partial measurement invariance is a prerequisite for meaningful comparisons across groups, multiple group invariance analyses were carried out to examine the measurement invariance of the scales across the Belgian and Norwegian samples. Although one can also test for the equivalence of measurement error terms, it is now widely accepted that this test is overly restrictive. Except in particular instances when, for example, it might be of interest to test for the equivalent reliability of an assessment measure across groups, the equality of error variances is probably of least importance (Byrne, 2004, 2010). Consequently, the error terms were not constrained equally between the Belgian and the Norwegian samples in any of these analyses. The results of the

Table 1. Scale, means (M), standard deviations (SDs), correlations, and Cronbach's alphas for all variables used in study 1

Variable	Scale	M	SD	1	2	3	4	5	6	7	8	9
1. Procedural justice	1–5	3.28	0.92	(0.89)								
2. Distributive justice	1–5	3.60	1.02	0.52*	(0.95)							
3. Performance climate	1–5	2.63	0.65	−0.43*	−0.37*	(0.79)						
4. Mastery climate	1–5	3.88	0.66	0.56*	0.37*	−0.34*	(0.85)					
5. Team identification	1–7	5.37	1.05	0.32*	0.27*	−0.30*	0.45*	(0.85)				
6. Social cohesion	1–9	6.30	1.49	0.18*	0.13*	−0.18*	0.23*	0.46*	(0.65)			
7. Task cohesion	1–9	6.39	1.55	0.32*	0.13*	−0.28*	0.44*	0.52*	0.34*	(0.70)		
8. Self-reported loafing	1–5	1.72	0.61	−0.11	−0.19*	0.26*	−0.33*	−0.41*	−0.18*	−0.26*	(0.72)	
9. Perceived social loafing	1–5	2.50	0.85	−0.34*	−0.17*	0.29*	−0.42*	−0.30*	−0.26*	−0.56*	0.26*	(0.81)

Note. * $P < 0.05$.

multiple group invariance analyses showed at least partial invariance for all scales across the Belgian and the Norwegian samples (for an overview of the results. see Appendices 1 and 2).

Because the CFA of the GEQ showed unacceptable fit indices, we highlight the analyses of the GEQ. Taking into account the suggestion that authors should retain the subscale structure, but also considering that poor subscales are usually dropped altogether when they do not function appropriately (Carron & Brawley, 2000), we decided to drop both subscales referring to individual attraction to the group (i.e., individual attraction to the group-social, and individual attraction to the group-task).

Furthermore, because of nonsignificant factor loadings, we had to delete one item of the group integration-social subscale (i.e., "Members of our team would rather go out on their own than get together as a team") and one item of the group integration-task subscale (i.e., "Our team members do not communicate freely about each athlete's responsibilities during competition or practice"). As a result, the final reliable, valid and partial invariant cohesion measurement consisted of eight items equally distributed over the group integration-social and the group integration-task subscales.

DEs

To help determine the effect of nesting on the results, and to determine if multilevel analyses were necessary, DEs were computed for each variable. DEs take into account both the intra-class correlation (ICC) and the within-group sample size [$DE = 1 + (\text{average within group sample size} - 1) \times ICC$]. Taking into account this sample size is important given that ICCs are likely to be inflated when there are many groups with few individuals within the groups (which is the case in study 1). A DE of 2 or more implies that team membership does have an effect on the responses of the individuals and therefore multilevel modeling should be conducted to account for the multilevel nature of the data (Sorra & Dyer, 2010). In the current study, the average DE for the team level was 1.71 (ranging from 1.01 to 2.43; see Appendix 3 for a complete overview).

In addition to these small DEs, there were also practical restrictions to perform multilevel structural equation modeling (SEM). Hox and Maas (2001) caution against using multilevel SEM when the number of groups is smaller than 100, especially if the ICCs turn out to be under 0.25. Furthermore, Meuleman and Billiet (2009) showed that to detect large (> 0.50) structural effects at the between level, at least 60 groups are required. Because only 45 teams participated in study 1 and given the small DEs, we decided not to perform multilevel SEM.

Confirmatory and exploratory path analysis

We decided to use path models without the latent scores of the measurement models, in order to decrease the ratio of cases (*N*) to the number of model parameters that require statistical estimates. An acceptable ratio of sample size/parameters should range between 20:1 and 5:1 (Kline, 2011). An SEM analysis with latent scores would contain 122 parameters that require statistical estimates. Using a path model with observed variables, we decreased the number of parameters to 23, which results in an acceptable ratio of sample size/parameters of approximately 11:1 in the current study.

The confirmatory path analysis indicated a poor fit of the hypothesized model [$\chi^2 = 100.65$, degrees of freedom (df) = 22, $P < 0.01$; goodness-of-fit index (GFI) = 0.92; comparative fit index (CFI) = 0.87; root mean square error of approximation (RMSEA) = 0.12]. Consequently, we performed an exploratory path analysis to add paths that would increase the model fit (Fig. 1). Taking into consideration only the Amos modification indices that significantly improved the fit of our model and were in line with previous research, we added two paths: (a) a positive path from mastery climate to task cohesion; and (b) a negative path from team identification to self-reported social loafing. The relation between a mastery climate and task cohesion is in line with previous research in which the concept of team identification was not included (Heuzé et al., 2006). However, we had expected that this relation would be mediated by players' team identification. The relation between team identification and social loafing was already suggested

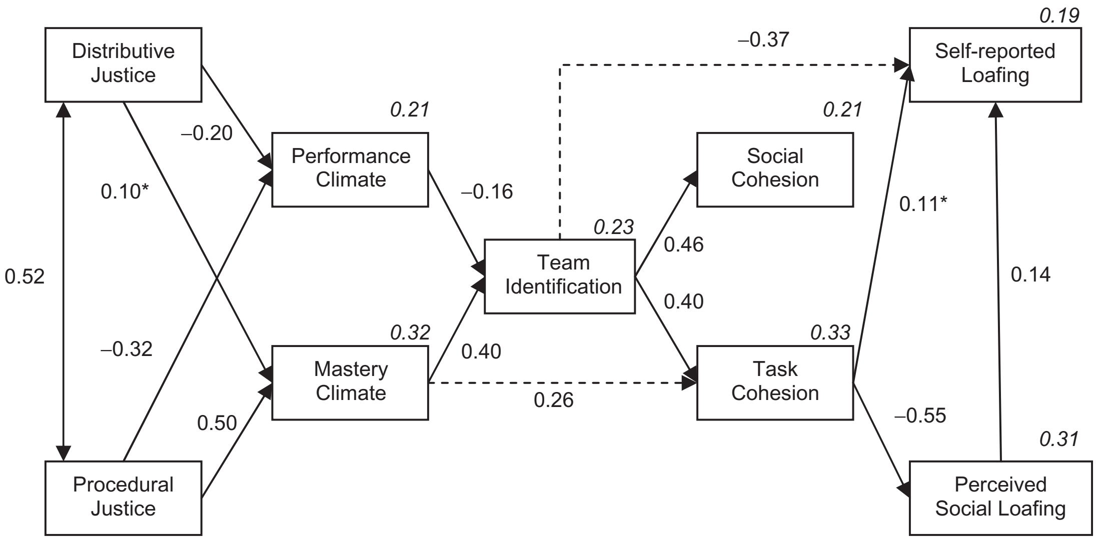

Fig. 1. Structural model best fitting the data of study 1 with the regression coefficients and the proportions explained variance. Legend: All coefficients presented are standardized and significant ($P < 0.05$) except for *$P > 0.05$. The dashed arrows represent the added paths based on Amos modification indices and in line with previous research.

by Karau and Williams (1993). Again, we had expected that this relation would be mediated by players' task cohesion.

Results of the exploratory path analysis provided a good fit of the revised model to the data ($\chi^2 = 50.26$, df = 20, $P < 0.01$; GFI = 0.96; CFI = 0.95; RMSEA = 0.08). The standardized regression path coefficients and the proportions explained variance are illustrated in Fig. 1. Examination of the revised model (see Fig. 1) reveals that only the paths between distributive justice and mastery climate and between task cohesion and self-reported social loafing were not significant. We also tested an additional model with two extra paths from social cohesion to both forms of social loafing. The results showed that social cohesion was not significantly related to any form of social loafing. Furthermore, because our CFAs showed significant correlations between mastery and performance climate and between task and social cohesion, we linked the error terms of these variables. Only the error terms of the mastery and performance climate ($r = -0.13$, $P = 0.05$) were significantly interrelated, while the error terms of task and social cohesion ($r = 0.10$, $P > 0.05$) were not significantly interrelated.

We followed the SEM approach by Holmbeck (1997) to test the mediational effects in the full structural model. More specifically, we tested the full structural model in two turns. First, we tested whether the relation between both forms of justice and team identification was mediated by the motivational climate (i.e., mastery and performance climate). Second, we tested whether the relation between team identification and both forms of social loafing was mediated by task cohesion. Based on our results, we can conclude that the relation between procedural justice and team identification was mediated by both motivational climates, while we found only an indirect relation from distributive justice to team identification via the performance climate. Furthermore, the relation between team identification and perceived social loafing was fully mediated by task cohesion, while we found a direct relation between team identification and self-reported social loafing (for a complete overview of the mediation tests, see Appendix 4).

Discussion

Our results partially supported Hypothesis 1a. They showed that only procedural justice of the coach was positively and substantially related to a mastery climate. More specifically, coach-related justice explained 32% of the variance of the mastery climate, which in turn positively predicted team identification. In addition, consistent with Hypothesis 1b, both distributive and procedural justice of the coach negatively predicted the performance climate. Perceived justice explained 21% of the performance climate, which in turn negatively predicted team identification. Taken together, coach-related justice and the coach-created motivational climate explained 23% of the variance of players' team identification.

Moreover, the results demonstrated that distributive justice is a smaller (or even a nonsignificant) predictor of the motivational climate and consequently team identification (Fig. 1). This finding is consistent with Tyler and Blader's (2003) group engagement model. This model hypothesizes that both procedural justice and distributive justice shape people's social identity within groups, which in turn influences attitudes and behaviors. According to this model, people are more strongly influenced by procedural justice judgments, while distributive justice has a supplementary, smaller influence on identity judgments.

The second aim of our research was to examine the relations between team identification, team cohesion, and social loafing. Consistent with previous research (De Backer et al., 2011), our model showed that team identification is an important predictor of both social and task cohesion (respectively, $\beta = 0.46$, $P < 0.01$; and $\beta = 0.40$, $P < 0.01$). In other words, identification seems to be the basic underlying predictor, while team cohesion is the behavioral consequence of this shared identity. Furthermore, in line with Høigaard et al. (2006) and Hypothesis 2a, we found that athletes with a strong team identification report higher levels of task cohesion and perceive less social loafing among their teammates. By contrast, Hypothesis 2b was not confirmed because task cohesion played no intervening role in the relation between team identification and self-reported loafing.

Study 2

To enhance the external validity of the research model, study 2 tested exact the same hypotheses on a sample of Norwegian world-class female handball players. These data were already collected in the context of previous research linking role ambiguity to athletes' social loafing (Høigaard et al., 2010). Taking into account that expertise is a factor that reduces the incidence of social loafing (Hardy & Crace, 1991), we expected low levels of social loafing among the world-class Norwegian handball players. However, at the top level, even the slightest amount of social loafing could have a serious impact on the team performance. Therefore, study 2 aimed to check if perceived justice of the coach can reduce and potentially eliminate social loafing in female world-class handball teams.

Furthermore, based on the findings that task reasons play a vital role for the group functioning of professional teams (Heuzé et al., 2006), we expected higher amounts of task cohesion in the Norwegian compared with the Belgian sample. Moreover, in line with Hypothesis 2a and 2b, but in contrast with the results of the Belgian sample, we expected that task cohesion would mediate the relation between team identification and both forms of social loafing in these world-class teams.

Method

Participants and procedures

The head coaches of the 12 Norwegian female first-division clubs were informed about the purpose of the study by telephone. Ten coaches allowed their teams to take part in the study, two coaches refused because of the high workload of their players. The final research sample consisted of 110 top-level female handball players (i.e., a response rate of 57%). The mean age of the participants was 22.8 years (SD = 4.0) and they trained on average 14.0 h per week (SD = 3.0). The number of years of experience at elite level ranged from 1 to 20 years (M = 4.0, SD = 3.6) and 18 participants had played for the Norwegian national team (i.e., Olympic champions 2008 and 2012) during the last 3 years.

Questionnaires were sent by post to a contact person within the club (i.e., a player or the team manager), who distributed them among the players together with an information letter. The players who agreed to participate were asked to put their completed questionnaires in a separate envelope, which was enclosed with the questionnaire. They sealed the envelope before returning it to the contact person within the club, who sent them to the authors. The APA ethical standards were followed in the conduct of the study. No rewards were given for participation; informed consent was obtained from all participants.

Measures

The same scales as in study 1 were used to measure the different concepts. We used the Norwegian versions of the PMCSQ, GEQ, PSLQ, and SRSLQ. The results of those scales proved to be valid and reliable in previous research with university team sport players (i.e., PMCSQ), junior male football players (i.e., GEQ), and top level female handball players (i.e., PSLQ and SRSLQ) (Roberts & Ommundsen, 1996; Høigaard et al., 2006, 2010; De Backer et al., 2011). Furthermore, the perceived justice and the team identification items were back-translated by two native-speaking Norwegian researchers.

Data analysis

The same analyses as in study 1 were performed, but we did not calculate DEs because our sample consisted of players from only 10 different teams, and it is not recommended to perform multilevel analyses based on such a small sample size at level two (Maas & Hox, 2005).

Results

Descriptive statistics, correlations, and scale reliabilities

Scales, means, SDs, and correlations for the variables of study 2 are provided in Table 2. Scale reliabilities (Cronbach's alphas) are provided on the diagonal. Furthermore, an analysis of variance showed that, compared with their Belgian counterparts, the Norwegian players reported significantly higher scores for performance climate, team identification, social cohesion, and task cohesion, and lower scores for perceived and self-reported social loafing.

CFA and multiple group measurement invariance testing

To construct identical scales for both studies, we used the same items as in study 1. All the Norwegian scales

Table 2. Scale, means (M), standard deviations (SDs), correlations, and Cronbach's alphas for all variables used in study 2

Variable	Scale	M	SD	1	2	3	4	5	6	7	8	9
1. Procedural justice	1–5	3.02	0.92	(0.85)								
2. Distributive justice	1–5	3.43	1.11	0.35*	(0.95)							
3. Performance climate	1–5	2.87	0.79	−0.45*	−0.38*	(0.85)						
4. Mastery climate	1–5	3.94	0.63	0.50*	0.44*	−0.44*	(0.78)					
5. Team identification	1–7	5.69	1.02	0.38*	0.17	−0.33*	0.44*	(0.85)				
6. Social cohesion	1–9	6.63	1.74	0.30*	−0.02	−0.28*	0.24*	0.52*	(0.78)			
7. Task cohesion	1–9	7.03	1.32	0.40*	0.07	−0.42*	0.50*	0.46*	0.48*	(0.70)		
8. Self-reported loafing	1–5	1.48	0.55	−0.26*	−0.03	0.39*	−0.26*	−0.31*	−0.37*	−0.33*	(0.74)	
9. Perceived social loafing	1–5	1.74	0.61	−0.28*	−0.14	0.29*	−0.34*	−0.31*	−0.19*	−0.40*	0.32*	(0.68)

Note. $^*P < 0.05$.

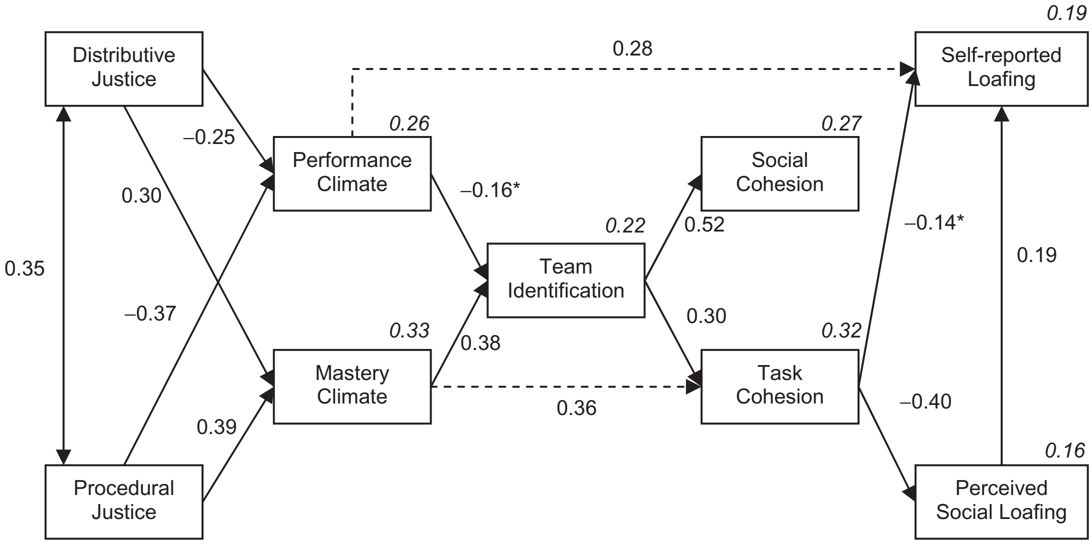

Fig. 2. Structural model best fitting the data of study 2 with the regression coefficients and the proportions explained variance. Legend: All coefficients presented are standardized and significant ($P < 0.05$) except for $^*P > 0.05$. The dashed arrows represent the added paths based on Amos modification indices and in line with previous research.

(except the GEQ) exhibited acceptable to good fit indices and significant factor loadings. Furthermore, all scales showed at least partial invariance across the Belgian and the Norwegian samples (for an overview of the results, see Appendices 1 and 2).

Confirmatory and exploratory path analysis

As in study 1, we used path models without the latent scores of the measurement models in order to decrease the ratio of cases (*N*) to the number of model parameters that require statistical estimates (see study 1). Furthermore, we are aware of the general rule of thumb that the minimum sample size should be no less than 200. However, the specific sample of top-level Norwegian female handball players has an upper limit lower than 200. Consequently, we followed the guideline of Hoyle and Kenny (1999) who stated that lower sample sizes can be used for models with no latent variables such as path models.

The confirmatory path analysis indicated a poor fit of the hypothesized model ($\chi^2 = 61.55$, df = 22, $P < 0.01$; GFI = 0.91; CFI = 0.84; RMSEA = 0.13). Consequently, an exploratory path analysis was performed using Amos's modification indices to add paths that would increase the model fit and were supported by previous research (Fig. 2). Two paths were added to our model: (a) a positive path from mastery climate to task cohesion; and (b) a positive path from performance climate to self-reported social loafing. The direct relation between mastery climate and task cohesion is in line with study 1 and a previous research of Heuzé et al. (2006). The positive relation between performance climate and self-reported social loafing is comparable with the findings of Høigaard and Ommundsen (2007), who found that a performance climate is positively related to social loafing. However, we had expected that this relation would run through team identification and cohesion.

Results of the exploratory path analysis provided an acceptable fit of the model to the data ($\chi^2 = 36.41$, df = 20, $P = 0.01$; GFI = 0.94; CFI = 0.94; RMSEA = 0.09). Examination of the revised model (Fig. 2) reveals that the paths between performance climate and team identification, and between task cohesion and self-reported social loafing were not significant. We also tested a model with two extra paths from social cohesion to both forms of social loafing. The results showed that social cohesion was not significantly related to social loafing (i.e., neither perceived nor self-reported social loafing). Furthermore, based on the significant correlations in the CFAs, we estimated and found the error terms of performance and mastery climate ($r = -0.22$, $P = 0.05$), and those of task and social cohesion ($r = 0.33$, $P < 0.05$) to be significantly related.

We followed the SEM approach by Holmbeck (1997) to test the mediational effects in the full structural model. More specifically, we tested the full structural model in two turns. First, we tested whether the relation between both forms of justice and team identification was mediated by the motivational climate (i.e., mastery and performance climate). Second, we tested whether the relation between team identification and both forms of social loafing was mediated by task cohesion. Based on our results, we can conclude that the relation between procedural justice and team identification was fully mediated by the mastery motivational climate, while we found only an indirect relation from distributive justice to team identification via this mastery climate. Furthermore, the relation between team identification and both perceived and self-reported social loafing was fully mediated by task cohesion (for a complete overview of the mediation tests, see Appendix 4).

Discussion

In line with Hypothesis 1a, the results among the Norwegian sample showed that coach-related justice was positively related to a mastery climate and explained 33% of its variance. This mastery climate in turn positively predicted athletes' team identification. Unlike the Belgian findings, the Norwegian results revealed a significant relation between distributive justice of the coach and the mastery climate. A possible explanation for this dissimilarity is the higher performance level of the Norwegian athletes. The Norwegian respondents were professional or semiprofessional athletes, who played in world-class teams in one of the strongest competitions of the world, whereas the Belgian athletes were amateurs, who practiced their sport at an elite level in the Belgian competitions. This difference is also reflected in the amount of training hours (i.e., 14.0 and 4.8 training hours per week, respectively). Consequently, the Norwegian players were probably more focused on performing and winning, and attached even greater importance to the selection of the best possible starting team and substitutes (i.e., distributive justice) than their Belgian counterparts. However, more research is necessary to understand the relations between justice and the motivational climate in greater depth.

Furthermore, coach-related justice was negatively related to a performance climate and explained 26% of its variance. In contrast with Hypothesis 1b, this performance climate was not significantly related to the athletes' team identification. We assume that this nonsignificance is caused by the smaller power of the Norwegian sample.

To conclude, both forms of coach-related justice were linked to the motivational climate (i.e., both mastery and performance climate) within the Norwegian teams, which in turn explained 22% of the variance in team identification. In line with study 1 and the group engagement model (Tyler & Blader, 2003), the path coefficients of distributive justice were clearly smaller than the path coefficients of procedural justice (see Fig. 2).

Supporting Hypothesis 2a, the results of the Norwegian sample showed a positive and substantial relation between team identification and task cohesion, which in turn negatively predicted perceived social loafing. These findings confirm the results of Everett et al. (1992) that task cohesion was negatively related to social loafing in female teams. However, in line with the Belgian results, the Norwegian ones did not confirm Hypothesis 2b. Task cohesion only indirectly predicted self-reported social loafing via its influence on perceived social loafing.

General discussion

To our knowledge, this is the first study that linked perceived fairness of the coach to a series of intervening variables leading to athletes' social loafing. In this general discussion, we draw some overall theoretical conclusions and provide implications for coaches in the field.

First, the results of both studies support the assumption that perceived fairness of the leader has the capacity to bridge the gap between leadership and followership (Haslam et al., 2011). For example, Tyler and Degoey (1995) showed that "the more fair community members perceived the authority to be, the more willing they were to forgo their own short-term individual gains for the benefit of the community as a whole" (Haslam et al., 2011, p. 115) and the more they were prepared to help other community members. Both the willingness to contribute to the community's effort and the readiness to help each other are important characteristics of a mastery climate. In line with these results, we demonstrated that the more a coach is perceived as fair, the more team players experience a mastery climate and the less a performance climate. In other words, when the coach is perceived as fair, players appear to focus less on their personal interests and the rivalry with their team members and more on the general good of the team by fulfilling their personal tasks within the team.

Furthermore, Haslam et al. (2011) emphasized the active nature of leadership. According to these authors, leaders have the ability to shape social identities (i.e., craft a sense of us) through the proactive process of shaping the social context. The results of both studies seem to confirm this statement and showed that fairness of the coach is an important antecedent of the motivational climate within the team, which in turn predicts athletes' team identification. One possible explanation for this positive relation may be the fact that a mastery climate facilitates cooperation between teammates. It fosters a sense that every member is respected and has an important role within the team (Heuzé et al., 2006). Based on our results and in line with the analysis by Haslam et al. (2011), we propose that coaches are entrepreneurs of the social identity of their players.

Second, consistent with previous research (Dutton et al., 1994; De Backer et al., 2011), our studies demonstrated that team identification positively predicts both forms of team cohesion. This can be attributed to the fact that team identification accentuates the similarities with other group members, which in turn leads to enhanced in-group cohesion (Dutton et al., 1994). In addition, both studies established that athletes' task cohesion was not only predicted by team identification, but also by a mastery climate. Consequently, coaches who successfully create a mastery climate, will foster a higher level of task cohesion within their team, both directly and indirectly (i.e., via an increased team identification).

Finally, in line with prior findings (Everett et al., 1992; Høigaard et al., 2006), both models suggest that in competitive sport teams, task and not social cohesion is a key predictor of athletes' social loafing. More precisely, our results indicated that task cohesion was negatively related to perceived social loafing, which in turn increased self-reported social loafing. By contrast, social cohesion did not emerge as a significant predictor of any form of social loafing at all. This differential impact of task vs social cohesion could possibly be explained by the fact that both samples consisted of high-level female athletes. Those players are focused on performance and winning. As a result, friendship relations among the players (i.e., social cohesion) may have only a minor impact on the players' effort on the field.

As with any research, the current studies had specific strengths and limitations. A first limitation is the cross-sectional nature of our data. Given that the coach–athlete relation is a dynamical process and that its quality fluctuates during the season, longitudinal studies are recommended to perform a more stringent test of the model.

A second limitation was the fact that based on Tyler and Blader's (2003) group engagement model, the current study focused only on distributive and procedural justice, while interactional justice was not measured. In order to examine the impact of interactional justice in sport teams, future research should extend our model with this form of justice.

Third, we had to rely on athletes' perceptions of social loafing apparent in other team members and on athletes' self-reported social loafing. Only few studies have been conducted on social loafing in interactive sports (Høigaard & Ommundsen, 2007; Høigaard et al., 2010), and to our knowledge, no valid objective measure of social loafing in such contexts has been reported. This gap in the literature can be explained by the fact that it is much harder to objectively determine individual effort in interactive sports (e.g., handball, volleyball). Nevertheless, it is important that future studies try to develop such (sport-specific) objective measures, in order to rule out the biases inherent to self-report scales.

Finally, the competitive level of the Norwegian athletes was considerably higher than the level of the Belgian athletes. Consequently, the differences in culture and competitive level were confounded. This overlap makes it hard to interpret the dissimilarities between the findings of both samples. Furthermore, because of the high workload of their players, some coaches did not allow their team to participate in our research. Consequently, we cannot exclude that these teams may not fit the proposed model. More research is necessary to examine the role of these potential confounds.

As a first strength, we want to underline that these studies were the first to examine the interrelations between OJT, the motivational climate (AGT), players' attitudes toward the team (SIA and cohesion), and behavioral outcomes (social loafing) in interactive sport teams. Our findings may offer important insights into the mechanisms that lead from coaches' justice to behavioral consequences among female senior team athletes.

A second strength is that almost all our hypotheses were confirmed in different cultural settings, at different competitive levels (including the absolute top-level), and in different interactive team sports (e.g., basketball, handball, soccer, and volleyball). In other words, the external validity of our findings seems to be quite high.

From a more practical standpoint, our comprehensive research model indicated that the coach can guide the group processes and attitudes of the players and consequently reduce social loafing, even in world-class teams. Previous research has shown that social loafing is responsible for a substantial portion of the decreased performance of groups (Karau & Williams, 1993). As a result, our model can be used to optimize team performance in (female) top-level interactive sport teams.

A first practical guideline that can be derived from our studies is that coaches should be aware of the importance of athletes' perceived justice. Research in the business context suggested that there are a number of strategies to increase employees' perception of fairness (Cropanzano & Greenberg, 1997). Two strategies that have been shown to be effective even when the person was disappointed with the outcomes he/she received are: (a) the application of Leventhal's rules; and (b) the provision of voice (Greenberg & Lind, 2000). Considering that coaches are not always able to provide the outcomes desired by an athlete (i.e., starting position), it seems valuable to transfer these strategies into team sport settings (for an overview of the sport-specific strategies, see Jordan et al., 2004).

A second guideline we recommend to coaches is to explicitly create a mastery climate. Highlighting the importance of every athlete's individual development within the team is one way to create such a mastery climate. It keeps them focused on improving their individual tasks. Furthermore, it reassures athletes that their efforts and individual improvements are valued by the team.

A third guideline is based on the fact that, particularly in world-class teams, increasing players' task cohesion seems crucial to decrease social loafing. Coaches should pay sufficient attention to the integration of the individual tasks in the overall team objectives by stressing the necessity of different individual contributions by individual team members with different qualities, in order to obtain the team goals. They should endorse the importance of cooperation to achieve the group's objectives (i.e., task cohesion) and persuade the players to set their individual goals for the benefit of the team goals.

Perspective

To conclude, our studies support the statement of Pensgaard and Roberts (2002) that coaches play an important role in the creation of a mastery climate in elite sport teams. More precisely, in two independent studies and in different team sports, we indicated that coaches' organizational justice strengthens a mastery climate. Furthermore, in line with previous research of Balaguer et al. (1999), we confirmed that such a mastery climate optimizes the functioning of athletes. Based on these results, sport psychologists can be advised to pay more attention to the concept of organizational justice. Obviously, also coaches should be aware of these results, as it seems that fair coach behavior can optimize the team performance as it creates an optimal motivational climate and reduces the amount of social loafing.

Key words: Coaching, cohesion, group dynamics, organizational justice, motivational climate, team identification.

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Appendix 1

Confirmatory factor analyses (CFA) and multiple group invariance testing

Because the establishment of partial measurement invariance is a prerequisite for meaningful comparisons across groups, multiple group invariance analyses were carried out to examine the measurement invariance of the scales across the Belgian and the Norwegian sample. To test for multigroup invariance of our measurement models we followed a series of hierarchical steps proposed by Byrne (2004, 2010). First, we performed CFAs to test whether the factorial structures of our measurement instruments are reliable and valid for the Belgian and the Norwegian sample separated and to check if both samples conceptualize the constructs in the same way (i.e., configural invariance). In this step, we added covariances between error terms if: (a) the content of the items associated with these error terms (partly) overlapped; and (b) this covariance significantly improved the fit of our model. As a result, we found slightly distinct models for the Belgian and the Norwegian sample (i.e., some covariances between the error terms differed between the samples). Subsequently, we used these separate baseline models in the following step of the multiple group invariance testing.

In the second step, we tested the validity of the baseline models across the two samples simultaneously (i.e., parameters are estimated for both groups at the same time). This is in contrast with step one, in which the validity of the baseline models was tested for each group separately. This simultaneously estimated model provides the baseline fit values against which all subsequently specified models were compared.

Third, we tested the invariance of the fully constrained model across Belgian and Norwegian players.

Regarding this application, we specified a model in which all factor loadings, all factor variances, and all factor covariances were constrained equal across Belgian and Norwegian players (i.e., strong factorial or scalar invariance). Testing for the equality of error variances across groups is considered to be excessively stringent (i.e., full measurement or strict invariance is unlikely to hold in practice) and therefore was not implemented in the current analyses. Partial measurement invariance allows appropriate cross-group comparisons, even if full measurement invariance is not obtained.

Finally, if the fully constraint model showed noninvariance across groups, we proceeded with a logically organized strategy to test which parameters were noninvariant. This strategy started with the testing of invariance of all factor loadings (i.e., test of weak factorial or metric invariance). Given evidence of noninvariance at this level, we tested the invariance of all factor loadings in each subscale (i.e., all factor loadings of one specific subscale). Given findings of noninvariance at the subscale level, we tested for the invariance of each factor loading separately. Due to the fact that the metric invariance is usually difficult to satisfy, researchers suggest that the cross-group comparisons can still be made if the noninvariant items constitute only a small portion of the model (Cheung & Rensvold, 2002). At last, we tested the invariance of the factor covariances (i.e., strong factorial or scalar invariance). Strong factorial invariance is a prerequisite for the comparison between groups, because it implies that the measurement scales have the same operational definition across groups (Cheung & Rensvold, 2002).

The absolute indices we used to evaluate overall model fit were: the normed chi-square (χ^2/df), the CFI, GFI, and RMSEA. A χ^2/df ratio of 3:1 or less indicates good fit; CFI and GFI values close to respectively .95 and .90 indicate good fit; RMSEA values close to .06 indicate good fit, and RMSEA values in the range of .08 to .10 indicate mediocre, but acceptable fit. Furthermore, the difference in chi-square between two nested models (i.e., χ^2 difference test) and the ΔCFI have been used as fit indices to calculate improvements over competing models. Significant results for the χ^2 difference test indicate that the model with smaller χ^2 has a statistically better fit. However, this test has the limitation that with large samples even very trivial differences yield a significant test result. Consequently, we used the difference test only as indicative of significant improvements. Based on a rigorous Monte Carlo study, Cheung and Rensvold (2002) indicated that ΔCFI is a robust statistic for testing the between-group invariance of CFA and that it is more reasonable to base invariance decisions on a difference in CFI rather than on χ^2 values. A value of ΔCFI smaller than or equal to .01 indicates that the null hypothesis of invariance should not be rejected.

Perceived justice

The CFA indicated a good fit of the hypothesized two-factor model of perceived justice for both the Belgian ($\chi^2 = 31.64$, df = 18, $P < .02$; χ^2/df = 1.76; GFI = 0.97; CFI = 0.99; RMSEA = 0.06) and the Norwegian sample ($\chi^2 = 16.10$, df = 19, $P < .65$; χ^2/df = 0.85; GFI = 0.96; CFI = 1; RMSEA = 0.00). In the Belgian sample we added a covariance between error 7 and error 8, while no covariances were added for the Norwegian sample.

The simultaneously estimated baseline model showed an excellent fit ($\chi^2 = 47.76$, df = 37, $P = 0.11$; χ^2/df = 1.29; GFI = 0.97; CFI = 0.995; RMSEA = 0.03). Consequently, we tested if the fully constrained model is invariant across the groups (i.e., test of strong factorial invariance). Goodness-of-fit statistics related to this fully constrained model also demonstrated excellent fit to the data ($\chi^2 = 59.93$, df = 46, $P < .08$; χ^2/df = 1.30; GFI = 0.96; CFI = 993; RMSEA = 0.03). Furthermore, both the χ^2 difference test ($\Delta\chi^2 = 12.17$, Δdf = 9, $P > 0.05$) and the ΔCFI (.002) indicated no significant difference between the baseline and the fully constrained models. Considering that the covariance between error 7 and error 8 in the Belgian sample was lacking in the Norwegian sample, we may conclude that the two-factor perceived justice scale is partial and strong factorial invariant across the Belgian and the Norwegian sample.

Motivational climate

The CFA indicated an acceptable fit of the hypothesized two-factor model of motivational climate for both the Belgian ($\chi^2 = 333.92$, df = 165, $P = 0.00$; χ^2/df = 2.02; GFI = 0.88; CFI = 0.894; RMSEA = 0.07) and the Norwegian sample separated ($\chi^2 = 250.902$, df = 165, $P = 0.00$; χ^2/df = 1.52; GFI = 0.80; CFI = 0.874; RMSEA = 0.08). In the Belgian sample we added covariances between errors 3 and 5, 4 and 6, 7 and 8, and 16 and 19. In the Norwegian sample we added covariance between errors 3 and 5, 4 and 6, 15 and 16, and 16 and 17.

The simultaneously estimated baseline model also showed an acceptable fit to the data ($\chi^2 = 585.38$, df = 330, $P = 0.00$; χ^2/df = 1.77; GFI = 0.85; CFI = 0.888; RMSEA = 0.05). Consequently, we tested if the fully constrained model is invariant across the groups (i.e., test of strong factorial invariance). Goodness-of-fit statistics related to this fully constrained model demonstrated acceptable fit to the data ($\chi^2 = 630.48$, df = 353, $P = 0.00$; χ^2/df = 1.79; GFI = 0.84; CFI = 0.878; RMSEA = 0.05). The χ^2 difference test ($\Delta\chi^2 = 45.1$, Δdf = 23, $P < 0.01$) indicated a significant difference between the baseline and the fully constrained model. As signed earlier, however, this test has some limitations and was used only as an indicative of significant improvements. Considering that ΔCFI (.01) indicated no significant difference between both models, we concluded that the two-factor motivational scale is partial and strong factorial invariant across the Belgian and the Norwegian sample.

Team identification

The CFA indicated an acceptable fit of the hypothesized one-factor model of team identification for the Belgian sample ($\chi^2 = 37.75$, df = 7, $P = 0.00$; χ^2/df = 5.39; GFI = 0.96; CFI = 0.949; RMSEA = 0.13) and a good fit for the Norwegian sample ($\chi^2 = 9.69$, df = 8, $P = 0.29$; χ^2/df = 2.21; GFI = 0.966; CFI = 1; RMSEA = 0.05). In the Belgian sample we added a covariance between error 1 and error 4, and between error 2 and error 4, while in the Norwegian sample only a covariance was added between error 1 and error 2.

The simultaneously estimated baseline model showed a good fit to the data ($\chi^2 = 47.41$, df = 15, $P = 0.00$; χ^2/df = 3.16; GFI = 0.95; CFI = 0.971; RMSEA = 0.08). Consequently, we tested if the fully constrained model is invariant across the groups (i.e., test of strong factorial invariance). Goodness-of-fit statistics related to this fully constrained model demonstrated also good fit to the data ($\chi^2 = 59.33$, df = 21, $P = 0.00$; χ^2/df = 2.83; GFI = 0.94; CFI = 966; RMSEA = 0.07). Furthermore, both the χ^2 difference test ($\Delta\chi^2 = 11.92$, Δdf = 6, $P > 0.05$) and the ΔCFI (.005) indicated no significant difference between the baseline and the fully constrained models. Consequently, we may conclude that the one-factor team identification scale is partial and strong factorial invariant across the Belgian and the Norwegian sample.

Team cohesion

The CFA indicated a poor fit of the hypothesized four-factor model of team cohesion for both the Belgian ($\chi^2 = 319.66$, df = 128, $P = 0.00$; χ^2/df = 2.50; GFI = 0.87; CFI = 0.767; RMSEA = 0.08) and the Norwegian sample ($\chi^2 = 227.29$, df = 127, $P = 0.00$; χ^2/df = 1.79; GFI = 0.78; CFI = 807; RMSEA = 0.09). Taking into account the suggestion that authors should retain the subscale structure, but also considering that poor subscales are usually dropped altogether when they do not function appropriately (Carron & Brawley, 2000), we decided to drop both subscales referring to individual attraction to the group (i.e., individual attraction to the group-social, and individual attraction to the group-task). Furthermore, based on nonsignificant factor loadings, we had to delete one item of the group integration-social subscale (i.e., "Members of our team would rather go out on their own than get together as a team") and one item of the group integration-task subscale (i.e., "Our team members do not communicate freely about each athlete's responsibilities during competition or practice"). As a result, the final measurement consisted of eight items equally distributed over the group integration-social and the group integration-task subscales. The CFA indicated a good fit of the adapted two-factor model of team cohesion for both the Belgian ($\chi^2 = 41.43$, df = 18, $P = 0.00$; χ^2/df = 2.30; GFI = 0.96; CFI = 0.930; RMSEA = 0.07) and the Norwegian sample ($\chi^2 = 17.63$, df = 17, $P = 0.41$; χ^2/df = 1.04; GFI = 0.96; CFI = 997; RMSEA = 0.02). In the Belgian sample we added a covariance between error 1 and error 2, while in the Norwegian sample a covariance was added between error 1 and error 5, and between error 2 and error 4.

The simultaneously estimated baseline model showed a good fit to the data ($\chi^2 = 67.54$, df = 35, $P = 0.00$; χ^2/df = 1.92; GFI = 0.95; CFI = 0.943; RMSEA = 0.05). Consequently, we tested if the fully constrained model is invariant across the groups (i.e., test of strong factorial invariance). Goodness-of-fit statistics related to this fully constrained model also demonstrated good fit to the data ($\chi^2 = 79.92$, df = 44, $P = 0.00$; χ^2/df = 1.82; GFI = 0.95; CFI = 937; RMSEA = 0.05). Both the χ^2 difference test ($\Delta\chi^2 = 12.38$, Δdf = 9, $P > 0.05$) and the ΔCFI (.006) indicated no significant difference between the baseline and the fully constrained models. As a result we may conclude that the adapted two-factor team cohesion scale is partial and strong factorial invariant across the Belgian and the Norwegian sample. Future research should examine in greater depth why the scores of the translated four-factor structure of the GEQ were unreliable and invalid with our Belgian and Norwegian samples. The translation of the individual items (especially those of both individual attraction to the group subscales) should be revised and retested. Until this happened, we suggest to use only the translated items of the group integration subscales.

Social loafing

The CFA indicated a good fit of the hypothesized two-factor model of social loafing for both the Belgian ($\chi^2 = 54.15$, df = 24, $P = 0.00$; χ^2/df = 2.26; GFI = 0.96; CFI = 0.955; RMSEA = 0.07) and the Norwegian sample separated ($\chi^2 = 40.98$, df = 25, $P = 0.02$; χ^2/df = 1.64; GFI = 0.92; CFI = 925; RMSEA = 0.08). In the Belgian sample we added a covariance between error 1 and error 4, and between error 7 and error 8, while for the Norwegian sample a covariance was added between error 6 and error 7.

The simultaneously estimated baseline model showed a good fit ($\chi^2 = 95.10$, df = 49, $P = 0.00$; χ^2/df = 1.94; GFI = 0.95; CFI = 0.946; RMSEA = 0.05). Consequently, we tested if the fully constrained model is invariant across the groups (i.e., test of strong factorial invariance). Goodness-of-fit statistics related to this fully constrained model also demonstrated a good fit to the data ($\chi^2 = 135.11$, df = 59, $P = 0.00$; χ^2/df = 2.29; GFI = 0.93; CFI = 910; RMSEA = 0.06). However, both the χ^2 difference test ($\Delta\chi^2 = 40.01$, $\Delta df = 10$, $P < 0.05$) and the ΔCFI (.036) indicated a significant difference between the baseline model (i.e., Model 1) and the fully constrained model. Consequently, we examined which factor loadings and covariances were noninvariant across the Belgian and the Norwegian sample. Turning to the table in Appendix 2, we see that Model 2A, with all factor loadings constrained equal, is significantly worse than Model 1 ($\Delta\chi^2 = 26.396$, $\Delta df = 7$, $P < 0.01$; $\Delta CFI = 0.023$). Given findings of noninvariance at this level, we proceeded to test for the invariance of all factor loadings comprising each subscale. Therefore, in Model 2B we placed equality constraints on all freely estimated factor loadings associated with the factor perceived social loafing (PSL). The comparison with Model 1 (see Appendix 2) advises us that all items of the perceived social loafing subscale are operating equivalently across the two groups of players. Provided with this information, we now know that the problematic items are housed in the self-reported social loafing subscale. Subsequently, we tested one factor loading at a time within this subscale (i.e., see model 2C, 2D, and 2E). Our findings reveal evidence of noninvariance related to item 6 and item 8 (see Appendix 2, results for Model 2C and 2E). From these findings we learn that items 6 and 8 of the self-reported social loafing scale are operating somewhat differently for the Belgian and Norwegian players. Item 6 suggests that the players try as hard as they could, and item 8 states that the players contributed less than they should. A possible explanation could be that the Belgian and the Norwegian players interpret "I try as hard as I could/should" different. It could be that the world-class female handball players were more severe for themselves when they reported their own level of social loafing than their Belgian counterparts. However, further research should examine these differences in greater depth. Based on the fact that the noninvariant items constitute only a small portion of the model, cross-group comparisons can still be made (Cheung & Rensvold, 2002).

After establishing the equivalence of the measurement model, the next step in the process is to test for invariance related to the structural portion of the model (i.e., strong factorial invariance). These tests can involve the factor variances as well as the factor covariances. However, in line with many researchers, we considered the latter of most interest. In particular, testing for the invariance of the factor covariance between PSL and SRSL addresses the similarity of the underlying theoretical structure across both samples. Results for this structural invariance test (see Appendix 2, Model 3), revealed that the factor covariance is equivalent across Belgian and Norwegian players. As a result, we can conclude that despite the two noninvariant items, the social loafing scale is partial and strong factorial invariant across the Belgian and the Norwegian sample.

Appendix 2

Table A1. Goodness-of-fit statistics for tests of multigroup invariance of the social loafing scale

Model description	Comparative model	χ^2	df	$\Delta\chi^2$	Δdf	Statistical significance	CFI	ΔCFI
1. Baseline model; no equality constraints imposed (i.e., configural invariance)	—	95.095	49	—	—	—	0.946	—
2. Measurement model (i.e., weak factorial or metric invariance); (Model A) All factor loadings constrained equal	2A vs 1	121.491	56	26.396	7	$P < 0.01$	0.923	0.023
(Model B) Factor loadings for only PSL constrained equal	2B vs 1	103.063	53	7.968	4	NS	0.941	0.005
(Model C) Model B with factor loading for Item 6 (SRSL) constrained equal	2C vs 1	111.276	54	16.181	5	$P < 0.01$	0.933	0.013
(Model D) Model B with factor loading for Item 7 (SRSL) constrained equal	2D vs 1	106.696	54	11.601	5	$P < 0.05$ $P > 0.01$	0.938	0.008
(Model E) Model B with factor loadings for Item 7 and 8 (SRSL) constrained equal	2E vs 1	112.893	55	17.798	6	$P < 0.01$	0.932	0.014
3. Structural model (i.e., strong factorial or scalar invariance); Model 2.D with covariance among PSL and SRSL constrained equal	3 vs 1	107.040	55	11.945	6	NS	0.939	0.007

Note. $\Delta\chi^2$, difference in χ^2 values between models; Δdf , difference in number of degrees of freedom between models; ΔCFI, difference in CFI values between models; PSL, perceived social loafing; SRSL, self-reported social loafing.

Appendix 3

Table A2. Intra-class correlations (ICC's) and design effects of Study 1

Construct	ICC	Design effect
Procedural justice	0.27	2.26
Distributive justice	0.00	1.00
Performance climate	0.19	1.91
Mastery climate	0.30	2.43
Team identification	0.03	1.16
Social cohesion	0.28	2.31
Task cohesion	0.15	1.73
Perceived social loafing	0.00	1.01
Self-reported social loafing	0.12	1.58

Note. Design Effect = 1 + [Average within group sample size − 1] × ICC. Average within group sample size = 5.76.

Appendix 4

Test of the mediational effects: Belgian sample

We followed the SEM approach by Holmbeck (1997) to test the mediational effects in the full structural model. More specifically, we did this in two turns (a) justice (i.e., distributive and procedural justice) – motivational climate (i.e., mastery and performance climate) – team identification; and (b) team identification – task cohesion – social loafing (i.e., perceived and self-reported social loafing).

Justice – motivational climate – team identification

We tested whether the relation between justice and team identification is mediated by the motivational climate. As first step, we tested a model estimating the direct paths from predictor to outcome variables. However, a model with only the direct relations between both justice concepts and team identification is just-identified. Just-identified models are scientifically uninteresting because they have no degrees of freedom and can never be rejected. Consequently, we have chosen for a model from distributive and procedural justice to team identification and from team identification directly to perceived and self-reported social loafing (i.e., the second part of the mediational tests). This model provided a good fit to the data ($\chi^2 = 7.51$, df = 3, $P = 0.06$; GFI = 0.99; CFI = 0.98; RMSEA = 0.08). The regression path between procedural justice and team identification was significant ($\beta = 0.25$, $P < 0.05$), but the regression path from distributive justice to team identification was not significant ($\beta = 0.13$, $P > 0.05$). As such, the mediational condition of significant associations between predictor and outcome variables was not satisfied for the relation between distributive justice and team identification.

In the second step of the Holmbeck's (1997) SEM approach, we estimated the model fit of a constrained model with no direct regression paths from predictor to outcome variables. This constrained model from justice via motivational climate to team identification showed a good fit ($\chi^2 = 5.13$, df = 3, $P = 0.04$; GFI = 0.99; CFI = 0.99; RMSEA = 0.05). Furthermore, all regression coefficients except the one from distributive justice to mastery climate were significant ($\beta = 0.10$, $P > 0.05$). Consequently, the mediational conditions of significant associations between the predictor and mediator and between the mediator and the outcome variables were satisfied, except for the relation between distributive justice and mastery climate.

In the third step of the Holmbeck's (1997) SEM approach, we examined an unconstrained model, allowing also direct regression paths between the predictors (i.e., procedural and distributive justice) and the outcome variable (i.e., team identification). The unconstrained model had an acceptable fit to the data ($\chi^2 = 3.93$, df = 1, $P = 0.29$; GFI = 0.99; CFI = 0.99; RMSEA = 0.11). However, the regression paths from distributive and procedural justice to team identification were nonsignificant (respectively $\beta = 0.06$, $P > 0.05$; and $\beta = 0.02$, $P > 0.05$).

The final step to determining mediation was to examine whether the unconstrained model provides a significant improvement in fit over the constrained model (Holmbeck, 1997). Taking into account the nonsignificant direct paths in the unconstrained model, this step was not strictly necessary. Nevertheless, we performed a chi-square difference test between the less (i.e., unconstrained) and the more restrictive (i.e., constrained) models. The chi-square difference test indicated no significant difference between the constrained and the unconstrained models [$\Delta\chi^2$ (2) = 1.2; $P > 0.05$]. Consequently, the constraint model was not improved by adding the direct paths from justice to team identification.

Based on these steps we can conclude that the relation between procedural justice and team identification is fully mediated by both mastery and performance climate. However, we found only an indirect relation from distributive justice to team identification via performance climate (i.e., there was no direct relation between distributive justice and team identification and the relation between distributive justice and mastery climate was not significant).

Team identification – task cohesion – social loafing

We tested whether the relation between team identification and both forms of social loafing is mediated by task cohesion. First, due to identification issues, we used the same model as in the first step of the previous mediation test (i.e., direct paths from distributive and procedural justice to team identification and from team identification directly to perceived and self-reported social loafing). This model provided a good fit to the data ($\chi^2 = 7.51$, df = 3, $P = 0.06$; GFI = 0.99; CFI = 0.98; RMSEA = 0.08). The regression paths between team identification and perceived and self-reported social loafing were significant (respectively $\beta = -0.22$, $P < 0.05$; $\beta = -0.37$, $P < 0.05$). As such, the mediational condition of significant associations between predictor and outcome variables was satisfied.

In the second step of the Holmbeck's (1997) SEM approach, we estimated the model fit of a constrained model with no direct regression paths from predictor to outcome variables. This constrained model from team identification via task cohesion to both forms of social loafing showed a moderate fit ($\chi^2 = 33.76$, df = 4, $P = 0.00$; GFI = 0.95; CFI = 0.89; RMSEA = 0.17). Furthermore, all regression coefficients were significant. Consequently, the mediational conditions of significant associations between the predictor and mediator and between the mediator and the outcome variables were satisfied.

In the third step of the Holmbeck's (1997) SEM approach, we examined an unconstrained model, allowing also direct regression paths between the predictor (i.e., team identification) and the outcome variables (i.e., perceived and self-reported social loafing). The unconstrained model had an excellent fit to the data ($\chi^2 = 2.86$, df = 2, $P = 0.24$; GFI = 1; CFI = 1; RMSEA = 0.04). However, the regression path from team identification to perceived social loafing and the path from task cohesion to self-reported social loafing were nonsignificant (respectively $\beta = -0.02$, $P > 0.05$; and $\beta = 0.04$, $P > 0.05$).

The final step to determining mediation was to examine whether the unconstrained model provides a significant improvement in fit over the constrained model (Holmbeck, 1997). We performed a chi-square difference test between the less (i.e., unconstrained) and the more restrictive (i.e., constrained) models. The chi-square difference test indicated a significant difference between the constrained and the unconstrained models [$\Delta\chi^2$ (2) = 30.9; $P < 0.05$]. Consequently, the constraint model was improved by adding the direct paths from team identification to social loafing (i.e., only the path between team identification and self-reported loafing was significant).

Based on these steps we can conclude that the relation between team identification and perceived social loafing is fully mediated by task cohesion (i.e., the direct relation between team identification and perceived social loafing was not significant). However, the relation between team identification and self-reported loafing was not mediated by task cohesion (i.e., when the direct relation between team identification and self-reported loafing was added, the path between task cohesion and self-reported loafing was no longer significant).

Test of the mediational effects: Norwegian sample

We followed the SEM approach by Holmbeck (1997) to test the mediational effects in the full structural model. More specifically, we did this in two turns (a) justice (i.e., distributive and procedural justice) – motivational climate (i.e., mastery and performance climate) – team identification; and (b) team identification – task cohesion – social loafing (i.e., perceived and self-reported social loafing).

Justice – motivational climate – team identification

We tested whether the relation between justice and team identification is mediated by the motivational climate. As first step, we tested a model estimating the direct paths from distributive and procedural justice to team identification and from team identification directly to perceived and self-reported social loafing. Analogous to the tests in the Belgian sample, we had to combine these models to become an over-identified model. This model provided a good fit to the data ($\chi^2 = 6.13$, df = 4, $P = 0.19$; GFI = 0.98; CFI = 0.96; RMSEA = 0.07). The regression path between procedural justice and team identification was significant ($\beta = 0.37$, $P < 0.05$), but the regression path from distributive justice to team identification was not significant ($\beta = 0.04$, $P > 0.05$). As such, the mediational condition of significant associations between predictor and outcome variables was not satisfied for the relation between distributive justice and team identification.

In the second step of the Holmbeck's (1997) SEM approach, we estimated the model fit of a constrained model with no direct regression paths from predictor to outcome variables. This constrained model from justice via motivational climate to team identification showed an acceptable fit ($\chi^2 = 8.96$, df = 3, $P = 0.03$; GFI = 0.97; CFI = 0.95; RMSEA = 0.14). Furthermore, all regression coefficients except the one from performance climate to team identification were significant ($\beta = -0.17$, $P > 0.05$). Consequently, the mediational conditions of significant associations between the predictor and mediator and between the mediator and the outcome variables were satisfied, except for the relation between performance climate and team identification.

In the third step of the Holmbeck's (1997) SEM approach, we examined an unconstrained model, allowing also direct regression paths between the predictor (i.e., procedural and distributive justice) and the outcome variables (i.e., team identification). The unconstrained model had an acceptable fit to the data ($\chi^2 = 5.07$, df = 1, $P = 0.02$; GFI = 0.98; CFI = 0.96; RMSEA = 0.19). However, the direct regression paths from distributive and procedural justice to team identification were nonsignificant (respectively $\beta = 0.06$, $P > 0.05$; and $\beta = 0.02$, $P > 0.05$).

The final step to determining mediation was to examine whether the unconstrained model provides a significant improvement in fit over the constrained model (Holmbeck, 1997). Taking into account the nonsignificant direct paths in the unconstrained model, this step was not strictly necessary. Nevertheless, we performed a chi-square difference test between the less (i.e., unconstrained) and the more restrictive (i.e., constrained) models. The chi-square difference test indicated no significant difference between the constrained and the unconstrained models [$\Delta\chi^2$ (2) = 2.89; $P > 0.05$]. Consequently, the constraint model was not improved by adding the direct paths from justice to team identification.

Based on these steps we can conclude that the relation between procedural justice and team identification is fully mediated by mastery climate. However, we found only an indirect relation from distributive justice to team identification via mastery climate (i.e., there was no direct relation between distributive justice and team identification and the relation between performance climate and team identification was not significant).

Team identification – task cohesion – social loafing

We tested whether the relation between team identification and both forms of social loafing is mediated by task cohesion. First, due to identification issues, we used the same model as in the first step of the previous mediation test (i.e., direct paths from distributive and procedural justice to team identification and from team identification directly to perceived and self-reported social loafing). This model provided a good fit to the data ($\chi^2 = 7.51$, df = 3, $P = 0.06$; GFI = 0.99; CFI = 0.98; RMSEA = 0.08). The regression paths between team identification and perceived and self-reported social loafing were significant (respectively $\beta = -0.31$, $P < 0.05$; $\beta = -0.24$, $P < 0.05$). As such, the mediational condition of significant associations between predictor and outcome variables was satisfied.

In the second step of the Holmbeck's (1997) SEM approach, we estimated the model fit of a constrained model with no direct regression paths from predictor to outcome variables. This constrained model from team identification via task cohesion to both forms of social loafing showed an acceptable fit ($\chi^2 = 9.19$, df = 4, $P = 0.06$; GFI = 0.97; CFI = 0.95; RMSEA = 0.11). Furthermore, all regression coefficients were significant. Consequently, the mediational conditions of significant associations between the predictor and mediator and between the mediator and the outcome variables were satisfied.

In the third step of the Holmbeck's (1997) SEM approach, we examined an unconstrained model, allowing also direct regression paths between the predictor (i.e., team identification) and the outcome variables (i.e., perceived and self-reported social loafing). The unconstrained model had a good fit to the data ($\chi^2 = 3.73$, df = 2, $P = 0.16$; GFI = 0.99; CFI = 0.98; RMSEA = 0.09). However, the regression paths from team identification to perceived ($\beta = -0.16$, $P > 0.05$) and self-reported social loafing ($\beta = -0.18$, $P > 0.05$) were nonsignificant.

The final step to determining mediation was to examine whether the unconstrained model provides a significant improvement in fit over the constrained model (Holmbeck, 1997). Taking into account the nonsignificant direct paths in the unconstrained model, this step was not strictly necessary. Nevertheless, we performed a chi-square difference test between the less (i.e., unconstrained) and the more restrictive (i.e., constrained) models. The chi-square difference test indicated no significant difference between the constrained and the unconstrained models [$\Delta\chi^2$ (2) = 5.46; $P > 0.05$]. Consequently, the constraint model was not improved by adding the direct paths from team identification to social loafing.

Based on these steps we can conclude that the relations between team identification and both perceived and self-reported social loafing are fully mediated by task cohesion.