I Z A Institute of Labor Economics

Initiated by Deutsche Post Foundation

DISCUSSION PAPER SERIES

IZA DP No. 12726

Female Workers, Male Managers: Gender, Leadership, and Risk-Taking

Ulf Rinne
Hendrik Sonnabend

OCTOBER 2019

I Z A Institute of Labor Economics

Initiated by Deutsche Post Foundation

DISCUSSION PAPER SERIES

IZA DP No. 12726

Female Workers, Male Managers: Gender, Leadership, and Risk-Taking

Ulf Rinne
IZA

Hendrik Sonnabend
University of Hagen

OCTOBER 2019

ISSN: 2365-9793

IZA – Institute of Labor Economics

Schaumburg-Lippe-Straße 5–9
53113 Bonn, Germany

Phone: +49-228-3894-0
Email: publications@iza.org

www.iza.org

ABSTRACT

Female Workers, Male Managers: Gender, Leadership, and Risk-Taking*

This study examines gender differences in risk-taking behavior among managers in a female-dominated industry. Using data from international top-level women's soccer, we provide evidence that male coaches show a lower level of risk-taking than female coaches on average. We also find a U-shaped age effect that is independent of gender, meaning that young and more mature individuals tend to take riskier decisions. Our main results therefore strongly contrast with the majority of previous studies on gender differences in risk preferences, and thereby emphasize the importance of considering the industrial environment. Underlying selection processes may play an important role. We find no correlation between the gender gap in risk-taking and female empowerment defined by national gender equality scores.

JEL Classification: D81, J16, J4, M12, Z29

Keywords: gender, risk-taking, leadership, management, female empowerment

Corresponding author:
Ulf Rinne
IZA - Institute of Labor Economics
Schaumburg-Lippe-Straße 5–9
53113 Bonn
Germany
E-mail: rinne@iza.org

* We thank Thomas Peeters, Jürgen Weibler, and participants in the IZA internal seminar in Bonn, as well as participants at the SESM Conference 2019, the EALE Conference 2019 and the VfS Annual Conference 2019 of for helpful comments and suggestions on earlier drafts. In addition, we thank Mathias Tirtasana and Julian Herwig for their assistance in preparing the data. Financial support from the *FernUniversität in Hagen* is gratefully acknowledged. All remaining errors are our own.

1 Introduction

Women and men behave differently, with important implications for economic outcomes. For example, gender differences in competitiveness are a prominent explanation for persistent gender gaps in wages, earnings and social positions as women "shy away from competition" (Niederle and Vesterlund, 2007). These differences may in turn be closely related to gender differences in risk preferences and risk-taking (e.g., Bertrand, 2011; Blau and Kahn, 2017). Indeed, many experiments or larger surveys – the latter often representative of the entire population – find that men are generally more willing to take risk than women (see, e.g., Croson and Gneezy, 2009; Eckel and Grossman, 2008; Falk et al., 2018).

However, gender differences in risk-taking may actually be very different in real-life situations, and in particular when considering a professional environment.[1] For example, results are very mixed at the management level, which has – among other things – triggered the so-called "Lehman Sisters hypothesis" controversy (van Staveren, 2014).[2] More generally, when acting in the same professional environment, any gender differences in risk preferences may actually be counteracted by expert knowledge in the respective field, the specific framing under which decisions are taken, and selection mechanisms that put women and men in leadership positions. Although Beckmann and Menkhoff (2008) reject this "expertise dominates gender" hypothesis for their sample of funds managers in the financial industry, it could hold in other contexts.

We analyze gender differences in risk-taking among managers in a female-dominated industry, which nonetheless exhibits the typical feature of a very low share of women in advanced leadership positions. Female leadership remains a rare phenomenon, even in professions in which women constitute the majority of workers. To give only one example, Hoss et al. (2011) document that in the US health care sector – in which 75 percent of the workforce is female – women represent only 12 percent of CEOs. This

[1]Moreover, Filippin and Crosetto (2016) conclude from their meta-analysis of Holt and Laury (2002) replications that gender differences in risk-taking are strongly task-specific.

[2]In fact, Eckel and Füllbrunn (2015) demonstrate experimentally that all-male asset markets show a higher tendency of price bubbles than all-female asset markets. Additionally, Barber and Odean (2001) provide evidence of a gender gap in risk-taking related to common stock investments.

is the case even though female leadership can be productivity-enhancing and women can have a leadership advantage in some contexts. For instance, De Paola et al. (2018) ran a field experiment with students and show that female-led teams perform significantly better than male-led teams. Delfgaauw et al. (2013) document that teams in sales competitions are more successful if managers and subordinate employees have the same gender.

We focus on a particular industry, namely international top-level women's soccer, where workers (i.e. players) are by definition only women, but managers (i.e. coaches) can be either women or men, to address two research questions:

- Is leadership in a female-dominated industry characterized by gender differences in the willingness to take risks?
- Are such gender differences country-specific and if so, can they be explained by a nation's female empowerment level?

Working with sports data is generally appealing as they offer a wealth of information, include highly standardized tasks, and strong incentives for individuals. In addition, Pieper et al. (2014) and van Ours and van Tuijl (2016) highlight similarities between head coaches and top managers in terms of age, stress resistance, media skills, and accountability by a large number of stakeholders.

Our specific setting offers two additional advantages. First, while most of the field studies on gender difference in sports competitions rely on data from single-sex contests, we study mixed-sex competitions among individuals (not on the pitch, but rather at the management level). This is important because "the gender gap in competitiveness is smaller or absent in single-sex competitions" (Niederle and Vesterlund, 2011, p. 615). Second, while other studies using sports data often analyze risk measures that are in fact the joint outcomes of players' and coaches' decisions (for example, in basketball, see Grund et al., 2013), we study outcomes that can be unambiguously linked to head coaches' decisions alone.

2 Institutional Background

2.1 Coaching a women's soccer team

Despite nowadays being widely established and recognized, women's soccer can still be considered as a relatively recent phenomenon. For example, the English Football Association (FA) enacted a ban in the 1920s, which remained in effect until 1971 (Grainey, 2012). In Germany, a similar ban remained in effect until 1970. Indeed, even when these bans were lifted, women's soccer still led a niche existence for a longer period. For instance, the first official international match involving the German women's national soccer team was held in 1982, and the German women's national soccer league was founded in 1990.

Leadership in women's soccer is male-dominated. This applies to positions in senior governance and senior operations, but also extends to the level of head coaches. For example, in 2014 about two-thirds of senior coaching positions in women's national senior teams (head and assistant coaches), under-19s and under-17s teams were held by (white) men, while the remaining one-third of positions of this kind were held by (white) women (Bradbury et al., 2014). A similar share of about one-third of all head coaches being female is present in women's college soccer in the United States (Wicker et al., 2019). This under-representation of female coaches cannot be explained by performance differences, with Gomez-Gonzalez et al. (2018) showing that gender does not have a significant effect on team performance.

How do individuals become women's soccer coaches? In order to gain insights into the selection process into leadership positions in this particular industry, we conducted a small-scale survey.[3] Accordingly, about half of the sixteen coaches who responded stated that they received an offer to coach a women's soccer team as the main reason why they became a women's soccer coach (43.75 percent). This could indicate that the selection process is largely demand driven. However, a substantial share of our surveyed

[3] More specifically, we conducted an online survey in autumn 2018 among the coaches who are part of our sample (see Section 3 for details). As expected, the response rate was rather low and it is therefore only possible to interpret some results in a predominantly qualitative manner.

coaches also indicated their preferences for being a women's coach over being a men's coach (37.5 percent). Only one coach answered that a position as a men's soccer coach was not available (6.25 percent). No coach responded that climbing the job ladder was easier as a women's coach.

All surveyed coaches indicated having played soccer by themselves, at either a professional (i.e. national team or first division; 31.25 percent), semi-professional (i.e. second or third division; 56.25 percent), or an amateur level (i.e. fourth division or below; 12.5 percent). For the vast majority of surveyed coaches, their job as a soccer coach is their major source of income (75 percent). However, weekly hours of work are quite heterogeneous, ranging from 40 hours per week and more (37.5 percent) to less than 15 hours per week (31.25 percent). The remaining surveyed coaches indicated working an intermediate number of hours per week.

In order to gain more insights into the selection process into the profession, we also conducted a biographical analysis of the women's soccer coaches who are included in our sample. Accordingly, we observe the following patterns. The vast majority of female coaches appear to be either former active players and/or to have a closely-related professional background (e.g., as a sports scientist). The same holds for male coaches, although very often additional factors come into play that may explain their interest in women's soccer (e.g., family ties). Moreover, quite a few male coaches had previously coached men's soccer teams, sometimes not very successfully.

Hence, the selection process into the profession indeed appears to be different for male and female coaches. Importantly, there are still hardly any women coaching a men's soccer team. It also appears while some (male) coaches become women's soccer coaches after having previously coached male soccer teams, the opposite virtually never occurs. However, it should be noted that the general requirements for a women's soccer coach are in fact very similar to a men's soccer coach (e.g., training style, qualifications). At the top level, for example, the "UEFA A Licence" is required as a coaching license, at least in Germany.[4]

[4]See, for example, https://www.thepfa.com/coaching/courses/qualifications for a description of the coaches' "qualification pyramid," which is generally the same for male and female coaches.

2.2 UEFA Women's Champions League

We focus on international top-level women's soccer played in the UEFA Women's Champions League, an international soccer competition for club teams from countries that are affiliated with the Union of European Football Associations (UEFA). The competition was first played in the 2001-02 season under the name "UEFA Women's Cup," and was relaunched in the 2009-10 season under its current name.

The competition is open to the previous season's national champions of all 55 UEFA associations, while the top eight associations (according to the ranking of UEFA league coefficients) are allowed to send two teams. The number of participating teams has varied over time (ranging from 51 starters in 2010-11 to 61 starters in 2017-18), because – for example – not all associations have had a women's soccer league. Due to the varying number of participants, the precise tournament mode has been slightly adapted over time, while the general setting has remained constant since 2009-2010. Table 1 shows the precise mode for each year:

- The competition comprises a group stage and a final stage.
- A varying number of teams (depending on the number of participants in a given year) play in the group stage. These teams play against each other in groups of four teams (one-off games). The location is the same for all games of a given group (one team in that group serves as the host).
- The group winners advance to the competition's final stage, and – depending on the total number of teams – in some years the best runner(s)-up can also continue.
- 32 teams reach the competition's final stage (which thus starts with the round of 32). Teams that qualified in the group stage are joined by top-seeded teams that only enter the competition at this stage. The final stage is a single-elimination tournament with two-legged games in each round, except for the final round, which is a one-off game.

The fact that the "UEFA A Licence" is required as a coaching license in Germany, is stated here: https://www.dfb.de/sportl-strukturen/trainerausbildung/qualifizierung/.

Table 1: Tournament mode of the UEFA Women's Champions League.

Season	Teams	Associations	Qualifying Round / Group Stage	Round of 32 to Semi-Final	Final
2009-10	53	44	28 teams in 7 groups, only winners advance to round of 32	Two-legged games	One-off game
2010-11	51	43	28 teams in 7 groups, winners and 2 best runners-up* advance to round of 32	Two-legged games	One-off game
2011-12	54	46	32 teams in 8 groups, winners and 2 best runners-up* advance to round of 32	Two-legged games	One-off game
2012-13	54	46	32 teams in 8 groups, winners and 2 best runners-up* advance to round of 32	Two-legged games	One-off game
2013-14	54	46	32 teams in 8 groups, winners and 2 best runners-up* advance to round of 32	Two-legged games	One-off game
2014-15	54	46	32 teams in 8 groups, winners and 2 best runners-up* advance to round of 32	Two-legged games	One-off game
2015-16	56	46	32 teams in 8 groups, only winners advance to round of 32	Two-legged games	One-off game
2016-17	59	47	36 teams in 9 groups, only winners advance to round of 32	Two-legged games	One-off game
2017-18	61	49	40 teams in 10 groups, winners and best runner-up* advance to round of 32	Two-legged games	One-off game

*To determine the best runner(s)-up, only group games played against teams finishing first and third are considered.

3 Data Set

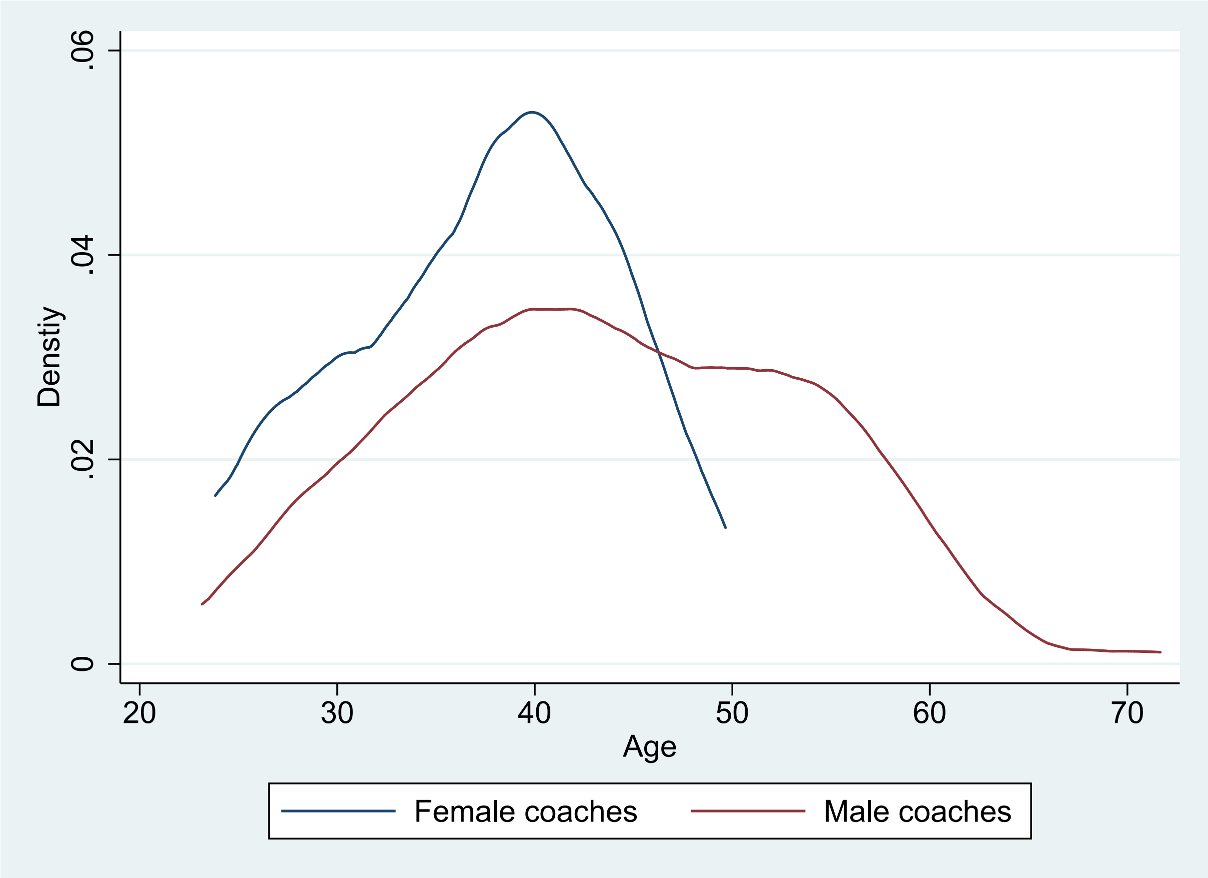
Our data stem from the UEFA Women's Champions League, including nine seasons from 2009-10 to 2017-18. They include information on 945 matches, in which 132 clubs from 49 different countries were involved. Along with detailed information on the match, club, and player level, we have information on coaches' name, gender, age, and nationality for 1,586 match-team-coach combinations, involving information on 3,538 substitutions.[5] With a share of 16.4 percent among the 214 coaches in our sample, the proportion of female coaches corresponds to the typical low share of women in leadership positions that can also be found in many other (female-dominated) industries.

A first obvious gender difference in our data relates to the coaches' age structure. As Figure 1 indicates, while female coaches are on average younger (mean: 36.82, SD: 6.99), the age of male coaches is more evenly distributed (mean: 42.98, SD: 9.73).[6] The maximum age of a female coach in our sample is 50 years, while male coaches are as old as 71 years. The different age structure by gender appears to be related to social stigma and the aforementioned bans, due to which women's soccer has only rather recently achieved a (semi-)professional level. As a result, the pool of former female

[5]This information was extracted from the website *us.soccerway.com*. Note that information on the player level also includes her primary position (i.e. goalkeeper, defender, midfielder, or striker). This position is assigned at a fixed date and hence there are no changes over time in this respect.

[6]Note that the density estimation is based on all observations in our data.

Figure 1: Gender differences in the head coaches' age structure.



soccer players – from which female women's soccer coaches are predominantly recruited – is still markedly younger than the corresponding pool of former male soccer players. However, since age has a significant effect on the willingness to take risks (see, e.g., Dohmen and Falk, 2011; Falk et al., 2018), and because this effect may be differently pronounced depending on gender (e.g., Flory et al., 2018), we need to carefully control for the coaches' age in our subsequent empirical analysis.

Finally, there are 24 individuals in our sample whose nationality differs from the country in which their club is located in at least one season. With four female coaches, the share of women is the same in this sub-sample as in the entire sample. Hence, with our data we cannot confirm the general impression that male individuals show a stronger tendency for work-related migration decisions (see, e.g., Altonji and Blank, 1999).

4 Empirical Analysis

According to Hvide (2002, p. 883), increasing the risk in a competition means inducing "a mean-preserving spread of Y_i through increasing the variance of ϵ_i," where the output Y_i of contestant i depends on the realization of a shock ϵ_i with $E\left(\epsilon_i^2\right) = \sigma^2$. In other words, a higher risk is associated with a greater variance σ^2.

In this paper, we follow Grund and Gürtler (2005) and focus on two moments of decision-making to analyze risk-taking behavior in soccer, namely the choice of players for the starting lineup, and the substitution of players within a match. We define starting lineup decisions as riskier to the extent that the concentration of offensive players is higher (*ex-ante* risk-taking). In the same way, a risky substitution means that a more offensive player enters the game (*within-game* risk-taking).

This approach is quite intuitive as it is conventional wisdom in sports that increasing the number of offensive players (which necessarily means reducing the number of defensive players) allows for more scoring opportunities for both one's one team and the opposition.

Indeed, our data shows that increasing the number of offensive players increases the total number of goals per game (estimated coefficient = 0.151, SE = 0.019). Nonetheless, increasing the number of offensive players does not significantly affect the probability of winning a game, whereas the probability of a draw significantly decreases.[7] We thus consider the condition for risk-taking as stated in Hvide (2002) to be met.

However, it should be kept in mind that the two moments of decision-making that are subsequently analyzed – the choice of players for the starting lineup and the substitution of players within a match – cannot be interpreted fully independently of each other. Importantly, whereas the choice of players for the starting lineup by the coach may affect the entire match, the substitution of players within a match can only adjust his or her initial risk-taking decision.

[7]These results were derived using a OLS/Poisson regression or Probit regression framework with the total number of goals per match or "victory" and "draw" dummies as dependent variables, respectively. In these regressions, the number of offensive players and a set of match-specific controls were added as independent variables.

4.1 Starting lineup decision

Formally, we estimate the model

$$fwd_{i,j} = \beta_0 + \beta_1 FEMALE_i + \beta_2 AGE_i + \beta_3 AGE_i^2 + \beta_4 AGE_i \cdot FEMALE_i + \beta_5 AGE_i^2 \cdot FEMALE_i + \gamma' X + \varepsilon_{i,j} \quad (1)$$

with season fixed effects to assess whether the decision on the number of offensive players in the starting lineup of match *j* by coach *i* ($fwd_{i,j}$) is affected by *i*'s gender.[8]

Furthermore, X is a vector of game and tournament-specific control variables such as playing at home or away, the tournament phase, the fixture being a first- or second-leg tie (in the qualifying and knockout phase), the attendance, a measure of (*ex-ante*) heterogeneity in strength, and the favorite status. More specifically, we use betting odds to calculate team-specific winning probabilities for each match. Subsequently, *HET* is the (absolute) difference between the winning probabilities of the opposing teams.[9]

In an extended framework, we also include an indicator ($GGGR$) measuring the annual degree of female empowerment for a given country. This indicator is obtained from the *Global Gender Gap Report* and assigned according to the home country of each coach *i*.[10]

[8]It is quite standard to also include team fixed effects. However, in our specific setting the association between teams and their coaches is rather strong, resulting in insufficient variation in this dimension. For instance, for about 58 percent of the clubs in our sample we observe only one coach during the entire observation period.

[9]See Deutscher et al. (2013) for a detailed description of how to derive winning probabilities from betting odds. Betting odds were extracted from the website *www.betexplorer.com*. Note that betting markets are considered to work efficiently (meaning that they incorporate all publicly-available information, see, e.g., Deutscher et al., 2018) and they are often used as a measure of (ex-ante) heterogeneity in sport contests (e.g., Deutscher et al., 2013; Bartling et al., 2015). As a second measure, we use the UEFA Club Coefficient Ranking score of the teams. This approach has a clear downside, because the score refers to achievements in past seasons and may have limited explanatory power for the actual strength. The results based on this second measure are qualitatively very similar.

[10]The *Global Gender Gap Report* is published by the World Economic Forum. With a minimum value of 0 and a maximum value of 1, this indicator combines imbalances between women and men in the fields of health, education, economy and politics are combined into a single index. Following Falk and Hermle (2018), we group index values into four categories using the 25th, 50th, and 75th percentile.

Finally, we include a dummy variable that equals 1 if the match is a mixed-sex competition (MSC), namely a team with a female coach plays against a team with a male coach.

Table 2 provides some descriptive statistics on the sample that is used in the following regressions.

Table 2: Summary statistics: Starting lineup

Variable	Mean	Std. Dev.	Min.	Max.	N
fwd (# offensive players)	2.49	1.06	0	6	1586
age (coach)	44.449	9.769	23.151	71.765	1445
female	0.163		0	1	1586
attendance	1585.031	3410.622	5	50212	1586
HET (betting odds)	0.158	0.258	0.001	1	1427
finalround	0.546		0	1	1586
GGGR index	0.737	0.049	0.588	0.881	1586
MSC	0.315		0	1	1586

The regression outputs in Tables 3 and 4 display our main results. The coefficients displayed are estimated using OLS (with SE clustered at the match level) and multinomial logistic regressions.[11]

We first observe that female coaches tend to choose more offensive players as starters, therefore showing a higher risk preference (see columns (1) and (2) in Table 3). This finding also holds in a different specification where we use the number of offensive players in a matchday squad as the dependent variable (see Table A.1 in the Appendix). As a further robustness check, we restrict our sample to matches in which *HET* does not exceed the 75th percentile (*HET*=0.81). This procedure ensures that our results are not driven by very unbalanced and hence less competitive match-ups. Table A.2 in the Appendix shows that this is indeed not the case, as the results are very similar.[12]

However, Table 4 reveals that this effect is mainly driven by a lower probability of "ultra-defensive" play by female coaches, involving none or only one offensive player in their starting lineup (category 1). A split sample estimation combined with a Chow test

[11]Note that the depending variable fwd is divided into four categories for the logistic regression method: 0 to 1, 2, 3, and more than 3 offensive players.

[12]Note also that using Poisson regressions instead of OLS and more conservative standard errors (clustered at the coach or team level) do not change the results qualitatively.

(Chow, 1960) shows that the overall effect is the same for female-coached underdog teams as for female-coached favorite teams (p-value = 0.4881).

Second, we find a U-shaped age effect on our measure of risk-taking, which is more pronounced for female coaches (columns (2) and (3) in Table 3).[13] Figure 2 illustrates estimated marginal mean effects. This finding is confirmed in additional regressions in which we restrict our sample to younger coaches.[14] In these regressions, young female coaches do not exhibit a higher risk preference than young male coaches.

This result of a higher level of risk-taking for female coaches – which additionally appears to be primarily driven by older female coaches – is in line with the specific selection process into the profession as a female soccer coach, which has also changed over time. More specifically, the coaches in our sample typically played soccer by themselves, mostly at a (semi-)professional level. However, given the evolution of women's soccer, for female coaches above a certain age the decision to play soccer by themselves was actually a risky decision. At that time, playing women's soccer was not well regarded by society and it was somewhat stigmatized. As stigmatization gradually disappeared, less risk-taking was necessary at that stage of the selection process into the profession, and young female coaches are thus recruited from a pool of former players that is presumably on average less willing to take risk than previous generations of female players (and thus potential coaches).

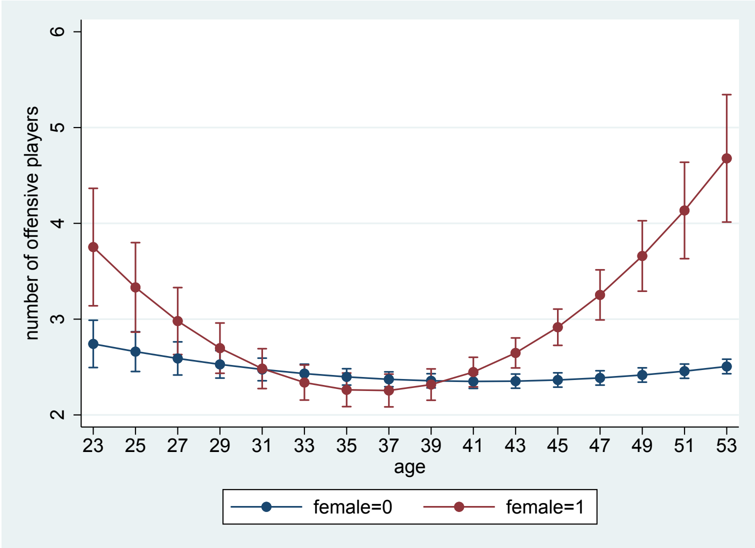
Third, column (4) of Table 3 indicates that a higher $GGGR$ indicator category of a coach's home country corresponds to a lower level of offensive players in the starting lineup. Nonetheless, there is no statistically significant relationship between gender and the $GGGR$ indicator category (column (5)). We therefore cannot confirm the result of Falk and Hermle (2018) – derived from global survey data – that the gender gap in risk-taking correlates with female empowerment in our specific setting.[15] As a possible explanation, we refer to the highly competitive international environment in the

[13]This can be more directly seen when using a split sample estimation (see Table A.3 in the Appendix).

[14]We restrict this analysis to coaches who are younger than the mean age in our sample, i.e. younger than 44.5 years. The results are displayed in Table A.4 in the Appendix.

[15]However, the estimated coefficient of the interaction term *female*index* is different from zero when we use the number of offensive players in a matchday squad as the dependent variable (Table A.1 in the Appendix).

Figure 2: Gender differences in risk taking according to age.



professional environment under consideration.

Fourth, column (6) as well as Table A.3 suggest that risk-taking behavior is not affected by whether coaches act in a same-sex or mixed-sex competition. This finding contrasts with experimental evidence from the lab suggesting that mixed-sex competitive environments affect the behavior (i.e. performance and entry decisions) of female subjects compared to other treatments (e.g., Niederle and Vesterlund, 2011).

Finally, estimations from probit regressions (with a ”victory” dummy as dependent variable) show that the number of offensive players does not influence the winning probability. Moreover, there are no gender differences related to winning probabilities (t-test, p-value = 0.2811) or the favorite status (t-test, p-value = 0.9125).[16] This result is in line with Gomez-Gonzalez et al. (2018), who report that gender does not explain team performance in their sample from top divisions in France, Germany, and Norway.[17]

4.2 Decisions on substitutions

As a second moment of decision-making involving risk, we use a model very similar to equation (1), but now including a dummy "offensive substitution: yes/no" as the depen-

[16]Moreover, female coaches are rather underrepresented in clubs from countries thab have dominated the tournament over the last decade (Germany, France, and Sweden).

[17]Note that Darvin et al. (2018) also find that the gender of a coach does not affect individual player performances in their sample from professional women's basketball.

dent variable. We define a substitution as offensive if a forward replaces a midfielder or a defensive player, or if a midfielder replaces a defensive player.

Note that we find that male coaches make substitutions more frequently than female coaches (two-sample Student's *t*-test, *p*-value = 0.0007). This difference can be explained by the fact that male coaches make substitutions more frequently during half-time (see Figure B.1 in the Appendix). Nonetheless, neither the number nor the timing of substitutions appear to be relevant for analyzing potential gender differences in risk-taking if the number of offensive players remains constant.

Since we now focus on the in-game level perspective, we include additional match-specific variables as controls. For example, these include the current score at the time of substitution and previous suspensions.[18] Table 5 displays summery statistics of our main variables used in this part of our analysis. Coefficients are estimated using probit regressions.

Table 6 displays the results of this exercise. First, columns (1) to (3) show that no gender differences in risk-taking can be identified when we control for the initial number of offensive players in the starting lineup (*n_fwd*) and the coach's age. Second, neither the coach's age nor the $GGGR$ indicator have an impact on a coach's propensity to make an offensive substitution. Third, we interpret the negative coefficient on HET as a form of *ex ante* "discouragement effect" (e.g., Baik, 1994; Szymanski, 2003; Konrad, 2009). This means that a stark heterogeneity reduces the coach's incentives for interventions changing the number of offensive players during the match.[19]

5 Conclusions

Differences in the willingness to take risks between men and women are often invoked as an explanation for persistent gender gaps in wages, earnings and social positions. Our paper shows that this "higher willingness to take risks" hypothesis does not serve as an explanation for the much higher share of male managers in the female-dominated

[18]For instance, Grund et al. (2013) show that the current score affects risk-taking behavior.

[19]Note that we do not find any gender differences in terms of this "discouragement effect".

industry under consideration. On the contrary, we find evidence that female coaches reveal a higher level of risk-taking than male coaches on average. In a general sense, we take this as a clear hint for the importance of considering the industrial environment when analyzing gender differences.

Our results appear to be primarily driven by a lower probability of female coaches starting a game with a very defensive lineup. Put differently, we find that male coaches are more likely have none or only one offensive player among their starters. In our view, this finding again emphasizes the importance of the underlying selection processes into the profession, especially when considering the fact that we still hardly observe any female coaches in men's soccer at a (semi-)professional level. Our results also indicate that these selection processes might change over time, as we do not find evidence of gender differences in risk-taking in a restricted sample of young coaches. Nonetheless, our findings might also show how masculine identity (Akerlof and Kranton, 2000) and stereotypes (or rather their confirmation) strongly depend on work environments (Nelson, 2015).

Table 3: Regression output: Starting lineup.

	(1) Model 1	(2) Model 2	(3) Model 3	(4) Model 4	(5) Model 5	(6) Model 6
female	0.175**	0.257***	0.486***	0.499***	0.465**	0.497**
	(0.0716)	(0.0754)	(0.0821)	(0.0806)	(0.192)	(0.196)
age		0.0103***	0.00809**	0.00681**	0.00677**	0.00675**
		(0.00298)	(0.00314)	(0.00320)	(0.00321)	(0.00322)
age^2		0.00113***	0.00114***	0.00120***	0.00120***	0.00121***
		(0.000238)	(0.000239)	(0.000239)	(0.000239)	(0.000238)
female*age			0.134***	0.133***	0.133***	0.133***
			(0.0184)	(0.0180)	(0.0182)	(0.0183)
female*age^2			0.00749***	0.00727***	0.00729***	0.00733***
			(0.00134)	(0.00133)	(0.00137)	(0.00137)
index (GGGR)				-0.0687**	-0.0708**	-0.0709**
				(0.0277)	(0.0317)	(0.0317)
female*index					0.0131	0.0133
					(0.0684)	(0.0686)
MSC						-0.0534
						(0.0711)
Constant	2.104***	2.187***	2.182***	2.401***	2.407***	2.452***
	(0.220)	(0.263)	(0.260)	(0.273)	(0.276)	(0.280)
Season FE	✓	✓	✓	✓	✓	✓
Further control	✓	✓	✓	✓	✓	✓
Observations	1586	1445	1445	1445	1445	1445
R^2	0.011	0.037	0.059	0.063	0.063	0.064

- Dependent variable: number of offensive players in the starting lineup.
- Coefficients are estimated with standard OLS.
- *age* was centered for ease of interpretation.
- Further controls: home/away, tournament phase, attendance, heterogeneity, favorite status
- Robust standard errors (clustered at the match level) in parentheses, * p<0.1, ** p<0.05, *** p<0.01

Table 4: Estimation results from multinomial logit model: Starting lineup.

	No. of offensive players (categories) Base outcome: two players		
	1	3	4
female	-0.530**	0.264	0.359
	(0.249)	(0.178)	(0.227)
age	-0.0301	-0.145***	-0.296***
	(0.0721)	(0.0531)	(0.0617)
age^2	0.000154	0.00166***	0.00350***
	(0.000814)	(0.000586)	(0.000674)
Season FE	✓	✓	✓
Further controls included?	✓	✓	✓
Observations	1445		
Pseudo R^2	0.031		

Standard errors in parentheses

* $p < 0.10$, ** $p < 0.05$, *** $p < 0.010$

Table 5: Summary statistics: Substitutions

Variable	Mean	Std. Dev.	Min.	Max.	N
offsub (offensive substitution)	0.257	0.437	0	1	3538
n_fwd (off. players in starting11)	2.480	1.038	0	6	3538
age (coach)	44.636	9.695	23.151	71.765	3223
female	0.158		0	1	3538
score_diff (goal difference at substitution)	0.115	3.028	-13	16	3538
attendance	1617.106	3558.806	10	50212	3538
home	0.508	0.5	0	1	3538
HET (betting odds)	0.585	0.258	0.001	1	3119
finalround	0.535		0	1	3538
GGGR index	0.735	0.05	0.588	0.881	3538
MSC	0.316		0	1	3538

Table 6: Probit regression output: Substitutions.

	(1)	(2)	(3)	(4)	(5)
female	-0.202**	-0.171**	-0.115	0.0481	0.125
	(0.0814)	(0.0806)	(0.0859)	(0.215)	(0.207)
HET	-0.201**	-0.131	-0.174*	-0.182**	-0.0960
	(0.102)	(0.0959)	(0.0941)	(0.0920)	(0.101)
n_fwd		-0.242***	-0.254***	-0.250***	-0.255***
		(0.0270)	(0.0280)	(0.0279)	(0.0292)
age			0.00454	0.00451	0.00538*
			(0.00288)	(0.00277)	(0.00299)
index (GGGR)				0.0216	0.0102
				(0.0238)	(0.0261)
female*index				-0.0490	-0.104
				(0.0746)	(0.0726)
Constant	-0.535	-0.0761	-0.230	-0.142	-0.281
	(0.419)	(0.398)	(0.411)	(0.161)	(0.423)
Season FE	✓	✓	✓	✓	✓
Further controls	✓	✓	✓	✓	✓
Observations	3119	3119	2913	2913	2435
Pseudo R^2	0.011	0.037	0.040	0.033	0.041

- Dependent variable: offensive substitution (FWD for MF or DF, MF for DF).
- Coefficients are estimated with probit regression.
- Column (5): Sample is restricted to matches where at least one player on the bench is coded as FWD.
- Further controls: MSC, home/away, tournament phase, attendance, favorite status, score dummies.
- Robust standard errors (clustered at the match level) in parentheses, * p<0.1, ** p<0.05, *** p<0.01

References

Akerlof, G. A. and Kranton, R. E. (2000). Economics and identity. *The Quarterly Journal of Economics*, 115(3):715–753.

Altonji, J. G. and Blank, R. M. (1999). Race and gender in the labor market. In Ashenfelter, O. C. and Card, D., editors, *Handbook of Labor Economics*, volume 3, chapter 40, pages 3143 – 3259. Elsevier / North-Holland, Amsterdam, Netherlands.

Baik, K. H. (1994). Effort levels in contests with two asymmetric players. *Southern Economic Journal*, 61(2):367–378.

Barber, B. M. and Odean, T. (2001). Boys will be boys: Gender, overconfidence, and common stock investment. *The Quarterly Journal of Economics*, 116(1):261–292.

Bartling, B., Brandes, L., and Schunk, D. (2015). Expectations as reference points: Field evidence from professional soccer. *Management Science*, 61(11):2646–2661.

Beckmann, D. and Menkhoff, L. (2008). Will women be women? analyzing the gender difference among financial experts. *Kyklos*, 61(3):364–384.

Bertrand, M. (2011). New perspectives on gender. In David Card, O. A., editor, *Handbook of Labor Economics*, volume 4B, chapter 17, pages 1543–1590. Elsevier, Amsterdam, Netherlands.

Blau, F. D. and Kahn, L. M. (2017). The gender wage gap: Extent, trends, and explanations. *Journal of Economic Literature*, 55(3):789–865.

Bradbury, S., Van Sterkenburg, J., and Mignon, P. (2014). The glass ceiling in European football: Levels of representation of visible ethnic minorities and women in leadership positions, and the experiences of elite level ethnic minority coaches. *Fare network report*. URL: `http://www.farenet.org/wp-content/uploads/2014/12/The-glass-ceiling-in-football-screen3.pdf`.

Chow, G. C. (1960). Tests of equality between sets of coefficients in two linear regressions. *Econometrica*, 28(3):591–605.

Croson, R. and Gneezy, U. (2009). Gender differences in preferences. *Journal of Economic Literature*, 47(2):448–474.

Darvin, L., Pegoraro, A., and Berri, D. (2018). Are men better leaders? An investigation of head coaches' gender and individual players' performance in amateur and professional women's basketball. *Sex Roles*, 78(7-8):455–466.

De Paola, M., Gioia, F., and Scoppa, V. (2018). Teamwork, leadership and gender. *IZA Discussion Paper No. 11861*.

Delfgaauw, J., Dur, R., Sol, J., and Verbeke, W. (2013). Tournament incentives in the field: Gender differences in the workplace. *Journal of Labor Economics*, 31(2):305–326.

Deutscher, C., Frick, B., Gürtler, O., and Prinz, J. (2013). Sabotage in tournaments with heterogeneous contestants: Empirical evidence from the soccer pitch. *The Scandinavian Journal of Economics*, 115(4):1138–1157.

Deutscher, C., Frick, B., and Ötting, M. (2018). Betting market inefficiencies are short-lived in german professional football. *Applied Economics*, 50(30):3240–3246.

Dohmen, T. and Falk, A. (2011). Performance pay and multidimensional sorting: Productivity, preferences, and gender. *American Economic Review*, 101(2):556.

Eckel, C. C. and Füllbrunn, S. C. (2015). Thar she blows? Gender, competition, and bubbles in experimental asset markets. *American Economic Review*, 105(2):906–20.

Eckel, C. C. and Grossman, P. J. (2008). Men, women and risk aversion: Experimental evidence. In Plott, C. R. and Smith, V. L., editors, *Handbook of Experimental Economics Results*, volume 1, chapter 113, pages 1061–1073. Elsevier, Amsterdam, Netherlands.

Falk, A., Becker, A., Dohmen, T., Enke, B., Huffman, D., and Sunde, U. (2018). Global evidence on economic preferences. *The Quarterly Journal of Economics*, 133(4):1645–1692.

Falk, A. and Hermle, J. (2018). Relationship of gender differences in preferences to economic development and gender equality. *Science*, 362(6412).

Filippin, A. and Crosetto, P. (2016). A reconsideration of gender differences in risk attitudes. *Management Science*, 62(11):3138–3160.

Flory, J. A., Gneezy, U., Leonard, K. L., and List, J. A. (2018). Gender, age, and competition: A disappearing gap? *Journal of Economic Behavior & Organization*, 150:256–276.

Gomez-Gonzalez, C., Dietl, H., and Nesseler, C. (2018). Does performance justify the underrepresentation of women coaches? Evidence from professional women's soccer. *Sport Management Review*. Online first, `https://doi.org/10.1016/j.smr.2018.09.00`.

Grainey, T. F. (2012). *Beyond Bend It Like Beckham: The Global Phenomenon of Women's Soccer*. University of Nebraska Press (USA).

Grund, C. and Gürtler, O. (2005). An empirical study on risk-taking in tournaments. *Applied Economics Letters*, 12(8):457–461.

Grund, C., Höcker, J., and Zimmermann, S. (2013). Incidence and consequences of risk-taking behavior in tournaments—evidence from the NBA. *Economic Inquiry*, 51(2):1489–1501.

Holt, C. A. and Laury, S. K. (2002). Risk aversion and incentive effects. *American Economic Review*, 92(5):1644–1655.

Hoss, M. A. K., Bobrowski, P., McDonagh, K. J., and Paris, N. M. (2011). How gender disparities drive imbalances in health care leadership. *Journal of Healthcare Leadership*, 3:59–68.

Hvide, H. K. (2002). Tournament rewards and risk taking. *Journal of Labor Economics*, 20(4):877–898.

Konrad, K. A. (2009). *Strategy and dynamics in contests*. Oxford University Press (UK).

Nelson, J. A. (2015). Are women really more risk-averse than men? A re-analysis of the literature using expanded methods. *Journal of Economic Surveys*, 29(3):566–585.

Niederle, M. and Vesterlund, L. (2007). Do women shy away from competition? Do men compete too much? *The Quarterly Journal of Economics*, 122(3):1067–1101.

Niederle, M. and Vesterlund, L. (2011). Gender and competition. *Annual Review of Economics*, 3(1):601–630.

Pieper, J., Nüesch, S., and Franck, E. (2014). How performance expectations affect managerial replacement decisions. *Schmalenbach Business Review*, 66(1):5–23.

Szymanski, S. (2003). The economic design of sporting contests. *Journal of Economic Literature*, 41(4):1137–1187.

van Ours, J. C. and van Tuijl, M. A. (2016). In-season head-coach dismissals and the performance of professional football teams. *Economic Inquiry*, 54(1):591–604.

van Staveren, I. (2014). The Lehman Sisters hypothesis. *Cambridge Journal of Economics*, 38(5):995–1014.

Wicker, P., Cunningham, G., and Fields, D. (2019). Head coach changes in women's college soccer: An investigation of women coaches through the lenses of gender stereotypes and the glass cliff. *Sex Roles*. online first, https://link.springer.com/article/10.1007/s11199-019-01022-2.

A Additional Tables

Table A.1: Regression output: Offensive players in matchday squad.

	(1) Model 1	(2) Model 2	(3) Model 3	(4) Model 4	(5) Model 5	(6) Model6
female	0.302***	0.340***	4.317*	4.215*	4.250*	4.256*
	(0.0924)	(0.0982)	(2.490)	(2.486)	(2.537)	(2.551)
age		-0.127***	-0.150***	-0.152***	-0.155***	-0.155***
		(0.0290)	(0.0314)	(0.0313)	(0.0313)	(0.0311)
age^2		0.00149***	0.00172***	0.00173***	0.00175***	0.00175***
		(0.000324)	(0.000344)	(0.000344)	(0.000344)	(0.000342)
female*age			-0.273**	-0.268**	-0.294**	-0.295**
			(0.131)	(0.131)	(0.136)	(0.137)
female*age^2			0.00428**	0.00422**	0.00454**	0.00454**
			(0.00170)	(0.00169)	(0.00177)	(0.00178)
index (GGGR)				-0.0184	-0.0487	-0.0487
				(0.0349)	(0.0402)	(0.0402)
female*index					0.196**	0.196**
					(0.0972)	(0.0972)
MSC						-0.00456
						(0.0925)
Constant	3.235***	5.921***	6.495***	6.599***	6.749***	6.755***
	(0.180)	(0.682)	(0.740)	(0.751)	(0.757)	(0.752)
Season FE	✓	✓	✓	✓	✓	✓
Further controls	✓	✓	✓	✓	✓	✓
Observations	1586	1445	1445	1445	1445	1445
R^2	0.053	0.070	0.079	0.079	0.082	0.082

- Dependent variable: Number of offensive players in matchday squad (starters + substitutes).
- Coefficients are estimated with standard OLS.
- Further controls: home/away, tournament phase, attendance, heterogeneity, favorite status
- Robust standard errors (clustered at the match level) in parentheses, * p<0.1, ** p<0.05, *** p<0.01

Table A.2: Regression output: Starting lineup (excluding very unbalanced match-ups)
.

	(1) Model 1	(2) Model 2	(3) Model 3	(4) Model 4	(5) Model 5	(6) Model 6
female	0.171**	0.298***	0.546***	0.546***	0.623***	0.705***
	(0.0869)	(0.0918)	(0.104)	(0.101)	(0.218)	(0.222)
age		0.0151***	0.0123***	0.0103***	0.0104***	0.0105***
		(0.00369)	(0.00384)	(0.00395)	(0.00398)	(0.00398)
age2		0.00114***	0.00121***	0.00128***	0.00128***	0.00131***
		(0.000307)	(0.000305)	(0.000303)	(0.000303)	(0.000302)
female*age			0.128***	0.127***	0.126***	0.128***
			(0.0221)	(0.0213)	(0.0214)	(0.0215)
female*age^2			0.00715***	0.00683***	0.00678***	0.00683***
			(0.00182)	(0.00179)	(0.00182)	(0.00181)
index (GGGR)				-0.0895***	-0.0847**	-0.0849**
				(0.0339)	(0.0389)	(0.0389)
female*index					-0.0298	-0.0306
					(0.0780)	(0.0784)
MSC						-0.132
						(0.0830)
Constant	1.898***	1.892***	1.920***	2.213***	2.197***	2.329***
	(0.492)	(0.491)	(0.488)	(0.492)	(0.496)	(0.502)
Season FE	✓	✓	✓	✓	✓	✓
Further control	✓	✓	✓	✓	✓	✓
Observations	1066	1001	1001	1001	1001	1001
R^2	0.015	0.049	0.069	0.076	0.076	0.079

- Dependent variable: Number of offensive players in the starting lineup.
- Coefficients are estimated with standard OLS.
- *age* was centered for ease of interpretation.
- Further controls: home/away, tournament phase, attendance, heterogeneity, favorite status
- Robust standard errors (clustered at the match level) in parentheses, * p<0.1, ** p<0.05, *** p<0.01

Table A.3: Regression output: Starting lineup per gender.

	Female Coaches	Male Coaches
age	-0.650***	-0.0987***
	(0.111)	(0.0215)
age2	0.00884***	0.00121***
	(0.00143)	(0.000233)
MSC	0.153	-0.0660
	(0.179)	(0.0824)
Constant	14.50***	4.090***
	(2.173)	(0.602)
Further Controls	✓	✓
Season FE	✓	✓
Observations	250	1195
R^2	0.218	0.041

- Dependent variable: Number of offensive players in the starting lineup.
- Coefficients are estimated with standard OLS.
- Robust standard errors (clustered at the match level) in parentheses, * p<0.1, ** p<0.05, *** p<0.01

Table A.4: Starting lineup (Sample restricted to young coaches).

	(1) Model 1	(2) Model 2	(3) Model 3	(4) Model 4
female	0.0866 (0.0843)	0.101 (0.0870)	0.101 (0.0867)	-0.0349 (0.220)
age		-0.0117* (0.00709)	-0.0121* (0.00709)	-0.0127* (0.00728)
index (GGGR)			-0.0420 (0.0362)	-0.0551 (0.0457)
female*index				0.0518 (0.0794)
Constant	2.224*** (0.281)	2.697*** (0.397)	2.853*** (0.422)	2.918*** (0.448)
Season FE	✓	✓	✓	✓
Further controls included?	✓	✓	✓	✓
Observations	761	761	761	761
R^2	0.031	0.035	0.036	0.037

- Dependent variable: Number of offensive players in the starting lineup.
- Coefficients are estimated with standard OLS.
- Further controls: home/away, tournament phase, attendance, heterogeneity, favorite status
- Robust standard errors (clustered at the match level) in parentheses, * p<0.1, ** p<0.05, *** p<0.01

B Additional Figures

Figure B.1: Density of substitutions over the course of time.