Searching for the Queen's Gambit: An Exploratory Analysis of Male-Female Ratings Gaps in U.S. Chess

Brian Kisida, University of Missouri, Columbia
Matthew Pepper, Basis Policy Research
Michael Podgursky, SCAER and University of Missouri, Columbia
Michelle Wickman, SCAER at St. Louis University

Abstract

We examine the origin and evolution of male-female rating gaps for young chess players using two decades of data from the U.S. Chess Federation, the national chess association that tracks competitive tournament play and provides ratings for U.S. chess players. An important feature of our research is that we examine male-female gaps across a broad range of chess ratings, from novice to expert. We find large gaps favoring males at entry across the entire distribution. Once players have an established rating, we find similar returns to experience for males and females. Although female players have higher attrition rates than males, the net effect of this differential attrition on ratings gaps is null (to slightly equalizing) because stronger female players are at least as likely as males to persist. We find some evidence that that the male-female rating gap at entry declines modestly as female participation in the home locale rises, with an effect that is generally stronger for weaker players. Overall, the key explanation for differences in U.S male female chess ratings is the gap at entry, which is large when first observed and persists over time.

Working Paper 25-01
4-7-2025

The authors acknowledge the support of the Sinquefield Center for Applied Economic Research (SCAER) for research support, and to the U.S. Chess Federation for providing the data used in this study. We thank Michael Nolan of USCF for patiently answering our numerous questions regarding these data and for Richard Pointer and Tony Rich of the Saint Louis Chess Club for answering our questions about chess ratings and tournament play. The usual disclaimers apply.

Introduction

Public curiosity about women and chess stretches back for decades. Much of it is motivated by the conspicuous lack of females in the highest ranks of the game. While male-female gaps in many sports can be attributed to differences in male-female physiology such as strength or lung capacity, these differences do not seem relevant to chess, which is entirely a matter of cognitive skills and spatial perception. One narrative suggests female underrepresentation may be due to barriers to access, such as social forces and structural barriers that deter females from taking up or advancing in the game (Brancaccio & Gobet, 2023). Counter-narratives suggest that innate characteristics affect male-female differences in performance. While the issue has festered for some time, the great popularity of the TV series Queen's Gambit (based on a novel of the same title), has increased attention on the matter. The story follows the rise of a young female chess player who overcomes significant personal and structural barriers and ultimately becomes the world champion.

There are several respected chess ratings systems in use with varying methodologies, but at their core these all rate competitive chess players with a continuous, quantitative score that is based on the outcomes of rated games and the strength of their opponents. In 2025, according to FIDE, the governing body of international chess, there are no female chess players in the top 100 active players; the highest female ranked at 118.[1] Based on a rigorous standard of tournament play and FIDE ranking, the very best players can attain the rank of "Grandmaster." Players who reach this top tier maintain the title for the rest of their lives, independent of tournament play. There are roughly 1,700 living grandmasters, of whom just 42 (2.5%) are female.[2]

This paper contributes to the empirical literature on male-female gaps by examining the emergence and evolution of male-female rating gaps for young chess players using two decades of data from the U.S. Chess Federation (USCF), the national chess association that tracks competitive tournament play and provides ratings for U.S. chess players.[3] The contribution of our study is three-fold. First, we use 19 years of administrative data from the chess federation of a country that accounts for a large share of the world's top players. Second, rather than cross section comparisons, we construct longitudinal data files for beginning US competitive players as soon as they reach what USCF considers reliable chess performance rating (an "established" rating, earned after a player's 25th rated game). Finally, as opposed to previous investigations that focus on chess gaps at the mean or the extreme right tail of the performance distribution, we examine male-female gaps along a broad range of the chess ratings distribution.

Numerous explanations have been posited to explain the presence of male-female chess performance gaps, including differences in ability at entry, differential attrition, and differences in returns to experience. Examining all these possibilities, we find that the male-female gaps are significant across the full distribution at entry and persist over time. Although attrition rates are higher for females than males, the net effect of this differential attrition on ratings gaps is minor because stronger female players are at least as likely as stronger males to persist. Leveraging the

---

[1] https://ratings.fide.com/
[2] https://en.wikipedia.org/wiki/List_of_female_chess_grandmasters
[3] Similar to FIDE ratings, USCF ratings rise and fall based on player performance in rated games.

longitudinal nature of our panel with an individual fixed-effects approach, we find no differences in returns to experience between male and female players. Simply put, the male-female gap is strong and present at every point of the distribution at the point of entry into competitive play and persists throughout youth and adolescence. Finally, we reexamine the Chablis-Glickman (2006) finding that a higher share of female players in the home locale tends to lower the male-female gap at entry. We replicate and extend this analysis on initial and subsequent gaps in performance at different points in the distribution. Like Chablis and Glickman, we find that higher female participation is associated with narrower male-female ratings gaps, particularly at lower levels of performance. However, the effects are small. We conclude that the primary reason for male-female gaps overall are gaps at entry. To narrow gaps in the population of rated active players, young females with stronger potential must be encouraged to enter and persist in tournament play.

Literature Review

A large literature explores possible explanations for male-female differences in chess performance. Some researchers have highlighted the large difference between male and female participation as a statistical factor that explains the dominance of males in the very top ranks of chess players. Male-female differences may be due in part to the fact that there are simply more extreme values in larger male samples than smaller female samples (e.g., Bilalic, et al. 2008; Howard, 2014). This, however, seems to be an anomaly exclusive to the small sample of elite players. When conducting Monte Carlo simulations for our large sample of non-elite players, we find no evidence that male-female differences are due to sampling variance.

Others point to the underlying distribution of chess ability. Because top performing players come from the right tail of chess ability distribution, fatter tails (or larger variance) for males would imply more top (and bottom) players, similar to related research on male-female gaps in math performance (Halpern, et.al. 2007). Chablis and Glickman (2006), however, find no evidence of a higher variance for male as compared to female players among non-elite players.

In addition to lower participation overall, females also tend to have higher attrition rates, and this has been posited as another potential explanation for male-female performance gaps. However, no published study has been found by the authors that examines how differential attrition affects male-female performance gaps.

Differences in style of play may also affect the chess performance of females versus males. Typically focusing on elite players, a growing literature uses game-level data to explore differences in chess playing styles. Some have found mixed evidence of “stereotype threats” in which females play worse against equally ranked male opponents (Smerdon, et.al. 2020; Stafford, 2018). Others have found that females display greater risk aversion than males in play, and that men employ more aggressive strategies when playing female opponents (Gerdes and Gransmark, 2010). Research has also documented that females underperform relative to males in response to pressures from time-control of games (Dilmaghani, 2020, 2021; Gransmark, 2012). A recent study examines male-female differences in response to “personal bests” in tournament play and rankings, finding that women increase their effort relative to men when approaching their personal best, but exert less effort once surpassing it (Gonzales-Diaz, 2021). A common theme of these studies is to exploit data on elite tournament chess play to test various psychological or behavioral economic theories. While some of the findings provide insights to potential explanations for slight differences in certain situations, they fall short of explaining the large magnitude of male-female gaps in chess performance.

Clearly more work is needed to explore the development of players and the mechanisms through which players progress to higher levels of chess play. An important contribution of this study is that we examine male-female gaps across the full range of experience and skill levels rather than just the mean or, in much of the literature, the extreme right tail of experienced players. Rather than study the behavior of elite experienced players, our study takes the opposite tack and focuses on novice players using two decades of U.S. Chess Federation administrative data on games played by young competitors, some of whom may eventually become elite. We analyze the size and structure of male-female ratings differences at entry (when ratings become reliable). We then analyze the effect of differential male-female attrition on gaps. Next, we examine potential differences in the returns to game experience for males and females. Finally, we extend Chablis and Glickman's (2006) analysis, where they find that the percent of female players in one's home locale is associated with narrower male-female gaps, suggesting a positive female peer effect. As we will see below, important insights emerge when we examine the effect of variables across the full range of performance percentiles.

Data description:

We use data from the United States Chess Federation (USCF) to examine possible male-female differences in USCF chess ratings among young players at officially sanctioned tournaments within the US. The USCF rating system uses an algorithm to determine a player's revised rating based on the outcome of a game and the relative strength of their opponent. In theory, this rating gives an approximation of players' ability relative to all other players in the USCF system. The anonymous player-level data includes player rating changes by game, sex, birth date, and game-level outcomes (win, loss, or draw).

We measure ratings for individuals at the end of each calendar year. Age is computed by subtracting a player's birth year from the year of play. Missing data points are imputed using the prior year's rating. Finally, experience is calculated as the number of calendar years during which a player played at least one tournament game.

Our analysis focuses on players who begin their USCF participation between 2000 and 2019, with observed starting ages ranging from 3 to 15 years. For inconsistent male-female assignments, we retain the modal value. In our USCF data on tournament competitions, females represent just 16% percent of all players in 2019 but represent 20% of players aged 5-15 in 2019. Our dependent variable is chess rating. This rating is not considered reliable by USCF until at least 25 games have been played. Thus, we restrict our sample to players who have at least 25 rated games.[4] This results in a final sample of 106,398 players who began playing rated games between the ages of 3 and 15 who played at least 25 games to receive an established rating.

The annual player count in our dataset increases over time as player histories expand and participation rises. We exclude players who began their chess participation prior to the year 2000, hence there is both an expanding age group (as there is no age cap) and an expanding group of eligible players. As depicted in Table 1, females typically exhibit lower average ratings, are half a year younger, and play approximately the same number of annual games as males.

(Table 1)

While females tend to have lower ratings on average, it is useful to understand the full distribution of differences. Hence, Figure 1 and Table 2 show the cumulative distribution of scores (CDF) controlling for years of tournament play. Panel 1 shows the CDF of players in their first year of playing rated games after having an established rating, while panels 2 and 3 respectively portray the ratings CDF of players in their 3rd and 5th years. As expected, the CDF's for both males and females shift to the right as years of experience increase and ratings rise. However, in all cases the horizontal gap between the male and female curves persists. Moreover, the gap is roughly the same across percentiles of the ratings distribution.

(Figure 1 and Table 2)

Estimated Gaps at Entry

In order to estimate more precisely the male-female gaps at entry across the range of the ratings distribution, we estimate quantile regression models at each performance rating percentile (Koenker and Hallock, 2001). The quantile regressions take the following form:

$$y_i = \beta_0 + \beta_1 X_i + \beta_2 Male_i + \varepsilon_i \quad (1)$$

Where X includes a set of individual characteristics including a player's age, games played over the past two years, tournament year, and information from a player's home zip code including median household income, percent of residents identifying as white, and the locale type. Male is an indicator variable taking the value 1 if the player is a male. β_2 in equation (1) is estimated at every percentile of the performance distribution between 5 and 95. In effect, it identifies the horizontal displacement of the CDF of males versus females at each percentile of the pooled performance distribution. Figure 2 reports the point estimates of β_2 at entry (year 1) and year 5. In year 1 regression-adjusted ratings gaps range from 125-150 points over most of the range of the distribution, with a positive slope in the lower and upper percentiles. By year 5 the gaps

[4] As noted, we have followed USCF convention and defined "entry" as the point at which an individual has played 25 tournament games, at which point USCF considers their ratings "established." However, the many players never attain the 25 game threshold, with high, but broadly similar rates of attrition for males and females. To check the robustness of findings, we examined the male-female ratings gaps for the under 25 game group. As in Figure 1, there is a consistent gap favoring males across ratings percentiles. Our analysis finds that even at the conclusion of the very first tournament, males are rated higher than females at all levels.

range from 125 to 175, and the curve has flattened out somewhat. Of course, the confidence bands are wider in year 5 because there are fewer players at any given percentile.

(Figure 2)

<u>Male-Female Differences in Attrition</u>

Analysis in the previous section shows that there is a large male-female gap in ratings at entry (i.e., when ratings reliably stabilize at 25 rated games). Whether these entry gaps widen or narrow as young chess players mature depends on two factors that can be measured in our data: differential attrition and differential returns to game experience.

We begin with attrition. While there has been considerable discussion of male-female differences in chess participation overall, there has been much less discussion of male-female differences in attrition, which plays an important role in overall participation gaps. Statistics on attrition are reported in Figure 3 and Table 3. In Figure 3, we report the overall population of male and female players by years of tournament play (left axis) and the percent female (right axis). For both males and females, there is large attrition over years of tournament play. However, the percent female reported on the right axis shows that this attrition rate is much larger for female players. Popular discussions of male-female differences in chess often point out this much larger attrition rate for females. In principle, the large difference in attrition could exacerbate or reduce the rating gap.

(Figure 3 and Table 3)

In order to explore the effect of differential attrition on the ratings gap, we exploit the panel nature of our data to estimate variants of a simple linear probability model of retention:

$$\Pr\,(Retention = 1)_{it} = \beta_0 + \beta_1 Age_{it-1} + \beta_1 Rating_{it-1} + \beta_2 Female_i + \varepsilon_{it} \quad (3)$$

where i denotes the i-th player and t denotes year of play. Estimates of this simple model are reported in Column (1) of Table 4. Conditioning on prior rating and age, the female retention rate is -.02 (2 percent) lower than that of males. Column (2) adds an interaction between the female indicator and lagged ratings. Here we see that there is no significant difference between males and females in the association between lagged ratings and attrition, meaning that differences in attrition rates, while substantial, have no effect on the male-female rating gap of stayers. Finally, Column (3) adds an interaction with players' age. This most robust model finds the female x ratings interaction coefficient turns positive and significant, meaning that an increase in ratings lowers the attrition more for females than males and thus narrows the ratings gap for stayers.

(Table 4)

Table 5 presents simulations based on the estimated retention model in Column (3). The retention rates for males and females are simulated at the sample means by the level of ratings for both Year 1 and Year 5. At the median rating, the predicted retention rates of males and females are very similar. However, as ratings rise, the retention rate of females becomes slightly higher than that of males. For example, at the 90th percentile for the Year 1 population, the predicted retention rate for females is .73 and for males is .70. A similar result holds for the population at Year 5. We conclude from this exercise that while the retention rate of females is much lower during the first ten years of tournament play, this does not in any significant way influence the observed male-female ratings gap in the population.

(Table 5)

Male-Female Differences in Returns to Experience

If differential attrition does not affect the observed male-female performance gap, what about the returns to experience? Do males benefit more than females for each game played? A substantial literature has developed around differences in male and female tournament play. Some mixed evidence exists for "stereotype threats" which in our chess context means that females perform worse than expected when competing against males (Smerdon, et.al, 2020; Stafford, 2018). This has led some to advocate for female-only tournaments, while others claim that these hurt female talent development.[5] Other studies have found evidence of differences in risk aversion of male and female players in tournament play (e.g., Dilmaghani, 2020, 2021, 2022).

For our purposes, the important issue is whether these game-level differences in play aggregate to observable differences in returns to experience. In order to assess this, we estimate the following panel data model:

$$y_{it} = \beta_0 + \beta_1 X_{it} + \beta_3\, Games_{i,t+(t-1)} \quad +$$

$$\beta_4\, Games_{i,t+(t-1)} \times Female_i + \mu_i + \varepsilon_{it} \; (2)$$

where X_{it} is a set of time-varying controls, $Games_{i,t+(t-1)}$ measures the sum of games played in the current and previous year, and μ_i is an individual fixed effect. The results of this model are reported in Table 6. The first column reports the results over the full sample (i.e., all ratings levels). Unfortunately, it was not feasible to estimate quantile regression models with thousands of individual fixed effects. In order to investigate whether the returns to experience differ by initial ratings percentiles, we present overall estimates and by ratings bands, where (P20-50) includes players in percentile 20-40, etc. Reading across the first row, we see that experience matters: each additional rated game is associated with an increase of roughly 2 rating points. It is the second row that is most relevant for our investigation. This coefficient measures whether the return to rated games played differs for females as compared to males. There is some indication that in lower percentiles (20-50) the return is lower. However, the point estimate of the gap in returns per game (.28 rating points) is small. Among stronger initial players, there is no significant male-female difference in returns to experience.

(Table 6)

[5] For example, https://www.chess.com/forum/view/general/mens-and-womens-chess-should-not-be-seperated

Further Analysis of Gaps at Entry

The above analysis shows that the key factor affecting ratings gaps in the population of young USCF chess players are ratings gaps at entry. What can USCF data tell us about factors affecting entry gaps? Analyzing earlier cohorts of young USCF chess players, Chablis and Glickman (2006) find that the male-female ratio of participants in the player's home zip code affects the male-female gap. Specifically, as the ratio approaches 50 percent (parity), the mean male-female gap declines. We explored the relationship between spatial male-to-female ratios and performance gaps in our panel.

We begin by restricting the sample to newly established (year 1) players. We extended model (1) to include zip code level social and economic variables, total chess participation, and the female share of chess participants ($ShareF_j$) in the home locale:

$$y_{ij} = \beta_0 + \beta_1 X_i + \beta_2 Female_i + \beta_3 ShareF_j + \beta_4 \, ShareF_j \times Female + \varepsilon_i \quad (2)$$

In the model specified in (2), the effect of changes in the share of females in the home locale is β_3 for males and $(\beta_4 + \beta_3)$ for females. Thus, the effect of changes in female share on the (typically negative) female-male gap is β_4**.** Our estimated values for these coefficients by percentile of the current rating for newly established male and female players are reported in Figure 4-6 for locales with at least 20 or more observations.[6] In Table 7 we report descriptive statistics for the various locale samples and estimated coefficients at the median and 80th percentile. Figure 4 reports the effect of zip code percent female on the female-male ratings gap. A positive coefficient means that increases in the female share narrows the gap. While all the point estimates are positive, in most cases the 95 percent confidence band includes zero.

(Figure 4 and Table 7)

Figure 5 reports the same β_4 estimates with female share measured at the county level. In this case, the coefficient is positive and significant up to roughly the 60th percentile. In Table 7 we see that a percentage point increase in the female county share narrows the female-male ratings gap by 1.17 ratings points at the median and .62 (insignificant) ratings points at the 80th percentile. The slope of the quantile estimate function is negative, indicating that the narrowing effect of rising female share is weaker for stronger players. In comparing Figures 4 and 5, it should be noted that a one percentage point increase in the percent female is a larger effect in absolute and relative magnitude as we move from zip to county. One percentage point is roughly one-tenth of a standard deviation measured at the zip code level, but one-sixth of a standard deviation measured at the county level. Finally, we report β_4 estimates at the MSA level. These are generally larger in absolute value than the estimates at the county level and are significantly above zero over the full range of ratings. As with county, a clearly negative pattern is visible – with larger effects for weaker players. As before, a one percentage point increase in the share of

[6] We have experimented with sample selection cutoffs of locales with 10 or more and 30 or more players as well. The results are very similar to what we report here.

females measured at the MSA level represents more females than a one percent increase at the zip code or county level. In addition, one percentage point is roughly one-fifth of a standard deviation measured at the MSA level.

(Figures 5 and 6)

In short, we do find evidence for the Chablis-Glickman hypothesis that greater female participation narrows female-male ratings gaps. However, the calculations in Table 7 suggest that the effects are fairly modest. For example, a ten percentage point increase in the female county share (i.e., 1.7 standard deviations) would reduce the ratings gap from 138 to 126 points. Among stronger players the effect is considerably smaller. At the 80th percentile, for example, the gap would narrow from 141 to 135 rating points.

Conclusion.

In this paper we have examined male-female ratings gaps in two decades of USCF data on young tournament players. We find large gaps favoring males at all rating percentiles. The male-female ratings gap observed among the entire population of young chess players could arise from three different factors: a) ability gaps at initial entry; b) differential attrition; or c) differences in the return to experience (tournament play). We find that the observed ratings gaps for young USCF players are derived almost entirely from ratings gaps at initial entry. Once young players are established (i.e., have completed 25 rated games) we find no differences in returns to additional play for males versus females. Female players have far higher attrition rates than males, but the net effect of this attrition difference on ratings gaps is null (to slightly equalizing) because as compared to males, stronger female players are at least as likely as males to persist.

In order to better understand factors affecting gaps at entry, we explored the effect of the female share of active tournament players in their home locale. This builds on earlier work by Chamblis and Glickman (2006) who, studying earlier cohorts of USCF players, find evidence that average ratings gaps in home zip codes with a higher share of females tend to be smaller. We expand on their work by examining the effect of local female participation rates across the full distribution of ratings, not just at the mean, and at different levels of aggregation. We find evidence that higher rates of local female participation are associated with narrower male-female ratings gaps, however, these effects are strongest for the lower ability players.

In short, our analysis of USCF data for young chess players finds large gaps in female-male ratings across the full range of performance, and these gaps seem to be driven primarily by gaps in ability at entry. Our findings here shed some light on competing narratives that attempt to explain the male-female performance gap in chess ratings (Brancaccio & Gobet, 2023). We have shown that, while female attrition is much higher than male attrition, the net effect of this differential attrition on performance gaps in the population is null. Lower initial participation rates, and higher attrition rates, could suggest structural or discriminatory barriers remain that deter females from chess play. At the same time, disparate outcomes is not necessarily evidence of discrimination, as it may partially reflect different male/female preferences. Indeed, our analysis shows that males and females have similar gains from experience, suggesting that females have a similar, innate ability as males to improve their game through practice. Instead, the apparent source of the male-female gap is found at the origin. Overall, our findings suggest that significantly narrowing overall performance gaps will require policies that address the mechanisms of recruitment of females into chess.

Further insight into how this might be accomplished may be gleaned from more detailed analysis of game level data (and subsequent retention effects). Finally, one area of public interest is narrowing gaps at the very highest levels of tournament play (e.g., female-male gaps in grandmasters or the very top ranks of world rankings). Our research has not directly addressed that question. However, we believe that analysis of large longitudinal files such the USCF data in this study can shed light on how chess “stars” are grown and cultivated.

References

Chablis, C. and Glickman, M. (2006). Sex Differences in Intellectual Performance: Analysis of a Large Cohort of Competitive Chess Players. *Psychological Science*. 17(12), 1040-1046.

Bilalic, M., Smallbone, K., McLeod, P, & Gobet, F. (2008) Why are (the best) women so good at chess? Participation rates and gender differences in intellectual domains. The Royal Society Proceedings B Dec. 23, 276 (1659): 1161-1165.

Brancaccio, A. Gobet, F. (2023) Scientific Explanations of the Performance Gender Gap in Chess and Science, Technology, Engineering and Mathematics (STEM). *Journal of Expertise*. 6(21): 81-106.

Dilmaghani, M. (2020) Gender differences in performance under time constraint: Evidence from chess tournaments. Journal of Behavioral and Experimental Economics, 89

Dilmaghani, M. (2021) A Matter of Time: Gender, Time Constraint, and risk taking among chess elite. Economics Letters. 208, 110085.

Dilmaghani, M. (2022). Chess Girls Don't Cry: Gender Competition of Games and Effort in Competitions Among the Super Elite" Journal of Economic Psychology. 89,

Dreber, A, Gerdes, C. & Gransmark, P. (2013). Beauty Queens and Battling Knights: Risk-Taking and Attractiveness in Chess. Journal of Economic Behavior and Organization. 90, 1-18.

Gerdes, C. & Gransmark, P. (2010) Strategic Behavior Across Gender: A Comparison of Female and Male Expert Chess Players. Labor Economics. 17, 766-775.

Gonzales-Diaz, J, Palacios-Huerta, I, Abuin Mosquera, J.M. (2021). Personal Bests and Gender. Review of Economics and Statistics. 106(2), 409-422.

Halpern, D., Benbow, C., Geary, D, Gur, R., Shibley Hyde, S Gurnsbacher, M. (2007). The Science of Sex Differences in Science and Math. Psychological Science in the Public Interest 8(1), 1-51.

Howard, R. (2014) Explaining male predominance at the apex of intellectual achievement." Personality and Individual Differences 68(October): 217-220.

Koenker, R. & Hallock, K. (2001) Quantile Regression. Journal of Economic Perspectives 15, 4, 143-156.

Smerdon, D., Hu, H., McLennan, A, von Hippel, W., Albrecht, S. (2020) Female Chess Players Show Typical Stereotype-Threat Effects: Commentary on Stafford (2018). Psychological Science, 31, 756-759.

Stafford T. (2018). Female chess players outperform expectations when playing men. Psychological Science, 29, 429–436.

Table 1: Young established player counts from 2000-2019

		Female	Male	%Female
Initial annual rating		498	676	-
Initial age		9.7	10.2	-
Initial games		16.69	16.70	-
Annual players				
	2000	125	722	15
	2001	582	3,508	14
	2002	1,115	7,073	14
	2003	1,607	10,020	14
	2004	1,867	11,724	14
	2005	2,096	12,616	14
	2006	2,355	13,844	15
	2007	2,535	14,767	15
	2008	2,730	15,644	15
	2009	2,730	15,921	15
	2010	2,761	16,436	14
	2011	2,778	17,038	14
	2012	2,993	18,009	14
	2013	3,214	18,949	15
	2014	3,490	20,325	15
	2015	3,833	21,411	15
	2016	4,239	22,343	16
	2017	4,619	23,052	17
	2018	4,709	22,739	17
	2019	4,748	22,075	18
	Total	55,126	308,216	15

Table 2: Male-Female Gaps in Performance Ratings by Percentile

Percentile	Year 1	Year 3	Year 5
10	155	173	151
20	171	182	164
50	186	189	208
80	195	208	191
90	198	211	167
95	205	197	150

Table 3: Player Retention and Ratings

Spell length	Players	%Female	Female rating (mean)	Male rating (mean)	Δ
1	106,368	17	498	675	177
2	78,699	16	635	822	187
3	54,549	15	759	947	188
4	37,419	14	876	1,059	183
5	25,603	14	985	1,161	177
6	17,844	13	1,079	1,258	180
7	12,480	12	1,171	1,350	179
8	8,930	11	1,244	1,419	175
9	6,322	10	1,308	1,487	179
10	4,564	10	1,372	1,549	177
11	3,212	9	1,454	1,604	150
12	2,294	8	1,524	1,644	119
13	1,618	6	1,551	1,673	123
14	1,160	5	1,628	1,691	63
15	847	4	1,592	1,702	110
16	589	4	1,641	1,728	87
17	391	5	1,616	1,763	147
18	236	3	1,725	1,809	85
19	106	4	1,509	1,821	312
20	21	5	1,745	1,968	223

Table 4: Linear probability model of annual retention by male and female

	(1)	(2)	(3)
Female	-0.02***	-0.02***	0.08***
	(0.00)	(0.00)	(0.01)
Rating (100)	0.02***	0.02***	0.02***
	(0.00)	(0.00)	(0.00)
Female x Rating (100)		0.0005	0.0049***
		(0.00)	(0.00)
Male x Rating (100)		-	-
Age	-0.03***	-0.03***	-0.03***
	(0.00)	(0.00)	(0.00)
Female x Age			-0.01***
			(0.00)
Male x Age			-
Constant	0.81***	0.81***	0.80***
	(0.00)	(0.00)	(0.00)
Observations	363,252	363,252	363,252

Table 5: Predicted retention of first year males and females at the sample mean

	Ratings Percentile	Retention Female	Retention Male	Δ
Year 1				
	50	0.63	0.62	0.01
	80	0.70	0.67	0.03
	90	0.73	0.70	0.03
	95	0.77	0.73	0.04
Year 5				
	50	0.62	0.62	0.00
	80	0.72	0.70	0.02
	90	0.77	0.74	0.03
	95	0.81	0.77	0.04

Note: Predicted values based on Model (3) in Table 4.

Table 6: Returns to experience by third year percentile

	All	P (20-50)	P (50-90)	P (90-100)
Games (t & t-1)	1.97***	2.28***	2.15***	1.70***
	(0.02)	(0.04)	(0.03)	(0.04)
Female x Games (t & t-1)	0.00	-0.29***	-0.05	0.06
	(0.05)	(0.09)	(0.05)	(0.18)
Age	-41.49***	-28.70***	69.67***	76.69***
	(1.43)	(2.56)	(0.66)	(2.44)
Year	Y	Y	Y	Y
Non-active	-21.89***	-27.94***	-13.48***	8.95
	(1.30)	(1.99)	(1.95)	(6.67)
Constant	210.15***	134.91***	148.81***	426.53***
	(9.03)	(13.71)	(11.89)	(50.43)
Observations	287,283	80,265	149,322	23,443
Players	54,549	17,055	26,390	3,139

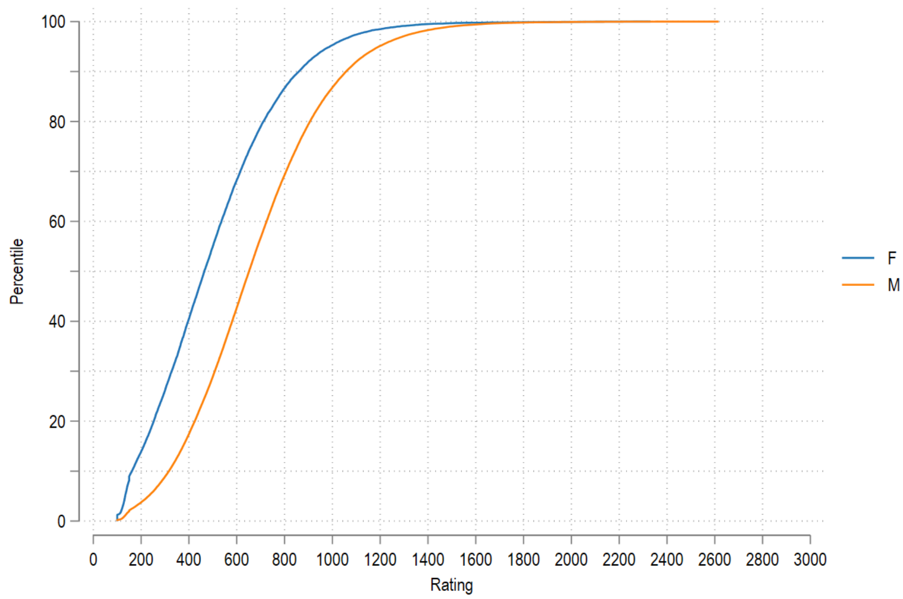
Note: A player's third year percentile is determined by their rating compared with the full sample over the entire period.

Table 7: Effect of more female chess players in locale on F-M ratings gap (locales with 20 or more players)

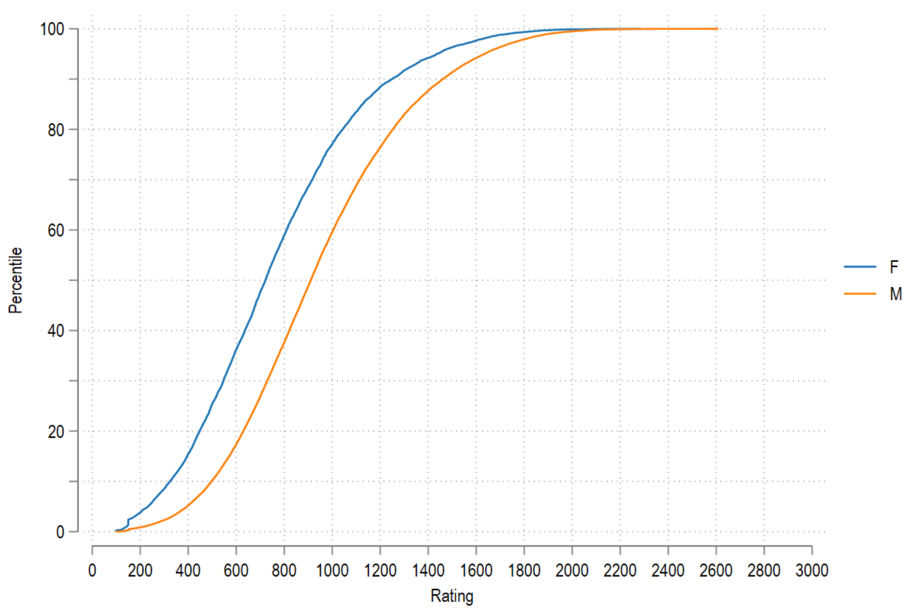
	Zip	County	MSA
Distinct Areas	452	284	164
Mean % Female	16.4	13.7	13.8
Std. Dev. % Female	9.9	5.8	4.9
F-M Gap at Median	-132.2	-138.1	-140.1
Effect of One Pct Increase in %Female on gap at Median	0.51	1.17***	2.21***
Gap at 80th Percentile	-137.4	-140.6	-144.1
Effect of One Pct Increase in %Female on gap at 80th Percentile	0.39	0.62	1.43***

Figures 1: Male-Female Gaps in Performance Ratings by Percentile

a. Year 1



b. Year 3



c. Year 5

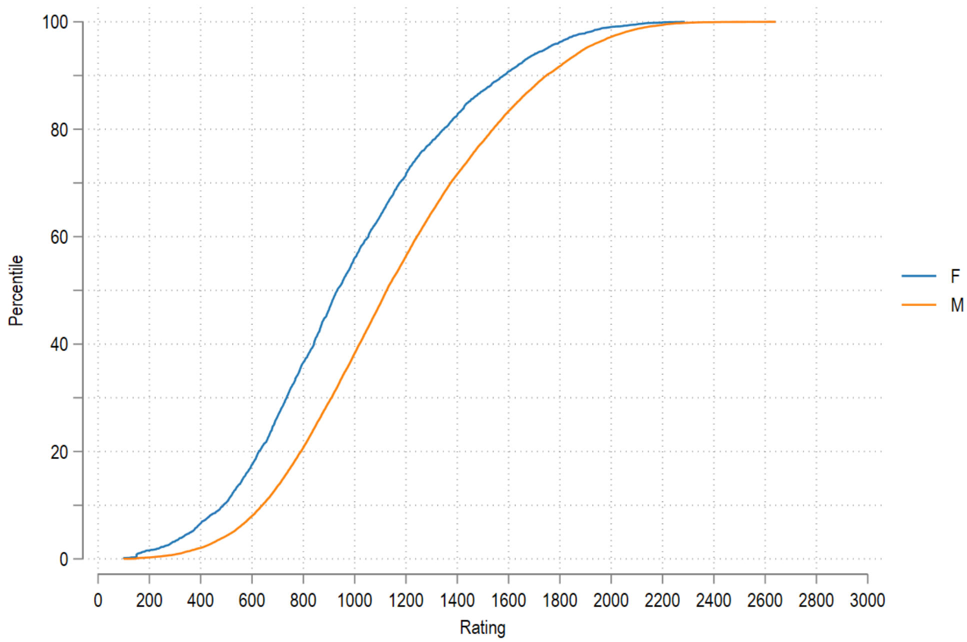
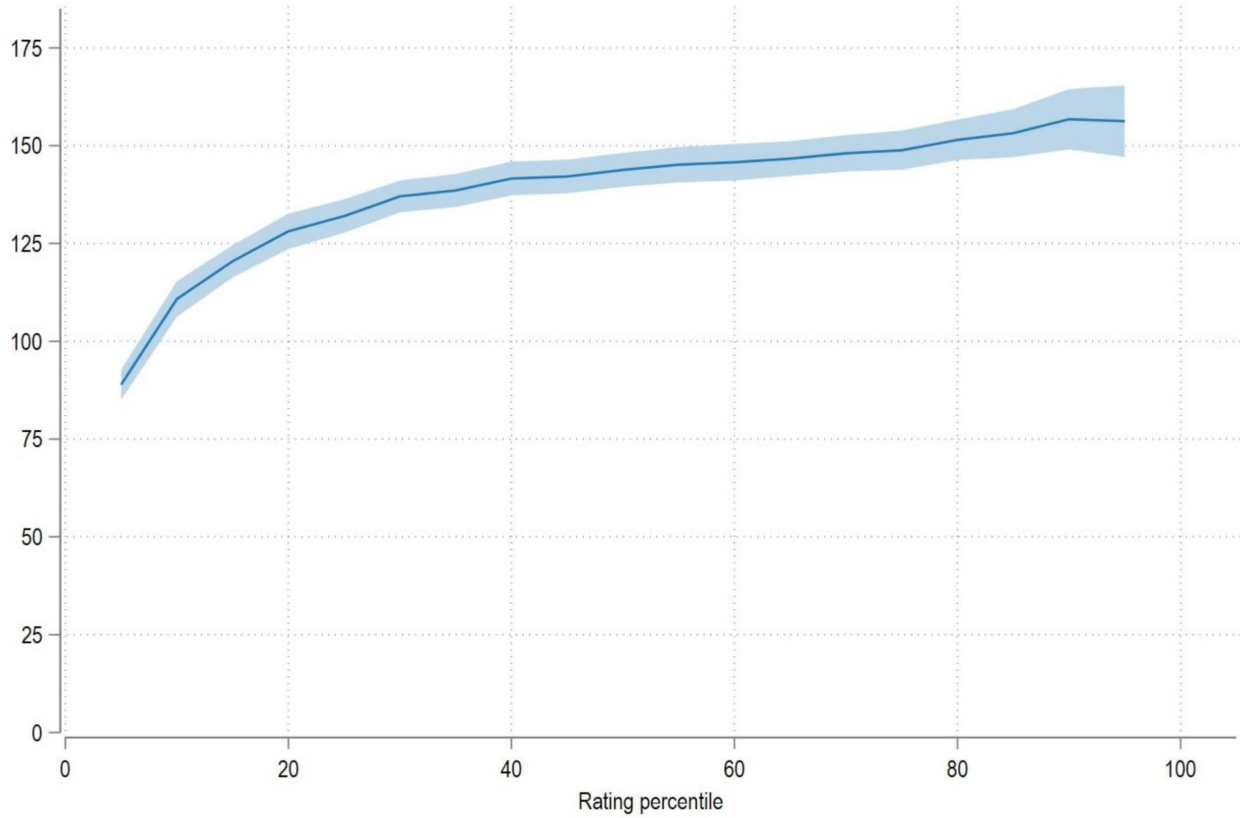

100
80
60
40
20
0
Percentile
0
200
400
600
800
1000
1200
1400
1600
1800
2000
2200
2400
2600
2800
3000
Rating
F
M

Figure 2: Quantile regression estimates of the male-female gap in ratings by rating percentiles at entry

a. Year 1



b. Year 5

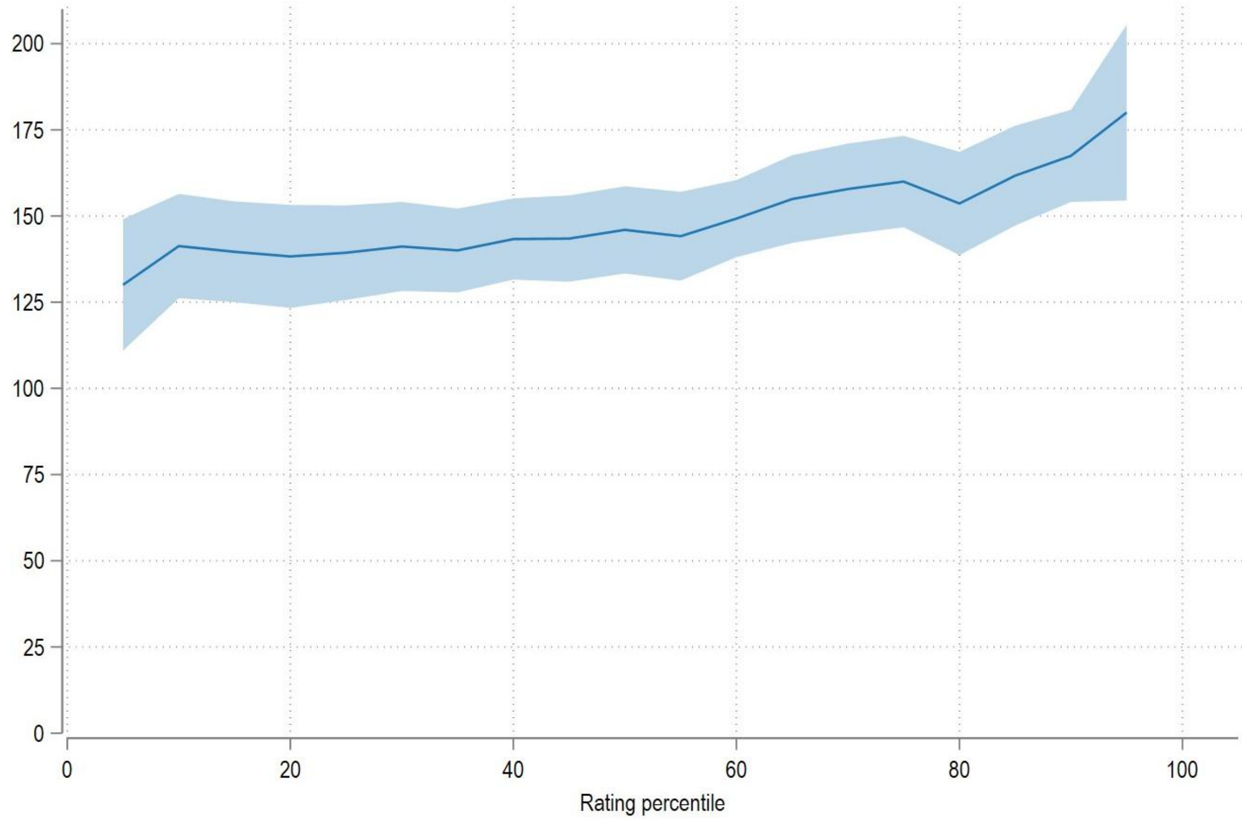

200
175
150
125
100
75
50
25
0
0
20
40
60
80
100
Rating percentile

Figure 3: Male and female retention by time in sample

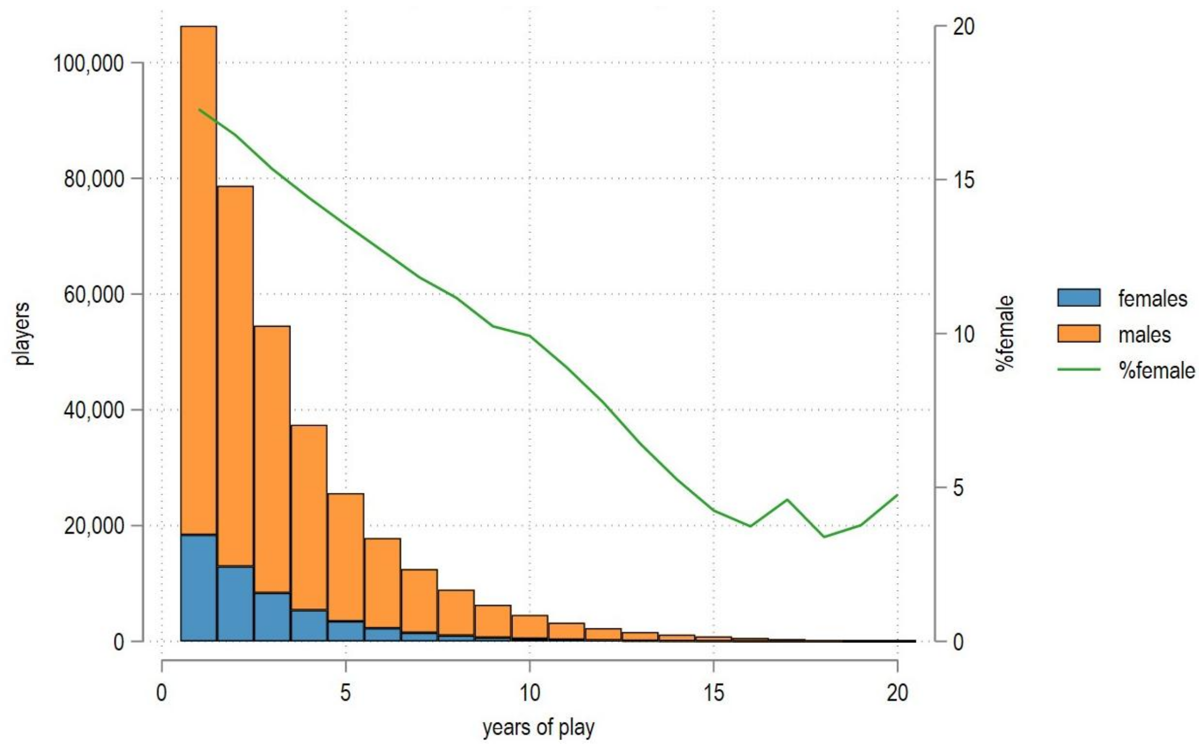


Figure 4: Effect of one percentage point increase in female zip code share on female-male gap for newly established players: zip codes with at least 20 players

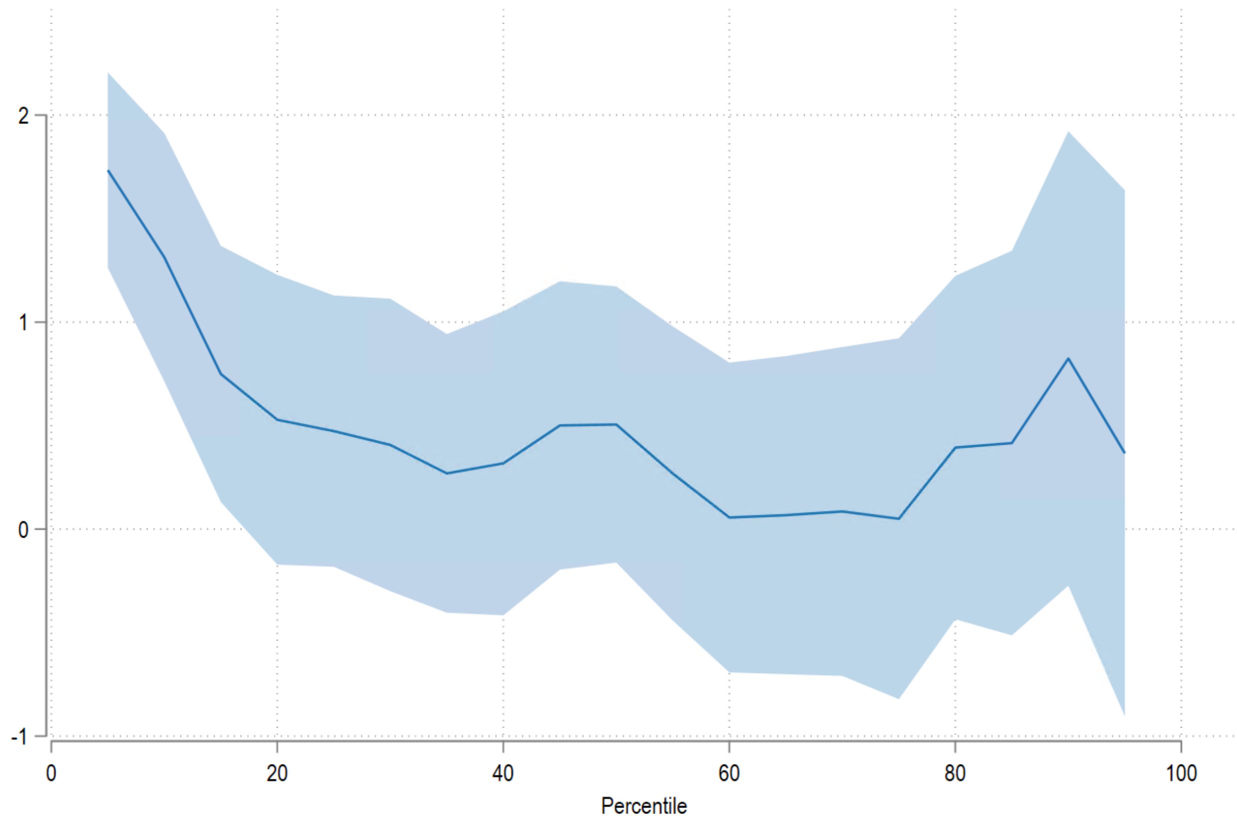


Figure 5: Effect of one percentage point increase in female county share on female-male gap for newly established players: counties with at least 20 players

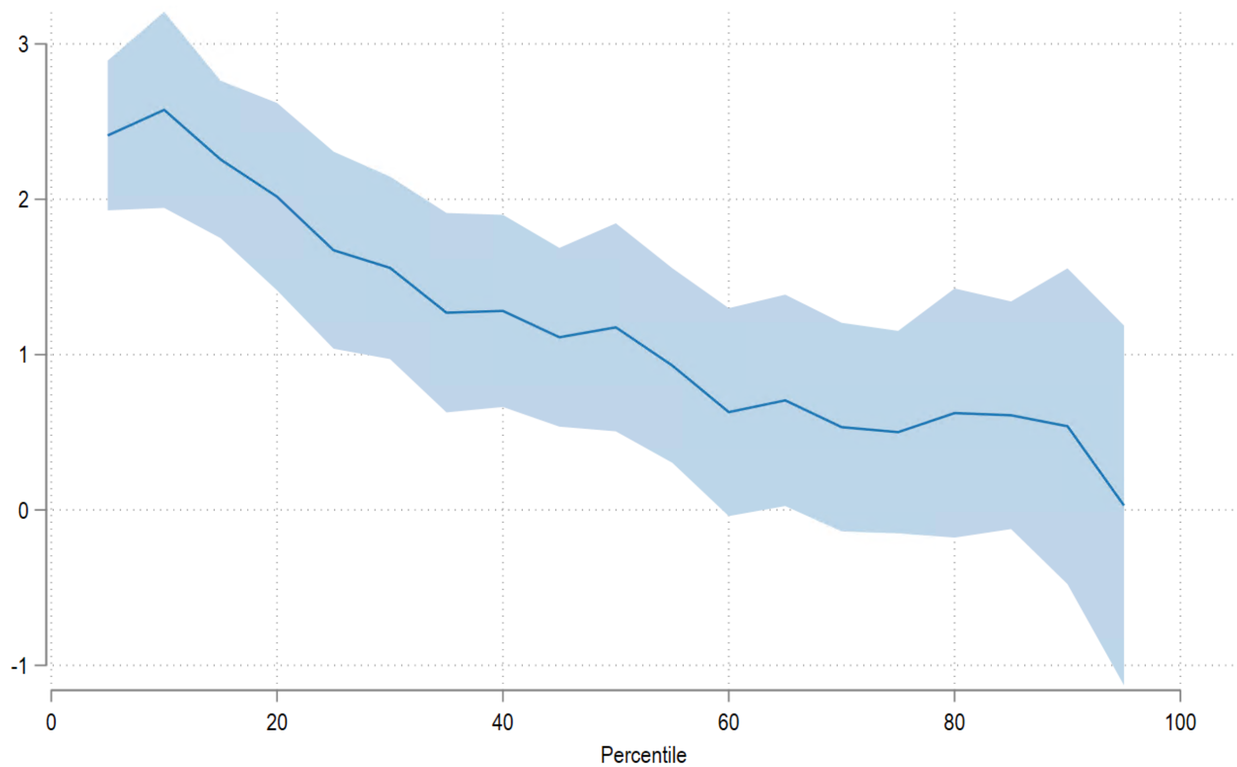


Figure 6: Effect of one percentage point increase in female MSA share on female-male gap for newly established players: MSAs with at least 20 players