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

Brian Kisida, Truman School of Government and Public Affairs, University of Missouri
Matthew Pepper, Basis Policy Research
Michael Podgursky, SCAER Saint Louis University and Department of Economics, University of Missouri
Michelle Wickman, SCAER, Saint Louis University

Abstract

We examine male-female rating gaps for young chess players using two decades of data from the U.S. Chess Federation. A contribution of our study is that we analyze the evolution of male-female gaps across a broad range of chess ratings, from novice to advanced. We find large gaps favoring males at entry and beyond across all percentiles of the ratings distribution. We find similar returns to tournament experience for males and females. Although female players have higher attrition rates than males, the net effect of this differential attrition on population ratings gaps is null. We find some evidence that that male-female rating gaps at entry narrow modestly as female participation in the home locale rises -- an effect that is generally stronger for weaker players. The key explanation for male-female differences in the population are ratings gaps at entry, which are large when first observed and persist over time.

Working Paper 25-01
Updated: 2-17-2026

The authors wish to thank the Saint Louis Chess Club and the Sinquefield Center for Applied Economic Research at Saint Louis University for supporting this research, and the U.S. Chess Federation for providing the data used in this study. We also thank Richard Pointer and Tony Rich for technical assistance, and Jeremy Compton, Alexey Root, and Ben Underwood for comments on earlier drafts. The usual disclaimers apply.

1. Introduction

According to the International Chess Federation (2023), more than twenty-five million school-aged kids worldwide play chess competitively. A growing body of research finds that chess instruction may confer academic and cognitive benefits to students (Sala and Gobet, 2016; Sala, Foley, and Gobet, 2017), and some schools now view the sport as a valuable educational tool with the capacity to strengthen executive function, concentration, and life skills instead of just a competitive game (Klein, 2023). However, males dominate the boards of most chess tournaments; in 2019 only 18% of the under 18 competitive chess players in the United States were female. The availability of data on large numbers of chess games has allowed researchers to explore key questions about the extent to which social norms influence female participation and actions during competitive and strategic pursuits. Ultimately, the male-female imbalance in chess may provide a real-world opportunity to explore barriers to female participation within competitive environments that are not inherently based on athletic ability.

Public curiosity about females 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 largely a matter of cognitive skill 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 The 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 to ultimately defeat the current world champion.

There are several respected chess ratings systems in use with varying methodologies, but at their core they 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 United States 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 20 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 a 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 (i.e., international tournament play), we examine male-female gaps along a broad range of the chess ratings distribution among young novice players.

Numerous statistical explanations have been posited to explain the presence of male-female chess performance gaps, including differences in ability at entry, differential attrition, differences in returns to experience, and the small relative female sample size. 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 null because, as compared to males with the same rating, stronger female players are at least as likely as weaker females to persist in tournament play. Leveraging the longitudinal nature of our panel with an individual fixed-effects approach, we find no differences in returns to tournament experience between male and female players. Using a simple Monte Carlo simulation, we find little evidence that male-female gaps within the far-right tail can be explained by the relatively small female sample size. 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, adolescence, and professional play. Finally, we reexamine the Chabris-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 Chabris 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.

In the next section we provide a brief overview of prior research on male-female performance differences in competitive chess and sports generally, and locates our study in the larger literature. We then describe our longitudinal data on young USCF players. Subsequent sections provide estimates of male-female ratings gaps and explore the hypotheses regarding these gaps. A final section discusses implications of our findings and potential future research.

2. Literature Review

A large literature explores male-female differences in sports performance. Much of this literature focuses not on male-female gaps in performance, per se, but rather on differences in competitiveness, risk-taking, responses to reward structures, and differential socialization. Given that chess is primarily a mental game, studies looking at sex differences stemming from psychological or cultural differences, rather than physiological differences, are especially relevant. For example, many studies have examined male-female differences in competitiveness using lab-based experiments, finding that males are more likely to embrace competitive styles of play (Niederle & Vesterlund, 2007; 2011). Using field data, Pikos and Straub (2020) examine male-female differences in competitive behavior and performance in ten-pin bowling, which has direct male-female competition, and find that males tend to outperform females. Frick (2011), however, finds that male–female performance gaps in ultramarathons have declined over time, likely reflecting increased returns to success for women (e.g., prize money) and evolving sociocultural conditions. Consistent with this pattern, Frick (2025) shows that male-female gaps in competitiveness, measured by coefficients of variation in elite track and field performance, continue to narrow.

Numerous studies explore 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 can explain the dominance of males in the very top ranks of chess players. Male-female differences in the extreme right tail 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). We examine and reject this hypothesis below using our USCF data.[4]

Others point to an 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). Chabris 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. This is another hypothesis we examine below.

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 (Bilen and Matros, 2021). Some have found mixed evidence of "stereotype threats" in which females play worse against equally ranked male opponents (Smerdon, et al. 2020; Stafford, 2018, Backus, et.al., 2023)). Using detailed tournament game-level data, Backus, et.al. (2023) find that female quality of play declines when playing males of similar rating. Maas, et.al. (2008) ran a small on-line chess experiment on a sample of Italian chess players. When female players knew that they were playing a male, performance declined relative to when the identity was disguised. Others have found that females display greater risk aversion than males in play, and that males employ more aggressive strategies when playing female opponents (Gerdes & Gränsmark, 2010). Research has also documented that females underperform relative to males in response to pressures from time-control of games (Dilmaghani, 2020, 2021; Gränsmark, 2012). A recent study examines male-female differences in response to "personal bests" in tournament play and rankings, finding that females increase their effort relative to males when approaching their personal best, but exert less effort once surpassing it (Gonzales-Diaz, et.al., 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 as to potential explanations for small 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 USCF 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 Chabris and Glickman's (2006) analysis, where they find suggestive evidence for a positive female peer effect in one's home locale. As we will see below, important insights emerge when we examine the effect of variables across the full range of performance percentiles.

3. Data description

We use data from the USCF to examine possible male-female differences in ratings among young players at officially sanctioned tournaments within the United States. 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 a player's ability relative to all other players in the USCF system. The anonymous player-level data includes player rating, sex, birth year, 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 ratings 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, in calendar 2019, females represent just 16 percent of all players but 20 percent of players aged 5-15 in that year. Our dependent variable is chess rating. This rating is not considered reliable by USCF until at least 25 rated tournament games have been played. Thus, we restrict our sample to players who have at least 25 rated games.[5] 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 more young cohorts are added. 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 ratings (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 CDFs 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)

4. Estimated Gaps at Entry and at Five Years

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 & 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 y_i is the rating of player i, and X_i includes: 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, percentage of residents identifying as White, and the locale type (city, suburb, town/rural). *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 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. Tables 3 and 4 report the estimated coefficients on *Male*, and selected regression statistics at selected percentiles in years one and five. It is interesting to note that the R^2 rises at all percentiles in year 5 as compared to year 1. One possible interpretation is that player ratings contain less "noise" as more tournament games are played and true player ability is more precisely measured.

(Figure 2 and Tables 3-4)

5. Male-Female Differences in Attrition

Analysis in the previous section shows that there is a large male-female gap in ratings at entry (i.e., when 25 rated games have been played). 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 5. 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 higher 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 5)

In order to explore the effect of differential attrition on the ratings gap, we exploit the panel nature of our data to estimate Cox proportional hazard models. The general structure of these models is shown in equation (2) below:

$$\text{h}\ (t; x))_{it} = h_0(t)\ \exp(\beta_0 + \beta_1 Age_{it-1} + \beta_2 Rating_{it-1} + \beta_3 Male_i + \beta_4\ Male_i\ x\ Rating_{it-1}) \quad (2)$$

where *i* denotes the *i*-th player and *t* denotes year of play. This is a semi-parametric model of the exit hazard (i.e., the conditional probability of exit given that the player has continued in tournament competition to year *t*-1). A non-parametric baseline hazard, $h_0(t)$, is fit to the entire sample. Covariates act to shift the baseline hazard up or down in a proportionate manner. Two key covariates – *Age* and *Rating* – vary over time for any individual player and across players. The sex of the player, indicated by the binary variable *Male*, is time invariant. Of particular interest for us is the effect of *Male x Rating*. This tells us what is happening to the male-female gap with attrition. A value of β_4 less than one tells us that higher rated males exit at a lower rate than similarly rated females. This would imply that the male-female gap among survivors widens. The estimated effects are reported in Table 6 below. Here we report the proportionate effect of a unit change in a covariate on the relative hazard rate. For example, a one hundred point gain in USCF ratings implies a reduction in the relative exit hazard to .92, or an 8 percent reduction in exit probability.

(Table 6)

The estimates in the first column of Table 6 are plausible. For our sample of young players, exit rates decline with ratings but rise with age. The key estimate for us, however, is whether the ratings effect differs between men and women (β_4). Here we see that the effect is not significantly different from unity, implying similar ratings effects by sex. This means that the gap between male and female ratings do not change among players who "survive" and continue in tournament play.

Estimates in the first column constrain the baseline hazard for males and females to be the same (allowing a proportional shift for males). Estimates in column (2) relax this assumption and allow different non-parametric baseline hazards for males and females. A likelihood-ratio test of equality of baseline hazards by sex is easily rejected. Overall, our coefficient estimates are very similar. In particular, the estimated value of β_4 is very close to unity. As before, we fail to reject the hypothesis that the effect of ratings on exits differs between males and females. This implies that the higher female attrition rate has no effect on the male-female ratings gap among survivors.

6. 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; Backus, et.al.,2023). This has led some to advocate for female-only tournaments, while others claim that these hurt female talent development.[6] 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:

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

$$\beta_4 \, Games_{i,t\,\&(t-1)} \times Female_i + \mu_i + \varepsilon_{it} \; (3)$$

where y_{it} is the rating of player *i* at the end of year *t*, X_{it} is a set of time-varying controls (including age in year *t*, and a set of calendar year dummy variables), $Games_{i,t\,\&(t-1)}$ 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 7. 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 ratings percentiles, we present overall estimates and by ratings bands based on a player's third-year rating, where (P20-50) includes players in percentiles 20-50, etc. We restrict the sample to players with at least three years of experience.

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 female return is lower. However, the point estimate of the gap in returns per game (.29 rating points) is small. Among stronger initial players, there is no significant male-female difference in returns to experience.

(Table 7)

This approach mirrors an educational production function (Hanushek, 1986) and, from the authors' search, is the first applied to chess ratings. Table 7 shows that, on average, a young US chess player gains 2.0 ratings points with each additional tournament game played.[7]

7. The Extreme Right Tail

Earlier sections explored patterns throughout the distribution; in this section we turn to the male-female gaps in the extreme right tail of the ratings distribution -- the 99th percentile and above. Considerable public interest has focused on the extreme right tail of the performance distribution. Some researchers posit an important role to the small relative female sample size on gaps in the extreme right tail, hypothesizing that achieving participation parity is a sufficient pathway to closing the male-female gap. For example, Bilalic, et al. (2008) attribute 96 percent of the gap among the top 100 German chess players to the smaller relative share of females. Our analysis, however, finds that the smaller relative participation rate of female players has little effect on the male-female ratings gap within the top 1 percent of our sample of young chess players.

We explored the male-female gaps in extreme values in our sample of young USCF players, with results shown in Table 8 below. The columns labelled "Data" report the coefficients on a male dummy variable from a quantile regression that includes only age as an additional regressor. The columns labelled "Simulation" are from a simple Monte Carlo simulation in which we pool the male and female subsamples and draw 10,000 random samples the same size as the actual male and female shares. We report the resulting gaps for these "pseudo-sex" samples at the various percentiles. These represent the right tail ratings gaps one would expect simply based on the smaller female population shares.

(Table 8)

We note that the quantile regression coefficients on the extreme right tail become less precise as we move further out in time (five versus ten years), as the sample of females is shrinking overall and relative to males. However, in all cases, males outscore females and the gap is statistically significant. We compare the simulated gap to the actual gap to assess the "share of females" effect and find the effect of female share is very small relative to the actual gap. Only in the case of the 99.5th percentile does the simulation reach ten percent of the actual gap.

This analysis indicates that our core findings for the lower 99 percentiles extend to the top percentile as well. The principal factor underlying the lower ratings of top female chess players relative to their male counterparts is a leftward shift in the overall distribution of female player ratings, rather than the smaller number of female participants.

8. Further Analysis of Gaps at Entry

The above analysis shows that the key factor affecting ratings gaps in the population of young USCF players are ratings gaps at entry. What can USCF data tell us about factors affecting entry gaps? Analyzing earlier cohorts of young USCF players, Chabris and Glickman (2006) find some evidence that the presence of female peers improves the relative performance of female chess players. They find that the female share of participants in the player's home zip code affects the male-female gap. Specifically, as the share approaches 50 percent (parity), the mean male-female gap is nearly zero. We explored the relationship between spatial measures of female participation rates and performance gaps in our panel of young chess competitors

We begin by restricting the sample to newly established (year 1) players. We extended model (1) to include the female share of chess participants ($ShareF_j$) in the home locale:

$$y_{ij} = \beta_0 + \beta_1 X_i + \beta_2 Female_i + \beta_3 ShareF_j +$$

$$\beta_4\, ShareF_j \times Female + \varepsilon_i \quad (4)$$

where y_{ij} refers to the rating of player i in home locale j. In the model specified in (4), 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. In our estimates below, we use three definitions for home "locale" that are available in our data: zip code, county, and MSA. 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.[8] In Table 9 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 9)

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 in Figure 6. 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 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 Chabris-Glickman hypothesis that greater female participation narrows female-male ratings gaps. However, the calculations in Table 9 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 for a player at the median rating 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 134 rating points.

9. Conclusion.

Chess is a sport in which performance and international rankings among tournament players are summarized in a single metric (an Elo rating). A large and persistent gap in ratings exists between top male and female players and a substantial literature has examined possible explanations for these gaps. This paper contributes to the literature by examining male-female ratings gaps using two decades of data from the USCF on young players who are just beginning tournament competition. 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 arise almost entirely from ratings gaps at initial entry, that is, when youth have played enough tournament games to get a reliable rating. 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 higher attrition rates than males, but the net effect of this attrition difference on ratings gaps is null because as compared to males, stronger female players are as likely to persist as weaker females.

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 Chabris 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 competitive chess play. At the same time, disparate outcomes are not necessarily evidence of discrimination, as this may partially reflect different male-female preferences. Indeed, our analysis shows that males and females have similar gains from experience, suggesting that females have 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 would require policies that address the mechanisms of recruitment of females into chess. Indeed, USCF and other organizations have made considerable efforts to recruit more young girls into competitive chess. The popularity of the Netflix series *The Queen's Gambit* may have stimulated interest as well.

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 our analysis of large longitudinal files such the USCF data in this study can shed light on how chess "stars" are grown and cultivated.

A contribution of this study is that we analyze a much broader segment of the U.S. chess-playing population, and arguably, a much larger population of young people, than most studies in the literature which rely on data from international tournament competitions. In this regard, it is important to recognize that players in these international competitions (sanctioned by the international chess federation, FIDE) are clearly in the far right tail of the chess ability distribution. This is important because many prior chess studies draw inferences about male and female differences (e.g., risk-aversion and other behavior characteristics) in other domains. Our USCF database is composed of young, novice players, most of whom will never participate in, or qualify for, an international FIDE tournament. So while we seek to contribute to the literature on male-female differences in competitive chess, it should be recognized that we are analyzing a very different group of players. Indeed, it is somewhat analogous to a study of young novice junior high or high school basketball players as compared to top collegiate or NBA players.

Finally, an important limitation of the USCF database, and hence this study, is that we only know the outcome of games played in rated tournaments (i.e., win, lose, draw) and consequent effects on player ratings. As discussed in our literature review, many recent studies use data on expert tournament games (e.g., Gerdes and Gränsmark, 2010; Backus, et.al. 2023) to analyze differences in male and female play. A valuable extension of the research in this paper would be to analyze differences in male-female play among the novice young players in our sample. To this end, the researchers are exploring the feasibility of collecting and analyzing detailed game level data on young scholastic competitors in USCF tournaments sponsored by a large U.S. chess club.

References

Backus, P., Cubel, M. Guid, M., Sanchez-Pages, S., & Manas, E.L. (2023). Gender, Competition, and Performance: Evidence from Chess Players. *Quantitative Economics.* 14(1), 349-380.

Bertoni, M., Brunello, G., & Rocco, L. (2015). Selection and the age-productivity profile: Evidence from chess players. *Journal of Economic Behavior and Organization. 110, 45-58.*

Bilalić, 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. *Proceedings of the Royal Society B: Biological Sciences, 276*(1659), 1161–1165.

Bilen, E., & Matros, A. (2021). The Queen's Gambit: Explaining the superstar effect using evidence from chess. *Journal of Economic Behavior & Organization, 215*, 307–324.

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*(2), 81–106.

CBS New York. (2025, April 10). Success Academy middle school girls win national chess championship title. *CBS News.* https://www.cbsnews.com/newyork/news/success-academy-national-chess-championship-title/

Chabris, C., & Glickman, M. (2006). Sex differences in intellectual performance: Analysis of a large cohort of competitive chess players. *Psychological Science, 17*(12), 1040–1046.

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*.

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*(C).

Dreber, A., Gerdes, C., & Gränsmark, P. (2013). Beauty queens and battling knights: Risk-taking and attractiveness in chess. *Journal of Economic Behavior & Organization, 90*, 1–18.

Frick, B. (2011). Gender Differences in Competitiveness: Evidence from Professional Distance Running. *Labour Economics* 18, 389-398.

Frick, B. (2025). Gender Differences in the Concentration of Individual Performance: Empirical Evidence from Track and Field Athletics. *Journal of Sports Economics*. 26 (1), 35-54.

Gerdes, C., & Gränsmark, P. (2010). Strategic behavior across gender: A comparison of female and male expert chess players. *Labour 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.

Gränsmark, P. (2012). Masters of our time: Impatience and self-control in high-level chess games. Stockholm University: Swedish Institute for Social Research. Working Paper 2/2012.

Halpern, D. F., Benbow, C. P., Geary, D. C., Gur, R. C., Hyde, J. S., & Gernsbacher, M. A. (2007). The science of sex differences in science and mathematics. *Psychological Science in the Public Interest, 8*(1), 1–51.

Hanushek, E. A. (1986). The economics of schooling: Production and efficiency in public schools. *Journal of Economic Literature, 24*(3), 1141–1177.

Howard, R. (2014). Explaining male predominance at the apex of intellectual achievement. *Personality and Individual Differences, 68*, 217–220.

International Chess Federation. (2023, April 25). Scholastic chess x FIDE: The goal is 50 million kids. *FIDE.* Retrieved June 20, 2025, from https://www.fide.com/news/2360

Klein, A. (2023, April 14). 'Everybody at school wants to play': Chess is trendy again. *Education Week.* https://www.edweek.org/teaching-learning/everybody-at-school-wants-to-play-chess-is-trendy-again/2023/04

Koenker, R., & Hallock, K. F. (2001). Quantile regression. *Journal of Economic Perspectives, 15*(4), 143–156.

Maass, A., D'Ettole, C., Cadinu, M. (2008). Checkmate? The role of gender stereotypes in the ultimate intellectual sport. *European Journal of Social Psychology*, 38(2), 231-245.

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

Niederle, M., & Vesterlund, L., (2011). Gender and Competition. *Annual Review of Economics* 3(1), 601-630.

Pikos, A.K., & Straub, A. (2019). Mind the Absent Gap: Gender-Specific competitive behavior in Nonprofessional Sports. *Journal of Sports Economics.* 21(3), 215-233.

Sala, G., Foley, J., Gobet, F. (2017). The Effects of Chess on Pupils' Cognitive and Academic Skills: State of the Art and Theoretical Challenges. *Frontiers in Psychology*. Feb. 23: 8: 328.

Sala. G, & Gobet, F. (2016). Do the Benefits of Chess Instruction Transfer to Academic and Cognitive Skills: A Meta-Analysis. *Educational Research Review. 18 (May): 46-57.*

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*(6), 756–759.

Stafford, T. (2018). Female chess players outperform expectations when playing men. *Psychological Science, 29*(3), 429–436.

Vishkin, A. (2022). Queen's Gambit Declined: The gender-equality paradox in chess participation across 160 countries. *Psychological Science.* 33(2), 276-284.

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: Year 1 Quantile Regression Results at Selected Quantiles

	Q10	Q20	Q50	Q80	Q90
Female	-110.64***	-127.27***	-142.16***	-148.03***	-151.40***
	(2.21)	(2.26)	(2.18)	(2.53)	(3.64)
Player controls	Y	Y	Y	Y	Y
Home area controls	Y	Y	Y	Y	Y
Constant	3.99	55.56***	146.93***	278.66***	344.54***
	(14.85)	(12.29)	(10.20)	(10.20)	(13.91)
Observations	105,153	105,153	105,153	105,153	105,153
R2	0.16	0.18	0.21	0.23	0.23

Notes: Robust standard errors are reported in parentheses. Player controls include sex (male or female), age, tournament year, and the number of games played in the prior two years. Home-area controls include median household income, the share of White residents, ZIP code locale (city, suburb, town, or rural), and the number of players and tournaments located in the player's home county; locale classifications are updated in 2015 and 2020. ZIP code–level characteristics are observed annually from 2011–2019; for observations from 2000–2010, ZIP code characteristics are fixed at their 2010 values.

Table 4: Year 5 Quantile Regression Results at Selected Quantiles

	Q10	Q20	Q50	Q80	Q90
Female	-140.43***	-140.26***	-146.25***	-147.21***	-161.53***
	(7.80)	(7.57)	(6.59)	(6.43)	(6.88)
Player controls	Y	Y	Y	Y	Y
Home area controls	Y	Y	Y	Y	Y
Constant	-115.81***	-111.27***	-46.20	87.21***	225.70***
	(28.75)	(30.22)	(28.17)	(32.30)	(37.31)
Observations	25,300	25,300	25,300	25,300	25,300
R2	0.24	0.27	0.32	0.35	0.34

Notes: Robust standard errors are reported in parentheses. Player controls include sex (male or female), age, tournament year, and the number of games played in the prior two years. Home-area controls include median household income, the share of White residents, ZIP code locale (city, suburb, town, or rural), and the number of players and tournaments located in the player's home county; locale classifications are updated in 2015 and 2020. ZIP code–level characteristics are observed annually from 2011–2019; for observations from 2000–2010, ZIP code characteristics are fixed at their 2010 values.

Table 5: 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 6: Tournament Play Exits: Cox Hazard Model Estimates

	Pooled	Stratified by sex
Rating (/100)	0.920***	0.917***
	(0.002)	(0.002)
Age (t-1)	1.113***	1.113***
	(0.002)	(0.002)
Games (t-1)	0.991***	0.991***
	(0.000)	(0.000)
Male	0.929***	
	(0.016)	
Male x Rating (/100)	0.997	1.000
	(0.002)	(0.003)
Log Likelihood	-838,849	-802,076
Observations	316,933	316,933

Notes: Standard errors are in parentheses. We estimate the models with right-censored duration data. Players are observed from entry to either failure or censoring at the end of 2019. There are 79,557 players with "failures" in the data and 26,811 players who are right censored.

Table 7: 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

Notes: A player's third year percentile is determined by their rating compared with the full sample over the entire period. Non-active is an indicator variable taking the value one if the player had an inactive year but then returned.

Table 8: Male-Female Ratings Gaps in the Right Tail

	99th		99.5th		99.9th	
Experience	Data	Simulation	Data	Simulation	Data	Simulation
year 5	140	1.2	95.5	5.5	125	2.4
	(21.5)	(20.9)	(20.2)	(24.7)	(40.8)	(40.5)
year 10	102.25	5.5	148	4.8	204	36.3
	(46.2)	(48.6)	(112.3)	(68.3)	(90.2)	(86.8)

Notes: Total player counts: Year 5: (male) 21,950 (female) 3,422; Year 10: (male) 4,064 (female) 447

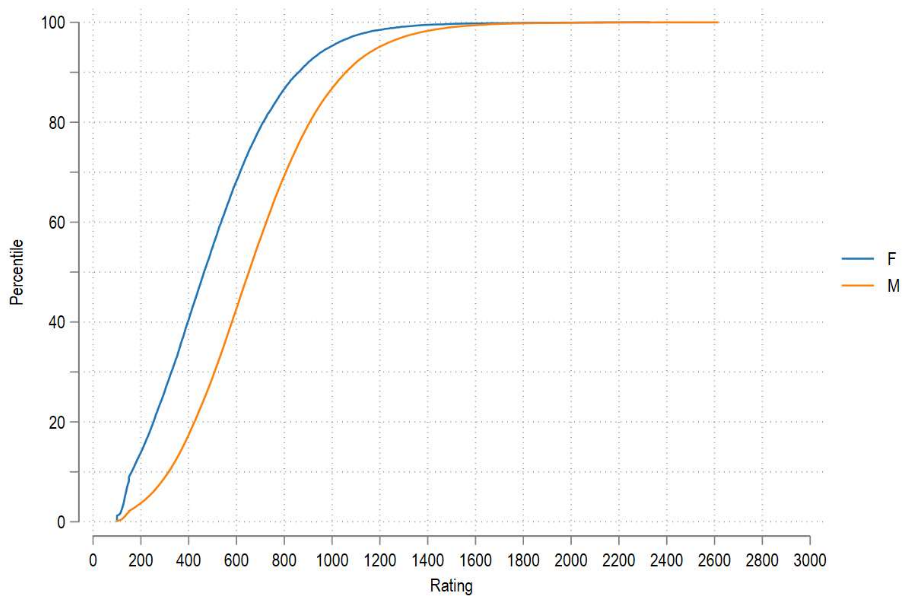
Table 9: Effect of More Female Chess Players in Home Locale on Female-Male Ratings Gap

	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	-130.6	-138.1	-140.3
Effect of One Pct Increase in %Female on gap at Median	0.22	1.18***	2.26***
Gap at 80th Percentile	-136.4	-140.8	-143.4
Effect of One Pct Increase in %Female on gap at 80th Percentile	0.34	0.74	1.39***

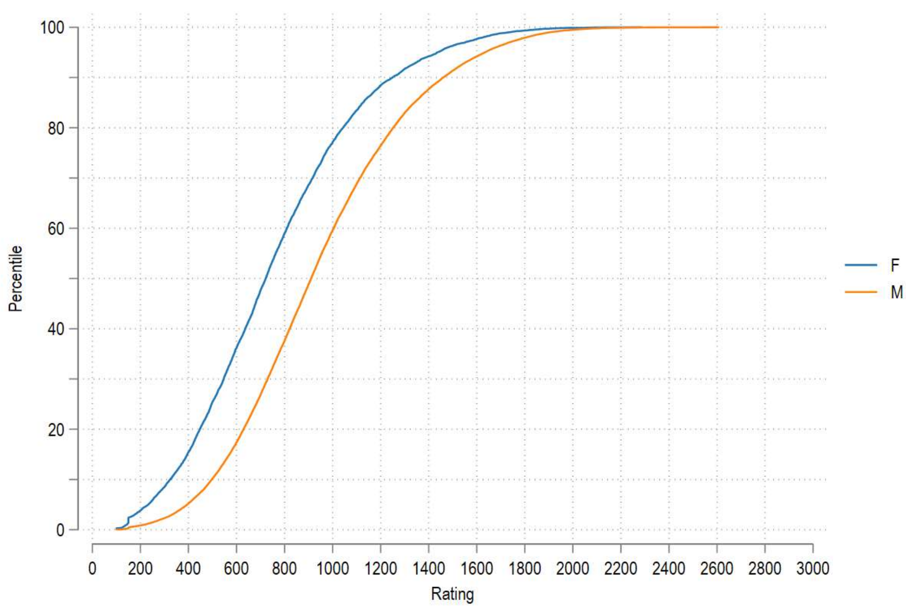
Note: Locales with at least 20 Players.

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

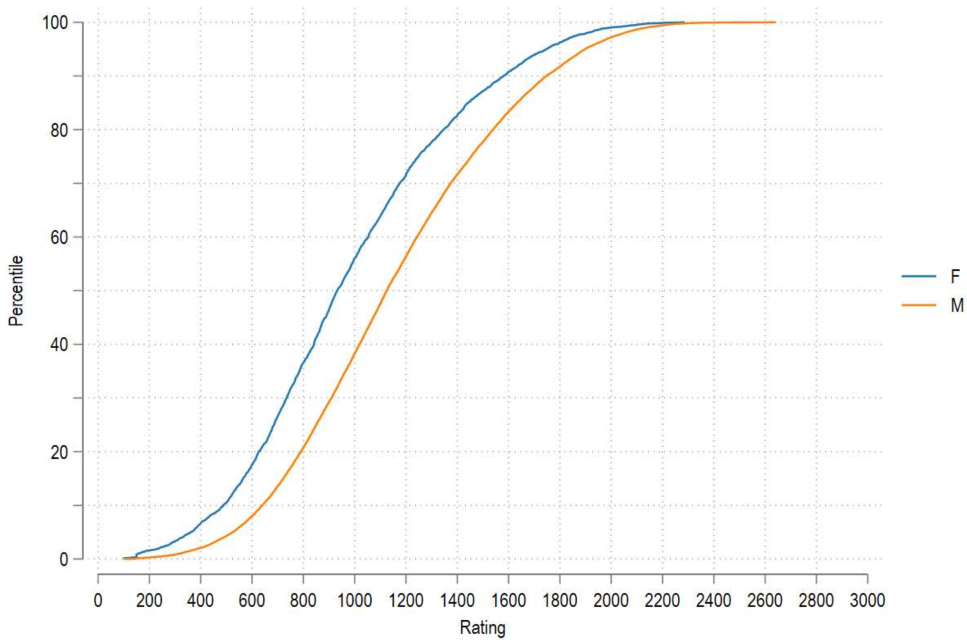
a. Year 1



b. Year 3



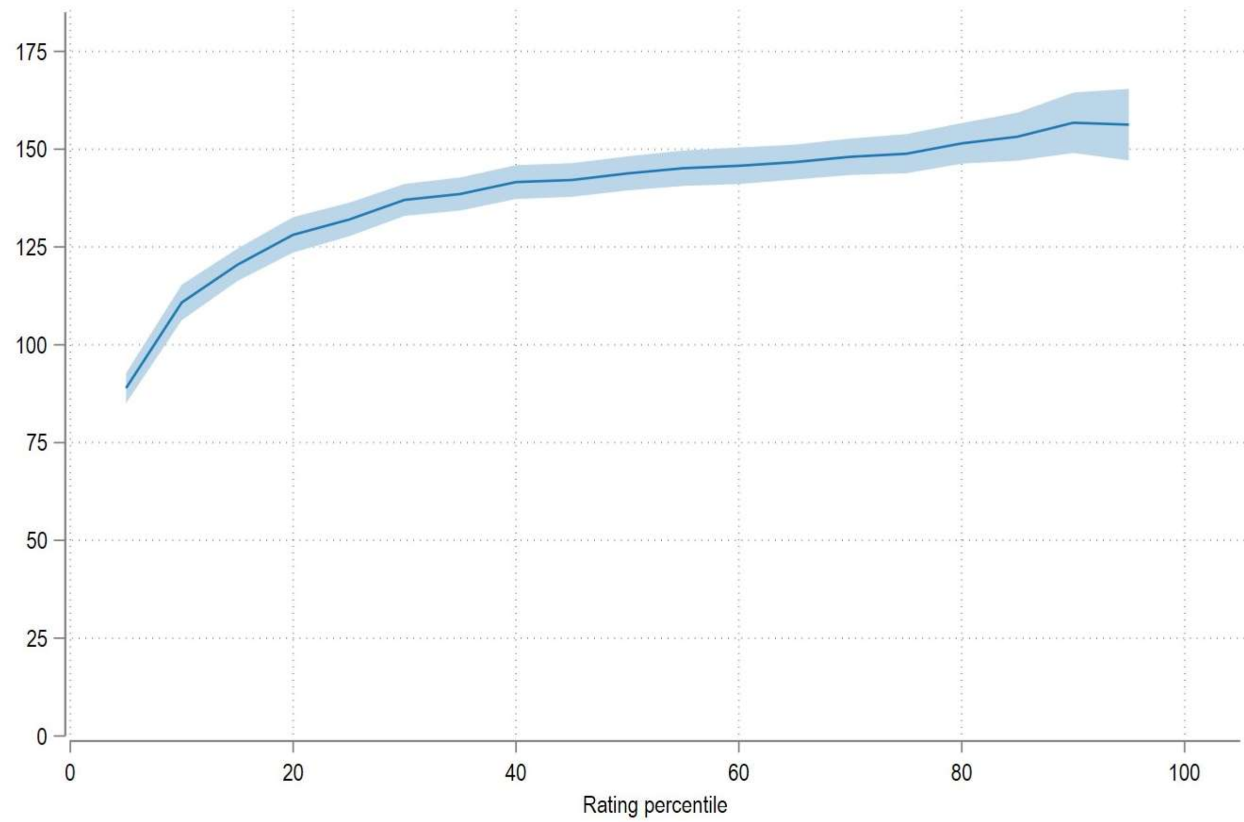
c. Year 5



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

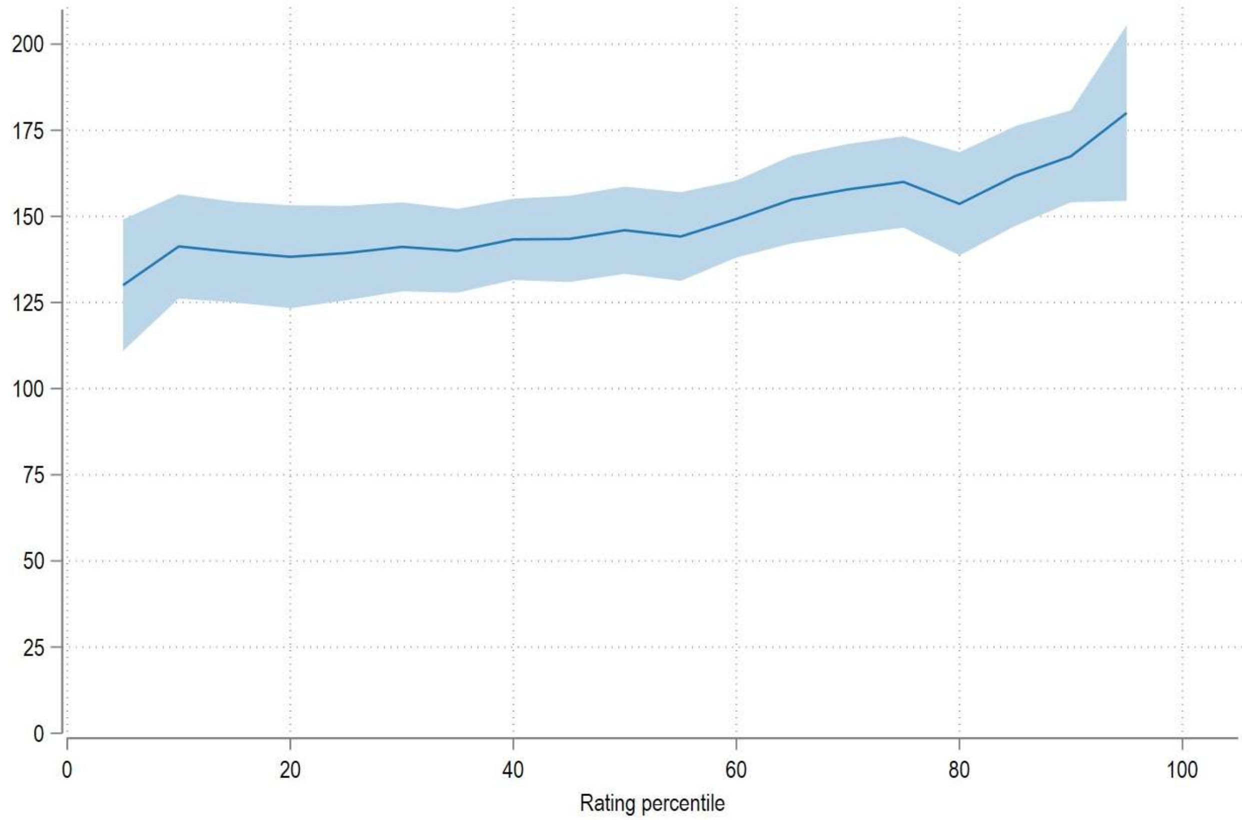
Figure 2: Quantile Regression Estimates of the Male-Female Ratings Gap by Rating Percentiles at Entry and Year 5

a. Year 1



Note: These are estimates of β_2 in Model (1) by Ratings Percentile. See text.

b. Year 5



Note: These are estimates of β_2 in Model (1) by Ratings Percentile. See text.

Figure 3: Male and Female Retention by Years of Tournament Play

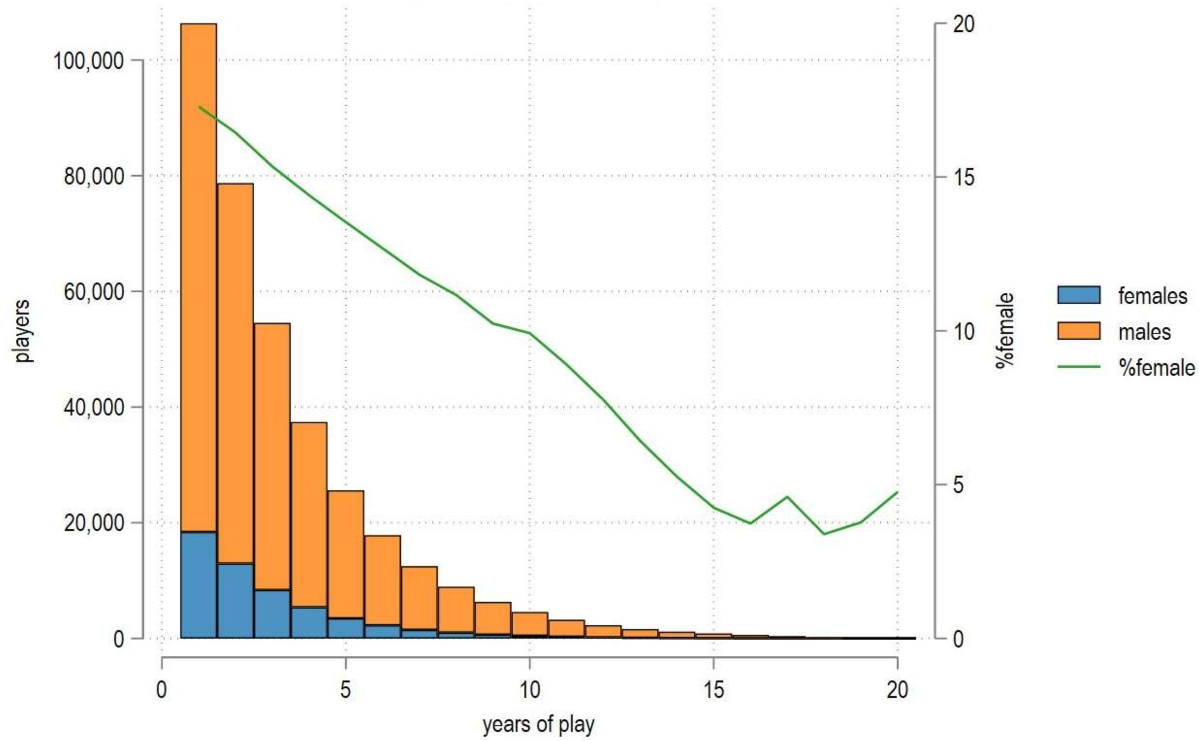
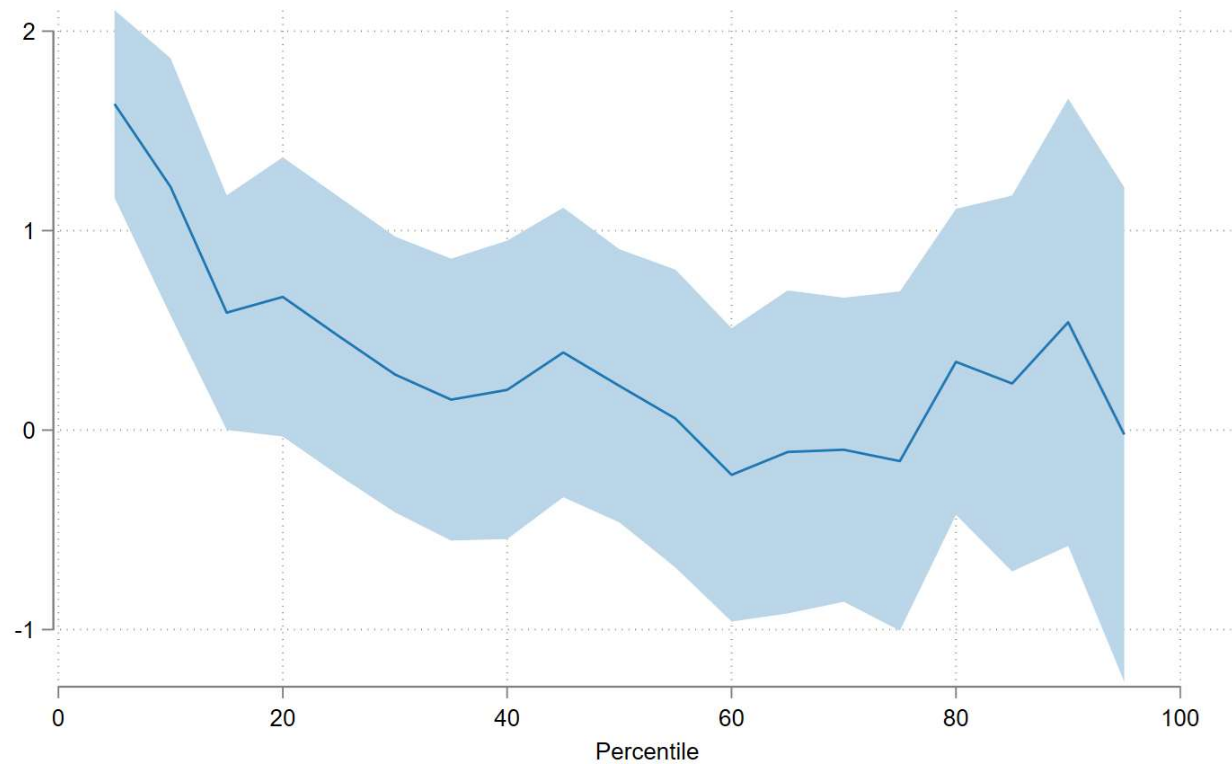
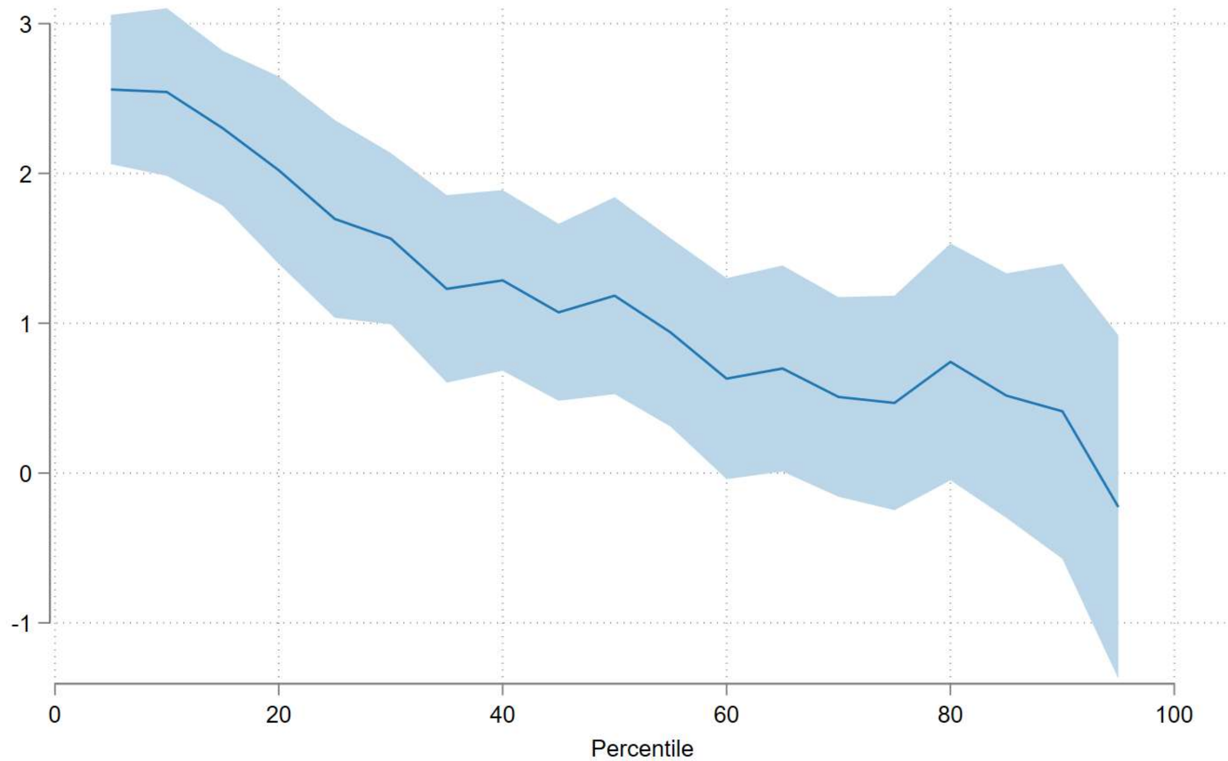


Figure 4: Effect of One Percentage Point Increase in Female Zip Code Share on Female-Male Gap for Newly Established Players



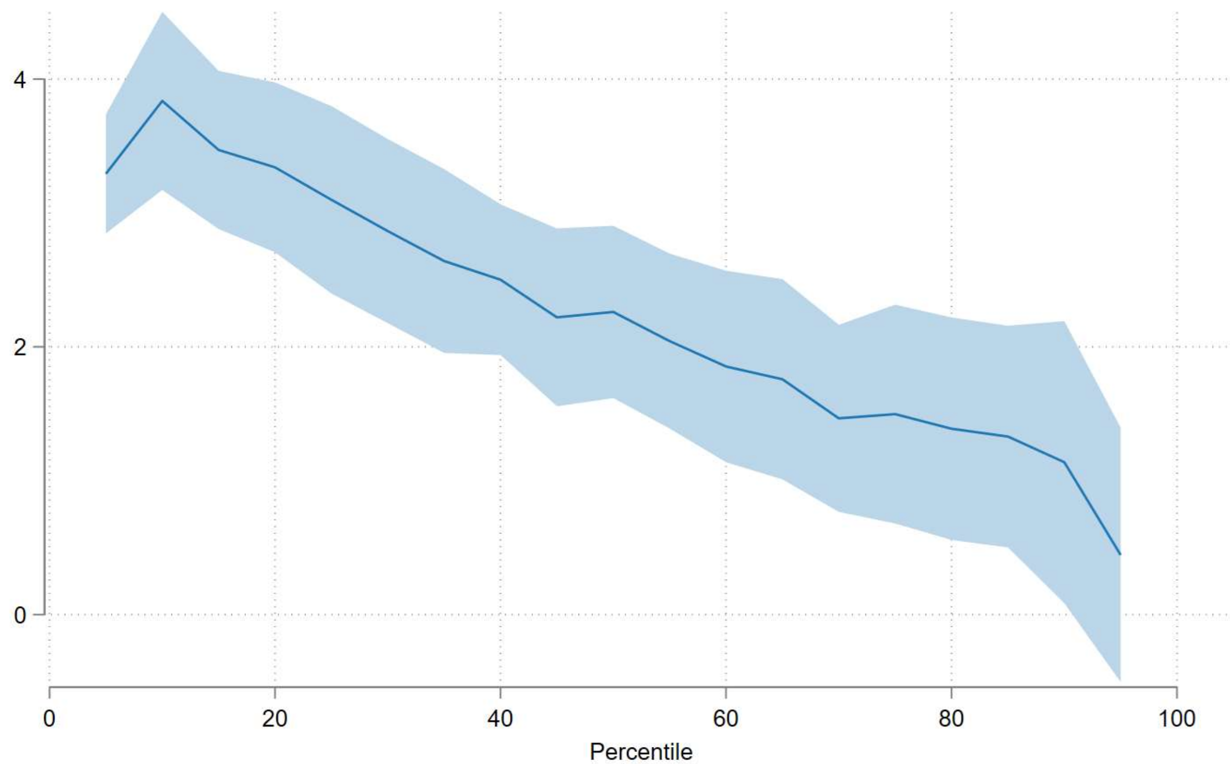
Note: Zip codes with at least 20 Players.

Figure 5: Effect of One Percentage Point Increase in Female County Share on Female-Male Gap for Newly Established Players



Note: 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



Note: MSAs with at least 20 Players.

Endnotes

---

[1] https://ratings.fide.com/
[2] https://en.wikipedia.org/wiki/List_of_female_chess_grandmasters. In recognition of the large gap in male-female chess performance, FIDE introduced a Woman's Grandmaster title in 1976. This sets a lower bar than the overall grandmaster title. Similarly, since 1937, the USCF has run a female-only national invitational tournament to determine the top female U.S. player (the U.S. Chess Women's Chess Championship). This is the exception, however, over 95 percent of the tournaments in our USCF database are not segregated by sex. Females competing in these tournaments can expect to play males and females.
[3] Similar to FIDE ratings, USCF ratings rise and fall based on player performance in rated games. Both of these are Elo scores and are intended to predict the winner of a game based on the prior history of play. For details see https://www.houseofstaunton.com/blogs/chess-facts/elo-chess-rating-system#:~:text=The%20Elo%20(%20Elo%20Rating%20System%20),US%20Chess%20Federation%20)%20and%20FIDE%20ratings.
[4] An analysis of international differences in chess participation (Vishkin, 2022) finds a curvilinear relationship between female chess participation and measures of overall social gender equality, with chess female participation highest in low and high equality countries
[5] 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, 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.

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

[7] Bertoni,et.al. (2015) explore the effect of attrition on estimates of the effect of age on chess performance. They find that differential attrition imparts an upward bias on the effect of age and performance. While a similar phenomenon may be present in estimating the return to experience (as measured by tournament chess games played), our focus is on male-female gaps in these returns and not on the level per se.
[8] 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.