Career Choices and the Evolution of the College Gender Gap

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Abstract

We propose an explanation for the evolution of the college gender gap that emphasizes the increase in men's opportunity cost of attending college due to an increase in the return of becoming a superstar. We present evidence from a natural experiment in European soccer markets (the so-called "Bosman ruling") that generated an exogenous increase in men's (and, crucially, not in women's) payoffs in a prominent superstar path. In line with our explanation, we find that countries more exposed to the Bosman ruling experienced a significant increase in the female-male ratio in college relative to less-exposed countries.

JEL Codes: I20, J16, J24.

Keywords: gender gap, superstars, education.

1 Introduction

It is a well-documented fact that women have been catching up with men in terms of educational achievement, and have actually overtaken them in many countries. Figure 1 plots the recent evolution of the average ratio of female to male tertiary enrollment for 147 countries from all over the world, and shows a clear upward trend in female college enrollments relative to male enrollments, with women catching up with men in the early 1990s, and consistently overperforming them in terms of college attendance ever since.

[FIGURE 1 ABOUT HERE]

As Goldin, Katz, and Kuziemko (2006, p. 153) point out, the puzzle here is "[w]hy have females surpassed males in college going and college completion and not simply caught up to them." Several explanations to this puzzle have been advanced.[1] For instance, changes in social norms and expectations about the roles of work, marriage, and motherhood for women plus greater protection for gender equality at work and the introduction of the pill, allowed greater participation of women in higher education (Goldin and Katz, 2002; Goldin, Katz, and Kuziemko, 2006). As argued in Goldin (2004), the fact that it is now easier to combine having a family and a career also encourages female participation. Incidentally, the largest gender gap among OECD countries is found in Scandinavian countries, the first to achieve this family-career balance (Vincent-Lancrin, 2008).

The decline in the within-family discrimination against girls may have also played a role. As explained in Buchmann and DiPrete (2006), the higher the parents' level of education and economic resources, the more likely they are to be relatively gender-neutral in their investments in education. According to this explanation, secular increases in education and income would tend to improve female relative college outcomes.[2] Moreover, exogenous changes in labor markets favoring women have led many parents to have higher educational expectations for their daughters than for their

[1]See Vincent-Lancrin (2008) for a recent summary of common explanations.

[2]According to this explanation, the college gender gap should be more pronounced in poorer, less-educated households. Goldin, Katz, and Kuziemko (2006), however, report that there are no significant differences in college completion rates by gender across the socioeconomic status distribution (at least in the United States).

sons (Chen et al., 2011), which might provide additional incentives for girls if, as argued by psychologists, girls deem pleasing adults more important than boys do (Fortin, Oreopoulos, and Phipps, 2013). Finally, decreases in family size may have also increased female college-going if, as evidence (Averett and Burton, 1996; Ono, 2004; Tansel, 2002) shows, the larger the family, the smaller the chance of girls of going to college.

In and of themselves, the previous demographic and sociological explanations can explain why women have caught up to men once barriers to female participation were removed, but are less well suited to explain the reversal of the gender gap. Explanations based on economic factors, like changes in the returns to higher education, seem more promising: if the returns of attending college increased more for women than for men, reversal could be explained (as long as women and men respond equally to the wage premium). Evidence of such a relative change in returns can be found in Dougherty (2005) and Fortin (2006) for the United States, although evidence for other countries is scarce (Vincent-Lancrin, 2008). The change is larger when considering a broader definition of returns (including the probability of getting and staying married, and of avoiding poverty), as in DiPrete and Buchmann (2006), since women graduates are less likely to be the head of a single-parent household and work part-time, and hence less likely to be poor.

Educational factors may also explain the reversal in the college gender gap. In particular, girls have been improving their academic preparation relative to boys over the past few decades, and this investment appears to be driven by economic (increase in women's returns to higher education) and demographic (increase in the age at first marriage) factors (see Goldin, Katz, and Kuziemko, 2006). Besides, women have greater aspirations than men, and aspirations have risen more rapidly for women (McDaniel, 2007; Goldin, Katz, and Kuziemko, 2006). As shown in Fortin, Oreopoulos, and Phipps (2013), gender differences in post-secondary expectations are the most important factor accounting for the disparities in academic achievement in high school.

All of the above explanations are both partial and complementary, and focus mostly on women. In this paper we propose an alternative explanation that puts the focus on men. Concern about the educational choices of men seem to be of more than just academic interest. Educators and college admissions officials in the United States have been recently voicing their concerns about the dwindling presence of men in colleges, and some are even speculating that some sort of affirmative action might be needed in the near future to deal with gender imbalances.[3]

We first develop a simple model of career choices in which an increase in the return of becoming a superstar increases the opportunity cost of pursuing college studies. Under the premise that superstar payoffs increase relatively more for men, our model predicts that men are more likely to stay longer in the superstar path to the detriment of their accumulation of human capital. Thus, we should observe that an increase in superstar earnings leads to an increase in the ratio of female to male enrollment in college.

We then provide causal evidence in support of our explanation from a natural experiment in a prominent superstar path, professional sports, in which the compensation gap grossly favors men. Nowadays, the best-paid athlete in each of the 182 countries for which the information is available is a man, and 97 of the world's 100 highest-paid athletes are men.[4] The fact that becoming an athlete superstar is more appealing for men is not new.[5] According to Forbes Magazine, in every year from 1990 to 2013 the top 10 paid athletes have been men (only one woman, tennis player Monica Seles, appeared 10th in 1992).[6]

[3] For a sample of views on the subject, check, for instance, http://usatoday30.usatoday.com/news/education/2005-10-19-male-college-cover_x.htm, http://www.collegepossible.org/, and http://www.postsecondary.org/. On admission criteria that favor men, check further Long (2007) and Bailey and Smith-Morest (2006).

[4] See http://sports.espn.go.com/espn/news/story?id=6391145 for the top paid athlete by country (accessed March 2014). The only three women that appear among the world top 100 are tennis players, the sport generally considered to be the one with the smallest gender pay gap among professional sports. Only one woman made it to the top 50: Maria Sharapova, ranked 22nd (source: http://www.forbes.com/sites/kurtbadenhausen/2013/06/05/the-worlds-highest-paid-athletes-2013-behind-the-numbers/, accessed March 2014).

[5] Another factor that can explain why the superstar path is more appealing for men is that men are more competitive than women. For robust experimental evidence on the subject, see Gneezy, Niederle, and Rustichini (2003), and Niederle and Vesterlund (2007). Buser, Niederle, and Oosterbeek (forthcoming) link gender differences in competitiveness to career choices.

[6] See http://www.topendsports.com/world/lists/earnings/forbes-index.htm (accessed March 2014).

For the last 40 years the relative salary of superstar athletes has been increasing dramatically, thus raising the opportunity cost of schooling. As an example, Figure 2 documents the evolution of weekly salaries in baseball relative to the weekly earnings of production workers in the United States in the period 1970-2011. Given that superstar athletes are mostly men, rather than women, we claim that this different increase in the opportunity cost of attending college can help explain the observed trends in the college gender gap depicted in Figure 1.

[FIGURE 2 ABOUT HERE]

We exploit changes in European soccer markets as a source of exogenous variation in men's opportunity cost of pursuing college studies. European soccer markets provide an ideal setting. Soccer is the most popular sport in the world and it is male-dominated: according to a 2006 FIFA (Fédération Internationale de Football Association, the world governing body of soccer) census of its 207 member associations, 90% of soccer players and all professional players were men. Professional players in the top European soccer leagues are among the highest paid athletes in the world, and top leagues employ an increasing fraction of players from all over the world. These features of the European soccer market can be seen, to a great extent, as consequences of the Bosman ruling, a European Court of Justice judgment in 1995 that established free agency and abolished existing limitations on the nationality of players. Whereas free agency shifted bargaining power from clubs to players, and resulted in a marked increase in wages for every professional player in a top European league, exposure to such a wage increase depended on the perceived chances of making it to those leagues, which vary strongly across countries. We exploit this cross-country variation in the exposure to the Bosman ruling to identify the impact of a change in men's opportunity cost of pursuing college studies on college enrollment.

Our main finding is that being in the group of countries that are more exposed to the Bosman ruling is associated with an increase in the ratio of female to male tertiary enrollment in the post-Bosman period in the range of 11.40 to 18.49 women for every 100 men relative to the rest of the countries in the sample – a significant increase of between 13 to 21 percent when compared to a sample mean of 86.30. This result is robust to controlling for GDP per capita growth and population, to using different samples of countries, and to exploiting alternative definitions of exposure. Consistent with our explanation of the evolution of the college gender gap, our results hold when we only consider countries in which attending college is likely to be an actual option for individuals, but not in countries in which college is less of an option.

Our paper is close to Atkin (2012) in that both of our papers highlight changes in the opportunity cost of schooling to explain trends in educational outcomes. In particular, Atkin exploits variation in the timing of factory openings across municipalities in Mexico to show that the creation of low-skill export manufacturing jobs raised the opportunity cost of schooling and lead to an increase in school dropout. In contrast, we put the focus on factors that affect the opportunity cost of schooling differentially between boys and girls. Charles and Luoh (2003) argue that greater uncertainty on returns for men would make risk-averse men study less and present data consistent with a larger increase in uncertainty for men relative to women over time; similar to our paper, the authors put the focus on men when explaining the college gender gap.

The remainder of the paper is organized as follows. In Section 2 we present a simple model that generates our main prediction about the educational attainment of men and women. We then confront this prediction with the data: we discuss our identification strategy in Section 3 and present our results in Section 4. Section 5 concludes.

2 A simple model of career choices, human capital accumulation, and lifetime earnings

2.1 The economic setup

Consider a population of individuals with two different innate skills. The first skill is a general skill that we call talent, θ, while the other skill is a specific skill that we will denote β. The two skills are ex-ante unknown to individuals, but it is common knowledge that talent is a random draw from a cumulative distribution function $G(\theta; \lambda)$ defined over $\Theta \subseteq \mathbb{R}$, with parameter λ (possibly a vector) and probability density function $g(\theta;\lambda) > 0$ for all $\theta \in \Theta$; and that the specific skill β follows a binomial distribution with parameter $p > 0$.

At each calendar time $t = 0, ..., T$, each individual has to decide between two alternative (and mutually exclusive) uses for her skills – we refer to these uses as "career paths": a "regular" path, R, and a "superstar" path, S. Each path leads to a different stream of lifetime earnings after time T. There is no discounting and there is no cost to individuals from changing from the S path to the R path, but once made, such a choice is irreversible.

Individuals are expected utility maximizers with utility function $u(\cdot)$. To simplify the exposition, we assume $u(\cdot)$ is the identity function. Thus, individuals are risk neutral and maximize expected earnings.[7] Expected lifetime earnings of individual i in the R path, w_i^R, are increasing in the individual's *effective* talent, η_i (i.e., $w_i^R = w(\eta_i)$, with $w' > 0$). For simplicity, we take $w(\cdot)$ to be linear in η_i:

$$w_i^R = w(\eta_i) = a + b\eta_i, \tag{1}$$

where $a, b > 0$ are known constants, and b measures the sensitivity of earnings to effective talent.

Effective talent is a function of the individual's innate talent and her human capital accumulated up to T. Human capital can only be accumulated by spending time in the R path. Letting τ_i denote the time at which individual i chose path R, $T - \tau_i$ measures the number of periods of human capital accumulation up to time T (i.e., total time spent on path R). Individual i's effective talent at time T is then given by $\eta_i = \theta_i x_{iT}$, where $x_{iT} = (1 + T - \tau_i)$.[8] For instance, if the individual chose R at time T, then $\tau_i = T$ (no human capital accumulation), $x_{iT} = 1$ and $\eta_i = \theta_i$ (effective and innate talent coincide).

In the other career path, S, only raw ability (i.e., innate skills θ and β), is important for expected lifetime earnings, w_i^S. In particular we assume

[7]Risk neutrality is not central to our results. All of them continue to hold with a general utility function as long as $u' > 0$, a mild condition. Proofs are available upon request.

[8]This formulation is akin to that of Gibbons and Waldman (1999), who use $\eta_i = \theta_i h(x_{iT})$, where $h' > 0$ and $h'' \leq 0$. We adopt a linear specification for h for simplicity.

that both skills are perfect complements in generating w_i^S:

$$w_i^S = \theta_i W \Longleftrightarrow \beta_i = 1, \tag{2}$$

where W is a fixed prize. In case $\beta_i = 0$, the individual is forced to exit the S path at $t = T$ once it becomes known that she does not have the specific skill β.[9] For these individuals $\tau_i = T$ and expected earnings are $w_i^R\big|_{\tau_i=T} = w(\theta_i)$.

A crucial difference between both types of skills is that learning of the realization of θ predates that of β: while θ becomes known at $t = 1$, individuals only learn β at $t = T > 1$, provided they stayed in path S until T. As an example, consider an individual with a natural talent for basketball. Such a characteristic (θ) is usually known at a young age. Whether the individual will have the necessary height to play professional basketball, however, can only be known later, say after adolescence. In the same vein, an individual can be talented for tennis, but never make it to the professional circuit because she discovers that she cannot handle pressure or stay away from home for a very long time. A somewhat different example is a young wannabe entrepreneur, who is selected by an angel investor (like the Thiel Foundation) to receive funds to develop her idea, but eventually discovers she is not the next Bill Gates or Steve Jobs, and fails in the endeavor.[10]

For simplicity, let $T = 2$ in what follows.[11] Summing up:

1. At $t = 0$, individuals decide which path, R or S, to follow, without knowledge of their skills, θ_i and β_i. Given that at this stage all

[9] Alternatively, $w_i^S = 0$ if $\beta_i = 0$, and the individual is allowed to choose which path to take after $t = T$. Given our assumptions, individual i would always choose path R under those circumstances, and earn $w_i^R = w(\theta_i)$.

[10] Since 2010, The Thiel Foundation selects 20-25 students under the age of 20 per year through a competitive process to receive a fellowship of \$100,000 over two years if they drop out of school to pursue scientific research or create a startup. Doubts about the success of the initiative are multiplying. See, e.g., http://www.forbes.com/sites/singularity/2013/09/11/peter-thiel-promised-flying-cars-instead-we-got-caffeine-spray/.

[11] Allowing for $T > 2$ is straightforward in our setup, and it would amount to letting individuals accumulate human capital for a longer period of time. Such an assumption would mechanically operate against our main result, and hence $T = 2$ can be seen as the most conservative assumption.

individuals have the same information and beliefs, they all make the same decision. We focus on the most interesting case in which all individuals choose the superstar path S at this time.

2. At $t = 1$, individuals learn their innate talent θ, but they are still uncertain about their specific skill β. Each individual then decides whether to exit or not from the S path. The decision to exit is irreversible. Individuals choosing path R accumulate human capital in this period.

3. At $t = 2$, individuals who decided to stay in the S path in the previous period learn their specific skill β. With probability p, an individual observes $\beta_i = 1$, and with probability $1 - p$, she observes $\beta_i = 0$ and is forced to exit to the regular path R. Payoffs accrue according to (1) and (2).

2.2 Equilibrium analysis and predictions

At $t = 2$, expected lifetime earnings of individuals who chose path R at $t = 1$ are given by:

$$w_i^R\big|_{\tau_i=1} = a + b\,\eta_i|_{\tau_i=1} = a + 2b\theta_i.$$

Among individuals that did not exit path S at $t = 1$, some learn that $\beta_i = 1$ at $t = 2$ and earn:

$$w_i^S = \theta_i W,$$

but those who learn that their $\beta_i = 0$ have expected lifetime earnings given by:

$$w_i^R\big|_{\tau_i=2} = a + b\,\eta_i|_{\tau_i=2} = a + b\theta_i.$$

At $t = 1$, an individual will decide to exit the S path if expected lifetime earnings are greater in the regular path R; i.e., individual i chooses path R if and only if:

$$a + 2b\theta_i > pW\theta_i + (1 - p)(a + b\theta_i),$$

or, put differently, if her talent (known at this time) is low enough:

$$\theta_i < \theta^* \equiv \frac{pa}{pW - (1+p)b}. \tag{3}$$

Provided that $pW > (1+p)b$, there exists an individual with $\theta_i = \theta^*$ who is just indifferent between career paths after learning her θ_i.[12] For all individuals with $\theta_i < \theta^*$, we have $\tau_i = 1$. Hence, the proportion of individuals in the population that exit the S path at $t = 1$ and start investing in their human capital is given by $G(\theta^*; \lambda)$. Intuitively, less gifted individuals compensate their lower *innate* talent with human capital accumulation, so that their *effective* talent ends up being larger than for more gifted individuals (i.e., individuals born with a higher θ), but who decided to stay longer in the S path.[13] Notice that $G(\theta^*; \lambda)$ is decreasing in W since $\frac{\partial \theta^*}{\partial W} < 0$ and $g(\theta; \lambda) > 0$ for all θ.[14]

Suppose there are two groups (M, F) of individuals in the population characterized as before, and that only differ in the W they face. Given the expected earnings of each group in the superstar path it is straightforward to compute the fraction of individuals in each group with higher capital accumulation, $G\left(\theta^*_M\left(W^M\right); \lambda\right)$ and $G(\theta^*_F\left(W^F\right); \lambda)$, where we have made explicit the dependence of θ^* on the W faced by each group. We interpret λ as capturing other determinants of human capital investments that affect both groups of individuals.

Suppose now that there is a shock σ that increases W^M but (crucially) not W^F, and which may also affect λ (in any direction). Such a shock has a direct negative effect on the human capital accumulation of group M (because $\frac{\partial G(\theta;\lambda)}{\partial \theta} > 0$ and $\frac{\partial \theta^*}{\partial W} < 0$ from equation (3)), but its total effect is uncertain because of the confounding effect of other determinants of human capital decisions that might be affected by the shock (i.e., when $\frac{\partial \lambda}{\partial \sigma} \neq 0$).

[12] In general, the assumption needed is that expected utility be more responsive to innate talent in the superstar path than in the regular path. See Gibbons and Waldman (1999) for similar assumptions.

[13] By the Law of Large Numbers, a fraction $1-p$ of individuals who exit at $t = 1$ also have $\beta = 0$. These individuals are strictly of lower talent than individuals like who did not exit, even when the specific skill is considered in the definition of "more gifted". In any case, since the specific skill is useless in R, the definition of "gifted" just in terms of θ is also warranted.

[14] This result does not hinge on the linearity assumption, but just on having W large enough. The proof is available from the authors upon request.

However, if $G(\cdot)$ is multiplicatively separable in θ and λ, we can obtain an unambiguous prediction for the effect of the shock to W^M on the human capital accumulation ratio of group F to group M.[15] Letting Ω denote this ratio and assuming $G(\theta;\lambda) = \lambda H(\theta)$, $H' > 0$ for all $\theta \in \Theta$, we have:[16]

$$\begin{aligned}
\Omega &= \frac{\lambda(\sigma) H\left(\theta_F^*\left(W^F\right)\right)}{\lambda(\sigma) H\left(\theta_M^*\left(W^M(\sigma)\right)\right)} \\
&= \frac{H\left(\theta_F^*\left(W^F\right)\right)}{H\left(\theta_M^*\left(W^M(\sigma)\right)\right)}. \qquad (4)
\end{aligned}$$

A simple comparative-statics exercise with respect to σ on Ω yields our main result:

Proposition 1 *Suppose there are two groups (M, F) of individuals in the population that differ only in the W they face. Then, a positive shock σ to the prize that group M faces in the superstar path, W^M, will increase the ratio of human capital accumulation Ω; i.e., $\frac{d\Omega}{d\sigma} > 0$.*

Proof. Totally differentiating equation (4) with respect to σ yields

$$\frac{d\Omega}{d\sigma} = \frac{\frac{\partial H\left(\theta_F^*\left(W^F\right)\right)}{\partial\sigma} H\left(\theta_M^*\left(W^M(\sigma)\right)\right) - \frac{\partial H\left(\theta_M^*\left(W^M(\sigma)\right)\right)}{\partial\sigma} H\left(\theta_F^*\left(W^F\right)\right)}{H\left(\theta_M^*\left(W^M(\sigma)\right)\right)^2}.$$

Since σ does not affect the prize faced by individuals in group F, $H\left(\theta_F^*\left(W^F\right)\right)$ remains unchanged: $\frac{\partial H\left(\theta_F^*\left(W^F\right)\right)}{\partial\sigma} = 0$. Therefore,

$$sgn\left(\frac{d\Omega}{d\sigma}\right) = -sgn\left(\frac{\partial H\left(\theta_M^*\left(W^M(\sigma)\right)\right)}{\partial\sigma}\right).$$

Given that $\frac{\partial H\left(\theta_M^*\left(W^M(\sigma)\right)\right)}{\partial\sigma} = \frac{\partial H\left(\theta_M^*\left(W^M(\sigma)\right)\right)}{\partial\theta}\frac{\partial\theta_M^*\left(W^M(\sigma)\right)}{\partial W^M}\frac{\partial W^M(\sigma)}{\partial\sigma}$, $\frac{\partial W^M(\sigma)}{\partial\sigma} > 0$ by definition, $\frac{\partial\theta_M^*\left(W^M(\sigma)\right)}{\partial W^M} < 0$ from equation (3), and $\frac{\partial H\left(\theta_M^*\left(W^M(\sigma)\right)\right)}{\partial\theta} > 0$, we can conclude that $\frac{\partial H\left(\theta_M^*\left(W^M(\sigma)\right)\right)}{\partial\sigma} < 0$, and hence $\frac{d\Omega}{d\sigma} > 0$. ■

[15] If $\frac{\partial\lambda}{\partial\sigma}$ turns out to be zero, assuming separability would imply no loss of generality since it would yield exactly the same comparative statics. However, if $\frac{\partial\lambda}{\partial\sigma} \neq 0$, separability allows for a clear prediction where not assuming it would not.

[16] Given that G and H are cumulative distribution functions, we are implicitly assuming that λ is such that for $G, H \in [0,1]$ for every θ.

Proposition 1 provides a potential explanation for the puzzle about why women are catching up to men, and have in many countries overtaken them, in terms of college enrollment and completion. Under the premise that payoffs in the superstar path (e.g., professional sports) increase for men, but not for women, our model predicts that men are more likely to stay longer in the superstar path to the detriment of their accumulation of human capital (say, by engaging sports rather than enrolling in college). Thus, we should observe that an increase in W^M leads to an increase in the ratio of female to male enrollment in college – a claim we will support with causal evidence in Section 4.

3 Natural experiment and data

As discussed in Section 2 the theoretical model has the implication that a positive shock to the prize for men in the superstar path (W^M) should lead to an increase in the ratio of women to men accumulation of human capital (Ω). To test our Proposition 1, we will treat the ratio of female to male enrollment in tertiary education as the empirical analog of Ω. In this section, we describe the setup for our empirical exercise and discuss how we can exploit changes in European soccer markets as a source of exogenous variation in the expected earnings for men associated with the superstar path.

3.1 The natural experiment

Soccer is indisputably the most popular sport in the world and it is played mostly by men.[17] According to the 2006 Big Count, a FIFA survey of its 207 member associations, 265 million players (professional, registered, and occasional) were actively involved in this sport. Of this grand total, 90% of players were male, and only men were involved in professional play.

[17] Many rankings of popular sports can be devised, each ranking using different criteria (viewership, players, revenue) and sources (facts, personal opinion, online votes). Soccer (association football) appears on top in every credible ranking that can be found on the Internet (see, e.g., http://www.topendsports.com/world/lists/popular-sport/analysis.htm). The FIFA World Cup's final game is the single most viewed sporting event: the 2010 match between Spain and The Netherlands was watched by an estimated 700 million people.

Professional soccer players are among the highest paid athletes of any sport in the world, and among the top 10 sports teams by average salary per player, seven are soccer clubs from Europe's Big Five leagues.[18] The Big Five leagues are the German Bundesliga, Italian Serie A, Spanish Liga, English Premier League, and French Ligue 1 – these are the leagues that generate the largest revenues and pay the highest salaries.[19] The highest paying team in the world is FC Barcelona (of the Spanish league) and has an average salary of $8.7 million a year, with each player earning about $167,000 a week.

The high salaries obtained in the Big Five leagues are hard to match outside of Europe: the average player for the Sporting Kansas City, the 2012-13 champion of America's Major League Soccer, made $106,836 a year as base salary in 2013, whereas minimum wages for professional players in South America are below $3,600 a year and strikes over unpaid wages are not uncommon.[20] While in 2012 top players Lionel Messi (FC Barcelona) and Cristiano Ronaldo (Real Madrid) made over $20 million just in salary, Kansas City's top paid player, Graham Zusi, made $350,000 in 2013, an amount the average player of FC Barcelona can earn in just two weeks.[21]

Soccer teams in Big Five leagues recruit talent from all over the world. At the start of the 2013-14 season, 51 percent of players in a given Big Five league were foreigners.[22] Comparing the figures for Big Five players to the FIFA Big Count suggests that an average of only one in every 100,000

[18] Along with the seven soccer clubs, three American teams – the Yankees, Los Angeles Lakers, and Philadelphia Phillies – are represented in the top 10. No NFL team in the league started 2011 with a payroll higher than 75th (Pittsburgh Steelers, who payed an average of $2.9 million per player). See http://espn.go.com/espn/story/_/id/7850531/espn-magazine-sportingintelligence-global-salary-survey-espn-magazine for the full story.

[19] See Deloitte's *Annual Review of Football Finance* at http://www.deloitte.com/view/en_GB/uk/industries/sportsbusinessgroup/sports/football/annual-review-of-football-finance/.

[20] Figures for the US are from the Major League Soccer Players Union and can be checked at http://www.mlsplayers.org/salary_info.html. Minimum wages for South American players were obtained from http://www.elsalario.com.ar/main/trabajo-decente/Informeslaborales/argentina-tiene-el-salario-minimo-mas-alto-de and http://www.agremiados.com.ar/faa/notas/2010/04/09/61700.html.

[21] See http://www.forbes.com/sites/christinasettimi/2013/04/17/the-worlds-best-paid-soccer-players/.

[22] The figures are from the website of Transfermarkt GmbH & Co. KG (http://www.transfermarkt.com/).

soccer players in the world eventually makes it to a professional Big Five league. There is, however, huge variation in this figure, with dozens of countries with zero players in Big Five leagues, and countries like Iceland and Uruguay in which over 10 in 100,000 players play in those leagues.

The landscape of European soccer salaries has not always been this bright: when soccer star Diego Maradona was transferred to FC Barcelona in 1982, his salary was a meagre $50,000 a year, and in the early 1990s the average player in England's top division made $138,000 to $147,000 a year.[23] Nor is it the case that foreigners have always been so present in European leagues: when Manchester United won England's Premier League in 2012-13, roughly 50% of its team roster and typical lineup were foreigners, whereas in 1995-96 the club had achieved the same feat with just five foreigners out of a squad of 31 players (16%), and no more than two in their typical lineup.

To a great extent, the changes in salaries and team composition can be traced back to an European Court of Justice (ECJ) judgment in December 1995 that established free agency and abolished existing limitations on the nationality of players (see, e.g., Downey, 2001; Frick, 2007; and Poli, 2006, 2010). Before the ECJ judgment, most players in a given club were national, thanks to a UEFA rule stating that no team could field more than three non-nationals at the same time.[24] When a player reached the end of his contract with a club, he did not become a free agent: he could ask for a transfer to another institution, but such a transfer would only proceed provided the new club could agree on a transfer fee with the old club; else, the player would be forced to re-sign with his old club or not play for the whole season.[25]

[23] In constant US dollars, Maradona's annual salary is less than 3/4 of what the average FC Barcelona's player earns per week. For English salaries, see http://www.sportingintelligence.com/2011/01/20/from-20-to-33868-per-week-a-quick-history-of-english-footballs-top-flight-wages-200101/.

[24] UEFA (Union des Associations Européennes de Football) is the governing body of European Soccer. The rule was colloquially referred to as the "3+2" rule, because it permitted each national association to limit to three the number of foreign players whom a club may field in any first division match in their national championships, plus two players who had played in the country of the relevant national association for an uninterrupted period of five years, including three years as a junior (http://eur-lex.europa.eu/smartapi/cgi/sga_doc?smartapi!celexplus!prod!CELEXnumdoc&lg=en&numdoc=61993J0415).

[25] Such a rule was similar to the reserve clause once common in North American

These two basic principles (limits on non-nationals and transfer fees for out-of-contract players) were challenged in court by a Belgian player, Jean-Marc Bosman, in 1990. When his contract with Belgian club RFC Liège expired, Bosman arranged for a transfer with Dunkerque, a French team. As both clubs could not settle on a transfer fee, RFC Liège refused to let Bosman go. The player refused to re-sign with his old club, was suspended and took his case to court. In the final judgment on the case (which became known as the Bosman ruling), the ECJ declared in December 1995 that both UEFA regulations (the transfer system and the nationality rule) were incompatible with article 48 (now article 39) of the Treaty establishing the European Union (EU), which guaranteed freedom of movement for workers.[26]

The resulting liberalization of the European transfer market had easily noticeable consequences for players' salaries and mobility. The abolition of the transfer fee for players out of contract meant a player would become a free agent at the expiration of his contract, and could now offer his services to the highest bidder. The free-agency regime shifted bargaining power from clubs to players, and resulted in a drastic increase in wages. Figure 3 shows the evolution of the gross salary expenditure in each of the Big Five leagues from 1995/96 (the last season played under the old rules) to 2001/02. Salaries had doubled by 1998/99 and more than tripled by 2001/02.[27]

professional sports.

[26] Article 39 (formerly 48) states: "1. Freedom of movement for workers shall be secured within the Community. 2. Such freedom of movement shall entail the abolition of any discrimination based on nationality between workers of the Member States as regards employment, remuneration and other conditions of work and employment. 3. It shall entail the right, subject to limitations justified on grounds of public policy, public security or public health: (a) to accept offers of employment actually made; (b) to move freely within the territory of Member States for this purpose; (c) to stay in a Member State for the purpose of employment in accordance with the provisions governing the employment of nationals of that State laid down by law, regulation or administrative action; (d) to remain in the territory of a Member State after having been employed in that State, subject to conditions which shall be embodied in implementing regulations to be drawn up by the Commission. 4. The provisions of this article shall not apply to employment in the public service" (http://eur-lex.europa.eu/LexUriServ/LexUriServ.do?uri=CELEX:12002E039:EN:HTML).
The full judgment can be found at http://eur-lex.europa.eu/ smartapi/cgi/sga_doc?smartapi!celexplus!prod!CELEXnumdoc&lg=en&numdoc=61993J0415.

[27] The impact of free agency on salaries comes as no surprise when comparing with

[FIGURE 3 ABOUT HERE]

By abolishing nationality requirements in the composition of teams at club level, the Bosman ruling also implied that every soccer club in the EU could now field as many non-nationals as they saw fit, provided the players held a EU passport, thereby opening up a new international transfer market. To begin with, every non-national slot occupied by a EU citizen by the time of the ruling immediately became a vacant slot for the 1996/97 season. But the ruling also markedly increased the value of holding a European passport, or obtaining one through European descent (as in the case, e.g., of Argentine players of Italian or Spanish descent) or residence (for instance, Belgium is usually considered to place the less stringent requirements to grant nationality to foreign residents).[28] As Figure 4 shows, the Bosman ruling had a deep impact on the number of foreign players in Big Five leagues, as Europe's search for soccer talent became a global phenomenon. While the number of foreign players in Big Five leagues had been around 400 in previous seasons close to the Bosman ruling (and growing at an average annual rate of less than 7%), that figure jumped to 635 in the 1996/97 season, the first under the new rules – a 35% increase. The number of foreign players had doubled by 1998/1999 and tripled by 2007/08.

[FIGURE 4 ABOUT HERE]

3.2 Data

The empirical analog of Ω in our test of Proposition 1 is the ratio of total female to male enrollment in tertiary education in public and private schools

previous experiences with free agency in US professional sports, beginning with baseball in 1976. For a general overview, see Kahn (2000). Player mobility between clubs also increased dramatically in the US with free agency; see, e.g., Goldberg (2008).

[28] Fraud can also be a source of EU passports: the fake passport scandal that began in the Italian Serie A in 2000, when as many as nine South American players were implicated in suspicions over the provision of false Italian and Portuguese passports to enable their clubs to field them as Europeans, illustrates the importance to non-EU players of obtaining a EU passport. Shortly after, other three South American players were found guilty of using fake European passports and banned from entering France. At an informal meeting in 2001, officials from France, Spain, Italy, Portugal, and the UK discussed the role of criminal organizations in the supply of the fake documents. For the full stories, see http://news.bbc.co.uk/sport2/hi/football/europe/962023.stm, http://sportsillustrated.cnn.com/soccer/news/2001/03/23/fake_passports_ap/, and http://news.bbc.co.uk/ sport2/hi/football/europe/1260498.stm.

(also known as the Gender Parity Index in tertiary level enrollment in the United Nations' Millennium Development Goals Indicators).[29] The World Bank, through its World Development Indicators collection, provides consistent yearly data on this indicator for 214 countries from 1970 to 2011.[30] We eliminated 7 countries for implausible data, and also every country for which the database does not provide at least one observation pre-Bosman and one observation post-Bosman, leaving us with 147 countries to work with.[31] Table 1 contains summary statistics for our sample.

[TABLE 1 ABOUT HERE]

Whereas the Bosman ruling implied an increase in salaries in the superstar path only for men from any country, conditional on playing for a team in a Big Five league, exposure to such change depended on the perceived chances of eventually making it to those leagues – chances that vary strongly across countries, due to the quality of their players and their ease of access to a European passport.[32]

To assess these chances, we begin by computing for every country in our sample the number of soccer players from each country in Big Five leagues in the 1994/95 season (the last full season prior to the Bosman ruling).[33] To classify a player in a Big Five league as a national or a non-national, we have used their first nationality (usually the country where the player was

[29] Tertiary education includes categories 5 and 6 of the 1997 International Standard Classification of Education (ISCED), a statistical framework for organizing information on education maintained by the United Nations Educational, Scientific and Cultural Organization (UNESCO), and refers basically to undegraduate and graduate studies leading to a degree. See http://www.uis.unesco.org/Education/Pages/international-standard-classification-of-education.aspx for details.

[30] The data are accesible at http://databank.worldbank.org/data/views/variable Selection/selectvariables.aspx?source=world-development-indicators. The series used is *Ratio of female to male tertiary enrollment*. The series was downloaded on 02/06/2014.

[31] The 7 countries are Afghanistan, Djibouti, Grenada, Guyana, Madagascar, Qatar, and St. Kitts and Nevis. As an example of implausible data, St. Kitts and Nevis reported a ratio of female to male enrollment of 267 in 1985, and then just 73 in 1986, while Djibouti had a ratio of only 44 in 1992 and jumped to 128 in 1993. All of our results are robust to the inclusion of these countries.

[32] Prior to the ECJ judgment, players with a German, Italian, Spanish, UK, or French passport could play for a team in the corresponding Big Five league without occupying a non-national slot. After the judgment, they could play for *any* team in a Big Five league as a national.

[33] The total number of players in the 1994/95 season in Big Five leagues by country was obtained from the website of Transfermarkt GmbH & Co. KG (http://www.transfermarkt.com/; accessed on 04/14/2014).

developed) rather than the nationality under which they are playing.[34] In the 1994/95 season, Brazil (with 30 players) and the Netherlands (25) come out on top when ranking countries by the absolute number of players in Big Five leagues. As it is not the same, in terms of career choices and perceived chances of success in the superstar path, to see 100 fellow countrymen succeeding in Big Five leagues in a country of 1,000,000 inhabitants than in another country with a population of 100,000,000, in the empirical exercise we divide the number of players by each country's population in 1994.[35] By this normalization, Iceland and Ireland become the top two countries.

4 Econometric methods and results

The previous section shows that the Bosman ruling increased expected earnings for young males in a group of countries by increasing both the earnings associated with the superstar path and the probability of achieving success on that path. In this section we exploit the Bosman ruling as a source of exogenous variation in the expected earnings of men in the superstar path. As discussed in Section 2 the theoretical model has the implication that an increase in the prize for men in the superstar path (W^M) should lead to an increase in the ratio of female to male accumulation of human capital (Ω). We test this prediction by estimating the effect of the Bosman ruling on the ratio of female to male tertiary enrollment. Formally, we estimate the following regression model:

$$\begin{aligned} \Omega_{it} &= \beta W_{it} + \gamma X_{it} + \alpha_i + \mu_t + \varepsilon_{it} \\ &= \beta \left(Bosman_t * Average\ players_i \right) + \gamma X_{it} + \alpha_i + \mu_t + \varepsilon_{it}, \end{aligned} \tag{5}$$

[34] For instance, Sergio Agüero was born in Argentina to Argentine parents, developed as soccer player in Argentina, plays for Argentina's national team, and holds both Argentine and Spanish passports. Thanks to his Spanish passport he did not occupy a non-EU slot at his former team, Atlético de Madrid, but is nevertheless counted as a foreign player in the Spanish Liga for our purposes. Players from Northern Ireland, Wales and Scotland have been considered nationals to the English Premier League, as they are all citizens of the United Kingdom.

[35] Population data were obtained from the World Bank's World Development Indicators (series name: *Population (Total)*, downloaded on 02/06/2014).

where Ω_{it} is the ratio of female to male tertiary enrollment in country i at time t, X_{it} is a set of controls, α_i is a time-invariant country effect, μ_t is a time-period effect common to all countries, and ε_{it} is the usual error term. W_{it} has variability both across time and across countries. The time variability is captured by a dummy variable that takes the value of one after 1995 ($Bosman_t$), and the cross-country variation is captured by $Average$ $players_i$, defined as the ratio of the number of country i's players in Big Five soccer leagues to the country's population, in 1994. The parameter of interest in equation (5) is β, which captures the interaction effect between the time variability and the country variability in the prize faced by men in the superstar path. In light of Proposition 1 we expect $\beta > 0$.

The use of the ratio of female to male enrollment as our dependent variable is central for our identification strategy, since it guarantees that all confounders that vary across countries and time, and that affect in a similar way female and male enrollment (the λs in Section 2) are not biasing our estimates of β.

The main empirical result of the paper is anticipated in Figure 5. The figure compares the evolution of the average ratio of female to male tertiary enrollment for countries that are more exposed to the Bosman ruling relative to the rest of the countries in the subsample of countries with ratios above the sample mean in 1994.[36] The levels and the trends of the ratio in the two groups of countries are remarkably similar in the pre-Bosman period and only start to diverge after the Bosman ruling.[37]

[FIGURE 5 ABOUT HERE]

More formally, the difference-in-differences model assumes that the change in the ratio of female to male enrollment in those countries that are less exposed to the Bosman ruling is an unbiased estimate of the counterfactual. While we cannot directly test this assumption, we can test whether time trends in the two groups of countries were the same in the pre-Bosman period. If time trends are the same in the pre-Bosman period, then it is

[36] The exposed group consists of the top 10 countries according to the ratio of the number of soccer players playing in Big Five leagues to the country's population (in 1994).

[37] Given that our data panel is unbalanced, the figure should be interpreted with caution, as composition effects might be at work. However, our formal results below show that the simple message the figure sends holds when appropriately dealing with those effects.

likely that they would have been the same in the post-Bosman period in the absence of the Bosman ruling. As in Galiani, Gertler, and Schargrodsky (2005), to test the hypothesis that the pre-Bosman time trends are not different in the two groups, we estimate a model like the one in (5) for the full sample of countries, but we exclude the interaction term and include separate year dummies for the exposed countries. We use only observations in the pre-Bosman period; that is, we use data for the years 1970 to 1995 for all the countries in the sample. When we run this regression (not shown, available upon request), 24 out of the 25 dummy variables capturing the interaction between the year effects and the dummy for the exposed countries are not significant, thus validating our difference-in-differences identification strategy. We have also tried with a linear trend and an interaction term between the linear trend and the exposure dummy and, again, the interaction term is not significant.[38]

Table 2 reports Ordinary Least Squares (OLS) estimates of equation (5). The coefficient associated to the interaction variable between *Bosman* and *Average players* is positive and statistically significant in our baseline regression in column (1). Indeed, the coefficient is not only statistically significant but also quantitatively substantial. An increase of one standard deviation in the number of players in Big Five soccer leagues (normalized by population) is associated with an increase of 3.48 women for every 100 men enrolled in college, or 4 percent of the sample mean. In column (3), we control for the set of covariates available (GDP per capita growth and population).[39] Again, the value of the coefficient of interest is positive and statistically significant.[40]

[TABLE 2 ABOUT HERE]

Players from the countries that host the Big Five leagues may be more exposed to the Bosman ruling than what their country's ratio of players to

[38] These results also hold when we consider the top 15 countries, and if we include controls in the regressions.

[39] Data for GDP per capita growth was obtained from the World Bank's World Development Indicators collection (series name: *GDP per capita growth (annual %)*, downloaded on 02/06/2014).

[40] The increase in the coefficient from column (1) to column (3) is due to sample selection: when we run the specifications in columns (1) and (2) on the same sample of columns (3) and (4), the values of the coefficients are similar to the ones with controls, suggesting that much of the drop in the value of the coefficient is due to differences in the samples.

population would suggest, because the Bosman ruling implied a jump in soccer salaries in their domestic leagues that has no parallel in players from other countries. As a robustness check, we have thus run our regressions excluding Germany, Italy, Spain, United Kingdom, and France from the sample. As reported in columns (2) and (4), the effect of interest is still positive and statistically significant when we exclude the countries hosting the Big Five leagues from the sample.

Assuming that all the change in the gender parity index comes from a reduction in male enrollment, back-of-the-envelope calculations suggest it only requires that 3-4 out of 10 potential college enrollees (from the under-18 subpopulation actively involved in soccer playing) decide not to attend college to generate the estimated effects in Table 2 (evaluated at the sample mean of the ratio of female to male enrollment).[41]

According to our measure of exposure (players in Big Five leagues over population, in 1994), the countries most affected by the Bosman ruling are all from Africa, Latin America and the Caribbean, and Europe.[42] As we show in columns (5) to (8) of Table 2, results are robust to limiting the sample to countries in these three regions. Restricting the sample in this way also makes control countries more comparable to exposed countries. All coefficients in columns (5) to (8) are remarkably similar to those in columns (1) to (4), implying an increase of between 3.54 and 4.70 females for every 100 males in college education after a one-standard-deviation increase in *Average players*.

To aid in the interpretation of the effect of an increase in W^M on the ratio of female to male enrollment, we have generated a discrete measure of exposure to the Bosman ruling, *Top countries*, an indicator variable that takes the value of one for the top 15 countries according to the ratio of the number of soccer players playing in Big Five leagues to the country's population (in 1994). As reported in columns (1) to (4) of Table 3 the coefficient on the interaction effect between *Bosman* and *Top countries* is

[41] We have estimated the number of males attending college from the World Development Indicators series (*School enrollment, tertiary, male*, downloaded on 02/06/2014) using the fraction of population aged 20-24 in OECD countries (http://stats.oecd.org/Index.aspx?DataSetCode=RPOP#; accessed on 04/11/2014). The figures for players under 18 come from the 2006 FIFA Big Count.

[42] The first country outside these regions to appear is New Zealand, ranked 22nd.

positive and statistically significant in the models with and without controls, whether or not the hosts of the Big Five leagues are included. Being in the group of countries that are more exposed to the Bosman ruling is associated with an increase in the ratio of female to male tertiary enrollment in the range of 11.40 to 13.27 women for every 100 men relative to the rest of the countries in the sample – 13 to 15 percent of the sample mean.

[TABLE 3 ABOUT HERE]

Even though the definition of *Top countries* might seem somewhat arbitrary, columns (5) to (8) of Table 3 show that our results are robust to an alternative definition that considers the top 10 countries according to *Average players*. The effect of being exposed to the Bosman ruling is an increase of 17.34 to 18.49 women for every 100 men in college enrollment relative to countries not exposed, a substantial increase of about 21% of the sample mean.

The model developed in Section 2 implicitly assumes that individuals face an actual choice between going to college or pursuing a superstar career. In practice, such a choice is more likely to be present in more developed, higher-income countries – where attending college is a feasible option for large portions of the population – than in less developed, lower-income countries in which college is not a real option. Hence, we would expect our results to hold in a subsample of the first group of countries, but not if we restrict our sample to the second group of countries. Table 4 shows the results of such a false experiment: when we restrict the sample to European (i.e., more developed) countries in column (1), the coefficient on the interaction between *Bosman* and *Average players* is positive, significant, and remarkably similar to those in Table 2, but the coefficient is statistically nonsignificant when we look only at African (i.e., less developed) countries in column (5).[43] Our results also hold if we split the countries by their 1994 income according the World Bank's classification.[44] Column (2) in Table 4 shows a positive and significant effect in middle- and high-income countries,

[43] The results for Europe are robust to excluding the countries hosting the Big Five leagues.

[44] See http://data.worldbank.org/about/country-and-lending-groups (accessed on 06/27/2014). The income series used is *GNI per capita, Atlas method (current US$)*, downloaded on 06/27/2014 from http:// databank.worldbank.org/data/views/variableSelection/selectvariables.aspx?source=world-development-indicators#.

whereas no effect shows up in column (6) for low-income countries.[45]

[TABLE 4 ABOUT HERE]

We can also proxy for the likelihood of having an actual option of attending college directly through our enrollment data. In columns (3) and (7), we have split countries according to whether male enrollment in 1994 was above or below the sample mean for that year; whereas in columns (4) and (8) we have restricted attention to countries in which the ratio of female to male tertiary enrollment in 1994 was above and below the mean ratio for the full sample in that year.[46] As expected, a positive and statistically significant effect only obtains in the group of countries where a tertiary education is more likely to be an option.[47]

5 Concluding remarks

In this paper we have provided an alternative explanation for the evolution of the gender gap in college attendance that puts the focus on the raising opportunity cost of pursuing a college degree for men – a raise due to the increase in the rewards to becoming a superstar in occupations typically dominated by men, like professional sports. We have then supported our explanation with causal evidence from a natural experiment in which we have exploited changes in European soccer markets as a source of exogenous variation in the expected earnings for men associated with the superstar path to identify the effect of interest. Consistent with our story, we have found a significant positive effect of an increase in male earnings in a superstar path on the ratio of female to male tertiary enrollment in college education.

If a college gap favoring females is a cause for concern, it is only natural

[45] Countries were classified as low-income countries in 1994 by the World Bank if their GNI per capita fell below $725; as middle-income countries if GNI per capita was between $725 and $8,955; and as high-income countries if GNI per capita was above $8,955.

[46] The series used in columns (3) and (7) is *School enrollment, tertiary, male*. The data are accesible at http://databank.worldbank.org/data/views/variableSelection/selectvariables.aspx?source=world-development-indicators. The series was downloaded on 02/06/2014.

[47] To avoid cluttering, we have not reported the results of the regressions without controls in Table 4, but they are all similar to the results displayed.

to guess that institutions may appear to deal with the causes. Athletic scholarships in US universities might be a case in point. By reducing the cost of failing to achieve superstardom in the sport of choice, such an institution could attenuate the effect highlighted in this paper. But as long as boys continue to neglect their studies in favor of sports to try to obtain a scholarship, instead of a direct success in the superstar path, the mechanism we have discussed could still be at work at a lower educational level, such as high school (which should eventually have an effect on the college gender gap). Furthermore, is it conceivable that the mere existence of the athletic scholarships strengthens individual incentives to enter and remain in the superstar path, which could exacerbate the problem.[48]

Even though our natural experiment looked at one particular professional sport, soccer, we expect the mechanics outlined in this paper to be at work in other markets where superstars are present, and in which the gender gap has been increasing in favor of men. For instance, changes in the return to becoming an entrepreneur might also contribute to the college gender gap, given that men are much more likely to become entrepreneurs than women.[49] An empirical investigation into these ideas must, however, await future research.

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[48]Besides, it is unclear how it would be possible to target scholarships to boys without violating Title IX, a law passed in United States in 1972 that requires equal funding for male and female athletes, including scholarship distribution.

[49]See http://www.oecd.org/publications/factbook/oecdfactbook2011-2012.htm and http://www.bls.gov/opub/mlr/2010/09/art2full.pdf (accessed April 2014).

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Table 1. Summary statistics

Variable	Mean	Standard deviation	Minimum	Maximum
Female to male tertiary enrollment	86.30	47.44	0	530.86
Players in Big Five soccer leagues in 1994	2.48	4.87	0	25.09
Average players	0.39	1.05	0.00	7.52
Population (millions)	33.42	121.91	0.05	1344.13
GDP per capita growth	1.93	5.88	-50.24	91.67

Notes: *Female to male tertiary enrollment*, *Population (millions)*, and *GDP per capita growth* correspond to yearly data from the World Development Indicators for the period 1970 to 2011. *Players in Big Five soccer leagues in 1994* is the number of players by country in 1994. *Average players* is the ratio of *Players in Big Five soccer leagues in 1994* to *Population* in 1994. The Big Five leagues are the German Bundesliga, Italian Serie A, Spanish Liga, English Premier League, and French Ligue 1.

Table 2. Main results

	Dependent variable: Female to male tertiary enrollment							
	All countries				Latin America and the Caribbean, Europe, and Africa			
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
Bosman x Average players	3.314**	3.431**	4.417***	4.571***	3.374**	3.557**	4.219***	4.474***
	(1.390)	(1.395)	(1.384)	(1.382)	(1.458)	(1.475)	(1.563)	(1.576)
Controls	No	No	Yes	Yes	No	No	Yes	Yes
Countries hosting the Big Five leagues	Yes	No	Yes	No	Yes	No	Yes	No
Observations	3,518	3,358	3,227	3,067	2,499	2,339	2,296	2,136

Notes: All models are estimated by OLS and include year fixed effects and country fixed effects. Standard errors clustered at the country level are in parentheses. The controls are the countries' population and GDP per capita growth. Models (5) to (8) only include countries from Latin America and the Caribbean, Europe, and Africa. The countries hosting the Big Five leagues are Germany, Italy, Spain, United Kingdom, and France. The Big Five leagues are the German Bundesliga, Italian Serie A, Spanish Liga, English Premier League, and French Ligue 1. **Significant at 5 percent level. ***Significant at 1 percent level.

Table 3. Alternative measures of exposure to the Bosman ruling

	Dependent variable: Female to male tertiary enrollment							
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
Bosman x Top countries	11.395*	11.807*	12.814*	13.267*	17.336**	17.748**	17.994**	18.494**
	(6.963)	(7.000)	(6.800)	(6.866)	(7.232)	(7.264)	(8.725)	(8.759)
Controls	No	No	Yes	Yes	No	No	Yes	Yes
Countries hosting the Big Five leagues	Yes	No	Yes	No	Yes	No	Yes	No
Observations	3,518	3,358	3,227	3,067	3,518	3,358	3,227	3,067

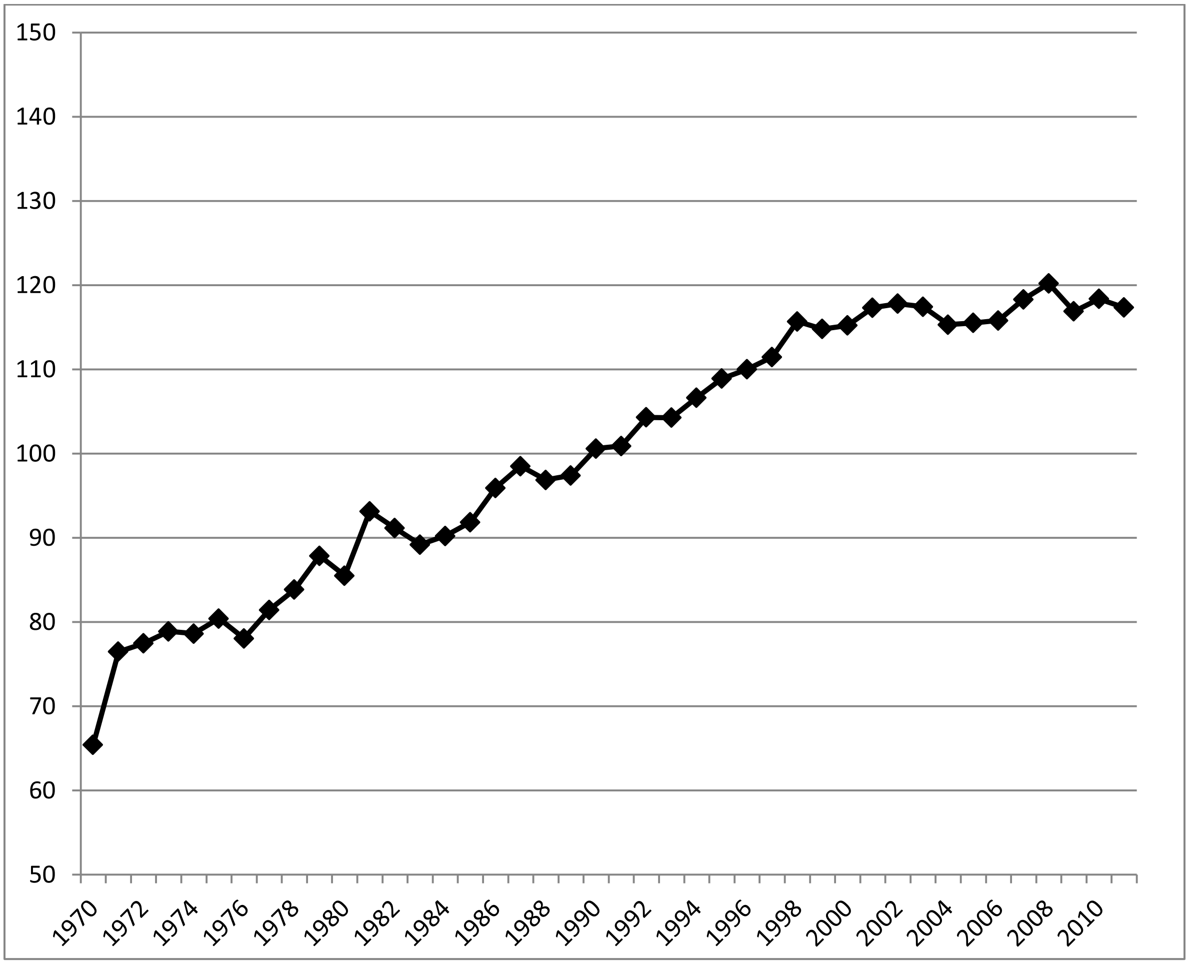
Notes: All models are estimated by OLS and include year fixed effects and country fixed effects. Standard errors clustered at the country level are in parentheses. The controls are the countries' population and GDP per capita growth. In models (1) to (4) the coefficient of interest is the interaction between Bosman and an indicator variable that takes the value of one for the top 15 countries according to the ratio of the number of soccer players playing in Big Five leagues to the country's population (in 1994). In models (5) to (8) the coefficient of interest is the interaction between Bosman and an indicator variable that takes the value of one for the top 10 countries according to the ratio of the number of soccer players playing in Big Five leagues to the country's population (in 1994). The countries hosting the Big Five leagues are Germany, Italy, Spain, United Kingdom, and France. The Big Five leagues are the German Bundesliga, Italian Serie A, Spanish Liga, English Premier League, and French Ligue 1. *Significant at 10 percent level. **Significant at 5 percent level.

Table 4. Additional results

	Dependent variable: Female to male tertiary enrollment							
	Tertiary education more likely to be an option				Tertiary education less likely to be an option			
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
Bosman x Average players	3.985**	3.274*	5.243***	4.607***	-1.962	-5.530	-5.952	-0.557
	(1.584)	(1.671)	(1.356)	(1.451)	(6.659)	(7.311)	(8.366)	(3.787)
Observations	1,052	2,328	2,083	2,309	795	899	1,144	918

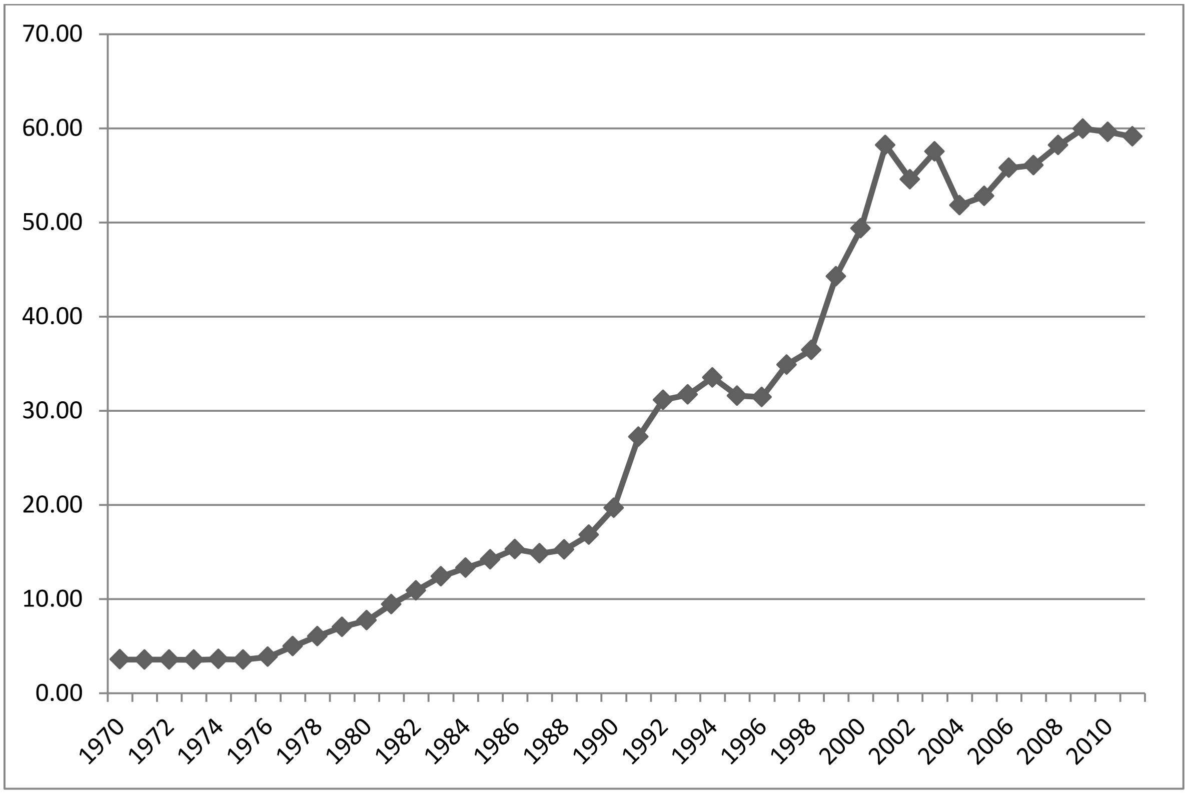
Notes: All models are estimated by OLS and include year fixed effects, country fixed effects, and controls (countries' population and GDP per capita growth). Standard errors clustered at the country level are in parentheses. Model (1) includes European countries only. Model (2) only includes middle- and high-income countries, according to the World Bank's classification in 1994 (GNI per capita above $725). Model (3) includes countries with male enrollment in 1994 above the sample mean of male enrollment in 1994. Model (4) includes countries with the ratio of female to male tertiary enrollment in 1994 above the sample mean of this ratio in 1994. Model (5) includes African countries only. Model (6) only includes low-income countries, according to the World Bank's classification in 1994 (GNI per capita below $725). Model (7) includes countries with male enrollment in 1994 below the sample mean of male enrollment in 1994. Model (8) includes countries with the ratio of female to male tertiary enrollment in 1994 below the sample mean of this ratio in 1994. *Significant at 10 percent level. **Significant at 5 percent level. ***Significant at 1 percent level.

Figure 1. Evolution of the ratio of female to male tertiary enrollment



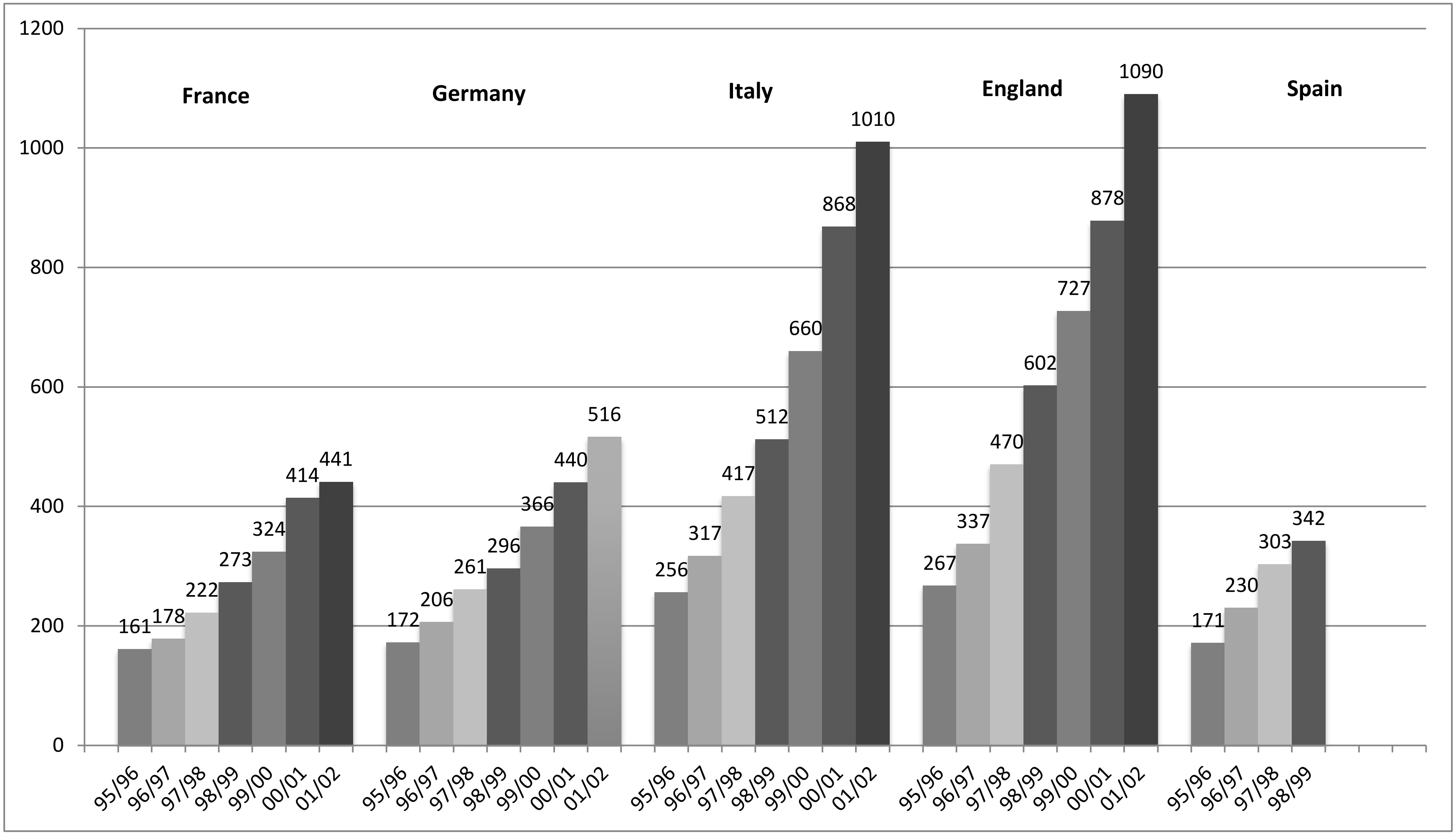
Note: The data corresponds to the yearly average ratio of female to male tertiary enrollment.

Figure 2. Evolution of the ratio of the average weekly salary of professional baseball players to the average weekly salary of production workers in the US



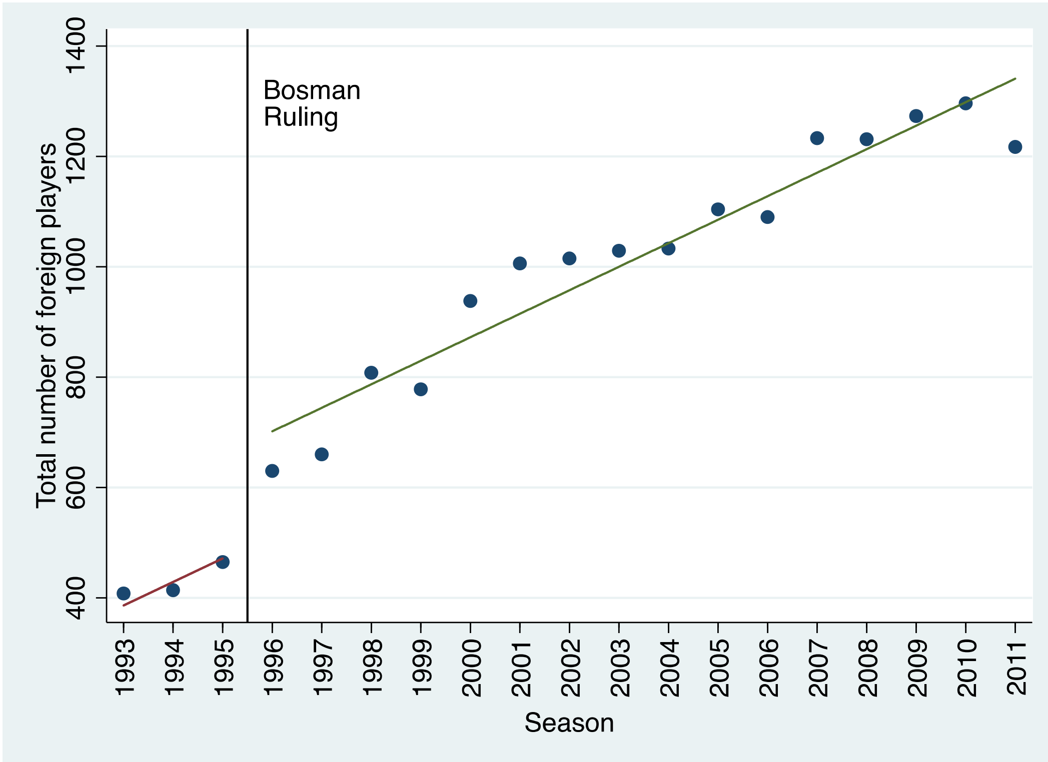
Notes: The series on average weekly salaries of professional baseball players (MLB) was built from annual data obtained through the websites of the Economic History Association (http://eh.net/encyclopedia/the-economic-history-of-major-league-baseball/) and USA Today (http://content.usatoday.com/sportsdata/baseball/mlb/salaries/team/), and dividing it by 52. The series on weekly salaries of production (blue-collar, hourly rated workers, or nonoffice) workers was built by multiplying data on hourly wages obtained from http://www.measuringworth.com/uswage/ by 40.

Figure 3. Salaries in the Big Five soccer leagues before and after the Bosman ruling



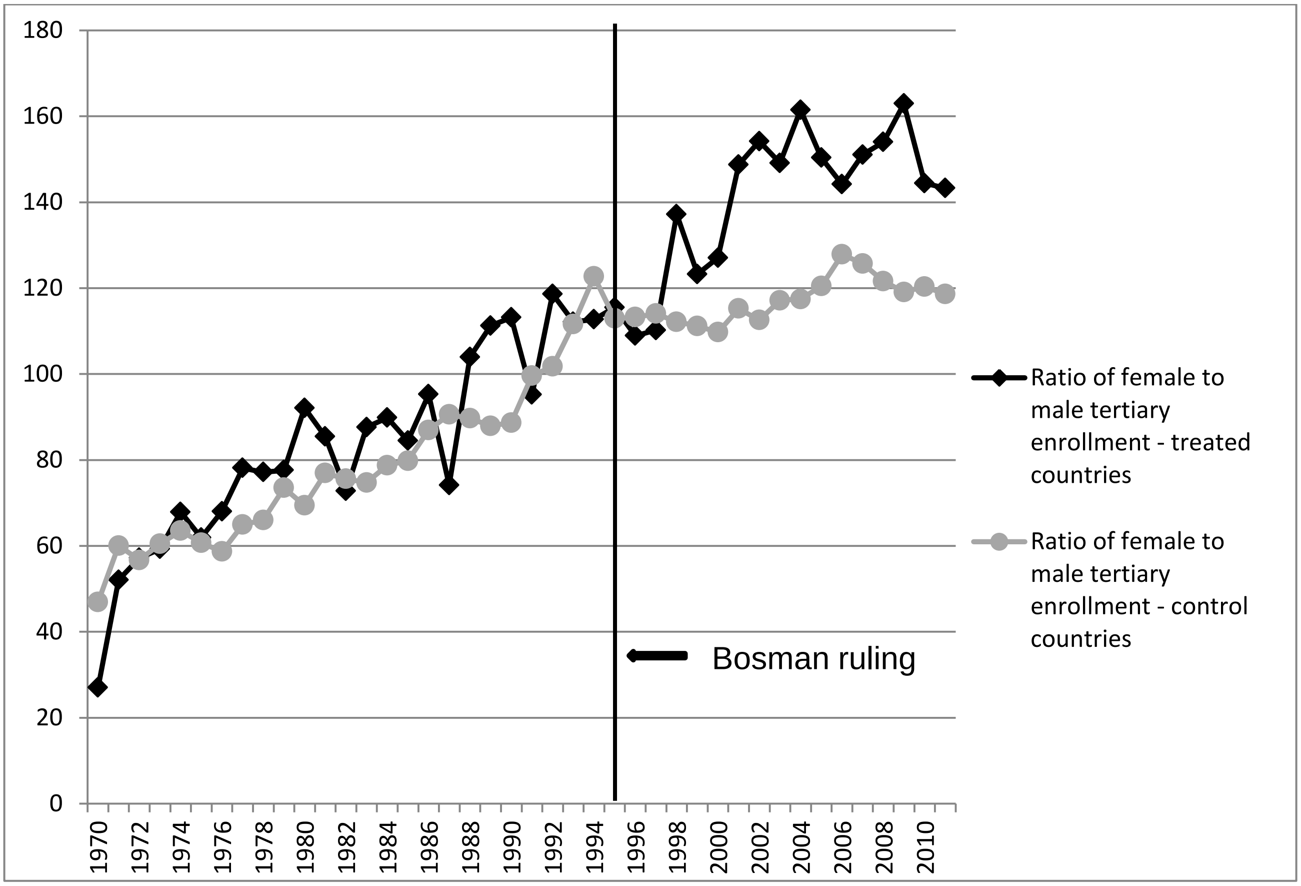
Notes: Data on gross salary expenditure (in millions of euros) by league and season were obtained from various issues of Deloitte's *Annual Review of Football Finance*. The horizontal axis displays soccer seasons in Europe. The Big Five leagues are the German Bundesliga, Italian Serie A, Spanish Liga, English Premier League, and French Ligue 1.

Figure 4. Evolution of the number of foreign players in Big Five soccer leagues before and after the Bosman ruling



Notes: The data points correspond to the total number of foreign players by year in the Big Five leagues, and were taken mostly from the website of Transfermarkt GmbH & Co. KG (http://www.transfermarkt.com/). Data for France in 1993-96 and Spain in 1993 and 1995 were estimated using data from Transfermarkt and http://www.national-football-teams.com/. The fitted lines through the data points are obtained from an OLS regression of the total number of foreign players on a time trend and a dummy variable that takes value 1 after the Bosman ruling (significantly different from zero at the 1 percent level).

Figure 5. Evolution of the ratio of female to male tertiary enrollment according to the countries' exposure to the Bosman ruling



Notes: The figure considers countries in which the ratio of female to male tertiary enrollment in 1994 was above the mean ratio for the full sample in that year. The treated group in the figure includes the top 10 countries according to the ratio of the number of soccer players in Big Five leagues to the country's population, in 1994. The Big Five leagues are the German Bundesliga, Italian Serie A, Spanish Liga, English Premier League, and French Ligue 1.