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Abstract

This paper assesses the impact of labor mobility on racial discrimination. We present an equilibrium search model that reveals an inverted U-shaped relationship between labor mobility and race-based wage differentials. We explore this relationship empirically with an exogenous mobility shock on the European soccer labor market. The Bosman ruling by the European Court of Justice in 1995 lifted restrictions on soccer player mobility. Using a panel of all clubs in the English first division from 1981 to 2008, we compare the pre- and post-Bosman ruling market to identify the causal effect of intensified mobility on race-based wage differentials. Consistent with a taste-based explanation, we find evidence that increasing labor market mobility decreases racial discrimination. (JEL J15, J31, J6, J71)

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1 Introduction

Race differentials in labor market outcomes continue to persist and evidence suggests that these differences are partly the consequence of racial discrimination.[1] "Why do people discriminate against one another, and how can we get them to stop?" (Gneezy and List, 2013, 5). Job-to-job mobility is a promising route to answer these questions. Constraints on mobility, such as quotas, work permits, or restrictive contracting rules on national labor, may limit the ability of workers to move from prejudiced firms to unprejudiced ones. When mobility is constrained, a firm is able to act on its prejudice because of the low cost of doing so.

This paper assesses the role of job-to-job mobility on racial discrimination. We present an equilibrium search model that reveals an inverted U-shaped relationship between labor mobility and race-based wage differentials. We explore this relationship empirically with an exogenous mobility shock on the European soccer labor market. The Bosman ruling by the European Court of Justice in 1995 lifted restrictions on soccer player mobility. Using a panel of all clubs in the English first division from 1981 to 2008, we compare the pre- and post-Bosman ruling market to identify the causal effect of intensified mobility on racial discrimination. Consistent with a taste-based explanation based on the work of Becker (1957), we find evidence that racial discrimination disappears with high levels of labor market mobility.

The European soccer market offers four important advantages for the study of mobility and discrimination. First, following the Bosman ruling, we observe large variation in labor mobility, making this market a valuable and visible laboratory to study the effect on discrimination. Specifically, the pre-Bosman era had two important restrictions on job-to-job mobility: (1) transfer fees needed to be paid for out-of-contract players and (2) the number of foreigners was restricted by a quota system.[2] The ruling, which came into effect in December 1995, removed the quota barriers for European

[1] See Altonji and Blank (1999), Lang and Lehmann (2012) and Charles and Guryan (2011) for reviews of the literature.

[2] In Europe, the number of foreigners allowed to play was governed by a "3+2" rule: 3 foreign European players and 2 non-European players.

Union (EU) nationals and the obligation to pay a fee for out-of-contract players.[3] Figure (1) illustrates the intensified mobility of the soccer market in the wake of the Bosman ruling.[4] In 2008, the ratio of foreigners in first league squads exceeded 50% in England and in other European countries such as Germany, Greece and Switzerland. Players can now field offers from potentially any country in the EU. This policy change creates a compelling quasi-experimental variation to identify the causal effect of mobility on racial discrimination.

Second, extensive data on the career paths of professional soccer players can be gathered for most countries over long time periods.[5]

Third, we can match this extensive individual data with racial information. This observable information enables economists to study racial discrimination issues in countries maintaining a "color-blind" model of public policy. France, for instance, does not collect data on the race or ethnicity of its citizens and has banned the computerized storage of race-based data.

Fourth, the soccer market offers a simple test for racial discrimination in salary setting (Szymanski, 2000). Assessing whether a group of workers is facing wage discrimination is typically difficult because we may not be able to capture productivity exactly.[6] With this caveat in mind, Szymanski's 'market test' is particularly elegant and parsimonious. Under the assumption that soccer is an efficient market,[7] a team's wage bill should perfectly reward the talent of its players. Discrimination can then be said to exist if clubs fielding

[3]Specifically, in June 1990, at the end of his contract, the former Belgian player Jean-Marc Bosman wanted to move from Liege in Belgium to Dunkirk in France. Because Dunkirk refused to meet the transfer fee demand, Liege refused to let him go. Bosman decided to take his case to the courts and won. As a result, the European Commission applied European law on worker mobility to the soccer labor market.

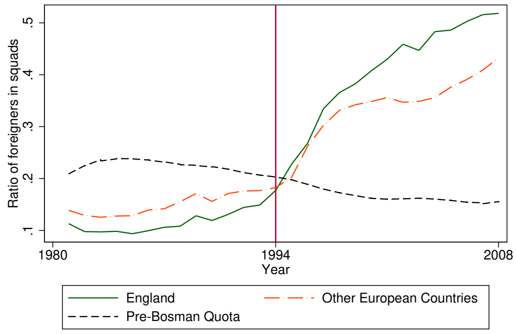
[4]The transfer rate of players from one club to another is statistically different before and after the Bosman ruling (p-value=0), with an average of 20% of players changing clubs each season over the 1996-2008 period compared with 13.5% for the 1981-1993 period.

[5]Kleven, Landais, and Saez (2013) compiled data of all first-league soccer players for 14 Western European countries since the eighties. We are very grateful to them for sharing their data with us.

[6]Using a Mincer-type wage equation, for instance, with a dummy indicating whether a person belongs to a particular group is plagued by a common estimation problem: how do we account for unobserved variables? (See Charles and Guryan, 2011).

[7]Unlike the professional sport labor markets in the US, there are no collective bargaining agreements, salary caps, or draft picks to maintain a competitive balance between teams.

Figure 1: Share of foreigners in club squads over time (1981-2008)



Notes: Other European Countries are Denmark, France, Germany, Greece, Italy, Netherlands, and Spain. The vertical axis indicates the pre-Bosman ruling year. The dashed line indicates the pre-Bosman quota of foreign players (3 EU + 2 non EU) in percentage of the size of the squads.

an above-average proportion of black players systematically outperform clubs with a below-average proportion of black players. This implies that, for a given wage bill, a team can improve its performance by employing a higher share of black players and *in fine* that black players are being paid less than their talent would warrant. Szymanski finds evidence of discrimination while performing his test on a panel dataset of professional English clubs between 1978 and 1993, i.e., *before* the Bosman ruling. We extend the analysis one step further. First, we theoretically demonstrate why discrimination can survive in equilibrium with labor mobility constraints. Second, we exploit the Bosman ruling shock and provide empirical evidence that discrimination disappears with high mobility. We briefly present these two contributions.

The first contribution consists of establishing a simple theoretical model to guide our empirical analysis. To derive equilibria in which group differentials persist, we merge ideas from search models of the labor market, à la Burdett and Mortensen (1998), with a Becker-style assumption of taste for discrimination.[8] Firms are heterogeneous in the talent of the workers whom they employ, and they offer workers two types of contracts: perfect and imperfect contracts. A job offer is considered to be "perfect" when a worker receives a wage that perfectly rewards his talent. In the case of an imperfect job offer, the worker earns only a fraction of what his talent is worth and thus

[8] Altonji and Blank (1999) reviews the pioneering works introducing search into taste-based theories of discrimination: Borjas and Bronars (1989), Black (1995) and Bowlus and Eckstein (2002).

searches on the job. Hence, the firm faces a job offer trade-off between diverting a share of the player's talent and seeing the player potentially poached by a rival, which induces a costly turnover.[9]

We assume that the taste parameter for discrimination affects the job offer trade-off. A prejudiced firm has lower disutility when it terminates an employment relationship with disliked workers, and thus, its probability of offering these workers a perfect job offer is lower. This results in race-based wage differentials. However, we find that this discriminatory behavior does not survive in all circumstances. The prejudice does not translate into discrimination when labor mobility is sufficiently low or sufficiently high. The intuition is straightforward. For a given talent, when mobility is sufficiently low, firms have high monopsony power and offer *imperfect* contracts to all workers, independent of their race.[10] By contrast, when mobility is sufficiently high, monopsony power is reduced, and firms offer *perfect* contracts to all workers to avoid costly turnovers. In between these extreme cases, the relationship between labor mobility and racial wage differentials follows an inverted U-shaped form.

Our model is close to Bowlus and Eckstein (2002), where firms are engaged in search.[11] In their model, disliked workers are (potentially) less talented on average and are sought less intensively by prejudiced firms; thus, these workers receive fewer offers and lower wages and have higher unemployment rates. We complement their analysis by considering that prejudiced firms do not search less intensively for disliked workers, who have the same distribution of talents as preferred workers. The prejudice applies to the separation rate. Prejudiced firms have a lower disutility in parting with disliked workers. Our results are consistent with Bowlus and Eckstein (2002)

[9] In Holden and Rosén (2009), firms also offer two types of contracts, high versus low, depending on a random parameter governing productivity of the job-worker match. In our setting, the probability of offering a perfect or imperfect contract is endogenized.

[10] Note that this statement does not imply that all wages are equal, as workers are heterogeneous in talent, and wages are proportional to talent.

[11] Black (1995) also introduces taste-based discrimination in a search model. However, our model deviates from Black's in three important ways. First, workers rather than firms are engaging in search. Second, prejudiced firms cannot refuse to hire disliked workers at any positive wage, and third, workers are also allowed to search on the job.

and long-term race-base wage differentials. However, as in simpler models, such as Becker (1957), we predict that discrimination is eliminated through intensified labor mobility, i.e., increased competition to attract talent.

The second contribution consists of using the Bosman ruling on the English soccer market to estimate the influence of intensified labor mobility on discrimination.[12] Specifically, we interpret the Bosman ruling as a shock to job-to-job mobility. Recall that (1) players whose contracts have expired can now change clubs freely and (2) EU players can move within the EU without taking up a valuable space in a foreign team's quota. These two effects make it easier to "poach" employees and influence firms' monopsony power.

By comparing the pre- and post-Bosman ruling situation, we identify the causal effect of intensified mobility on racial discrimination.[13] We find empirical evidence that increasing mobility decreases apparent racial discrimination.[14] This result could be important for public policy. If we consider restrictive contracts to be an important component of the typically nebulous "labor market frictions," making it easier for disliked workers to change jobs and reducing frictions could lower discrimination.

Our result regarding intensified mobility is consistent with Becker's argument that intensified product market competition can reduce race-based differentials caused by prejudice. Using regulatory reforms in U.S banking, Levine, Levkov, and Rubinstein (2008) show that the exogenous intensifica-

[12]See Kahn (2000) for a survey on the use of sports to estimate the extent of discrimination in a much more detailed way than in other industries.

[13]The Bosman shock has been exploited in other contexts (see Binder and Findlay, 2012, or Kleven, Landais, and Saez, 2013). Kleven *et al.* (2013) look at changes in the tax rates and the response by players in order to estimate the effects of higher marginal tax rates on high-skilled labor. They find that players responded to changes in tax incentives, especially after the Bosman ruling.

[14]In a fascinating book documenting how soccer can help economics, Palacios-Huerta (2014) reviews Szymanski's paper. Concomitant to our work, Palacios-Huerta extends Szymanski's sample from 1993 to 2008 and confirms the absence of discrimination in this period. His intuition is that the emergence of a market for corporate control of English professional clubs since the early 1980s has increased the competitiveness of English soccer. The added competition would have been able to drive discriminating firms out of the market. This effect may complement our mobility mechanism. However, a potential issue is that a club experiencing very poor performance simply moves down from one division to another and is thus not driven out of the soccer market. Our hypothesis is that the labor market became more competitive primarily because of relaxed constraints on mobility.

tion of competition among non-financial firms has reduced the manifestation of racial prejudices in the demand for labor and has raised the wages for black workers to approach those of equally productive white workers.[15]

Our paper is also related to more recent works examining the influence of labor market tightness on discrimination (Biddle and Hamermesh, 2013, Baert *et al*, 2014). These studies emphasize that employers discriminate less in labor markets with a small number of job seekers relative to vacancies.[16] By offering more job opportunities, our mobility shock can be interpreted as a decrease in the ratio of employed job seekers to job offers, as there will be more competition between firms to attract workers. This situation should discourage employers from indulging in discriminatory tendencies.[17]

The rest of the paper is as follows. In section (2), we describe the context of our analysis and the competitive soccer market. In section (3), we set out a theoretical model to guide our empirical analysis and to explain how labor mobility affects discrimination. In section (4), we present the identification strategy and the specifications of the market test for discrimination. Our empirical results on discrimination in the English soccer league are presented in section (5). The most important result is that discrimination disappears after the Bosman Ruling. Section (7) concludes.

[15]Other studies trace the impact of competition on relative wages within a single industry. Heywood and Peoples (1994) and Peoples and Talley (2001) find that the deregulation of transportation industries increased the relative wages of black workers.

[16]Biddle and Hamermesh (2013) establish that gender discrimination is lower when the ratio of job seekers to vacancies decreases. The evidence for racial discrimination is less conclusive, however. Baert *et al* (2014) sent resumes to firms in industries with different labor market tightness. In sectors with few available workers and a large number of vacancies, the difference in call-back rates between resumes with Flemish-sounding names and those with Turkish-sounding names was almost zero, whereas it was significantly lower in "loose" labor markets.

[17]The Bosman ruling shock may have also increased competition among workers by offering firms a larger pool of job seekers. However, to the extent that firms pay a turnover cost when workers quit, their monopsony power is reduced by the increased mobility of the labor market.

2 The competitive market for soccer players

We have already discussed the four important advantages offered by the soccer market to the study of mobility and discrimination: (1) large observed variation in labor mobility following the Bosman ruling; (2) an extensive collection of data on the career paths of professional soccer players, (3) a match between individual data and race information; and (4) a simple test for racial discrimination in salary setting (Szymanski, 2000). We now describe the competitiveness of the soccer market.

The data are similar to those of Szymanski (2000) but differ in scope and time. In terms of scope, Szymanski uses a panel of 39 clubs from four divisions in the English soccer league over the 1978-1993 period. Thanks to Kleven, Landais, and Saez (2013), our dataset contains all professional soccer players, regardless of nationality, from the first league in England.[18] In terms of time, we extend Szymanski's sample to cover the post-Bosman era, from 1994 to 2008. Despite differences in scope and time, the most salient features remain.

First, league competition is hierarchical, focused on league rankings without play-offs. Each year, approximately 20 teams participate in the English first league. At the end of each season the worst-performing teams swap places with the highest-ranked teams in the second league. There are no collective bargaining agreements, salary caps, or draft picks to maintain a competitive balance between teams.

Second, clubs are heterogeneous in wage bills. To illustrate this idea, we use wage bills from the Companies House website, a British government agency that collects annual reports from registered companies. Precise wage data are provided for almost all the English clubs in our sample.[19] These wage data are considered reliable because they are obtained from audited annual

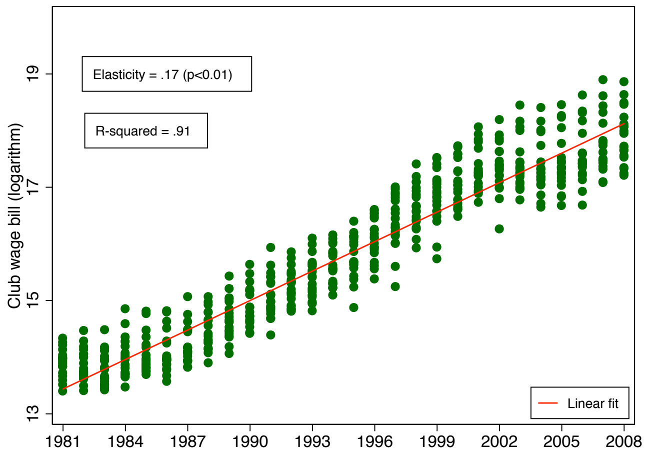
[18]We only have access to information on professional players from the top league, but our number of clubs is fairly similar to that in Szymanski's sample: 41 teams from 1981 to 1993.

[19]We are missing some data from clubs who have gone bankrupt during the season, such as Crystal Palace in 1998 or Leicester City in 2001, or from clubs that did not report wage bills in their financial accounts.

accounts.[20] One problem, however, is that we do not know what proportion of the pay is incentive related (e.g., bonuses for performing well in a cup competition) and what proportion is fixed. Another problem is that the wage bill is given for all staff at a soccer club, not only the players. Salaries for scouts, statisticians, physiotherapists, and coaches are also included in the wage bill. However, this practice is unlikely to be a problem because the pay for most of these employees is small compared with player wages and likely accounts for a similar share of the wage bills in all clubs.

In Figure (2), we report the log of the clubs' wage bills in the English first league from 1981 to 2008. Three facts are worth mentioning. First, each season (year), we observe variation in the wage bills across clubs. Second, we note that wage bills are linearly increasing over the years for all clubs. Third, remarkably, this linear positive trend is not affected by the Bosman ruling. A common explanation for this trend is the increasing price of talent.

Figure 2: Wage Bills in the English First League (1981-2008)

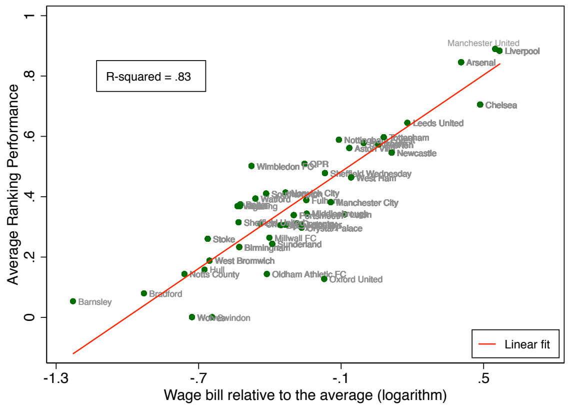


Third, soccer is a competitive market for talent. Players earn wages that reflect their talent, and the club's wage bill, which is a sum of all this

[20]Unfortunately, the reports are not homogeneous, as they sometimes use different starting and ending dates. In situations in which the wage bill reported was for a period of twelve months, we left the wages unchanged. Over the 30-year period, some companies changed the ending date on their company accounts. Such changes led to some annual accounts reporting thirteen or more months of data, in which case the data were adjusted on a pro rata basis.

talent, explains the club's ranking quite well. Using clubs' wage bill data and computing an index of performance based on league rankings (see 4.3 for detailed computation), we confirm that the English soccer market is quite competitive. Wage expenditures on players and performance are heavily correlated in the English league between 1981 and 2008. This correlation is depicted in Figure (3) and holds despite the surge in player remuneration (see Figure 2).

Figure 3: Average Wage bill and Performance in the English First League (1981-2008)



3 Theoretical framework

We establish an equilibrium labor market search model in the style of Burdett and Mortensen (1998), in which firms are homogeneous in size because the number of players per team is fairly rigid. Firms are instead heterogeneous in their wage bills (see section 2). We assume that a club's budget is given and is spent searching for talent and paying wages. In this search setting, we introduce racial discrimination. We assume that some employers hold a 'taste' for racial discrimination, meaning that there is a disamenity value of employing disliked workers.[21] If this assumption holds, then disliked workers

[21]In the final discussion (see section 7) we present some evidence of the persistence of this 'taste' or prejudice over time.

should 'compensate' prejudiced employers by being more productive at a given wage or, equivalently, by accepting a lower wage for an identical level of productivity. However, Becker (1957) argues that prejudice against disliked workers does not necessarily result in economic discrimination and race-based wage differentials. In other words, a taste for discrimination does not mean economic discrimination at the margin. Without some market failure, these wage differentials should be eliminated with competition. Using a job search model, we study the role of constraints on job-to-job mobility as a market failure. Specifically, we show that limited mobility, as was the case before the Bosman ruling, may explain race-based wage differentials. This result is formalized below.

We first present worker behavior (3.1), the flow conditions (3.2) and firm behavior (3.3) before introducing the role of taste discrimination(3.4). We solve the model analytically and present some simulations as an illustration of our results on wage discrimination (3.5) and job turnover (3.6).

3.1 Workers

The mass L of workers is divided into two types according to their appearance, A and B. Type B could be discriminated against by employers. All workers are heterogeneous in talent independent of their type, and each type has the same distribution of talent. The decision problem faced by a worker in a traditional job search model is simple (see e.g. Burdett and Mortensen, 1998); he maximizes utility over an infinite horizon in continuous time by adopting a reservation wage strategy that is state dependent. At any moment in time, each worker is either unemployed (state 0) or employed (state 1). Firms are engaged in search, and, at random time intervals, workers receive information about new or alternative jobs. This information is encapsulated in the parameter of the Poisson arrival process, λ, which denotes the arrival rate of job offers. This parameter reflects the general state of the labor market, specifically contracting rules, institutional constraints on and barriers to mobility. This market parameter also depends on the worker's current situation (employed or not). We thus assume that workers search more intensively while unemployed (state 0) than while employed (state 1), i.e. $\lambda_0 > \lambda_1$. Job-worker matches are destroyed at an exogenous positive rate, δ.

Workers are assumed to be risk neutral, with the discount rate r. Workers must respond to offers as soon as they arrive. The wage (ω) that they receive is function of their talent, t, such that $\omega = kt$, where $0 < k \leq 1$. Workers accept the job offer if it pays a higher wage while employed or if the instantaneous utility of being unemployed is lower than that of being employed.

In this wage posting framework, firms have monopsony power.[22] When a firm does not exercise its monopsony power, $k = 1$, and the worker receives a wage that perfectly rewards his talent, t, such that $\omega = t$. Because this perfect job offer is the best offer that the can receive with talent t, the worker remains with the employer. By contrast, when a firm exercises its monopsony power, $k < 1$, and the worker receives an "imperfect" or bad job offer. This implies that the worker receives a wage that is equal to only a fraction of his talent. Because his talent is imperfectly rewarded and because soccer is a market for talent (see section 2), the worker searches on the job. However, the probability of receiving a perfect job offer, γ, depends on the state of the labor market and on firm behavior (which is described in subsection 3.3).

Given this framework, the expected discounted utility of a job-seeker when unemployed, U, can be expressed as follows:

$$rU = b + \lambda_0(\gamma W_P + (1 - \gamma)W_I - U), \tag{1}$$

where W_P and W_I are the discounted values of filling a perfect and an imperfect offer, respectively. Equation (1) is rather standard (e.g., Pissarides, 1990). Being unemployed is similar to holding an asset. This asset pays a dividend of b, the unemployment benefit, and it has a probability $\lambda_0(1 - \gamma)$ of being transformed into a bad match, in which case the worker obtains W_I and loses U. It also has a probability $\lambda_0\gamma$ of being transformed into a good match, yielding a capital gain of W_P.

[22] Evidence of monopsony power in sport is documented in Kahn (2000).

For the sake of simplicity, we derive the discounted present value of employment in an imperfect match, W_I, for a given value of k rather than a continuous distribution as follows:

$$rW_I = kt + \lambda_1 \gamma (W_P - W_I) + \delta (U - W_I), \tag{2}$$

and for a perfect match as

$$rW_P = t + \delta (U - W_P). \tag{3}$$

Equations (2) and (3) have an intuition similar to that of equation (1). If a job seeker finds a perfect job, then he accepts the offer and remains in the job until an exogenous separation process moves him to unemployment (equation 3). If the job that he finds is bad or imperfect, then he rejects it if he is already employed, whereas if he is unemployed, he accepts the job and continues to search on the job. He remains in the job until either quitting to obtain a better job or an exogenous separation (equation 2).

Observe also that those equations are written under the assumption that U is always smaller than W_I (and W_P). Because this condition holds, there is no "waiting" behavior in this model, as an individual who receives a bad offer cannot hold out to receive a good offer. His reservation wage is such that it is always beneficial to accept an imperfect match.[23]

3.2 Flow conditions

As jobs are identical apart from the wage associated with their talent, employed workers move from lower- to higher- paying jobs as the opportunities arise. Workers also move from employment to unemployment and vice versa. We use standard equilibrium conditions (e.g., Mortensen 1988) to solve for the steady-state equilibrium labor supply. In the steady state, all flows must be balanced for there to be a stable equilibrium. Thus, we find three equilibrium conditions to determine the equilibrium shares of unemployed (u) and employed workers in perfect (P) or imperfect (I) jobs. Flows in and out of

[23] Given the high wages of soccer players, this assumption does not seem to be outlandish.

unemployment have to balance (i), as well as flows in and out from imperfect (ii) and perfect (iii) jobs:

i. Unemployment flows have to balance:

$$\lambda_0 u = \delta(1-u) \Rightarrow u = \frac{\delta}{\lambda_0+\delta}. \tag{4}$$

ii. Number of flows out of imperfect matches are equal to number of flows into imperfect matches

$$I(\delta+\lambda_1\gamma) = \lambda_0(1-\gamma)u \Rightarrow I = \frac{\lambda_0(1-\gamma)\delta}{(\delta+\lambda_1\gamma)(\delta+\lambda_0)}. \tag{5}$$

iii. Number of flows out of perfect matches are equal to number of flows into perfect matches

$$\delta P = \lambda_1\gamma I + \lambda_0\gamma u \Rightarrow P = \frac{\lambda_1\gamma\lambda_0(1-\gamma)}{(\delta+\lambda_1\gamma)(\delta+\lambda_0)} + \frac{\lambda_0\gamma}{(\lambda_0+\delta)}. \tag{6}$$

3.3 Firms

Based on the firm's behavior, we endogenize the probability of offering a perfect job offer, γ. Firms maximize utility (a function of profits) by offering a wage for workers depending on their talent (the role of appearance will be determined in the next subsection). Their reason for offering a perfect or imperfect job offer arises from the tension between diverting a share $(1-k)$ of the worker's talent and seeing the worker potentially poached by a rival firm at the rate $\lambda_1\gamma$, which induces a turnover cost $c > 0$. In this model, we do not detail how workers and firms are matched. For the sake of simplicity, we assume a random match. Once a random match is made, the firm's expected profits or value functions for offering a perfect (J_P) or imperfect (J_I) contract for a given talent can be determined.

$$
\begin{aligned}
rJ_I &= (1-k)t + (\delta + \lambda_1\gamma)(V - J_I - c) \\
&\Leftrightarrow \;\; J_I = \frac{(1-k)t + (\delta + \lambda_1\gamma)(V - c)}{r + (\delta + \lambda_1\gamma)},
\end{aligned} \tag{7}
$$

$$
rJ_P = \delta(V - J_P - c(t)) \Leftrightarrow J_P = \frac{\delta[V - c]}{r + \delta}. \tag{8}
$$

The equation for the imperfect job offer (7) states that the value of choosing the offer is equal to what the firm can divert, $([1-k]t)$, minus the loss that occurs from the player being poached $(V - J_I - c)$ with V the value of a vacancy. We assume that when a worker leaves the firm, a vacancy is created; however because the turnover cost is positive, the firm incurs a loss whenever a worker leaves. A worker leaves a firm for two reasons: either through the exogenous separation process, which occurs at rate δ, or because a rival firm has poached the worker, which occurs at rate $\lambda_1\gamma$. When the firm offers a perfect job offer, the worker can leave only through the exogenous separation process.

From the equations above, we can easily find that:[24]

$$
\begin{aligned}
J_P > J_I \;\; &\Leftrightarrow \;\; \frac{\delta[V - c(t)]}{r + \delta} > \frac{(1-k)t + (\delta + \lambda_1\gamma)(V - c(t))}{r + (\delta + \lambda_1\gamma)} \\
&\Leftrightarrow \;\; c(t) > \frac{(1-k)t(r+\delta)}{\lambda_1\gamma r}.
\end{aligned} \tag{9}
$$

We assume that c follows a Pareto distribution with a lower turnover cost bound $\tilde{c}$ and shape parameter $\alpha \geq 0$. This assumption implies a distribution of turnover cost draws given by

$$
G(c) = \left(\frac{c}{\tilde{c}}\right)^{-\alpha}, \qquad c \in [\tilde{c}, \infty].
$$

[24] We use here a result that is standard in the search literature: in a steady-state equilibrium, free entry ensures that the value of a vacancy is zero. Thus: $V = 0$.

The shape parameter α indexes the dispersion of turnover cost draws. The Pareto parametrization of c is intuitive because most turnover costs are low, but as α increases, the relative number of high turnover costs increases, and the cost distribution is more concentrated at these higher cost levels. Assuming that c is distributed Pareto and convex in talent, such that $c(t) = ct^2$, yields a simple closed-form solution for γ. The probability of receiving a perfect offer is such that $J_P > J_I$:

$$\gamma(t) = \left(\frac{\tilde{c}\lambda_1 \gamma r t}{(r+\delta)(1-k)}\right)^{\alpha} = \left(\frac{\tilde{c}\lambda_1 r t}{(r+\delta)(1-k)}\right)^{\frac{\alpha}{1-\alpha}}. \qquad (10)$$

As shown in equation (10), the probability γ depends primarily on three important variables: λ_1, k, and t. Everything else being equal, λ_1 determines the strength of the firm's monopsony power. For a high value of λ_1, firms are less likely to propose low wage offers to employees because they anticipate that other firms can poach these workers. Thus, a high λ_1 not only directly decreases the number of employees that are stuck in imperfect matches (see equation 5) but also reduces the probability of being in an imperfect match by increasing γ. Equation (10) gives an indirect channel through which a high λ_1 can decrease wage differences (and thus discrimination) among workers.

The effect of the other parameters is straightforward: as k increases, firms can divert a smaller share of the monetary value of a worker's talent; therefore giving an imperfect offer to a worker is less attractive. Regarding talent t, we assume that workers who are more talented are more costly to replace. This assumption seems reasonable but depends on the convexity of the turnover cost.

3.4 Employer prejudice

We assume that some firms hold a 'taste' for discrimination. In this spirit, Bowlus and Eckstein (2002) and Bowlus *et al.* 2001 consider that the arrival rates (λs) and the job destruction rate (δ) vary according to the type of worker (A and B). Recall that employers can discriminate against type B. Accordingly, the arrival rates of job offers for type B workers from prejudiced firms are lower than those for A workers. It seems natural to assume that, if an employer does not like a particular type of worker, then lower efforts would be made by the employer and the worker to meet one another. This implies that $\lambda_{i,A} \geq \lambda_{i,B}$ for $i = 0, 1$. Despite this natural assumption, we simplify the analysis and assume that λs and δ do not vary according to prejudice, such that $\delta_A = \delta_B$ and $\lambda_i = \lambda_{i,A} = \lambda_{i,B}$ for $i = 0, 1$.

This simplification has two main advantages. The first is abstracting from explicit discriminatory hiring practices. The λ parameters reflect the general state of the labor market, and thus, the potential constraints on labor mobility. Holding everything else constant, lower constraints on job-to-job mobility are associated with a higher arrival rate of job offers. Therefore, in our context, the end of both transfer fees for out-of-contract players and quotas within the EU imply that players receive a higher number of job offers from a larger number of firms after the Bosman ruling than they did before. It is thus reasonable to assume that λ_1 is higher after the ruling than before the ruling because of the different states of the labor market.[25] Furthermore, assuming also different values of λ_1 for types A and B would add a complication by introducing discriminatory hiring practices related to different search intensities (i.e., $\lambda_{1,A} \neq \lambda_{1,B}$).

The second reason for this simplification is that models based on Mortensen (1988) have unrealistic, right-skewed wage distributions. This problem does not occur in our model, as the wage distribution depends on the distribution of talent.[26]

In our approach, we assume that the two groups of workers, A and B, differ in their probability of drawing a perfect match, γ. These two groups have different Pareto distributions governing the turnover cost. In particular, consider the case in which $\tilde{c}_B = \tilde{c}_A - d$, where d represents the taste for discrimination and $\tilde{c}$ represents the lower bound cost, such that the Pareto distribution of the turnover cost for B workers is shifted to the left. In our particular labor context, this shift to the left reflects the taste for discrim-

[25] We may also consider different values of λ_0 before and after the Bosman ruling, but we abstract from this complication to focus our attention on the role of λ_1.

[26] Because talent and productivity are often Pareto distributed, adopting the same approach would give us a more plausible left-skewed distribution for wages.

ination d and the idea that a type B player who is unlikely to be a fan or manager favorite will be less costly to let go or easier to replace. This assumption gives us a lower equilibrium value of γ for individuals in group B compared with those in group A.

From this assumption on γ, we can ground Szymanski's market test. Assume that a firm faces a budget constraint.[27] Then revenues (R), from the overall team performance, are equal to the club's wage bill (Ω). We can write this down as:

$$\Omega = R = (1-\mu)(\gamma_A t_A + (1-\gamma_A)kt_A) + \mu(\gamma_B t_B + (1-\gamma_B)kt_B),$$

where μ is the share of black players in the squad, and t_i is the talent of an $i = A, B$ type player. Consider two teams with different shares μ of black players but with same revenues. Because γ_B is lower, they will have a lower wage bill for the same talent. Firms that hire more black players increase their performance compared with those that do not, which is exactly what Szymanski's market test aims to estimate.

3.5 Wage discrimination

Our perspective is that the Bosman ruling modifies the general state of the labor market. In particular, this modification affects the job-to-job mobility parameter, i.e., the arrival rates of job offers for employed workers (λ_1). We thus consider the effect of different values of λ_1 on wage discrimination.

We define wage discrimination as the difference in expected wages between individuals in groups A and B, such as:

$$\mathrm{E_A}(\omega|t) - \mathrm{E_B}(\omega|t) = bu_A + ktI_A + tP_A - bu_B - ktI_B - tP_B. \qquad (11)$$

After some algebraic manipulation (reported in appendix A.1) and given that both groups have the same likelihood of being out of work, equation (10)

[27] This budget constraint is not necessarily binding in the sort-run, since many clubs post a deficit in a any given year. However, in the medium-run, clubs who overspend can go bankrupt, for instance Crystal Palace in 1999, Leicester in 2001, Leeds United and Portsmouth in 2010.

reduces to

$$\mathrm{E_A}(w|t) - \mathrm{E_B}(w|t) = t\lambda_0\delta(\gamma_A - \gamma_B)\left(\frac{(1-k)(\lambda_1+\delta)}{(\delta+\lambda_1\gamma_A)(\delta+\lambda_1\gamma_B)(\delta+\lambda_0)}\right). \quad (12)$$

In this expression, there is clearly no wage discrimination if $\lambda_1 = 0$ or $\gamma_A = \gamma_B$, i.e., if job-to-job mobility is null or if the probability of receiving a good job does not depend on the type of worker.

We now investigate how this wage discrimination varies when we modify the job-to-job mobility parameter. With everything else held equal, the effect of a higher λ_1 on discrimination is twofold: it not only directly decreases the steady-state share of individuals in bad or imperfect matches I (see equation 5) but also increases the likelihood of receiving a perfect offer, γ (equation 10).

Overall, through γ, the effect of λ_1 on discrimination is non-linear. For values of λ_1 close to 0, firms anticipate that their players cannot be poached and both A and B players ultimately receive imperfect wage offers. In this extreme case, there is no discrimination because A and B players with the same talent earn the same wage. As λ_1 increases, discrimination also begins to increase because all firms endogenize the taste for discrimination of prejudiced firms (i.e., $\gamma_A \geq \gamma_B$). Thus, as job-to-job mobility increases, A players receive better offers than B players, and wage discrimination increases for a given level of talent. Then, as mobility continues to increase, the process is reversed. B players begin receiving a high number of perfect job offers. As a result, the gap narrows and even disappears when mobility is sufficiently high. Formally we state this in Proposition (1).

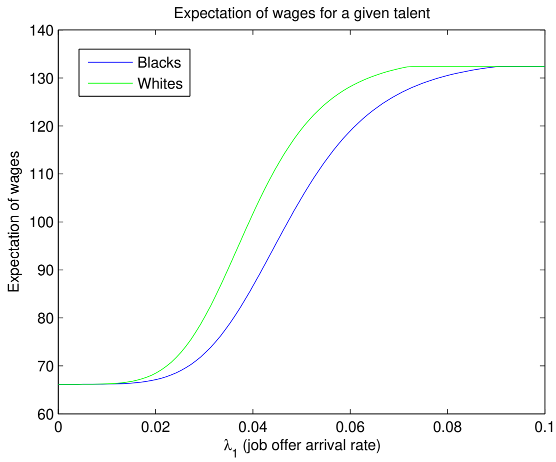
Proposition 1. *The difference in expected wage is a parabola that has a single maximum.*

Proof. Appendix (A.2) presents the proof. □

To gauge the effect of an increase in job-to-job mobility on wage discrimination, we simulate the effect of an increase in λ_1 on $\mathrm{E}(\omega|t)$ using plausible parameter values.[28]

[28] Although our labor market differs greatly from the market analyzed by Bowlus, we

Figure 4: Wage Expectation



As shown in figure (5), an increase in job-to-job mobility (captured by higher values of λ_1) can either decrease or increase discrimination in our model, even if it improves outcomes for workers of both type A and type B as depicted in figure (4).

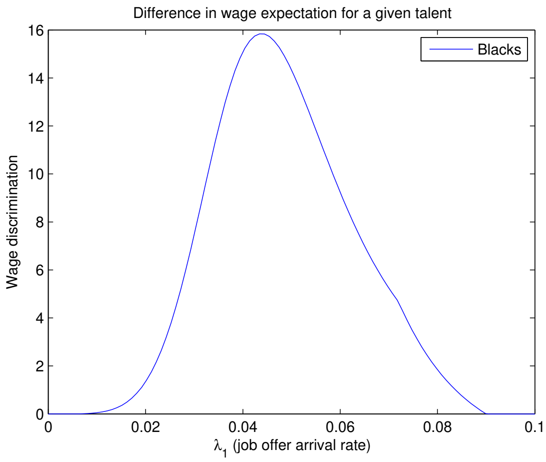
3.6 Job turnover

Our model also offers strong predictions regarding the turnover of type A and B workers, which affects the wage discrimination pattern. We establish in appendix (A.3) that there exists an optimal job offer arrival rate, such that type A workers change firms more often than B workers when below this rate, and less often when above it.

In our model, it is easy to find the expression for job turnover in a given period: $\lambda_1 \gamma I$. For each group, A and B, this expression yields an inverse U-shaped curve linking λ_1 and job turnover. There are two reasons that job turnover is low at both extremes of this curve. First, for low values of λ_1, few firms offer contracts to employed workers; thus, workers have a small likelihood of moving between firms. Second, there is an additional effect of

use the same values for the parameters employed in both papers. We hope that this methodological choice will allow the reader to better gauge the fit of our model. The parameter values for the simulations are tabulated in appendix C.

Figure 5: Wage discrimination



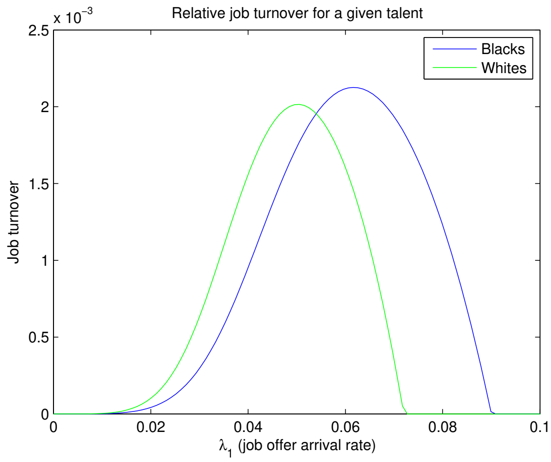
λ_1 on γ. When λ_1 is low, γ is also low; hence, even when employed workers receive a job offer, it is likely to be imperfect and thus rejected. Then, as λ_1 and γ increase, players move to take advantage of the new perfect job offers. However, as λ_1 increases, a "stock" effect comes into play: workers are less likely to move simply because those who accept the offer to change jobs (I) are much less numerous. Moreover, because a higher λ_1 implies a higher γ, workers are also likely to receive better offers. These two effects explain why mobility follows an inverse U-shaped curve. As λ_1 increases, the pool of workers who want to change clubs (I) decreases, as does the number of moves ($\lambda I \gamma$). The differentiated turnover between A and B thus derives from the differences in γ: the "stock" effect emerges earlier for A type workers. This pattern is stated formally in Proposition (2).

Proposition 2. *An increase in job-to-job mobility causes a racially differentiated change in job turnover.*

Proof. Appendix (A.3) presents the proof. □

As depicted in Figure (6), our simulation illustrates Proposition (2) and what was explained above.

Figure 6: Job turnover



3.7 Model conclusions

Our model provides a plausible explanation of why wage discrimination could have decreased post-Bosman despite the prejudice of some firms: if the job offer arrival rate increases, wages rise through two different channels. First, workers are more likely to receive "perfect" job offers. Second, the monopsony power of firms decreases because their employees are more likely to be poached.

Another interesting aspect of our model is that it predicts an unambiguous rise in wages following an increase in efficiency on-the-job search. In fact, wages have skyrocketed following the Bosman ruling, as depicted in a previous graph (see Figure 2).

In conclusion, our model fits a large number of empirical facts that were or will be stated in the next sections: an increase in wages, a racially differentiated effect on job turnover and a decrease in discrimination should follow an increase in job-to-job mobility.

4 Empirical strategy

This section describes how we move from our theory to empirics. To do so, we first present the market test to detect discrimination and discuss our data on racial information, its sources and its limitations (subsection 4.1). We then explain how we use the Bosman shock and present some descriptive statistics (subsection 4.2). Finally, we present the equations that we estimate to evaluate the extent of race-based wage discrimination (subsection 4.3).

4.1 The market test for discrimination

Our empirical results are based on Szymanski's market test. The intuition behind this test is that a team's performance is a good indicator of the team's talent. Thus, if all individual talent is perfectly rewarded, i.e., $\omega = t$ as in our model, then the team's wage bill should perfectly reflect its overall performance. Crucially, this performance should be independent of the team's racial composition when we control for the wage bill. By contrast, for a given wage bill, if teams fielding an above-average proportion of black players systematically outperform clubs with a below-average proportion of black players, then the labor market may be unfair toward black players (i.e., their talent is not fully rewarded and they face wage discrimination). Szymanski's market test is perfectly compatible with our theoretical framework (see subsection 3.4).

This test requires race information that we can match with extensive individual data. The race information was coded from an examination of players' photographs into categories of either black or not black (which we refer to as white). This method might sound arbitrary because we code players as "black" if they appear to be "black". However, this method is actually a good way to model the potential for discrimination because discriminators prejudge an individual based on appearances (Palacios-Huerta, 2014).[29] These pictures were obtained primarily from the reputable website `transfermarkt.de`, and when pictures from that site were not available, we

[29]For an explanation of why this appearance-based method is appropriate, Palacios-Huerta (2014) considers the case of the legendary Manchester United player Ryan Giggs. He appears to be Caucasian, and it was unlikely that he faced discrimination as a professional player during his career because discriminators prejudge an individual based on appearances. However, after he became famous, he publicly revealed that he had been victim of racism as a child because of his father's skin color. This revelation came as a surprise to his fans.

conducted Internet searches. We obtained pictures for nearly all the players in our sample. The players whose photos were missing were primarily youth team players who had had little game time and could thus be discarded from our analysis.

In addition to information on race, we added precise data on nationality, age, the number of matches played, the number of goals scored, national team selections (and their level - youth, A, ...), and whether a player participated in the World Cup. We use these last measures to create an objective, albeit imperfect, measure of quality of the players. Moreover, we added ranking and attendance information[30] to the club's wage bill and performance data (see section 2).

4.2 The Bosman shock and the identification strategy

We apply the market test for discrimination to a panel of all English clubs in the top league from 1981 to 2008, and we explore the Bosman ruling as an exogenous mobility shock to the European soccer labor market. The Bosman ruling was decided on December 15, 1995, by the European Court of Justice. This important decision lifted restrictions on soccer player mobility based on the European Community Treaty of the free movement of labor (article 39). This decision had a profound effect on transfers in the European soccer market by banning restrictions on foreign EU players within national leagues and by allowing players in the EU to move to another club at the end of a contract without a transfer fee being paid. Though this decision came into force in December 1995, it could have been anticipated because this case had been submitted to the Court on October 6, 1993. Thus, in December 1993, the European Union of Football Associations amended the regulations governing the *Status and Transfer of Football Players.* This amendment provided that a player may enter into a contract with a new club when the contract between him and his club has expired, has been rescinded or will expire within six months. However, the two clubs were still forced to agree on a transfer fee with a specific action in case of disagreement. To prevent

[30] From the European soccer statistics website.

any contamination of the results caused by a possible anticipation, we omit the 1994-1995 and 1995-1996 seasons. We thus compare the pre-Bosman era (1981-1993) to the post-Bosman era (1996-2008) to identify the causal effect of intensified mobility on racial discrimination.

Exploiting this shock and using our extensive player data, we constructed some informative descriptive statistics. We have information on 13,507 players who participated in the first English league during our two period; 77% of those players are English.[31] It is worth noting that this number was higher before the Bosman ruling (94.7%) than after (61.9%) it. In total, 12.7% of players are English and black; this number is fairly stable before (11.9%) and after (13.5%) Bosman. However, the number of black non-English players has skyrocketed from 0.7% before 1995 to 13% after the ruling. The consequence is that the number of white English players decreased from 82.8% before the ruling to 48.5% after the ruling.

Our strategy amounts to comparing black English to white English players to avoid comparing players of different nationalities. Table (1) presents information on the average number of matches per player, player quality and the age of the players across the two groups, both before and after the Bosman ruling. Among the different characteristics displayed in Table (1), we observe in the pre-Bosman period that black players did not play more matches or were not less tenured than white players. By contrast, we observe that these players were slightly more qualified and one year younger. In the post-Bosman era, the notable changes as follows: the quality difference is reduced (with $p < 0.1$), and black players have 6 fewer months of tenure. Interestingly, all the differences in differences are significant. In the pre-Bosman period, black players played more matches, performed better in terms of quality, were 6 months younger and had 6 fewer months of tenure. Those statistics clearly show that there was a Bosman effect on players and that this change differed along racial lines.

The difference between the pre- and post-Bosman eras is also observed

[31]We use the term English for the sake of comparison with Szymanski (2000), but a minority of those players are Irish, Scottish or Welsh. This naming makes sense in our context because these players were not considered foreigners in the English soccer market and were thus not subject to the foreign quotas.

Table 1: Individual differences in means

	Pre-Bosman (1981-1993)			Post-Bosman (1996-2008)			Overall
Variable	Black	White	Diff.	Black	White	Diff	D in D
Player's number of matches	20.77	20.24	.54	18.28	17.70	.59	2.48^a
			(.48)			(.43)	(.25)
Player's quality level	1.94	1.74	.20^a	2.26	2.16	.10^c	-.42^a
			(.05)			(.05)	(.03)
Age of players	24.07	25.01	-0.94^a	24.92	25.59	-0.67^a	-.56^a
			(.15)			(.15)	(.09)
Tenure (in years)	2.55	2.49	.06	2.65	3.06	-.41^a	-.48^a
			(.09)			(.07)	(.04)
Observations	904	5938		980	3531		

Notes: we compare here black and white English players in the first league. Diff. means Difference in means and D in D means Difference in Difference. Standard errors in parentheses, with ^{a} denoting significance at the 1% level.

in Table (2) using club statistics. All differences in terms of budget, transfer fee records, average attendance and club quality levels are highly significant both economically and statistically. The only exception is the share of black English players, which is constant in both eras, as shown above.

Table 2: Club differences in means

Variable	Pre-Bosman (1981-1993)	Post-Bosman (1996-2008)	Difference
Budget (in millions)	2,9	35,5	3,2^a
			(1,5)
Transfer fee record (in millions)	1,31	8,96	7,65^a
			(0,38)
Average attendance (in thousands)	21,5	33,2	11,7^a
			(0,9)
Share of black players	0.13	0.14	0.01
			(0.01)
Club's quality level	23.5	31.1	7.56^a
			(.64)
Observations	304	258	

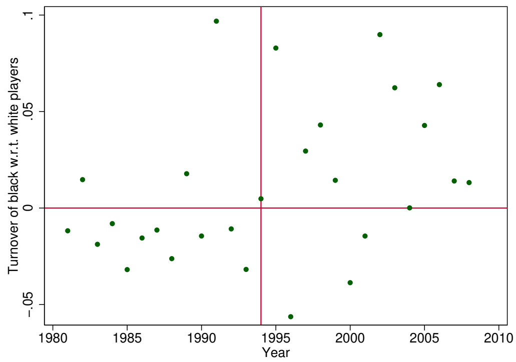
Notes: we compare here English first league clubs. Standard errors in parentheses, with ^{a} denoting significance at the 1% level.

As mentioned in the model, we should also expect a racially differentiated change in job turnover to accompany a change in the job offer arrival rate. In our empirical context, we define job turnover as club transfers, i.e., moving from one club to another during a given season. Figure (7) contrasts the turnover of black (B) and white (A) English players by comparing their share in the total number of transfers with respect to their share in the total population. The variable analyzed is the following:

$$\text{Turnover} = \frac{\text{Share in transfers}_B}{\text{Share in population}_B} - \frac{\text{Share in transfers}_W}{\text{Share in population}_W}$$

This variable is positive if black players change clubs more often in a given year than their white colleagues and negative if they do not.

Figure 7: Relative turnover of black English players



As shown in Figure (7), before the Bosman ruling, white players tended to change clubs more often than black players, but this tendency was reversed after Bosman. However, it is unclear when precisely this trend reversed.[32] We also find evidence, presented in part (B) of the appendix, that young black players took advantage of the new ruling to change clubs when they could have been discriminated against. How did all these changes affect wage discrimination? In the next subsection, we present the specifications that will enable us to estimate whether discrimination is present in the soccer labor market.

4.3 Estimated equations

To apply Szymanski's test, relating a club's performance to its wage bill (or budget) and its share of black English players, we use two different mea-

[32]This may be because we imperfectly observe club transfers. We have data on the first division only, so what we are measuring are transfers within the first division. This may explain our outlying data point for the season 1991-1992, a season where four clubs were promoted and whose transfers we do not perfectly capture.

sures of performance. The first measure is based on league rankings as in Szymanski's paper. The second is based on match results. In both cases, we corroborate Szymanski's finding of apparent wage discrimination in the *pre*-Bosman era. However, consistent with our theoretical predictions on the effect of relaxed mobility constraints, we find that wage discrimination has disappeared in the *post*-Bosman era.

The league performance specification

$$\begin{aligned}\text{League Performance}_{it} = \alpha_i &+ \beta_1(\text{WageBill}_{it} - \overline{\text{WageBill}}_t) \\ &+ \beta_2(\text{PlayersNb}_{it} - \overline{\text{PlayersNb}}_t) \\ &+ \beta_3(\text{Shareblack}_{it} - \overline{\text{Shareblack}}_t) + \epsilon_{it}. \qquad (13)\end{aligned}$$

In specification (13), our dependent variable, League Performance$_{it}$, is computed based on the final ranking of club's i at the end of season t.[33] α_i is a team i fixed effect, capturing permanent team-specific characteristics affecting performance, such as a location effect that allows clubs to underpay players, and ϵ_{it} is the usual error term. The team's wage bill ($\text{WageBill}_{it} - \overline{\text{WageBill}}_t$) is measured as the log difference of the club wage bill relative to the annual average ($\overline{\text{WageBill}}_t$). The relative number of players ($\text{PlayerNb}_{it} - \overline{\text{PlayerNb}}_t$) is computed as the difference between the number of players used in a season t and a club i relative to the average. This variable controls for "bad luck," as high turnover typically reflects a high level of injuries sustained. Finally, the relative share of black players ($\text{Shareblack}_{it} - \overline{\text{Shareblack}}_t$) is measured as the share of black players' appearances for a team in a given season t relative to the annual average ($\overline{\text{Shareblack}}_t$). We first compute this ratio based only on the share of black English players (see footnote 31). The share of black English players is relatively stable over time, but the share

[33]Note that as opposed to US team sports, there are no play-offs, and competition is solely about the ranking in the league. Specifically, League Performance is equal to $\frac{Ranking_{it} - min_t}{max_t - min_t}$, where min and max are the lowest and highest rankings possible for clubs at the end of the season. This transformation changes the sign of the discrimination effect compared to Szymanski's paper (see below), but not the main conclusions.

of black non-English players is constantly rising, which would lead us to estimate discrimination toward different individuals before and after Bosman. We later discuss results based on both black and white non-English players.

The coefficient of interest to us is β_3, the effect of the share of black players on performance. In case of race based wage discrimination, we expect β_3 to be positive.[34]

The match performance specification

One drawback of the previous measure of team performance based on final rankings is the relatively limited number of observations. We thus consider a new measure based on match results. Because the result of any given match is subject to some randomness or luck, there is a great deal of unexplained variance. However, using this measure, we obtain many more observations and can condition the performance between team i and team j on match fixed effects (ξ_{ij}). Our specification on match performance is:

$$\begin{aligned}\text{Match Performance}_{ijt} = \xi_{ij} &+ \beta_1 \log(\text{WageBill}_{it}/\text{WageBill}_{jt}) \\ &+ \beta_2(\text{PlayersNb}_{it} - \text{PlayersNb}_{jt}) \\ &+ \beta_3(\text{Shareblack}_{it} - \text{Shareblack}_{jt}) + e_{ijt}, \end{aligned} \tag{14}$$

Our new dependent variable, Match Performance$_{ijt}$ between team i and team j in year t, is simply the goal difference in the match between the two teams. Using this new variable, we apply the same idea to test for discrimination: for a given difference in the wage bill between teams i and j, a team i with a higher share of black players should not consistently outperform (in terms of goals) a team j with a lower share. By contrast, we expect large differences in wage bills that explain large differences in talent, to lead to large differences in performance and thus to large differences in goals scored. Our wage bill variable is the log difference between the budgets of the two clubs

[34]Compared with Szymanski, our transformation of the dependent variable implies, more intuitively, that in case of discrimination, an increase in the relative share of black players raises team's performance.

$\big(\log[\text{WageBill}_{it}/\text{WageBill}_{jt}]\big)$.[35] We also add the difference in the number of players used ($\text{PlayersNb}_{it} - \text{PlayersNb}_{jt}$). e_{ijt} is the usual error term. Finally, our variable of interest is the difference between the two teams in the share of matches played by black English players ($\text{Shareblack}_{it} - \text{Shareblack}_{jt}$).[36] The results are presented below.

5 Empirical results

We estimate models (13) and (14) with different panel data techniques in subsections (5.1) and (5.2), respectively. The combination of different specifications and estimators reinforces the robustness of our results.

5.1 Discrimination market test on league performance

We first use the ordinary least squares (OLS) estimator and the fixed-effects "within" estimator to eliminate the individual effect (α_i), with standard errors robust to clustering (because clubs are likely to be highly dependent across years). We then instrument for the wage bill by relying on the within-IV approach and the GMM estimator (Arellano and Bond, 1991). The latter estimator is useful, first, because it also eliminates the club fixed effects (through first-differencing) and, second, because it allows for a wide panel of instruments at the expense of removing some observations from our sample.

Estimation by OLS and Within

In Table (3), we first show our pre- and post-Bosman results using the discrimination market test on league rankings (equation 13) without instrumenting for the wage budget.[37] The relative wage bill variable has a positive

[35] Log differences are necessary because of the large increase in wage bills over time. Otherwise, we would not be able to pool our observations together.

[36] Unfortunately, we do not have team sheet data that would enable us to control the number of black players that are on the pitch at the match level.

[37] We exclude one year from the beginning estimation to ensure the use of the same sample period data for the instruments and the Arellano-Bond estimation. We also exclude the years 1994-1995 and 1995-1996 to avoid any anticipation effect. Therefore, this estimation is performed for the 1981-1982 to 1993-1994 seasons for the pre-Bosman period and for 1996-1997 to 2008-2009 seasons for the post-Bosman period.

effect on performance, which is economically and statistically significant. Unsurprisingly, this effect is larger if we do not control for the club fixed effect. The relative number of players used exhibits a negative effect in line with "bad luck" because high turnover typically reflects a high level of injuries sustained. As expected, we find contrasting results across the two periods for our estimate of interest: the estimates for the share of black English players. Controlling for the number of players used and the team's wage bill, we find apparent discrimination in the pre-Bosman era (columns 1 and 2). During this period, the performance depends significantly on the team's racial composition. Therefore, teams fielding an above-average proportion of black players outperform clubs with a below-average proportion of black players. After Bosman, the apparent wage discrimination disappears. Performance is now independent of the racial composition of teams.

Table 3: Market-test: League Performance and Discrimination - OLS and Within

Dependent Variable:	Relative League Performance			
Sample:	*Pre*-Bosman (1981-1993)		*Post*-Bosman (1996-2008)	
Estimator:	OLS	Within	OLS	Within
	(1)	(2)	(3)	(4)
Relative log wage bill	0.509^a	0.396^a	0.474^a	0.157^b
	(0.05)	(0.12)	(0.04)	(0.08)
Share of black English players	0.577^a	0.491^b	-0.130	-0.000
employed	(0.18)	(0.25)	(0.15)	(0.20)
Relative number of players used	-0.025^a	-0.031^a	-0.026^a	-0.027^a
	(0.01)	(0.01)	(0.01)	(0.01)
Observations	262	259	258	251
R^2	0.438		0.606	
Club fixed effect	No	Yes	No	Yes

Notes: the dependent variable Relative League Performance is equal to $\frac{Ranking_{it}-min_t}{max_t-min_t}$. Robust standard errors in parentheses, clustered by club, with ^{a}, ^{b} denoting significance at the 1% and 5% level respectively.

How economically meaningful is the estimate of discrimination? Let us compare the 1993 situation of two clubs that are identical except in their share of black English players. The "prejudiced" club does not employ black players, either because it cannot poach them or because it cannot retain them, whereas the "unprejudiced" club employs the 1993 average number of black players, i.e., 3.7 players. Based on the within estimates of column (2), we find that to obtain the same performance, the prejudiced club should pay

800,000 pounds more than the unprejudiced club. This value amounts to 15% of the average wage bill in 1993.

Estimation by Within-IV and GMM

Are our results plagued by endogeneity problems? Two possible problems are worth mentioning: (1) the potential mismeasurement of the wage bills, and (2) the fact that bonuses result in reverse causation because a higher performance may induce higher bonuses and thus a higher wage bill (if salary is incentive based). To address these problems, which could bias the estimate of the share of black players, we use an instrumental variable (IV) approach. We instrument the wage bill with the lagged performance on cups, lagged attendance, and relative record transfer fees.[38] The key identifying assumption is that these variables do not affect current final league performance apart from the wage bill.

We briefly discuss the relevance of our three excluded instruments. First, good performance in cups in the previous season should generate higher revenues in the contemporaneous season. Thus, there should be a strong link between the lagged year attendance in the stadium and the contemporaneous year income and wage bill. Finally, we use record transfer fees to capture potential buyouts by rich owners. New owners often break the club's transfer fee record, while these purchases often have little effect on final performance. Many record purchases often prove to be poor value for money, and the effect of transfer fees on performance are insignificant when controlling for the budget.

Te instruments are constructed as follows. The relative lagged attendance is measured as the one-year lag of a club's attendance relative to the annual average. Performance in cups is measured as the one-year club performance in the Football Association Challenge Cup and the League Cup. The relative transfer fee record variable is measured as the difference in a club's transfer fee record in a given season relative to the annual average.

The first-stage results reported in Table 9 (see appendix D) show that these instruments have a significant effect on the relative log wage bill, except

[38] Data are constructed from newspaper articles and online sources.

for the lagged cup performance in the post-Bosman era.

Table 4: Market-test: League Performance and Discrimination - IV and GMM

Dependent Variable:	Relative League Performance			
Sample:	*Pre*-Bosman (1981-1993)		*Post*-Bosman (1996-2008)	
Estimator:	Within-IV	AB	Within-IV	AB
	(1)	(2)	(3)	(4)
Relative log wage bill	0.335	0.576^a	0.200^c	0.472^a
	(0.25)	(0.10)	(0.12)	(0.04)
Share of black English players employed	0.501^b	1.364^b	-0.111	-0.116
	(0.25)	(0.54)	(0.18)	(0.48)
Number of players used	-0.030^a	-0.036^c	-0.027^a	-0.010
	(0.01)	(0.02)	(0.01)	(0.01)
Observations	259	262	247	257
Number of clubs	38	41	31	39
Number of instruments	3	28	3	30
AR1 p-value		0		0
AR2 p-value		0.70		0.37
Hansen p-value	0.72	0.47	0.15	0.48

Notes: the dependent variable Relative League Performance is equal to $\frac{Ranking_{it}-min_t}{max_t-min_t}$. AB means Arellano-Bond. Robust standard errors in parentheses, clustered by club, with ^{a}, ^{b}, and ^{c} denoting significance at the 1%, 5%, and 10% percent level, respectively.

The within-IV results are presented in columns (1) and (3) of Table (4). These results confirm the post-Bosman change. Whereas the coefficient for the share of black players employed is negative and significant before the Bosman ruling, it is positive and insignificant post-Bosman.

These results could be affected by a weak instrument problem. If the instruments correlates only weakly with the endogenous explanatory variable, then statements of statistical significance may be misleading. However, the first stage F-statistics on the excluded instruments are above the recommended threshold of 10 (see Table 9 in appendix D). It is also reassuring that the standard errors on the second-stage estimates are not much larger than those in the within model of Table (3). Moreover, the instruments pass standard validity assessments. The F-test of joint significance of the excluded exogenous variables is rejected at the 1% level. The test of overidentifying restrictions for the excluded instruments is also passed and the Angrist-Pischke first-stage chi-squared statistics reject the null of underidentification (Angrist and Pischke, 2009).

In columns (2) and (4) of Table (4), we use the two-step generalized method of moments (GMM) approach of Arellano and Bond (1991). This estimator differences away time-invariant club specific effects. It relies on the dynamic structure of the model for identification by using lagged levels of the independent variables as instruments for current differences. A problem with GMM estimators is that their validity is subject to the use of a relatively small or large number of instruments. A large number generates implausibly low values of Hansen tests of instruments exogeneity (Roodman, 2009), while using too few instruments is likely to generate a weak instruments problem and to deliver inaccurate estimates. Following Roodman's (2009) rule of thumb, the number of instruments is strictly lower than the number of clubs (groups) in the sample. This strikes a balance between estimate consistency and test validity. The diagnostic tests (Hansen and first and second order autocorrelation) presented at the bottom of the table reveal no evidence against the validity of the instruments used by the GMM estimator.

The GMM estimates of the share of black English players employed produce the same result as the other estimators: the apparent discrimination appears significant before the Bosman ruling but not after the ruling.

5.2 Discrimination Market-test on Match Performance

To test the robustness of our previous results, we employ a new methodology to detect discrimination based on a different performance variable: the discrimination market test on matches (see equation 14). For the sake of simplification, we use our two preferred estimators for this purpose: 'Within' and 'Within-IV'. This approach implies that for all estimations we control for the pair of clubs involved in the match. We are thus exploiting the time series variation in our panel by measuring the effect of differences in the racial composition of the teams on the difference in results. The pre- and post-Bosman results are reported in Table (5).

Even when controlling at the match level, we find the same evidence that discrimination is present before Bosman and disappears after the ruling. The estimates of the difference in the share of black players can be interpreted as follows. Consider two teams with the same wage bill and zero black players. The teams' expected game result is a draw. However, before Bosman (columns 1 and 2), a team that switched all its players for black players could expect to win the next match by a single goal. A similar effect could be achieved by doubling the wage budget. By contrast, the non-significant coefficients of the difference in the share of black players, reported in columns (3) and (4), suggests that wage discrimination has disappeared post-Bosman.

Table 5: Market-test: Match Performance and Discrimination - Within-IV and GMM

Dependent Variable:	Match Performance			
Sample:	*Pre*-Bosman (1981-1993)		*Post*-Bosman (1996-2008)	
Estimator:	Within	Within-IV	Within	Within-IV
	(1)	(2)	(3)	(4)
Difference in log wage bill	0.701[a]	1.469[a]	0.498[a]	0.340
	(0.17)	(0.42)	(0.15)	(0.28)
Difference in share of black players	0.854[a]	1.185[a]	-0.145	0.070
	(0.33)	(0.43)	(0.31)	(0.40)
Difference in number of players used	-0.048[a]	-0.054[a]	-0.029[a]	-0.030[a]
	(0.01)	(0.01)	(0.01)	(0.01)
Observations	4494	3129	4402	3264
Hansen p-value		0.79		0.22
Pair of clubs fixed effects	Yes	Yes	Yes	Yes

Notes: the dependent variable is the goal difference in the match. Robust standard errors in parentheses with [a], [b], and [c] denoting significance at the 1%, 5%, and 10% percent level, respectively. Standard errors are two-way clustered by receiving club and visiting club.

As before, we instrument the wage bill in order to account for possible measurement errors and reverse causality. In columns (2) and (4), we instrumented the difference in wage bills with the differences in the instruments used above (see subsection 5.1). The result of the first-stage estimates are reported in Table (10) in Appendix D. The instruments pass the standard validity assessments (see the bottom of Table 10). The F-test of joint significance of the excluded exogenous variables is rejected at the 1% level and above the recommended threshold of 10. The test of overidentifying restrictions for the excluded instruments is also passed and the Angrist-Pischke first-stage chi-squared statistics reject the null of underidentification (Angrist and Pischke, 2009). Moreover, it is again reassuring that the standard errors on the second-stage estimates (col. 2 and 4) are not much larger than those in the within model (col. 1 and 3, respectively).

6 Robustness analysis

We run two robustness checks: (1) on the differences between English and foreign black players (6.1) and (2) on the length of contracts (6.2).

6.1 Differences between English and foreign black players

We investigate the effect of the Bosman ruling on different categories of foreign players. Although the Bosman ruling lifted quotas for EU players, non-EU players are still subject to restrictive contracting conditions. For instance, to obtain a UK work permit, non-EU players must fulfill a set of stringent conditions.[39] As a consequence, despite a general increase in labor mobility, this increase should be relatively lower for non-EU players after the Bosman ruling. We find evidence of this pattern by performing the market test after Bosman on different shares of players: black English, EU black, non-EU black, and non-EU white.[40] Results are reported in Table (6). We find that the coefficients for the share of non-EU black players are significant and positive after Bosman, even if its statistical significance is lower when we introduce club fixed effects (col. 3 and 4). Those coefficients imply that wage discrimination did not disappear for the non-EU black players in the English first league.

6.2 Length of contracts

Although the Bosman ruling greatly liberalized the soccer market, players are not simply able to change clubs at will. If players remain under contract, then they cannot move to a new club without a transfer fee being paid to

[39]The rule is that the player must have played at least 75% of his national team's competitive matches over the last two years and that his national team mus be in the top 70 countries in the world. The appeals process allows for some flexibility in the rules, but the non-EU nationals playing in the Premier League are still expected to be of high quality.

[40]Non-EU players are non-member players in the common market or the EFTA zone. We do not report results for the pre-Bosman period, as there were few non-English black players playing in England (see section 4.2).

Table 6: Are non-EU players still discriminated? *Post*-Bosman (1996-2008)

Dependent Variable:	Relative League Performance			
Estimator:	OLS	IV	Within	Within-IV
	(1)	(2)	(3)	(4)
Relative wage bill	0.461^a	0.475^a	0.144^b	0.111
	(0.03)	(0.03)	(0.07)	(0.13)
Share of black English players employed	-0.060	-0.051	0.095	0.096
	(0.13)	(0.13)	(0.20)	(0.20)
Share of EU black players employed	0.056	0.029	0.125	0.174
	(0.20)	(0.20)	(0.31)	(0.32)
Share of ***non-EU black*** players employed	0.550^a	0.566^a	0.587^c	0.613^c
	(0.20)	(0.20)	(0.34)	(0.34)
Share of non-EU white players employed	0.082	0.073	0.256	0.267
	(0.17)	(0.17)	(0.17)	(0.16)
Relative number of players used	-0.026^a	-0.025^a	-0.026^a	-0.026^a
	(0.01)	(0.01)	(0.01)	(0.01)
Observations	157	157	251	249
Club fixed effect	No	Yes	No	Yes

Notes: robust standard errors in parentheses, with ^{a}, ^{b}, and ^{c} denoting significance at the 1%, 5%, and 10% percent level, respectively. Standard errors are clustered by club. The first-stage for the Within-IV (col. 4) is quite comparable with the one reported in Table (9) of Appendix (D) and available upon request with the corresponding usual statistics.

his old club. This observation may still be relevant in the aftermath of the Bosman ruling because some players may have signed contracts under the pre-Bosman market environment, and may therefore still face discrimination in the post-Bosman era until their contract comes to an end. Therefore, we should still observe some discrimination at the beginning of the Bosman period, with a decreasing trend in subsequent years as player contracts expire. We test for this possibility by interacting the share of black English players with a year trend in our regression. This interaction will capture whether the regression coefficient for the share of black players is indeed decreasing year by year as contracts expire in the immediate post-Bosman period.

We do not have data on contracts, but it appears that most contracts do not last more than five years. We thus run specifications (13) and (14) with the above interaction on a five-year period after the Bosman ruling: from 1996-1997 to 2000-2001. We use our two preferred estimators here: 'Within' and 'Within-IV'. The first stages of the Within-IV are qualitatively identical to those reported in Appendix (D) and are available upon request with the usual corresponding statistics. Results of this robustness check are reported in Table 7.

Table 7: Do long contracts delay the impact of the Bosman Ruling? *Post*-Bosman (1996-2000)

Dependent Variable: Team Performance based on	Rankings		Goal Difference	
Estimator:	Within	Within-IV	Within	Within-IV
Relative wage-bill	0.420^a	0.352	0.980^a	-0.868
	(0.07)	(0.32)	(0.08)	(1.04)
Share of black English players employed	0.823^c	0.926^c	0.251	1.626
	(0.43)	(0.51)	(0.64)	(1.04)
Year-trend	-0.290^b	-0.314^b	-0.204	-0.726^b
	(0.13)	(0.14)	(0.20)	(0.28)
Observations	99	94	1768	1202

Notes: columns (3) and (4) use explanatory variables in differences as in Table 5. Robust standard errors in parentheses, with ^{a}, ^{b}, and ^{c} denoting significance at the 1%, 5%, and 10% percent level, respectively. Standard errors are clustered by club in columns (1) and (2) and are two-way clustered by receiving club and visiting club in columns (3) and (4).

We find that there is indeed a negative and significant year trend in the immediate aftermath of the Bosman ruling, which reinforces our idea that the decrease in wage discrimination results from different contracting rules. Given that our regression is for 5 years only and applies to a low number of observations, the finding that these results are not very robust is expected. The significance of the year trend for instance, disappears when we introduce into the regression on rankings (col. 1 and 2) the number of players used, but the point estimates remain similar.

7 Conclusion and discussion

In this paper, we find strong evidence that wage discrimination has become insignificant following a decrease in labor market frictions. Given the empirical results, we feel confident stating that the Bosman ruling decreased and even eradicated the black-white wage gap for black English players in the Premier League. As shown in our model, a decrease in labor market frictions can erode the monopsony power of firms, potentially leading to a decrease in apparent discrimination. Our model appears to fit the empirical facts quite well. A heartening interpretation of our results is that the proper labor market conditions can cause wage differentials between white and black employees to disappear even if racist attitudes remain.

A potential objection to our findings is that what we are measuring may not truly be the effect of the Bosman ruling but may be a change in attitudes toward racism and racial discrimination. However, as Lang and Lehmann (2012) note, the wage gap between blacks and whites has been relatively stable over the period of the shock. The data reject the idea of a dramatic reduction in discrimination over our period of analysis for other professional categories as well. Racial discrimination remains an important issue today, as shown by Bertrand and Mullainathan (2004). These authors sent similar CVs with different names, and the more "black-sounding" names were half as likely as the others to receive a callback offer.

Moreover, racist incidents in soccer, whether from fellow soccer players, owners, managers or supporters continue to make the headlines of English newspapers: in 2011, the English captain John Terry was accused of racially abusing Anton Ferdinand, and in 2012, Bolton Wanderers striker Marvin Sordell was racially abused by a Millwall fan. In August 2014, it came to light that the Cardiff City manager Malky Mackay shared racist e-mails and texts with the director of soccer in charge of transfers, Iain Moody.[41] In November 2014, the owner of Wigan Athletic FC, Dave Whelan stated while defending his decision to hire Malky Mackay that, "I think Jewish people do chase money more than everybody else." All these incidents suggest first that racist attitudes are still present at all levels of English soccer, and, second, that the decrease in discrimination is more likely to be attributed to an increase in job-to-job mobility, in the wake of the Bosman ruling, rather than to a dramatic change in attitudes of prejudiced employers in 1995.

[41] For instance, the Daily Mail report the following: "On August 16, 2012, a list of players proposed by a French agent is forwarded, stating to Mackay that "he needs to rename his agency the All Blacks." A separate text in reference to a list of French players states the following: "Not many white faces amongst that lot but worth considering."

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A Proofs and derivations

A.1 A formula for wage discrimination

In this appendix we detail how we obtain the formula for wage discrimination (12). Let define the wage discrimination as the difference in expected wages between individuals in groups A and B, such as:

$$D = \mathrm{E_A}(\omega|t) - \mathrm{E_B}(\omega|t) = bu_A + ktI_A + tP_A - bu_B - ktI_B - tP_B. \quad (15)$$

We can simplify this expression since both types have the same likelihood of being out of work. Thus, the expression for discrimination D is:

$$\begin{aligned}
D &= kt\left(\frac{\lambda_0(1-\gamma_A)\delta}{(\delta+\lambda_1\gamma_A)(\delta+\lambda_0)} - \frac{\lambda_0(1-\gamma_B)\delta}{(\delta+\lambda_1\gamma_B)(\delta+\lambda_0)}\right) \\
&\quad +t\left(\frac{\lambda_1\gamma_A\lambda_0(1-\gamma_A)}{(\delta+\lambda_1\gamma_A)(\delta+\lambda_0)} + \frac{\lambda_0\gamma_A}{(\lambda_0+\delta)} - \frac{\lambda_1\gamma_B\lambda_0(1-\gamma_B)}{(\delta+\lambda_1\gamma_B)(\delta+\lambda_0)} - \frac{\lambda_0\gamma_B}{(\lambda_0+\delta)}\right). \\
&= kt\left(\frac{(\delta+\lambda_1\gamma_B)\lambda_0(1-\gamma_A)\delta - (\delta+\lambda_1\gamma_A)\lambda_0(1-\gamma_B)\delta}{(\delta+\lambda_1\gamma_A)(\delta+\lambda_1\gamma_B)(\delta+\lambda_0)}\right) \\
&\quad +t\left(\frac{(\delta+\lambda_1\gamma_B)\lambda_1\gamma_A\lambda_0(1-\gamma_A) - (\delta+\lambda_1\gamma_A)\lambda_1\gamma_B\lambda_0(1-\gamma_B)}{(\delta+\lambda_1\gamma_A)(\delta+\lambda_1\gamma_B)(\delta+\lambda_0)} + \frac{\lambda_0(\gamma_A-\gamma_B)}{(\lambda_0+\delta)}\right) \\
&= kt\left(\frac{\lambda_0\delta(\gamma_B-\gamma_A)(\lambda_1+\delta)}{(\delta+\lambda_1\gamma_A)(\delta+\lambda_1\gamma_B)(\delta+\lambda_0)}\right) \\
&\quad +t\left(\frac{\lambda_1\lambda_0((\gamma_B-\gamma_A)(\lambda_1\gamma_B\gamma_A-\delta)+\delta(\gamma_B^2-\gamma_A^2))}{(\delta+\lambda_1\gamma_A)(\delta+\lambda_1\gamma_B)(\delta+\lambda_0)} + \frac{\lambda_0(\gamma_A-\gamma_B)}{(\lambda_0+\delta)}\right) \\
&= kt\left(\frac{\lambda_0\delta(\gamma_B-\gamma_A)(-\lambda_1-\delta)}{(\delta+\lambda_1\gamma_A)(\delta+\lambda_1\gamma_B)(\delta+\lambda_0)}\right) \\
&\quad +t\left(\frac{\lambda_0\delta(\gamma_A-\gamma_B)(\lambda_1+\delta+\lambda_1\gamma_B+\lambda_1\gamma_A)+\lambda_0\lambda_1\delta(\gamma_B^2-\gamma_A^2)}{(\delta+\lambda_1\gamma_A)(\delta+\lambda_1\gamma_B)(\delta+\lambda_0)}\right) \\
&= t\lambda_0\delta(\gamma_A-\gamma_B)\left(\frac{(k*(\delta+\lambda_1))+(\lambda_1+\delta+\lambda_1\gamma_B+\lambda_1\gamma_A)-\lambda_1(\gamma_B+\gamma_A)}{(\delta+\lambda_1\gamma_A)(\delta+\lambda_1\gamma_B)(\delta+\lambda_0)}\right),
\end{aligned}$$

which amounts to the equation (12)

$$D = \mathrm{E_A}(\omega|t) - \mathrm{E_B}(\omega|t) = t\lambda_0\delta(\gamma_A - \gamma_B)\left(\frac{(1-k)(\lambda_1+\delta)}{(\delta+\lambda_1\gamma_A)(\delta+\lambda_1\gamma_B)(\delta+\lambda_0)}\right).$$

A.2 Proof of proposition 1

> Proposition 1. The difference in expected wage is a parabola that has a single maximum.

Proof. We look at the maximum and minimum of the equation D (12), derived in appendix (A.1), by taking the derivative of this function with respect to λ_1:

$$\begin{aligned}
\frac{\partial D}{\partial \lambda_1} &= t\frac{\lambda_0\delta\alpha}{(1-\alpha)\lambda_1(\delta+\lambda_0)}(\gamma_A-\gamma_B)\Big(\frac{(1-k)(\lambda_1+\delta)}{(\delta+\lambda_1\gamma_A)(\delta+\lambda_1\gamma_B)}\Big) + t\frac{\lambda_0\delta(1-k)}{(\delta+\lambda_0)}(\gamma_A-\gamma_B) \\
&\times \left(\frac{\delta^2 - (\frac{\alpha}{1-\alpha})\lambda_1\delta(\gamma_A+\gamma_B) - \lambda_1^2(\frac{\alpha+1}{1-\alpha})\gamma_B\gamma_A - \delta\left(\delta(\frac{1}{1-\alpha})(\gamma_A+\gamma_B) + \lambda_1(\frac{2\alpha}{1-\alpha}+2)\gamma_B\gamma_A\right)}{[(\delta+\lambda_1\gamma_A)(\delta+\lambda_1\gamma_B)]^2}\right) \\
&= t\frac{\lambda_0\delta}{\lambda_1(\delta+\lambda_0)(1-k)}(\gamma_A-\gamma_B)\left(\frac{(\frac{\alpha}{1-\alpha})(\lambda_1+\delta)(\delta^2+\delta\lambda_1(\gamma_A+\gamma_B)+\lambda_1^2\gamma_A\gamma_B)}{[(\delta+\lambda_1\gamma_A)(\delta+\lambda_1\gamma_B)]^2}\right. \\
&\quad \left. + \frac{\lambda_1\left[\delta^2 - (\frac{\alpha}{1-\alpha})\lambda_1\delta(\gamma_A+\gamma_B) - \lambda_1^2(\frac{\alpha+1}{1-\alpha})\gamma_B\gamma_A - \delta\left(\delta(\frac{1}{1-\alpha})(\gamma_A+\gamma_B)+\lambda_1(\frac{2\alpha}{1-\alpha}+2)\gamma_B\gamma_A\right)\right]}{((\delta+\lambda_1\gamma_A)(\delta+\lambda_1\gamma_B))^2}\right) \\
&= t\frac{\lambda_0\delta(1-k)}{\lambda_1(\delta+\lambda_0)}(\gamma_A-\gamma_B)\left(\frac{\delta^2\left((\lambda_1+\delta)\frac{\alpha}{1-\alpha}+\lambda_1\right) - \delta^2\lambda_1(\gamma_A+\gamma_B) + \lambda_1^2\gamma_A\gamma_B(\frac{\alpha\delta-\lambda_1-2\delta}{1-\alpha})}{[(\delta+\lambda_1\gamma_A)(\delta+\lambda_1\gamma_B)]^2}\right). \qquad (16)
\end{aligned}$$

In order to study the sign of this expression, we only need to study the sign of the second fraction of 16, since we consider cases where $\gamma_A \geq \gamma_B$. Consider the numerator of the second fraction:

$$\begin{aligned}
&\delta^2\left((\lambda_1+\delta)\frac{\alpha}{1-\alpha}+\lambda_1\right) - \delta^2\lambda_1(\gamma_A+\gamma_B) + \lambda_1^2\gamma_A\gamma_B\left(\frac{\alpha\delta-\lambda_1-2\delta}{1-\alpha}\right) \\
= \;& \frac{1}{1-\alpha}\left(\delta^2\left[\lambda_1(1-(1-\alpha)(\gamma_A+\gamma_B))+\delta\alpha\right] + \lambda_1^2\gamma_A\gamma_B\left[(\alpha-2)\delta-\lambda_1\right]\right).
\end{aligned}$$

We need to study the variation of the following equation in order to understand the sign of the derivative of equation (12):

$$\delta^2\lambda_1\left(1-(1-\alpha)(\gamma_A+\gamma_B)\right) + \lambda_1^2\gamma_A\gamma_B\left[(\alpha-2)\delta-\lambda_1\right] + \delta^3\alpha. \qquad (17)$$

Rewriting γ_A as $\left(\lambda_1^{\frac{\alpha}{1-\alpha}}y_A\right)$ we get:

$$\delta^2\lambda_1(1-(1-\alpha)\lambda_1^{\frac{\alpha}{1-\alpha}}(y_A+y_B)) + \lambda_1^{\frac{2\alpha}{1-\alpha}+2}y_Ay_B((\alpha-2)\delta-\lambda_1)) + \delta^3\alpha.$$

Taking the derivative once more, we get

$$\delta^2 - \delta^2\lambda_1^{\frac{\alpha}{1-\alpha}}(y_A + y_B)) - \frac{2}{1-\alpha}\lambda_1^{\frac{2\alpha}{1-\alpha}+1}y_A y_B(2-\alpha)\delta - \frac{3-\alpha}{1-\alpha}\lambda_1^{\frac{2\alpha}{1-\alpha}}y_A y_B.$$

This function is positive when λ_1 is 0, and then monotonically decreases as λ_1 increases. We should therefore expect function 17 to increase and then decrease. Notice, that when λ_1 is 0, equation 17 is positive and equal to $\delta^3\alpha$. For higher values of λ_1 (e.g. $\lambda_1 = 1$), this value is then negative. We can then establish that the function 17 is first positive, then negative at some value of λ_1. It follows that there will be a single value for which the derivative is equal to 0, and that below this value, equation 16 will be positive, and negative above it. Hence, the difference in expected wage will be a parabola with a single maximum. □

A.3 Proof of proposition 2

> Proposition 2. The difference in expected wage is a parabola that has a single maximum.

Proof. We need to find how the difference in moves evolves with λ_1, i.e., study the sign of: $DM = \lambda_1\gamma_B I_B - \lambda_1\gamma_A I_A$

$$\begin{aligned}
DM &= \frac{\lambda_1\gamma_B\lambda_0(1-\gamma_B)\delta}{(\delta+\lambda_1\gamma_B)(\delta+\lambda_0)} - \frac{\lambda_1\gamma_A\lambda(1-\gamma_A)\delta}{(\delta+\lambda_1\gamma_A)(\delta+\lambda_0)} \\
&= \frac{\lambda_1\gamma_B\lambda_0(1-\gamma_B)\delta(\delta+\lambda_1\gamma_A) - \lambda_1\gamma_A\lambda_0(1-\gamma_A)\delta(\delta+\lambda_1\gamma_B)}{(\delta+\lambda_1\gamma_B)(\delta+\lambda_0)(\delta+\lambda_1\gamma_A)} \\
&= \frac{\lambda_1\delta\lambda_0[\gamma_B(1-\gamma_B)(\delta+\lambda_1\gamma_A) - \gamma_A(1-\gamma_A)(\delta+\lambda_1\gamma_B)]}{(\delta+\lambda_1\gamma_B)(\delta+\lambda_0)(\delta+\lambda_1\gamma_A)} \\
&= \frac{\lambda_1\delta\lambda_0[\gamma_B\delta - \gamma_B^2\delta - \lambda_1\gamma_A\gamma_B^2 - \gamma_A\delta + \gamma_A^2\delta + \lambda_1\gamma_B\gamma_A^2]}{(\delta+\lambda_1\gamma_B)(\delta+\lambda_0)(\delta+\lambda_1\gamma_A)}.
\end{aligned}$$

If we want to find the crossing point, either λ_1 is 0, or $[\gamma_B\delta - \gamma_B^2\delta - \lambda_1\gamma_A\gamma_B^2 - \gamma_A\delta + \gamma_A^2\delta + \lambda_1\gamma_B\gamma_A^2]$ is 0. This gives us the formula for the crossing point:

$$\lambda_1^* = \frac{\delta[\gamma_B(\gamma_B - 1) + \gamma_A(1-\gamma_A)]}{\gamma_B\gamma_A(\gamma_A - \gamma_B)}.$$

Considering the case where $0 < \gamma_A < \gamma_B < 1$, above λ_1^* group B individuals change firms more often than individuals from group A and vice versa. The crossing point is above 0 whenever $\frac{\gamma_A}{\gamma_B} > \frac{1-\gamma_B}{1-\gamma_A}$. □

B Further evidence of discrimination

The age profile of black players before and after the Bosman ruling also presents some evidence of discrimination. Figures (8) and (9) depict the age density of black players in the squads of "discriminating" and "non-discriminating" teams before and after the Bosman ruling respectively. We consider discriminating firms as the teams whose proportions of black players in the squad is lower than 75% of the other squads, and non-discriminating firms as those where the proportions of black players is higher than 25% of the other squads. Of course, these measures of discriminating and non-discriminating firms are far from perfect, but they fit in with our empirical strategy and our model. In our model, firms that discriminate are more likely to have their players poached by rival firms. We should therefore expect these firms to have less black players.

In Figure (8), we observe that the age densities of black players pre-Bosman are quite similar in discriminating and non-discriminating clubs, but there is a huge change post-Bosman (Figure 9): the age density of black players is much more left-skewed in discriminating clubs. Why is this interesting? If we consider that mobility was constrained before Bosman, then we can propose the following explanation: when job-to-job mobility is high, players that were employed in discriminating clubs want to leave as soon as they get the chance. When players are young, they tend to play for their local clubs or simply for any club that wants them, whether these discriminate or not. Therefore, in discriminating clubs, the black players we find are mostly young: older players leave as soon as they get the chance. However, as job-to-job mobility was lower before Bosman, black players were less likely to gain from this strategy, hence the similar age profiles between discriminating and non-discriminating clubs.

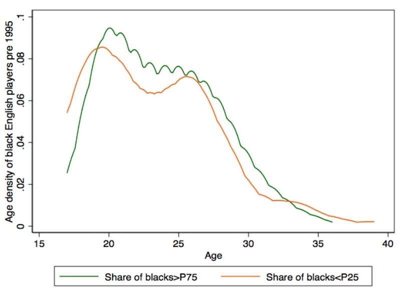

Figure 8: Pre-Bosman

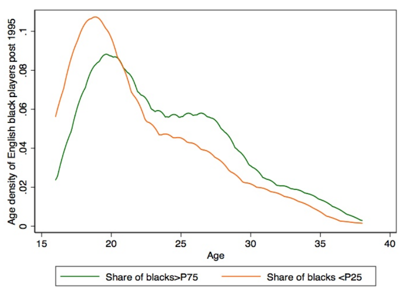

Figure 9: Post-Bosman

C Parameters for the simulations

Even though our labor market is very different to the one analyzed by Bowlus and Eckstein (2002), we use the same values for the parameters common in both papers. Our specification imposes an α between 0 and 1. Although the expectation of c is not defined in this case, we do not use this expectation in our analysis since firms directly observe the realization of c.

Table 8: Parameter values for the simulations

Parameters	Values
k	0.5
λ_0	0.04
δ	0.004
t	150
α_W	0.8
α_B	0.8
$\tilde{c}_W$	0.05
$\tilde{c}_B$	0.04
r	0.05

D First stage estimations

Table 9: Market-test: League Performance and Discrimination - 1st stage

Dependent Variable:	Relative Log Wage Bill	
Sample:	Pre-Bosman (1981-1993)	Post-Bosman (1996-2008)
Share of black English players	0.009	-0.003
	(0.140)	(0.140)
Number of players used	0.007^c	-0.002
	(0.004)	(0.003)
Lagged log attendances	0.248^a	0.522^a
	(0.008)	(0.098)
Lagged cup performance	0.011^b	0.004
	(0.005)	(0.004)
Relative record transfer fee	0.125^a	0.023^a
	(0.036)	(0.004)
Club fixed effects	Yes	Yes
Observations	259	268
F test of excluded instruments	14.83^a	28.69^a
Angrist-Pischke underidentification $\chi^2(3)$	46.41^a	83.99^a
Test of overidentifying restrictions	6.50	3.85
$\chi^2(2)$ p-value	0.72	0.15

Notes: robust standard errors in parentheses, clustered by club, with ^{a}, ^{b} and ^{c} denoting significance at the 1%, 5% and 10% level respectively.

Table 10: Market-test: Match Performance and Discrimination - 1st stage

Dependent Variable:	Difference in Log Wage Bill	
Sample:	Pre-Bosman (1981-1993)	Post-Bosman (1996-2008)
Difference in share of black English players	0.035	0.116
	(0.081)	(0.123)
Difference in number of players used	0.009	-0.001
	(0.003)	(0.002)
Difference in lagged attendances	0.010^{a}	0.009^{a}
	(0.003)	(0.003)
Difference in lagged cup performance	0.009^{a}	0.004^{b}
	(0.003)	(0.002)
Relative record transfer fee	0.088^{a}	0.027^{a}
	(0.032)	(0.045)
Pair of clubs fixed effects	Yes	Yes
Observations	3129	3264
F test of excluded instruments	11.74^{a}	17.09^{a}
Angrist-Pischke underidentification $\chi^2(3)$	36.49^{a}	53.46^{a}
Test of overidentifying restrictions	0.46	3.03
$\chi^2(2)$ p-value	0.79	0.22

Notes: robust standard errors in parentheses, clustered by pair of clubs, with ^{a} and ^{b} denoting significance at the 1% and 5% level respectively.